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ABSTRACT

Marital Sorting, Household Labor Supply, and Intergenerational Earnings Mobility across Countries*

We present comparable evidence on intergenerational earnings mobility for Denmark, Finland, Norway, the UK and the US, with a focus on the role of gender and marital status. We confirm that earnings mobility in the Nordic countries is typically greater than in the US and in the UK, but find that, in contrast to all other groups, for married women mobility is approximately uniform across countries when estimates are based on women's own earnings. Defining offspring outcomes in terms of family earnings, on the other hand, leads to estimates of intergenerational mobility in the Nordic countries which exceed those for the US and the UK for both men and women, single and married. Unlike in the Nordic countries, we find that married women with children and with husbands from affluent backgrounds tend to exhibit reduced labor supply in the US and the UK. In these countries, it is the combination of assortative mating and labor supply responses which weakens the association between married women's own earnings and their parents' earnings.

JEL Classification: J3, J62

Keywords: assortative mating, intergenerational mobility, joint labor supply

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1. Introduction

Family background is an important determinant of welfare in adulthood. Children of rich parents not only inherit more wealth than children of poor parents, they also tend to have higher earnings themselves and to marry more affluent partners. Economic inequalities are transferred across generations through numerous channels and in ways which vary across countries. Hence, the extent to which family inequalities persist will not in general be the same in all countries. A large empirical literature studying the intergenerational transmission of earnings from fathers to sons shows that that earnings persistence tends to be relatively strong in the United States, while the Nordic welfare states stand out with low intergenerational earnings persistence; see Björklund and Jäntti (2000), Solon (2002), Jäntti et al. (2006), and Corak (2006). The empirical evidence is much thinner on the intergenerational transmission of earnings for women. While US mobility estimates based on *individual* earnings typically are similar for men and women (e.g., Mazumder, 2005), evidence from other countries suggests that the earnings of daughters exhibit a lower correlation with parental earnings than is found for sons; see, e.g., Couch and Dunn (1997).

The economic well-being of adults is, however, not determined by individual earnings alone, but also by the earnings of partners. Chadwick and Solon (2002), for example, show that in the United States, the elasticity of daughters' *family* earnings with respect to her parents' income is of the same magnitude as that typically found for *individual* earnings of sons and fathers. Moreover, the study finds that the individual earnings of husbands and wives are equally highly correlated with the incomes of their own parents as they are with incomes of their parents-in-law. In the same spirit, Ermisch et al. (2006) argue that assortative mating - i.e., the tendency of men and women with similar socioeconomic characteristics to marry - plays an important role in explaining intergenerational earnings persistence in Germany and Britain, and assert that 40 to 50 percent of the covariance between parental and own permanent family incomes can be attributed to the person to whom one is married. Hence, marital sorting seems to play a key role in shaping intergenerational family income persistence.

The purpose of the present paper is to present comparable evidence from several countries on intergenerational earnings mobility with a focus on mobility among women. As an integral part of the analysis, we examine the implications of using family earnings rather than individual earnings for the assessment of differences in intergenerational earnings mobility across countries, for women and men, respectively. Earnings variation reflects, in part, differences in hours worked, and the interplay between assortative mating and household labor supply will have important implications for mobility. From a simple framework based on economic theory, we argue that country-differences in intergenerational family earnings persistence are potentially driven by four factors. These are cross-country differences in: (i) individual wage persistence across generations, (ii) labor supply responses with respect to own wage, (iii) the degree of assortative mating, and (iv) labor supply responses with respect to the spouse's wage, arising from joint labor supply decisions in households. Assortative mating and family labor supply decisions are important determinants of the intergenerational persistence of earnings, *even* if we were to focus only on individual earnings. Although the last of these factors has been virtually ignored in the existing empirical literature, partly because of the focus on full-time employees, we show that labor supply responses play an important role in explaining country-differences in intergenerational family earnings persistence for women.

We use data sets from five different countries; Denmark, Finland, Norway, the United Kingdom, and the United States. The samples have been constructed to be as comparable as possible for the assessment of intergenerational earnings correlations. This has been achieved by tailoring the administrative data from the three Nordic countries to key properties of the survey data from United Kingdom (the National Child Development Study) and the United States (the National Longitudinal Survey of Youth). Assessments of earnings persistence are in most cases based on estimated earnings elasticities. There are substantial differences in the functional forms of the intergenerational earnings equations across countries, however, implying that linear regressions of log offspring earnings on log parent earnings may produce a misleading foundation for cross-country comparisons, (Bratsberg et al., 2007). Hence, rather than comparing estimates from (mis-specified) log-linear equations, we let the data from each country determine properties of the functional form and compare elasticities evaluated at the 10^{th}, 50^{th}, and 90^{th} percentiles of the earnings distribution in each country. This approach also makes it possible to examine cross-country differences in the structure of earnings persistence across each country's earnings distribution.

In line with existing evidence (e.g., Fernández et al., 2005), we find significant marital sorting in all countries considered, but that the degree of assortative mating is stronger in the United States and the United Kingdom than in the Nordic countries. It is notable that for most groups – though, importantly, not all, as we shall see – the elasticities of *individual* earnings tend to be of similar magnitude as the elasticities of *family* earnings with respect to own parents' earnings. This somewhat surprising result is explained by strong marital sorting which ensures that the earnings of a spouse typically are as closely correlated with own parents' earnings as are own earnings. There are two conspicuous and important exceptions from this pattern, however: for married women in the UK and the US, the elasticity of own earnings with respect to parents' earnings is much lower than that of family earnings and is similar to that of the Nordic countries. The reason is that women marrying rich men respond to the high wage of their husband by working fewer hours or by withdrawing from the labor market. And in the UK and the US, this cross-wage labor supply response turns out to be stronger than the direct labor supply effect arising from the fact that women marrying rich men also tend to have high earnings potential themselves. In the Nordic countries, the latter effect dominates. Our findings also confirm that the overall intergenerational earnings persistence is significantly stronger in the United States than in the Nordic countries, with the United Kingdom somewhere between. This conclusion applies for the intergenerational transmission of actual family earnings as well as for the intergenerational transmission of family earnings potential, as reflected in the spouses' combined human capital (educational attainment).

The rest of the paper is structured as follows. The next section provides a brief overview of relevant existing empirical evidence. Section 3 offers a theoretical discussion of the various sources of intergenerational family earnings persistence, and also introduces some important notation. Section 4 describes the data, and section 5 presents our empirical analysis. Section 6 concludes.

2. Related Literature

The evidence on intergenerational earnings mobility among women remains scarce, and comparative cross-country studies have typically focused on males. Prior US studies that examine female mobility, such as Altonji and Dunn (1991), Peters (1992), Couch and Dunn (1997), and Mazumder (2005), report estimates of intergenerational mobility based on *individual* earnings that are relatively similar for men and women. In these studies, the offspring are observed early in their working career and there is a concern that any lifecycle bias might impair the validity of findings. For example, in their study of current and lifetime earnings of men, Haider and Solon (2006) conclude that observing earnings of offspring in their twenties will lead to severe downward bias in estimates of intergenerational earnings elasticities. For Germany, Couch and Dunn (1997) find a very low, and even negative, intergenerational elasticity for women, indicating that labor supply effects may be important.

Dearden et al. (1997) report higher UK estimates for the intergenerational elasticities among daughters compared to sons, but their samples are restricted to offspring aged 33 in full-time employment. Similar patterns are found by Blanden *et al*. (2004), although they also report elasticities adjusted for changes in inequality (i.e., correlation coefficients) that are almost identical for men and women (when earnings are observed around age 30).

Other contributions have focused explicitly on the role of assortative mating. Both Chadwick and Solon (2002) and Ermish et al. (2006) conclude that assortative mating is equally important for men and women in understanding how family earnings are transmitted from one generation to the next. The Ermisch et al. study avoids complications that arise from variation in labor supply by examining occupational status instead of offspring earnings (UK and Germany) or by restricting their earnings samples to full-time employees. According to the Chadwick-Solon study, intergenerational mobility among married couples is somewhat higher with respect to the wife's parents than with respect to those of the husband, whether based on individual or combined earnings. In light of the US household labor supply literature, which has uncovered important gender differences in the allocation of home and market hours of work, a surprising finding in the Chadwick-Solon study is that "elasticities (with respect to parental income) for own earnings and spouse earnings are similar" (p. 343) in their samples of married daughters and sons. In other words, the earnings elasticity of wives with respect to the income of *his* parents is nearly equal in magnitude to the husbands' own intergenerational earnings elasticity. Based on a sample of individuals aged 33 to 41 drawn from the NLSY, Jäntti et al. (2006) find that US mobility in terms of individual earnings is much higher among women than among men.

Scandinavian evidence suggests that intergenerational earnings mobility is somewhat greater for women than for men, when measured by individual earnings; see Österberg (2000), Österbacka (2001), Bratberg et al. (2005; 2007), and Jäntti et al. (2006). Recent studies based on Swedish data confirm this pattern, as both Holmlund (2006) and Hirvonen (2006) find individual earnings elasticities for women to be lower than those of men. Hirvonen replicates the Chadwick-Solon study with Swedish data and finds weaker assortative mating than in the US study, but concludes that the differences in earnings mobility in Sweden compared to the US cannot be explained by "factors that affect the marriage match" (p. 31).

The relationship between assortative mating, family labor supply behavior and cross-sectional income inequality has been the subject of a substantial literature focusing on changes in inequality over time. Juhn and Murphy (1997), consistent with earlier findings by Cancian et al. (1993), report that a rising correlation between the earnings of husbands and wives played a significant role in explaining the rising inequality in family incomes in the US between 1979 and 1989. Juhn and Murphy reveal that, despite these changes, both employment rates and annual hours worked for married women are decreasing in husbands' wages in the upper half of the wage distribution even in the 1989 period. For the UK, Harkness et al. (1996) show that wives' participation rates grew most rapidly in lower and middle income families between 1979 and 1991, producing a very shallow gradient between wives' participation and family income in the upper part of the distribution; as in the US, participation declines at the very top. Cancian and Reed (1998) use March CPS data for 1979 and 1989 to study the impact of the distribution of wives' earnings on family income. They show that the impact varies with the inequality measure employed, reporting that wives' earnings reduce inequality compared to a counterfactual of no wives' earnings. Hyslop (2001) develops an intertemporal labor supply framework to examine the link between the dispersion in individuals' rates of pay and the inequality of family well-being using PSID data on two-earner families for the period 1979 to 1985. Hyslop finds evidence of strong positive assortative matching and reports that the cross-correlations between the wages of spouses is greater than that between their earnings, implying that income effects in labor supply act to mitigate the effects of assortative matching. In contrast, Aslaksen et al. (2005), using data for Norway between 1973 and 1997, report that women's labor earnings increase inequality in family income, particularly towards the end of their sample period.[1] In conclusion, while the evidence suggests that family labor supply choices mitigate the effects of assortative mating on family income inequalities in the US and, to some extent in the UK, this is less apparent in Nordic countries. This evidence is consistent with our own subsequent finding that in the US and the UK, in contrast to the Nordic results, the combination of assortative mating and labor supply responses weakens the association between married women's own earnings and those of their parents.

3. Theoretical Framework

As we have discussed, previous empirical studies indicate that the presence and size of gender differences in earnings mobility depend on the outcome measure used. Mobility tends to be lower for males than for females when estimates are based on offspring earnings and samples are not restricted to fulltime employees, but estimates of female earnings mobility vary widely across studies and countries. We now present a framework that helps us understand why mobility, particularly among women, differs across countries, and why it is important to take into account the implications of both assortative mating and family labor supply.

Intergenerational earnings mobility will be influenced by the transmission of human capital, labor supply behavior, and the sorting of individuals into partnerships. In a two-adult household, family earnings (Z_i) consist of a person *i*'s own earnings (Y_i) and those of the spouse (Y_i^S), i.e., $Z_i = Y_i + Y_i^S$. We write annual earnings as the product of an (average)

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[1] Studies using Nordic data from the 1960s, 1970s, or 1980s, such as Björklund (1992) and Cancian and Schoeni (1998), tend to conclude otherwise, that wives' earnings have an equalizing effect on the distribution of family income in these countries.

hourly wage (W_i) and hours worked (L_i) during the year. Lower-case letters denoting logarithms, $y_i = w_i + l_i$.

Social scientists have established the existence of a strong relationship between human capital of offspring and parents, typically measured by educational attainment or parental earnings; see, e.g., Haveman and Wolfe (1995). To represent this intergenerational association, we assume that the log wage of the offspring is a function of log parental earnings (y_{pi}) plus a term (ε_i) capturing the combined effect of factors that are orthogonal to parental earnings;

(1) $$w_i = \alpha + \lambda y_{pi} + \varepsilon_i, \quad 0 < \lambda < 1 \ .$$

Various theoretical perspectives suggest that the marginal effect of parental earnings is positive (i.e., λ>0).[2] One example is the 'mechanical' transmission of genetic and cultural traits that contributed to the parents' high earnings in the first place (Becker and Tomes, 1979). Moreover, economists often argue that, when capital markets are not perfect, high-earning parents will invest more in their children's human capital acquisition than parents with fewer financial resources (Becker and Tomes, 1986). Empirically, educational attainment turns out to be highly correlated with the financial resources of parents, but some authors, such as Carneiro and Heckman (2002), have argued that the causal mechanisms are related to long-run economic conditions of families rather than the capability to finance education beyond post-compulsory schooling. Among economists, the jury is still out on whether money really matters and, if so, why. Identification of a causal effect is complicated since financial resources and borrowing constraints are typically not observed, and because parental earnings are strongly correlated with other (partly inheritable) determinants of offspring human capital. Further, high persistence in (parental) earnings makes it difficult to isolate the effects of long-

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[2] Note that λ is assumed to be equal for men and women; this is justified in our empirical section below.

run family economic conditions and constraints at the time of entry into post-compulsory schooling. Whatever the underlying reasons, the empirical literature has identified significant positive associations between offspring human capital and parental income in all of the countries covered by the present study.[3]

Partnerships are not products of random events alone. Both love and money are likely to matter when individuals are sorted into households. Matching takes place in numerous private and professional environments. For our purpose, it is useful to summarize the outcome in terms of correlated earnings *potentials* determined by resemblance of factors like education, health, occupation, industry, and personality. Similarity with respect to family background provides one part of the explanation. Thus, assortative mating may result in part from a causal mechanism, reflecting, for example, that mating is motivated by the economic resources and/or the risk insurance it provides (Hess, 2004), and in part from a non-causal sorting mechanism, reflecting, for example, that educational institutions are important meeting places where the density of potential spouses is high (Blossfeld and Timms, 2003) and search costs are low. Evidence from different countries indicates that about one in five meet their coming partner in school, college, or university (Skyt-Nielsen and Svarer, 2006; Lewis and Oppenheimer, 2001). Finally, assortative mating may also arise if individual traits and skills are complements in household production (Becker, 1973). We capture the degree of assortative matching by means of wage resemblance within partnerships:

$$(2) \qquad w_i^S = \pi w_i + (1-\pi)\overline{w^s} + \xi_i \,, \quad 0 < \pi < 1$$

where ξ_i represents factors orthogonal to the persons own wage (which on average may deviate from zero). This simplified reduced-form matching function expresses that the log wage of the spouse is made up from a weighted average of one's own log wage and the

[3] See, e.g., McIntosh and Munk (2007) for evidence from Denmark, Pekkala and Lucas (2007) for Finland, Aakvik et al. (2005) for Norway, Ermisch and Francesconi (2001) for the UK, and Carneiro and Heckman (2002) for the US.

average log wage in the pool of potential matches, plus a residual term. Thus, the parameter π in equation (2) captures the extent to which high-skill individuals tend to match with other high-skill individuals. When π equals zero earnings potentials are unrelated within couples. In this, we depart from studies such as Chadwick and Solon (2002) and Lam and Schoeni (1993; 1994) where matching is assumed to take place on the basis of *realized* earnings.[4] Empirically, measures of assortative mating based on realized earnings are likely to be misleading. When spouses make joint decisions with respect to labor supply, even a close matching on potential earnings (i.e., wage rates) will not necessarily imply that actual earnings are highly correlated as a higher spousal wage induces a negative own labor supply response.

Let log hours worked be a linear function of own and spouse's log wages;

(3) $$l_i = l(w_i, w_i^S) = \eta w_i - \eta^S w_i^S + \kappa_i \,,$$

where η denotes the (uncompensated) elasticity of labor supply with respect to one's own wage, $-\eta^S$ is the (uncompensated) cross elasticity with respect to the wage of the spouse and κ_i is an individual labor supply component orthogonal to wages (which on average may differ from zero). We build into the notation the assumption that the cross elasticity is non-positive and that the parameter η^S is a non-negative number. Standard theory predicts specialization within households and that individual labor supply will typically depend positively on one's own wage and negatively on the partner's wage. A higher wage for the spouse induces a substitution effect that alters the hours-distribution within the household as well an income effect that raises the demand for leisure.

Empirical studies typically find that own wage labor supply elasticities are close to zero for men, but clearly positive for married women; see the surveys in Killingsworth (1983)

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[4] Our approach is thus more similar to that of Ermisch et al. (2006), who specify the matching function on the basis of individual human capital aquisition.

and MaCurdy and Blundell (1999). The empirical evidence also documents a considerably larger cross-response to an increase in partner's wage for women than for males, presumably reflecting that home production, including child care, is a closer substitute for time spent in the labor market for women than for men.[5] In fact, several studies suggest that the labor supply of married men is quite inelastic and affected by neither the own nor the partner's wage. Due to a limited number of studies and the variety of model specifications, it is difficult to establish a firm pattern of differential labor supply elasticities for men and women across countries. Our own evidence in section 5, however, indicates that countries differ in this regard.

In order to facilitate interpretation, let us for the moment consider the female case and generalize the notation by letting the labor supply elasticities differ by gender. Combining equations (1) to (3), we can express log (individual) female earnings (y_f) in terms of the log earnings of her parents (y_{ip}) and all other factors (K_i) ;

(4) $$y_{fi} = \beta_f y_{pi} + K_i, \; where \; \beta_f = \left((1+\eta_f) - \pi\eta_f^S\right)\lambda \, .$$

The coefficient β_f denotes the intergenerational own earnings elasticity. As the equation shows, the elasticity of female earnings with respect to her parents' earnings is first of all determined by the intergenerational transmission of earnings capacity (λ). A positive own labor supply elasticity (η_f) reinforces the effect of parental earnings on the offspring's wage. With marital sorting ($\pi > 0$), the elasticity is lowered by (the absolute value of) the cross elasticity of women's labor supply with respect to her husband's wage (η_f^S). When women from high-income families also marry partners with high wages, their own hours of work will be reduced. This result illustrates that assortative mating and family labor supply decisions are

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[5] Relevant studies include Lundberg (1988), Juhn and Murphy (1997), Eckstein and Wolpin (1989), Devereux (2004), and Blau and Kahn (2006) for the United States; and Dagsvik and Zhiyang (2006) for Norway.

important components of the intergenerational persistence of earnings, *even* if we were to focus only on individual earnings. While labor supply responses to own wages will boost the intergenerational elasticity, assortative mating and cross labor supply responses will tend to moderate *individual* earnings persistence across generations. An implication is that the role of assortative mating cannot be inferred from the association between the partner's outcome and own family background alone. With non-zero cross labor supply responses, the partner's earnings will relate systematically to own parents' earnings even in the complete absence of assortative mating.[6]

The residual term in equation (4), K_i, is made up from a composite of residuals and constant terms from equations (1)-(3);

$$K_i = \kappa_i + \left((1+\eta) - \pi\eta_f^S\right)\left(\alpha + \varepsilon_i\right) - \eta_f^S\left((1-\pi)\overline{w} + \xi_i\right).$$

The specification hints at two potential sources of bias in estimates of the intergenerational earnings elasticity. First, because K_i contains the residual from the labor supply function, any positive correlation between parental earnings and labor supply above and beyond that captured by the intergenerational transfer of earnings potential (equation 1) will lead to overstatement of the earnings elasticity. One might argue that the labor supply mechanism also operates through intergenerational persistence in preferences (Altonji and Dunn, 2000), implying that offspring hours are positively correlated with parental earnings. Second, the residual from the matching function (equation 2) enters K_i with a negative sign. In other

[6] Ermisch et al. (2006) argue that identification can be achieved by using measures of occupational status that are highly correlated with hourly wages, or by restricting samples to full-time workers. Such a sample restriction is, however, likely to induce sample-selection bias in the estimates. And, even in the absence of unobserved sorting into full-time work, earnings may be a bad proxy for the hourly wage, as hours worked vary substantially even conditional on full-time status. To illustrate, consider the samples of full-time workers drawn from German GSOEP data used by Ermisch et al. In these samples, the gender log earnings differential is 0.47 (Table 1, p 688), which far exceeds recent estimates of the raw gender wage gap in Germany (e.g., 0.22 as reported by Olivetti and Petrongolo, 2006). The large gender earnings gap can only be understood in terms of systematic gender differences in hours worked, even in samples that are restricted to full-time workers.

words, any matching related to parental earnings (again, beyond that captured by the link between the offspring wage and parental earnings) will impart a negative bias in estimates of the intergenerational earnings elasticity through labor supply effects.

Turning next to the association between the earnings of the husband and those of the wife's parents, equations (1)-(3) imply that the elasticity of the husband's earnings (y_{if}^{S}) with respect to *her* parents can be written

(5) $$y_{fi}^{S} = \beta_{f}^{S} y_{pi} + K_{i}^{S}, \text{ where } \beta_{f}^{S} = \left\{\pi(1+\eta_{m}) - \eta_{m}^{S}\right\}\lambda \, .$$

Seen from her husband's point of view, β_{f}^{S} represents the elasticity of own earnings with respect to the earnings of his parents-in-law. Equation (5) shows that this elasticity depends crucially on the degree of marital sorting. Although the association between the husband's earnings and parents-in-law earnings will be reduced by any cross effect of wives' wages on husbands' hours worked (η_{m}^{S}), marital sorting gives rise to a positive relationship that is further reinforced by any own-wage male labor supply response.

Indeed, asymmetry in cross labor supply effects may well generate the result that, for married women, the "in-law elasticity" of husband's earnings with respect to the earnings of *her* parents exceeds her own earnings elasticity. It follows from equations (4) and (5) that

$$\beta_{f}^{S} > \beta_{f} \Leftrightarrow \pi(1+\eta_{m}+\eta_{f}^{S}) > 1+\eta_{f}+\eta_{m}^{S} \, .$$

Thus, the in-law elasticity might be the larger of the two if the degree of marital sorting is sufficiently strong, if the wife's labor supply is highly responsive to her husband's wage while less influenced by her own wage, and if the husband's labor supply responds more strongly to his own wage than to the wage of his wife.

The association between parental earnings and potential earnings of partners makes the case for focusing on the earnings of families rather than individuals. In fact, studies of the generational earnings mobility of women typically consider family or combined earnings (e.g., Chadwick and Solon, 2002). The elasticity of combined earnings with respect to the parental earnings of the wife (μ_{fi}) is a weighted average of own and spouse elasticities:[7]

(6) $$\frac{dz_{fi}}{dy_{pi}} = \mu_{fi} = (1-\theta_i)\beta_f + \theta_i \beta_f^S, \quad 0 \leq \theta_i = \frac{Y_i^S}{Y_i + Y_i^S} \leq 1.$$

In practice, the weight (θ_i) on the spouse elasticity will typically exceed one half for women as their husbands on average work longer market hours, and frequently at a higher wage. Note that the subscript for individual i follows from variation in wage levels and labor supply that are orthogonal to parental earnings.

Focusing on the *average* combined earnings elasticity for women with respect to the earnings of her parents, we substitute from (4) and (5) to get

(7) $$\mu_f = \left\{(1-\theta)(1+\eta_f) - \theta\eta_m^S + \pi\left[\theta(1+\eta_m) - (1-\theta)\eta_f^S\right]\right\}\lambda,$$

where θ is defined as the average of the husbands' shares of household earnings. The combined earnings elasticity contains four elements that modify the offspring human capital transfer component. The first element consists of the own labor supply elasticity weighted by the female share of household earnings. The second is a negative component that arises from any reduced spouse labor supply in response to a higher own wage. Marital sorting gives rise to two effects with opposite signs. A more favorable family background typically implies a partner with a higher wage (and possibly working more hours, *ceteris paribus*). Finally, a spouse with a higher wage will tend to reduce the person's own hours worked.

The elasticity of combined earnings with respect to the earnings of the husbands' parents is of course similar in structure;

(8) $$\mu_m = \left\{\theta(1+\eta_m) - (1-\theta)\eta_f^S + \pi\left[(1-\theta)(1+\eta_f) - \theta\eta_m^S\right]\right\}\lambda$$

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[7] Note that this expression holds for any value of y_p but that the empirical estimate of μ will not necessarily coincide with the sum based on individual earnings elasticities and the mean share of spouse earnings.

Marital sorting affects the combined elasticity for both men and women. The elasticities in equations (7) and (8) are increasing in π under the following conditions;

$$\frac{\theta}{1-\theta} > \frac{\eta_f^S}{1+\eta_m}$$

and

$$\frac{1-\theta}{\theta} > \frac{\eta_m^S}{1+\eta_f} \ .$$

In the absence of cross wage effects on labor supply (i.e., if η_f^S and η_m^S are zero), a higher degree of marital sorting will unambiguously reduce mobility. In general, a high husband share of family earnings will increase the likelihood that the first condition holds. Chadwick and Solon (2002), for example, report that θ equals .7 in their US samples. Likewise, low responsiveness of hours worked among married men to their wives' wages will raise the probability that the second condition holds.

When the division of labor is unequal within households and labor supply responses to wage changes differ by gender, the intergenerational elasticity of combined earnings will generally differ for men and women. Combining (7) and (8) we have

(9) $$\mu_m - \mu_f = (1-\pi)\left\{(2\theta-1) + \theta(\eta_m + \eta_m^S) - (1-\theta)(\eta_f + \eta_f^S)\right\}\lambda$$

In a "gender-blind" society, with $\theta = 0.5$ and equal male and female labor supply elasticities, mobility is the same for men and women given that transmission of human capital is the same for both sexes. In other words, under such conditions the combined earnings elasticity would be the same with respect to his or her parents' earnings. This point illustrates that any differences in earnings mobility between men and women will relate to gender-specific effects of parental earnings on skill formation, to uneven division of labor within the household, or to wage discrimination by gender. Even if wages did not affect labor supply at all, earnings persistence would be greater with respect to the husband's parents simply because the weight on his own elasticity typically exceeds one half, i.e., $\theta > 0.5$. Larger labor supply elasticities for women than for men will tend to equalize mobility when measured in terms of combined earnings. Naturally, the combined elasticity with respect to the wife's parents is more sensitive than that with respect to the husband's parents to daughters' labor supply responses.

Far from everyone lives in two-adult households and intergenerational earnings mobility is expected to differ for single and married women. For example, a negative cross labor supply elasticity with respect to the husband's wage applies to married women only, implying that the (own) intergenerational earnings elasticity of single women will be larger than that of married women. Furthermore, the process of household formation (and destruction) might generate selective subsamples of married and single individuals. To illustrate, suppose the probability of being single relates negatively to own earnings potential.[8] Positive selection into marriage will generate a negative correlation between the conditional expectation of the error term (of equation 1) and parental earnings, rendering a negative bias in within-group estimates of the intergenerational elasticity. The intuition is that the within-group conditional expectation of the residual is higher among those with a low-income than those with a high-income background, simply because the probability of being married increases with parental earnings. As such, both within-group estimates may be lower than the intergenerational earnings elasticity estimated from the pooled sample.[9] In other words, by conditioning on marital status, a 'channel' through which parental earnings affect offspring earnings is excluded.

[8] Our own data show that married men on average have higher earnings, greater educational attainment, and higher parental earnings than single men in all countries considered. For women, such differences by marital status are smaller. See Table 1.

[9] The reasoning is similar to that of Hertz (2004) who argues that the within-group intergenerational persistence of blacks and whites in the US is lower than the overall persistence.

Within our framework of earnings capacity transmission, labor supply, and assortative mating, the “deep” parameters λ and π can be identified from intergenerational earnings elasticity estimates if we make further (and restrictive) assumptions. For example, if the cross-wage labor supply elasticity for males is zero (i.e., male labor supply is unaffected by the spouse wage), we can retrieve the assortative mating parameter from the ratio of two intergenerational elasticities for males;

$$(10) \qquad \frac{\beta_f^S}{\beta_m} = \frac{\left\{\pi(1+\eta_m) - \eta_m^S\right\}\lambda}{\left\{(1+\eta_m) - \pi\eta_m^S\right\}\lambda} = \pi \quad if \quad \eta_m^S = 0 .$$

If, on the other hand, $\eta_m^S > 0$, the ratio in equation (10) will yield a downward biased estimate of the degree of assortative mating as

$$(11) \qquad \frac{\beta_f^S}{\beta_m} < \pi \Leftrightarrow (\pi^2 - 1)\eta_m^S < 0$$

which holds when $\pi < 1$. Identification of the intergenerational transmission of earnings capacity (λ) is more difficult and can be achieved from intergenerational earnings elasticity estimates only under the strong assumption that (male) hours worked do not respond to wage changes (i.e., $\eta_m^S = \eta_m = 0 \quad \Rightarrow \quad \lambda = \beta_m$).

This section has illustrated that measures of earnings mobility are affected by the transmission of human capital across generations, by family labor supply decisions, and by marital sorting on earnings potential. Thus, potential mechanisms that lie behind cross-country differences in earnings mobility extend beyond differences in the association between skills and family background.

4. Data

The empirical analyses draw on intergenerational earnings data from five countries: Denmark, Finland, Norway, the United Kingdom, and the United States. While the UK and US data are based on household surveys, the Nordic data are collected from administrative registers. As a guiding principle behind our adaptation of data from the different countries, we exploit the flexibility and richness of register data to ‘mimic’ the data generating processes behind the survey-based data sets. In terms of birth cohorts and ages at which we observe parental and own adult earnings, the UK National Child Development Study (NCDS) acts as a baseline to which we adapt the other data sets. The NCDS sampled all children born during the week of March 3rd- 9th, 1958, and is the data source used in the studies of Dearden et al. (1997) and Blanden (2005). The most recent sweep is that from 1999/2000, providing information on offspring’s earnings, employment, and marital status at age 41 or 42. As demonstrated by Böhlmark and Lindquist (2006) for Sweden and Haider and Solon (2006) for the US, (at least for males) observing earnings at this age offers the advantage that they correlate highly with lifetime earnings. Furthermore, from the 1974 sweep we obtain information on parental earnings at a time when the children were about 16 years of age.

Included in the UK analysis sample are offspring with valid data in the 1999/2000 sweep on economic activity or pay for oneself and, if married or cohabiting, the partner. The UK earnings data used in this study refer to net weekly pay, as this is the only earnings information available for partners and parents in the NCDS. The survey did not record pay information from self employment, and self-employed individuals are therefore excluded from analyses of earnings but are included in analyses of employment status. Because the age of the father was recorded only in the initial survey (in 1958), the sample is limited to those living with their biological father in 1974. The temporary reduction of the working week during 1974 to three days is a potential concern, but Grawe (2004b) finds no indication that *usual* weekly earnings are misreported by parents interviewed during this period and concludes that his study supports the use of NCDS income data for 1974. Parental earnings are computed as the sum of father's and mother's pay,[10] inflated from the interview month (which ranges from January 1973 to February 1975) to 2000 currency. Excluded from the sample are those whose father or mother report being employed at the time of the 1974 sweep but failed to provide pay information. Also excluded are families where neither the father nor the mother provides any pay information at all. Finally, we drop some observations from the analysis samples where the reported pay appears to be an extreme outlier.[11]

Our US data are drawn from the National Longitudinal Survey of Youth (NLSY79).[12] In order to obtain a sample of reasonable size, while maintaining comparability with the data from the other countries, we include respondents born between 1957 and 1964 in the US sample. Parental income refers to family income from all sources in 1978 and 1979, when the children were between 13 and 21 years old. Sample inclusion requires that we can link children and fathers, either because the respondent or a (younger) biological sibling lived in the same residence as the father at the time when we observe parental income. The earnings of cohort members (and their partners if married or cohabiting) refer to annual wage or salary income collected from the survey year in which the cohort member was 41 years of age.[13, 14]

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[10] For each of the parents, the 1974 sweep reports net pay within weekly or monthly intervals. We assign a value to each interval applying the algorithm of Blanden (2005), who used the within-band mean pay for each gender obtained from comparable families in the 1974 Family Expenditure Survey. Sensitivity checks based on data from the other countries in which we mimic the income bracketing of the UK data yield results that are very similar to those reported below based on non-bracketed parental earnings data.

[11] Sample exclusion is based on a regression of log parental earnings on home ownership, father's and mother's occupations, educational attainments, and ages, plus interaction terms. We drop observations when the absolute value of the residual exceeds three times the root mean squared error of the regression. From the sample of 5,327 parents with offspring participating in the 1999/2000 sweep, the algorithm dropped 67 observations with extremely low values and one observation with an extremely high value for parental earnings. Further details are available from the authors upon request.

[12] Below, we report correlations in educational attainment involving parents and parents-in-law. Because the NLSY does not provide information about the spouse's parents, these correlations are computed from a comparable sample (in terms of birth years) drawn from the PSID.

[13] All employment, earnings, and marital status data for the offspring generation are collected from one of the 1998, 2000, 2002, or 2004 interviews of the NLSY. The algorithm that picks survey year seeks the year closest to the respondent's 41st birthday, but selects the closest alternative year if key data for the cohort member or the spouse are missing in the preferred interview.

[14] The US earnings and income data are subject to top coding. Parental income (from 1978 and 1979) is top coded at $75,000 and offspring earnings are coded so that the top two percentile earners receive the same earnings value. For top-coded parents, we impute income as 1.5 times the income cut-off, and for offspring we use the conditional mean earnings for the top-coded group supplied in the survey data. Dropping top-coded

For the cohort members, hours worked are computed from the annual work history data the year of the earnings observation, while employment status of the spouse refers to activity the week prior to the interview.

The Norwegian sample is based on the complete 1958 birth cohort drawn from population registers, matched with the earnings records from 1971 and 1976 of their biological parents. Earnings records for both parent and offspring generations are drawn from the pension register, and include incomes from all work-related sources such as wage and salary income, self-employment earnings (net of interest payments), unemployment benefits, and long-term sickness benefits. Offspring and their partner's earnings refer to total earnings during the 1999 calendar year, and employment status (i.e., part or full time) is that reported by their main employer that year. For Denmark, we also use the 1958 birth cohort, with biological parents' earnings measured in 1980 and 1981 and offspring earnings and employment measured in 1999. The earnings information comes from tax registers and covers total earnings from all employers paid to the worker during the year. The earnings data are considered to be of high quality as they are used by tax authorities to assess each employee's earnings. The wage records also constitute deductible labor costs for employers, so firms have a strong incentive to provide accurate and timely information to authorities.[15] During the period studied, there was no change in the construction of the earnings variable.

For Finland, the matching of parents and children is based on household records taken from the quinquennial census panel covering the period 1970 to 2000. Because the sample from census records covers only about five percent of the population, we include in the

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parents (or, alternatively, not imputing their income) raises elasticity estimates, while dropping top-coded offspring leads to lower estimates, but the changes in elasticity estimates are only around 0.01-0.02. Hence, the substantive conclusions drawn in this paper are unaffected by how we handle top-coded earnings and income data.

[15] Unfortunately, the Danish earnings records do not include self-employment income. We therefore drop from the earnings data observations for which one of the parents, the cohort member, or the spouse are registered as self employed. We also drop a few observations with registered earnings below $100. As in the UK sample, self-employed individuals are included in employment regressions but are excluded from earnings regressions.

analysis offspring born between 1956 and 1960 out of sample size concerns. Earnings of parents are observed in 1970 and 1975 and earnings of offspring the year they turned 41 (in other words, earnings, employment, and marital status of the oldest cohort members are measured in 1997 and the youngest in 2001). The main source for the earnings data is tax records, which across the relevant years had quite similar definitions of earnings, and consist of all wages and salaries including both farm and non-farm self-employment income as defined for purposes of taxation.

All earnings measures used in this paper are reported in terms of US dollars in year 2000 prices. We convert from national currencies using the 2000 PPP exchange rates taken from World Development Indicators and use national CPI price indices to inflate nominal amounts into the year 2000 prices. To reduce estimation bias from measurement error in parental earnings, we base as a rule our analyses on two years of parental earnings. An exception is the UK where parental earnings data are available from the 1974 sweep only, but even for the other countries we rely on only one year of parental earnings whenever two observations are not available. This concern primarily applies to the US sample, where 31 percent of the parental earnings records draw on a single observation.[16] Prior studies based on US data, such as Solon (1992) and Mazumder (2005), conclude that using only one year of parental earnings data leads to attenuation bias in intergenerational elasticity estimates. As such, one might speculate that our UK, and even the US, estimates contain a larger negative bias compared to those from the other countries.[17] Finally, because estimation of the

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[16] The most common reason for a single observation in the US sample is that of missing information on earnings in one of the survey years or that the youth moved out of the family home between the first and the second observation year. In Norway, 94 percent of the sample builds on two observations of parental earnings; here a single earnings observation is typically due to the father being younger than the sample cut-off (35 years of age) in 1970.

[17] Sensitivity analyses, in which we drop one year of parental earnings records, or, alternatively, add up to ten years of parental earnings observations when such data are available, yield results that are consistent with attenuation bias in our estimates. These experiments indicate bias of an order similar to Mazumder (2005) — in the “balanced” subsample with two observations, the estimate of the father-son earnings elasticity falls by 11 percent in Norway and 13 percent in the US when based on only one year of parental data. Attenuation is

intergenerational earnings elasticity is based on log earnings regressions, those with zero earnings are excluded from individual earnings regressions but are included in analyses of combined (family) earnings.

Table 1 reports descriptive statistics for our analyzed samples, separately by gender and marital status of the cohort member. For married and cohabiting cohort members, we also list statistics for their spouses.[18] The table includes descriptive statistics for the parents of the cohort member for all five countries, and for Denmark and Norway, the two countries where parent-child matches are based on birth registers, we are also able to link cohort members and their parents-in-law. The age of fathers at the time we observe their earnings is fairly similar across countries, on average about 46 to 47, except for Denmark where parental earnings are measured when the offspring is older than in the other countries. Thus, any lifecycle bias arising from non-representative measures of parental earnings is likely to be small and not systematically different between the US, the UK, Finland, and Norway. The table shows familiar gender differences in earnings, with male earnings exceeding those of females in all samples. An interesting pattern, that is particularly relevant for the present study, is that marital status matters differently for men and women. While married men earn more than single men in all countries considered, unmarried women earn more than married women in the United Kingdom and in the United States. Similar patterns arise for employment status. While married men are more likely to be employed and are less likely to have part-time jobs than unmarried men in all five countries, labor market attachment of married women tends to fall below that of unmarried women. In particular, the likelihood of holding a part-time job is

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expected to cause downward bias in all of our samples. If the bias relates to the proportion of the sample with a single parental earnings observation, the problem will be greatest in the UK data, followed by the US, and least severe in the Nordic samples; see also the discussion in Bratsberg et al. (2007).

[18] We observe both those married with a spouse present and those cohabiting with an unwed partner in four of the five data sets. For Norway, the data allow classification of marital status (including whether separated, divorced, or widowed), but cohabiting partners are identified only if they have a child in common. Thus, for Norway our subsamples of singles will include some misclassified individuals.

higher for married than for unmarried women in all of the countries for which data on part-time employment is available. The importance of such gender differences in household labor supply behavior forms a key topic of interest in the empirical analyses of intergenerational earnings mobility that follow in the next section.

- Table 1 around here –

5. Empirical analysis

5.1. Intergenerational Earnings Mobility.

We start out this section by presenting our basic measures for intergenerational earnings persistence, being the elasticities of various measures of offspring earnings with respect to own parents' earnings. We focus on three different intergenerational elasticities with respect to own parents' earnings. These are the elasticity of; (i) own earnings (β), (ii) spouse earnings (β^s), and (iii) combined (family) earnings (μ). Given that the relationship between offspring and parental log earnings is strongly non-linear in some of the countries (Bratsberg et al., 2007), we do not impose constant elasticity assumptions. Instead, we model the log of each offspring earnings measure as a polynomial function of log parental earnings.[19] The order of the polynomial is then selected separately for each country on the basis of a Bayesian information criterion (BIC). As a main basis for cross-country comparisons, we report the estimated earnings elasticities evaluated at the median of the parental earnings distribution for each country; see Table 2. For the countries with nonlinearities in the intergenerational earnings elasticity estimates (Denmark, Finland, and Norway), we also report in the Appendix estimated elasticities evaluated at the 10^{th} and at the 90^{th} percentiles of the parental earnings distribution; see Table A-1.

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[19] See, e.g., Grawe (2004a) and Bratberg et al. (2005) for other studies which use flexible specifications that allow for nonlinear relationships between parental and offspring earnings.

- Table 2 around here –

A key finding reflected in Table 2 is that, in general, married women have higher earnings mobility than married men in all countries. While this holds for both individual and combined earnings, the gender difference in economic mobility is much larger for individual earnings than for combined earnings. This indicates that assortative mating is a relatively more important component of observed earnings mobility for women than for men. A second observation is that the differences in own-earnings mobility between single and married females are much larger in the UK and the US than in the Nordic countries. This finding suggests that joint (within-family) labor supply decisions are more important for mobility in the former than in the latter group of countries. A third observation is that, with the exception of Denmark, the estimates of the intergenerational earnings elasticity for married men exceed those for single men. The pattern is suggestive of differences in the own-wage labor supply elasticity by marital status, perhaps because of a greater scope for within-household substitution between market and non-market work for married men. But, the pattern could also result from downward bias in estimates based on the subsample of single men.[20] A fourth observation is that earnings persistence tends to be significantly stronger in the United States than in the Nordic countries, with the United Kingdom somewhere in-between. This pattern applies regardless of whether we look at singles or married couples or at individual or combined earnings, with one notable exception, namely that of married women's own earnings. Indeed, the estimated elasticity of married women's own earnings with respect to parental earnings is remarkably similar across countries and is no larger in the US than in the other countries. As we show below, this apparently high earnings mobility among married US females results from a combination of strong marital sorting and labor supply responses to the

[20]A close inspection of the samples reveal that, among those with high parental earnings, divorced men tend to have substantially lower own earnings than married men. Such differences by marital status are smaller among men with low parental earnings. As emphasized in Section 3, selectivity into marriage and divorce might trigger bias in within-group coefficient estimates.

husband's wage. An early indication of the importance of marital sorting and household labor supply responses can be obtained by looking at the elasticities of own earnings with respect to those of parents-in-law (also reported in Table 2). While men's earnings display a much stronger positive association with in-laws' earnings in the United States than in the other countries, there is no such cross-country pattern for females.

Table A-1 in the Appendix provides a description of the nonlinearities in intergenerational earnings elasticity estimates in the Nordic countries (for the US and the UK, linear models were selected on the basis of the BIC). For all of the Nordic countries and for each estimated elasticity, we find that earnings persistence is low at the bottom of the parental earnings distribution, and then increases sharply as we move towards the top of the distribution. However, even at the 90th percentile most of the estimated elasticities for the Nordic countries are firmly below those reported in Table 2 for the US. The finding of high economic mobility at the bottom of the earnings distributions in the Nordic countries is in line with results in Bratsberg et al. (2007), and suggests that expected adult earnings in these countries do not fall beyond a certain level, a pattern presumably generated by the educational systems and by the relatively high minimum wages.

In order to interpret the country differences in Table 2, we need to look more closely into the various mechanisms that shape persistence in combined earnings. From the discussion in section 3, we know that the own earnings elasticity (β) relates directly to the intergenerational transmission of own earnings potential (λ), the elasticity of labor supply with respect to ones own wage (η), the degree of marital sorting (π), and – to the extent that there is marital sorting – to the elasticity of labor supply with respect to the spouse wage (η^s). The elasticity of spouse earnings (β^s) also depends on the intergenerational transfer of earnings potential (λ), and/or on the spouse's cross labor supply response. With marital sorting (i.e., $\pi > 0$), the spouse earnings elasticity also depends on the spouse's own labor supply elasticity. Finally, the elasticity of combined earnings (μ) will reflect each of the factors underlying the own and spouse elasticities discussed above, with weights given by the relative size of the two incomes.

5.2. The Intergenerational Transmission of Earnings Potential.

We first examine the intergenerational transmission of earnings potential (λ). Unfortunately, data on hourly wage rates are in general not available in our datasets, and can in any case only be observed for individuals who work. It is therefore not obvious how the earnings potential of the offspring should be measured. An important determinant of the hourly wage, however, is educational attainment, and in Table 3 we examine the extent to which educational attainment depends on family background. Since the intergenerational transmission of earnings potentials cannot be assumed to occur exclusively through realized economic wealth in the parent generation (see Section 3), we first look at the correlation in years of schooling between offspring and parents; see panel A. The correlation coefficients tend to be slightly higher in the US than in the other countries. Further, men's years of schooling seem to be more strongly correlated with the attainment of their father than their mother in all countries, while for women the correlation is more or less the same with respect to both parents. Otherwise, there are only minor differences across genders.

Panel B of Table 3 reports results from regressions of offspring years of schooling on log parental earnings, with effects evaluated at the median of the parental earnings distribution. The listed coefficients display very similar patterns for men and women. The relative importance of parental earnings for educational attainment is greatest in the UK, and, somewhat surprisingly, the coefficients (evaluated at median parental earnings) reported in Table 3 do not indicate that educational mobility – at least in terms of years of schooling – is any lower in the US than in the other countries. Apparently, the low degree of US economic mobility revealed in Table 2 for both men and women does not stem from a particularly high impact of family economic resources on the offspring's years of schooling. However, as shown in the Appendix Table A-2, the relationship between years of education and log parental earnings is strongly non-linear in all countries except for the US. And once again, the typical pattern is that the estimated effects are small at the bottom of the distribution (smaller than in the US) and large at the top of the distribution (larger than in the US). Hence, the results indicate that growing up in a poor household has a larger bearing on educational opportunities in the US compared to the other four countries considered.

- Table 3 around here -

The economic returns to education vary significantly across countries, and the impact of educational attainment on earnings capacity clearly depends on the rate of return. In panel C of Table 3, we therefore present returns-adjusted elasticities of educational attainment with respect to parental earnings. The earnings-adjustment implemented here is quite primitive—it is based on log-linear earnings regressions using only the subsamples of full-time earners in each country, including dummy variables for years of schooling as the only right-hand-side variables. These models are then used to predict the earnings capacity for all of the offspring based on their educational attainment. The estimates presented in the table are the elasticities of predicted earnings capacity with respect to parental earnings.[21] Comparing the elasticities of returns-adjusted education reported in panel C with the unadjusted coefficients in panel B, we note that the returns adjustment tends to affect the elasticity estimates more for the US and Finland than for the other countries. These findings are consistent with existing evidence showing that the returns to education are particularly high in the US and Finland and low in Denmark and Norway (Harmon et al., 2003). The patterns even correspond to cross-country

[21] The standard errors have been adjusted for the extra sampling error that is generated by the fact that we use estimates of the earnings-adjusted educational attainment. The adjusted standard errors exceed the naive, unadjusted standard errors by 15 to 57 percent.

studies of wage inequality such as Blau and Kahn (2005) who argue that differences in skill prices are more important than variation in skills for explaining the higher wage dispersion in the US compared to many European countries.

An important point to note from the returns-adjusted elasticities in panel C is that there are virtually no differences between men and women in the intergenerational transmission of earnings capacity in any of the countries, in sharp contrast to the gender differences revealed in the intergenerational transmission of own earnings in Table 2. Across countries, earnings capacity seems to be subject to much stronger intergenerational persistence in the US and the UK than in the Nordic countries, which is in line with the descriptive pattern for males presented in Table 2.

5.3. Marital Sorting and Educational Attainment of In-Laws.

As discussed in Section 3, we examine the marital sorting mechanism (π) in terms of similarities in educational attainment and in earnings potentials (returns-adjusted educational attainment) rather than in terms of actual earnings. The reason for this is that actual earnings are as likely to *be caused* by marital sorting (through joint labor supply decisions) as they are to cause marital sorting. To a certain extent, this caveat also applies to educational attainment, since early marriages may have had an impact on the partner's years of schooling. For this reason, we also look at similarities in educational attainment between parents and parents-in-law. Table 4 reports these indicators of assortative mating. Panel A lists sample correlation coefficients of years of schooling between the two partners, between each partner and their parents-in-law, and, for Denmark, Norway, and the US, between parents and in-laws.[22] Significant marital sorting is prevalent in all countries. A general finding is that educational

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[22] The correlation coefficient is a consistent estimate of the elasticitity π when the variance of educational attainment is the same for both groups (i.e., men and women) within a generation.

attainment is more strongly correlated between spouses than between parents and offspring (see Table 3, panel A). All of the listed correlation coefficients indicate that the degree of assortative mating is strongest in the United States, but also stronger in the UK than in the Nordic countries**.** For US males, it even turns out that own educational attainment is as highly correlated with the attainment of their wife's father as their own father. The cross-country evidence is consistent with Fernández et al. (2005), who present a theoretical framework in which "..[an] increase in inequality increases sorting by making skilled workers less willing to form households with unskilled workers" (p. 282). The tradeoff in the matching process when a skilled individual gets a favorable match with an unskilled individual depends on the wage skill premium, which is larger in the US and the UK than in the Nordic countries.

- Table 4 around here -

An alternative estimate of the degree of assortative mating can be retrieved from the intergenerational elasticities for males, under the assumption that hours worked of married men are unaffected by changes in the spouse's wage; see equations (10) and (11). Panel B of Table 4 shows the corresponding estimates of π, which are larger than those based on years of schooling in all five countries. This is exactly as expected if marital sorting is also influenced by earnings capacity determinants orthogonal to educational attainment. The implied estimates confirm some of the earlier cross-country patterns, as sorting is stronger in the US than in the Nordic countries and also weaker in Denmark and Finland compared to Norway. The UK estimate, however, is surprisingly low compared to those of the Nordic countries, although it is almost identical to the UK estimate in Ermisch et al. (2006). An important caveat is that these estimates are downward biased if hours worked of husbands respond to their wives' wage, or if splitting the sample according to single status imparts bias; see the discussion in section 3.

Together, the two processes of intergenerational transfer of individual earnings capacity and marital sorting determine the intergenerational mobility of economic opportunities for married individuals. Panel C of Table 4 reports the elasticities of the partner's and of the combined (family) returns-adjusted education with respect to parental earnings. These numbers may be compared with the elasticities of own earnings capacity reported in panel C of Table 3. The panel shows that economic mobility in terms of earnings capacity is significantly lower in the US than in the Nordic countries, with the UK somewhere in-between (though closer to the US than to the Nordic countries). This pattern is similar to the pattern of male mobility in observed earnings. However, in sharp contrast to the mobility pattern for earnings, there seem to be no differences between men and women with respect to mobility of earnings capacity. Thus, in terms of overall economic opportunities, mobility is similar for men and women in all countries considered. The gender differences in observed earnings mobility revealed in Table 2 must therefore be driven by differences in the exploitation of these opportunities, i.e., in the labor supply decision.

5.4. Household Labor Supply.

Given the intergenerational transmission of human capital and given the marital sorting process, the degree of combined earnings mobility is determined by the spouses' joint labor supply decisions, as reflected by the labor supply elasticities. Again, our ability to examine these decisions is limited by the fact that in the majority of the data sets we do not observe either hourly wages or hours worked. What we do observe, however, is whether the offspring works or not, and (except for Denmark and Finland) whether a job is full-time or part-time. The extent to which labor supply is affected by own and spouse wages can indirectly be examined by looking at how labor supply relates to the earnings of own parents and the earnings of parents-in-law, respectively. Panel A of Table 5 reports the marginal effects of an increase in log parental earnings on the probability of working full-time, based on parameters from ordered probit models. For Denmark and Finland, we have not been able to distinguish properly between part-time and full-time work. In lieu of such data, in the Appendix Table A-3, we report probit models corresponding to those reported in Table 5, but this time with employment as the dependent variable. For comparison, these probit regressions are estimated for all five countries.

- Table 5 around here –

The ordered probit model does not give us estimates of the labor supply elasticities and the coefficients will reflect supply responses at the extensive as well as at the intensive margin. Provided that human capital is transmitted across generations $(\lambda > 0)$, the models in Table 5 may be viewed as reduced form labor supply functions. For men, there is a general pattern in the table that labor supply relates positively to the earnings of parents and parents-in-law. The former of these effects reflects a direct positive labor supply response with respect to the own wage. In other words, for men the direct substitution effect must dominate both the income effect arising from the higher own wage and (given that $\pi > 0$) the negative within-household substitution and income effects arising from a higher spouse wage. The positive effect of in-laws' earnings can only be understood in light of marital sorting. In the absence of marital sorting, the earnings of parents-in-law would enter into the labor supply function as a proxy for the spouse's earnings capacity, and a rise in the partner's earnings capacity should – ceteris paribus – cause an unequivocal reduction in own labor supply (both through an income effect and a within-family substitution effect). Our rather robust finding of the opposite effect for males arises from the fact that the earnings of parents-in-law not only serves as a proxy for the spouse's earnings capacity, but also – through assortative mating – as a proxy for own earnings capacity.

The pattern of positive effects of parental earnings on labor supply also appears for married women in the Nordic countries. In the samples of married females in the US and the UK, however, labor supply turns out to be *negatively* associated with the earnings of both own parents and in-laws. Hence, for women in the US and the UK, the marital sorting process appears to reverse the relationship between own earnings capacity and labor supply. Due to marital sorting, own earnings capacity also serves as a proxy for the spouse's earnings capacity. And for married females in the US and the UK, the income and within-family substitution effects with respect to partner's wage seem to dominate the direct effect of own earnings capacity.

The role of specialization within the family is also confirmed by the estimates reported in Panel B of Table 5 (Model 2), where the impact of log parental earnings on female labor supply is interacted with the presence of children (below 15 years of age) in the family. (For men, in all countries, the marginal effect of parental earnings on labor supply turned out to be unaffected by the presence of children, so these results are omitted.) Focusing first on couples without children, we find that the negative impact of parental earnings on US and UK female labor supply disappears in the samples of married women. In other words, the negative relationship observed in the top panel is entirely generated by joint labor supply decisions in households with children. In such families, it is likely that labor supply decisions are made collectively, with relative wages reflecting each spouse's bargaining power. As the wages of both spouses increase with parental earnings, the scope for specialization with respect to market and household work also rises with the earnings of parents and in-laws. This effect is, however, not constant across the earnings distribution. In panel C, we report estimates from models where the labor supply responses are estimated separately in the upper and lower halves of the parental earnings distribution. As the panel shows, the overall negative effect of in-laws' earnings on female labor supply observed in the US sample turns out to stem from labor supply responses among married women with children and in the upper half of the (in-laws') earnings distribution. A similar pattern arises with respect to own parents' earnings, although in this case the underlying positive association (for females without children) is stronger than that with respect to in-laws' earnings. In the bottom half of the earnings distribution it is actually the case that female labor supply rises with the earnings of both own parents and in-laws' even in the US, suggesting that in this part of the earnings distribution the direct work incentives associated with higher own wages dominate any income effects.

In the Nordic countries, we fail to uncover significant interaction terms between the presence of children in the household and parental earnings on female labor supply. Children do, however, affect female labor supply in all countries. But again, there are some interesting differences between the UK and the US, on the one hand, and the Nordic countries on the other. Note first from Table 5 that the presence of children reduces the mother's probability of working *full time* in all countries included, but that the effect is larger in the UK and the US than in Norway. Moving on to Table A-3 (Appendix), we see that while the presence of children reduces mothers' overall employment propensities in the UK and the US, this is not the case for any of the Nordic countries. In the Nordic countries, having children does not reduce female labor supply at the extensive margin, but lowers supply significantly at the intensive margin. Even in the UK, the impact of children on women's labor supply is much smaller at the extensive than the intensive margin.

Incentives for allocation of time between the household sector and the labor market differ across countries, and such differences may account for the patterns of female labor supply uncovered in Table 5. First, empirical studies show that part-time employment involves a wage penalty in the United States and in the United Kingdom, while the wage differential between full-time and part-time workers is negligible in the Nordic countries (Blank, 1990; Ermisch and Wright, 1993; Hardoy and Schøne, 2006). Recent studies point out that the US penalty is largely explained by measurable worker and job characteristics, as well as by unobserved worker heterogeneity (Hirsch, 2005), and that the UK penalty can be attributed to occupational segregation (Manning and Petrongolo, 2006). In the words of Manning and Petrongolo, "… women (in the UK) who want to move from FT (full-time) to PT (part-time) work are often forced to change employer and/or occupation and, on average, make a downward occupational move. This seems to occur even when they have the necessary skills and experience to do the higher-level job" (p.17). One might speculate that access to jobs that match the skills of highly educated women and also offer opportunities for working reduced hours is generally better in the Nordic countries, both because of the larger public sector and because of enhanced worker rights from legislation or collective agreements.

Second, cross-country differences in taxation may affect incentives for labor force participation of secondary earners. For example, joint taxation of families implies that the economic returns from participation will be low because the earnings of the secondary worker are taxed at the marginal rate of the bread winner (Piggott and Whaley, 1996). Empirical studies of tax reforms in the United States and Canada suggest that the "jointness" of the tax system affects household labor supply. LaLumia (2005) concludes that the US introduction of joint taxation in 1948 reduced participation among married women, and Crossley and Jeon (2006) show that the tax reform of 1988, which reduced the "jointness" of the Canadian tax system, led to increased participation among married women with high-income husbands. In the Nordic countries and in the UK, spouses are taxed as individuals although there are certain tax credits for married persons with a non-working spouse.

Third, the price, availability, and quality of child-care services are likely to affect labor supply decisions of mothers with pre-school children. With more subsidized and readily available child care in the Nordic countries compared to the UK and the US (Del Boca and

Locatelli, 2006), individual returns from market participation will differ and this may explain the finding that the impact of young children on female labor supply in the Nordic countries is lower than that in the US and in the UK. Moreover, assume that quality of child care is a ‘normal good’ and that care provided by highly educated parents and top-quality institutions are considered to be close substitutes. Given that the relative price of the market-based alternative is much lower in the Nordic countries than in the United States, affluent parents with high willingness to pay for quality child-care are more likely to leave their children with professional child-care workers in the Nordic countries.

Finally, the cross-country labor supply patterns in Table 5 could reflect underlying differences in preferences and attitudes towards work. Antecol (2000), for example, shows that there are large differences in male and female labor force participation rates across immigrant groups in the United States, and attributes the variation in the gender pattern to cultural factors. Along similar lines, in a study of second-generation American women, Fernandez (2007) finds that hours worked correlate with cultural proxies such as female labor force participation and survey-based evidence on attitudes towards women’s work in the country of ancestry (see also Fernandez and Fogli, 2005). While such evidence does not explain why parental background influences female labor supply differently across the countries included in the present study, it points to the possibility that American women with an affluent family background value non-market activities differently from their Nordic sisters.

6. Concluding remarks

In this paper, we have presented comparable evidence on intergenerational earnings mobility for Denmark, Finland, Norway, the UK and the US, with a focus on the role of gender and marital status. Married women experience greater mobility than men, both with respect to own and to combined (family) earnings; though the gender difference is greater in own earnings. We also confirm that earnings mobility in the Nordic countries is typically greater than that in the US, with the UK position intermediate. This is true for single men, for single women, and for married men. It is also true for married women if we base our estimates on either the husband’s or combined family earnings. However, in stark contrast, we find that when estimates are based on married women’s own earnings, then intergenerational mobility is approximately uniform across the US, the UK, and the Nordic countries. Similarly, while we find that mobility is typically greater for married women than for single women in all countries, the difference is greater in the US and the UK than in the Nordic countries.

We discuss and show empirically how patterns of mobility by gender and across countries are attributable to differences in intergenerational transmission of earnings capacity, in family labor supply and in the extent of assortative mating. Results show that the intergenerational transmission of human capital, when measured in terms of educational attainment, is fairly similar across the five countries. But, because the returns to schooling differ across countries, the channel of human capital transmission leads to greater intergenerational earnings persistence in the UK and the US than in the Nordic countries. It is also striking that, in terms of earnings capacity, mobility is similar for men and women in all countries considered.

As evidence of assortative mating, we examine the educational attainment of partners and find strong correlations within couples in all countries. Marital sorting tends to be slightly more prevalent in the UK and the US. We also find that there is a strong correlation between own education and that of parents-in-law; with variation both by gender and across countries, the parents-in-law educational correlation being especially strong for males in the US. This pattern is reflected in our finding that men’s earnings are highly correlated with those of parents-in-law, especially in the US; in contrast, the correlation between own earnings of married women and earnings of parents-in-law tends to be weaker in the US than elsewhere.

Our main conclusion is that the patterns we have uncovered in intergenerational mobility also reflect family labor supply decisions and variations in these across countries. We attribute the relatively weak correlation between married women's earnings and those of their parents-in-law in the US to the fact that the effects of assortative mating are being offset by the effects of household labor supply decisions. As evidence of this, we present findings on the influence of the earnings of parents and of parents-in-law on the labor supply of married men and women. In all countries, own parental earnings are positively associated with labor supply for men and single women. The same positive association holds for married women in the Nordic countries. In the US and in the UK, however, married women's labor supply is negatively associated with earnings both of own parents and of parents-in-law. Interestingly, this effect can be traced to relatively affluent households with children: in these households, joint labor supply decisions seem to follow traditional gender patterns of within-family specialization more than is the case in the Nordic countries. For married women, measures of mobility based on own individual earnings are misleading indicators of the intergenerational transfer of welfare for the purposes of comparisons both within and across countries; better measures would be based either on combined family earnings or on mobility in earnings potential.

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Table 1: Descriptive sample statistics

	Denmark	Finland	Norway	UK	US
Married and cohabiting female cohort members					
log earnings	10.04	9.58	9.82	5.40	9.89
log spouse earnings	10.42	9.99	10.46	6.17	10.68
log combined earnings	10.92	10.51	10.91	6.50	11.05
Age	41.0	41.0	41.0	41.4	40.7
Education	12.0	12.7	12.1	12.4	14.0
Employed	.870	.873	.915	.827	.820
Of which, part-time	NA	NA	.463	.476	.217
Child below 15 (18 in Finland)	.721	.834	.779	.655	.704
Spouse age	43.7	43.2	43.8	43.7	42.3
Spouse education	12.1	12.4	12.4	12.2	14.0
Spouse employed	.922	.924	.970	.922	.973
Of which, part-time	NA	NA	.045	.024	.040
log parents earnings	10.30	9.77	10.20	6.14	11.00
Father age	52.4	45.7	47.7	45.5	46.7
Father education	10.7	10.1	9.8	10.0	12.3
Mother education	10.0	9.8	9.0	10.0	11.8
log parents-in-law earnings	10.27		10.15		
Father-in-law age	52.9		50.0		
Father-in-law education	10.7		9.8		
Mother-in-law education	10.0		8.9		
Observations	15910	6433	19002	2148	1597
Single female cohort members					
log earnings	9.92	9.57	9.84	5.56	9.95
Age	41.0	41.0	41.0	41.4	41.0
Education	11.7	12.5	11.9	12.3	13.5
Employed	.719	.817	.858	.780	.867
Of which, part-time	NA	NA	.282	.256	.074
Child below 15 (18 in Finland)	.432	.495	.418	.384	.348
log parents earnings	10.28	9.79	10.16	6.15	10.82
Father age	52.3	45.8	47.6	45.6	46.8
Father education	10.8	10.1	9.7	10.2	11.4
Mother education	10.1	9.8	8.9	10.1	11.6
Observations	4975	2102	8007	481	594

Table 1, continued.

	Denmark	Finland	Norway	UK	US
Married and cohabiting male cohort members					
log earnings	10.43	9.98	10.46	6.25	10.71
log spouse earnings	10.00	9.49	9.76	5.30	9.89
log combined earnings	10.91	10.50	10.91	6.52	11.07
Age	41.0	41.0	41.0	41.4	40.9
Education	12.1	12.4	12.4	12.5	13.6
Employed	.927	.925	.977	.945	.947
Of which, part-time	NA	NA	.044	.020	.025
Child below 15 (18 in Finland)	.799	.823	.766	.715	.698
Spouse age	38.5	39.0	38.6	39.3	38.7
Spouse education	12.1	12.7	12.3	12.4	13.7
Spouse employed	.857	.862	.905	.776	.742
Of which, part-time	NA	NA	.469	.517	.198
log parents earnings	10.31	9.78	10.20	6.15	10.99
Father age	52.3	45.5	47.6	45.5	46.5
Father education	10.7	10.1	9.9	10.0	12.1
Mother education	10.0	9.8	9.0	10.0	11.8
log parents-in-law earnings	10.36		10.24		
Father-in-law age	49.4		45.6		
Father-in-law education	10.8		10.0		
Mother-in-law education	10.2		9.3		
Observations	17290	7281	18746	2092	1658
Single male cohort members					
log earnings	10.03	9.63	10.11	6.03	10.10
Age	41.0	41.0	41.0	41.4	40.7
Education	11.3	11.8	11.7	12.3	12.8
Employed	.708	.753	.846	.813	.853
Of which, part-time	NA	NA	.066	.043	.059
log parents earnings	10.23	9.68	10.13	6.11	10.87
Father age	52.4	46.6	47.7	45.6	47.2
Father education	10.6	10.0	9.6	10.0	11.5
Mother education	9.9	9.7	8.8	10.0	11.4
Observations	6569	2310	9305	538	655

Note: Sample size refers to number of observations with nonzero parental earnings; valid number of observations of other characteristics (such as employment status and own, spouse, or in-law earnings) may vary slightly from listed figure. Earnings of offspring are gross annual wage and salary incomes, except for in the UK where they refer to net weekly pay. Parental earnings are combined earnings of father and mother except for the US (total family income); annual except for the UK (weekly). All earnings are adjusted to 2000 prices and then converted to 2000 international US dollars using OECD PPP exchange rates. The cohort samples consist of individuals born in 1958, except for Finland (1956-60) and the US (1957-64).

Table 2: Intergenerational earnings elasticity estimates

	Denmark	Finland	Norway	UK	US
Female cohort members					
All: Own earnings	.190	.197	.186	.270	.252
wrt *her* parents	(.011)	(.018)	(.013)	(.047)	(.047)
Singles: Own earnings	.266	.169	.192	.382	.435
wrt *her* parents	(.031)	(.038)	(.026)	(.110)	(.081)
Couples: Own earnings	.170	.206	.183	.239	.197
wrt *her* parents	(.012)	(.020)	(.014)	(.051)	(.055)
Couples: Husband's earnings	.164	.213	.209	.266	.371
wrt *her* parents	(.011)	(.018)	(.012)	(.052)	(.036)
Couples: Combined earnings	.192	.226	.210	.325	.359
wrt *her* parents	(.009)	(.015)	(.009)	(.046)	(.035)
Couples: Combined earnings	.204		.237		
wrt *his* parents	(.009)		(.009)		
All: Combined earnings per	.203	.206	.210	.326	.408
adult wrt *her* parents	(.010)	(.015)	(.010)	(.043)	(.034)
Male cohort members					
All: Own earnings	.261	.284	.269	.415	.482
wrt *his* parents:	(.011)	(.016)	(.011)	(.036)	(.033)
Singles: Own earnings	.280	.186	.192	.282	.373
wrt *his* parents:	(.030)	(.042)	(.026)	(.082)	(.070)
Couples: Own earnings	.248	.307	.286	.438	.473
wrt *his* parents:	(.010)	(.016)	(.011)	(.040)	(.036)
Couples: Wife's earnings	.121	.182	.146	.203	.109
wrt *his* parents:	(.012)	(.021)	(.016)	(.068)	(.056)
Couples: Combined earnings	.233	.267	.252	.405	.431
wrt *his* parents:	(.009)	(.014)	(.009)	(.047)	(.033)
Couples: Combined earnings	.203		.217		
wrt *her* parents:	(.009)		(.009)		
All: Combined earnings per	.246	.251	.238	.374	.433
adult wrt *his* parents	(.010)	(.014)	(.010)	(.042)	(.031)

Note: "Combined earnings per adult" are set equal to combined earnings divided by two for couples and own earnings for singles. Standard errors are reported in parentheses. Estimates are based on linear specification in the UK and the US, second order polynomial in Finland, and third order polynomials in Denmark and Norway. In nonlinear models, elasticity estimates are evaluated at median parental earnings. Regressions control for father's age and its square.

Table 3: Intergenerational human capital transfers

	Denmark	Finland	Norway	UK	US
A. Intergenerational correlations in educational attainment					
Female cohort members					
Couples: Self vs. *her* father	.270	.270	.389	.377	.420
	(.008)	(.012)	(.007)	(.020)	(.023)
Couples: Self vs. *her* mother	.273	.263	.381	.386	.436
	(.006)	(.011)	(.007)	(.020)	(.023)
Singles: Self vs. *her* father	.239	.267	.349	.375	.334
	(.013)	(.020)	(.011)	(.043)	(.039)
Singles: Self vs. *her* mother	.242	.241	.349	.412	.385
	(.012)	(.019)	(.010)	(.042)	(.038)
Male cohort members					
Couples: Self vs. *his* father	.314	.323	.381	.373	.434
	(.007)	(.011)	(.007)	(.021)	(.022)
Couples: Self vs. *his* mother	.287	.275	.331	.328	.404
	(.006)	(.011)	(.007)	(.021)	(.023)
Singles: Self vs. *his* father	.262	.272	.301	.460	.362
	(.011)	(.019)	(.010)	(.039)	(.037)
Singles: Self vs. *his* mother	.242	.237	.286	.360	.316
	(.010)	(.018)	(.010)	(.041)	(.038)
B. Parental earnings and years of schooling					
Female cohort members					
All: Own education wrt	1.42	.79	1.72	1.79	1.23
log earnings of *her* parents	(.04)	(.04)	(.03)	(.11)	(.09)
Couples: Own education	1.42	.83	1.79	1.80	1.21
log earnings of *her* parents	(.05)	(.04)	(.04)	(.12)	(.11)
Singles: Own education	1.41	.70	1.55	1.80	1.15
log earnings of *her* parents	(.09)	(.07)	(.06)	(.25)	(.17)
Male cohort members					
All: Own education wrt	1.47	1.00	1.47	2.03	1.48
log earnings of *his* parents	(.04)	(.04)	(.03)	(.13)	(.09)
Couples: Own education wrt	1.49	1.03	1.56	2.03	1.39
log earnings of *his* parents	(.04)	(.04)	(.04)	(.15)	(.11)
Singles: Own education wrt	1.36	.83	1.21	1.96	1.72
log earnings of *his* parents	(.07)	(.07)	(.06)	(.28)	(.16)

Table 3, continued.

C. Elasticity of returns-adjusted educational attainment wrt parental earnings					
Female cohort members					
All: Own attainment	.091	.094	.091	.167	.159
wrt earnings of *her* parents	(.003)	(.005)	(.002)	(.014)	(.013)
Couples: Own attainment	.093	.099	.091	.167	.158
wrt earnings of *her* parents	(.004)	(.005)	(.003)	(.016)	(.016)
Singles: Own attainment	.086	.082	.081	.169	.142
wrt earnings of *her* parents	(.007)	(.009)	(.004)	(.033)	(.024)
Male cohort members					
All: Own attainment	.109	.125	.088	.141	.176
wrt earnings of *his* parents	(.003)	(.005)	(.002)	(.014)	(.013)
Couples: Own attainment	.111	.134	.093	.140	.171
wrt earnings of *his* parents	(.004)	(.006)	(.003)	(.016)	(.016)
Singles: Own attainment	.098	.091	.072	.144	.168
wrt earnings of *his* parents	(.006)	(.009)	(.004)	(.033)	(.022)

Note: Standard errors are reported in parentheses. Estimates in Panel B are based on second order polynomial specifications in the UK and the US, third order in Finland, and fourth order polynomials in Denmark and Norway. In Panel C, the dependent variable is the predicted value from a regression of log earnings on 5 (Denmark), 6 (Finland), 12 (Norway), 11 (UK), and 12 (US) indicator variables for educational attainment, estimated from the sample of full-time/full-year workers. Estimates in Panel C are based on linear specification in the UK and the US, second order polynomial in Finland, and third order in Denmark and Norway. Standard errors in Panel C have been adjusted for the fact that the dependent variable is the estimated returns-adjusted attainment. The adjustment is based on the first-step error variance and its correlation with the second-step errors. In nonlinear models, elasticity estimates are evaluated at median parental earnings. Regressions control for father's age and its square.

Table 4: Empirical indicators of marital sorting

	Denmark	Finland	Norway	UK	US
A. Correlations in educational attainment					
Female cohort members					
Couples: self vs. husband	.433	.407	.495	.533	.562
	(.005)	(.011)	(.006)	(.019)	(.021)
her father vs. *her* mother	.476	.447	.489	.571	.627
	(.006)	(.011)	(.006)	(.018)	(.020)
her father vs. *his* father	.153		.244		.359*
	(.009)		(.008)		(.059)
her father vs. *his* mother	.153		.211		.261*
	(.008)		(.008)		(.061)
her mother vs. *his* father	.153		.209		.322*
	(.008)		(.008)		(.060)
her mother vs. *his* mother	.153		.188		.292*
	(.008)		(.008)		(.060)
husband vs. *her* father	.223	.210	.295	.326	.446
	(.007)	(.012)	(.007)	(.021)	(.023)
husband vs. *her* mother	.207	.187	.279	.289	.401
	(.007)	(.012)	(.007)	(.021)	(.023)
Singles: *her* father vs. *her* mother	.446	.446	.490	.559	.615
	(.012)	(.019)	(.010)	(.038)	(.033)
Male cohort members					
Couples: self vs. wife	.438	.407	.462	.545	.560
	(.005)	(.010)	(.007)	(.019)	(.020)
his father vs. *his* mother	.470	.477	.486	.580	.669
	(.006)	(.010)	(.006)	(.018)	(.018)
his father vs. *her* father	.169		.234		.398*
	(.008)		(.008)		(.058)
his father vs. *her* mother	.160		.196		.312*
	(.007)		(.008)		(.060)
his mother vs. *her* father	.152		.200		.363*
	(.007)		(.008)		(.059)
his mother vs. *her* mother	.153		.185		.355*
	(.007)		(.008)		(.059)
wife vs. *his* father	.206	.233	.281	.322	.363
	(.007)	(.012)	(.007)	(.021)	(.023)
wife vs. *his* mother	.216	.191	.260	.247	.337
	(.006)	(.011)	(.007)	(.021)	(.023)
Singles: *his* father vs. *his* mother	.464	.508	.473	.578	.551
	(.011)	(.018)	(.009)	(.035)	(.033)

Table 4, continued.

B. Identification from intergenerational earnings elasticities					
β_f^S / β_m males in couples	.66	.69	.73	.61	.78
C. Elasticity of returns-adjusted educational attainment wrt parental earnings					
Female cohort members					
Couples: Husband's attainment	.084	.101	.085	.100	.185
wrt earnings of *her* parents	(.004)	(.006)	(.003)	(.014)	(.017)
Couples: Combined attainment	.089	.104	.088	.126	.174
wrt earnings of *her* parents	(.004)	(.007)	(.003)	(.016)	(.018)
Male cohort members					
Couples: Wife's attainment	.069	.082	.064	.156	.141
wrt earnings of *his* parents	(.003)	(.006)	(.003)	(.018)	(.016)
Couples: Combined attainment	.095	.116	.082	.144	.160
wrt earnings of *his* parents	(.004)	(.007)	(.003)	(.018)	(.017)

*Computed from PSID data.

Note: Standard errors are reported in parentheses. For model specifications and dependent variable in Panel C, see note to Table 3.

Table 5: Employment status and parental earnings, ordered probit estimates

	Norway	UK	US
A: Model 1			
Females			
Wives of male cohort members			
log earnings of *his* parents	.034	-.065	-.058
	(.006)	(.029)	(.021)
Married female cohort members			
log earnings of *her* parents	.056	-.017	-.040
	(.005)	(.028)	(.021)
Single female cohort members			
log earnings of *her* parents	.073	.230	.081
	(.008)	(.061)	(.026)
Males			
Husbands of female cohort members			
log earnings of *her* parents	.010	.050	.048
	(.003)	(.019)	(.010)
Married male cohort members			
log earnings of *his* parents	.012	.050	.027
	(.003)	(.017)	(.011)
Single male cohort members			
log earnings of *his* parents	.028	.134	.042
	(.006)	(.051)	(.026)
B: Model 2, with child interaction			
Wives of male cohort members			
log earnings of *his* parents	.051	.027	.142
	(.014)	(.060)	(.041)
log earnings * child < 15	-.017	-.101	-.255
	(.015)	(.068)	(.046)
child below 15	-.098	-.274	-.213
	(.011)	(.022)	(.025)
Married female cohort members			
log earnings of *her* parents	.054	.101	.020
	(.011)	(.054)	(.041)
log earnings * child < 15	.007	-.113	-.068
	(.012)	(.063)	(.048)
child below 15	-.075	-.225	-.192
	(.009)	(.023)	(.024)
Single female cohort members			
log earnings of *her* parents	.065	.293	.095
	(.010)	(.084)	(.034)
log earnings * child < 15	.024	-.062	-.039
	(.015)	(.120)	(.052)
child below 15	-.073	-.232	-.086
	(.011)	(.043)	(.037)

Table 5, continued.

	Norway	UK	US
C: Model 3, top vs. bottom of parental earnings distribution			
Wives of male cohort members			
Top:			
log earnings of *his* parents	.171	.053	.029
	(.071)	(.144)	(.093)
log earnings * child < 15	-.105	-.069	-.208
	(.074)	(.161)	(.102)
child below 15	-.111	-.279	-.277
	(.017)	(.027)	(.033)
Bottom:			
log earnings of *his* parents	.029	.099	.247
	(.018)	(.132)	(.077)
log earnings * child < 15	-.020	-.163	-.146
	(.019)	(.154)	(.107)
child below 15	-.091	-.227	-.083
	(.015)	(.028)	(.049)
Married female cohort members			
Top:			
log earnings of *her* parents	.145	.104	.129
	(.061)	(.143)	(.084)
log earnings * child < 15	.019	-.129	-.184
	(.064)	(.160)	(.095)
child below 15	-.096	-.256	-.213
	(.014)	(.028)	(.031)
Bottom:			
log earnings of *her* parents	.024	.023	.057
	(.013)	(.106)	(.088)
log earnings * child < 15	.011	.010	-.003
	(.016)	(.135)	(.105)
child below 15	-.068	-.179	-.192
	(.012)	(.027)	(.051)
Single female cohort members			
Top:			
log earnings of *her* parents	.053	.093	.395
	(.046)	(.204)	(.121)
log earnings * child < 15	.100	.370	-.386
	(.066)	(.280)	(.151)
child below 15	-.050	-.298	-.138
	(.016)	(.060)	(.059)
Bottom:			
log earnings of *her* parents	.050	.193	.057
	(.012)	(.187)	(.062)
log earnings * child < 15	-.014	-.075	.100
	(.020)	(.270)	(.107)
child below 15	-.099	-.181	-.054
	(.016)	(.063)	(.051)

Note: Dependent variable is an employment index set equal to 0 if not employed, 1 if part-time, and 2 if full-time work. Listed coefficients are marginal effects on the probability of full-time work. Effect of indicator variable for dependent child at home is evaluated at mean log parental earnings. “Married” includes cohabiting cohort members. “Top” and “bottom” refer to parental earnings above and below the sample mean. Regressions control for father’s age and its square.

Appendix Table A-1: Nonlinearities in intergenerational earnings elasticity estimates in Nordic samples (evaluated at 10th and 90th percentiles of parental earnings)

	Denmark		Finland		Norway	
	10th	90th	10th	90th	10th	90th
Female cohort members						
All: Own earnings	.047	.283	.097	.252	.105	.251
wrt *her* parents	(.014)	(.020)	(.016)	(.026)	(.012)	(.021)
Singles: Own earnings	.084	.376	.082	.217	.147	.221
wrt *her* parents	(.038)	(.054)	(.033)	(.054)	(.024)	(.046)
Couples: Own earnings	.035	.261	.099	.265	.089	.260
wrt *her* parents	(.015)	(.021)	(.018)	(.030)	(.013)	(.024)
Couples: Husband's earnings	.041	.250	.102	.275	.103	.293
wrt *her* parents	(.014)	(.020)	(.016)	(.027)	(.011)	(.019)
Couples: Combined earnings	.051	.290	.112	.289	.106	.293
wrt *her* parents	(.011)	(.017)	(.013)	(.022)	(.008)	(.015)
Couples: Combined earnings	.051	.312			.118	.332
wrt *his* parents	(.011)	(.017)			(.008)	(.015)
All: Combined earnings per	.061	.298	.105	.262	.127	.275
adult wrt *her* parents	(.013)	(.018)	(.013)	(.021)	(.009)	(.017)
Male cohort members						
All: Own earnings	.080	.377	.129	.369	.146	.365
wrt *his* parents:	(.013)	(.019)	(.016)	(.024)	(.010)	(.018)
Singles: Own earnings	.115	.373	.159	.201	.144	.223
wrt *his* parents:	(.035)	(.055)	(.037)	(.063)	(.021)	(.043)
Couples: Own earnings	.066	.369	.112	.416	.134	.406
wrt *his* parents:	(.013)	(.018)	(.017)	(.025)	(.010)	(.018)
Couples: Wife's earnings	.062	.153	.100	.228	.070	.206
wrt *his* parents:	(.015)	(.021)	(.022)	(.032)	(.015)	(.026)
Couples: Combined earnings	.077	.333	.131	.343	.123	.356
wrt *his* parents:	(.011)	(.016)	(.014)	(.021)	(.008)	(.015)
Couples: Combined earnings	.081	.276			.105	.308
wrt *her* parents:	(.012)	(.016)			(.009)	(.015)
All: Combined earnings per	.088	.344	.141	.312	.134	.318
adult wrt *his* parents	(.012)	(.018)	(.014)	(.021)	(.009)	(.017)

Note: Standard errors are reported in parentheses. Estimates are based on second order polynomial specification in Finland and third order in Denmark and Norway. See also note to Table 2.

Appendix Table A-2: Nonlinearities in parental earnings-years of schooling relationship (evaluated at 10th and 90th percentiles of parental earnings)

	Denmark		Finland		Norway		UK		US	
	10th	90th	10th	90th	10th	90th	10th	90th	10th	90th
Female cohort members										
All: Own educ	.31	2.16	.03	1.51	.41	3.10	.45	3.11	1.01	1.41
wrt *her* parents	(.05)	(.07)	(.05)	(.08)	(.04)	(.07)	(.22)	(.24)	(.11)	(.15)
Couples: Own educ	.26	2.24	.06	1.30	.39	2.79	1.13	2.46	1.03	1.25
wrt *her* parents	(.05)	(.08)	(.10)	(.15)	(.07)	(.13)	(.48)	(.50)	(.21)	(.29)
Singles: Own educ	.45	1.91	.02	1.61	.41	3.23	.25	3.32	.96	1.42
wrt *her* parents	(.10)	(.15)	(.06)	(.09)	(.05)	(.08)	(.25)	(.28)	(.14)	(.17)
Male cohort members										
All: Own educ	.40	2.20	.07	1.86	.38	2.62	.44	3.58	1.11	1.78
wrt *his* parents	(.04)	(.06)	(.05)	(.08)	(.04)	(.07)	(.27)	(.30)	(.10)	(.14)
Couples: Own educ	.40	2.19	.18	1.34	.34	2.13	.87	3.03	.64	2.61
wrt *his* parents	(.05)	(.07)	(.09)	(.16)	(.06)	(.12)	(.54)	(.62)	(.18)	(.28)
Singles: Own educ	.32	2.13	.02	2.00	.39	2.79	.27	3.76	1.18	1.56
wrt *his* parents	(.07)	(.13)	(.06)	(.09)	(.05)	(.08)	(.32)	(.35)	(.13)	(.16)

Note: Standard errors are reported in parentheses. Estimates are based on second order polynomials in the UK and the US, third order in Finland, and fourth order in Denmark and Norway. See also note to Table 3.

Appendix Table A-3: Employment and parental earnings, probit estimates

	Denmark	Finland	Norway	UK	US
A: Model 1					
Females					
Wives of male cohort members					
log earnings of *his* parents	.024	.006	.008	-.042	-.026
	(.003)	(.006)	(.003)	(.028)	(.019)
Married female cohort members					
log earnings of *her* parents	.027	.004	.021	.002	.001
	(.003)	(.005)	(.003)	(.024)	(.018)
Single female cohort members					
log earnings of *her* parents	.060	.022	.042	.172	.086
	(.009)	(.010)	(.005)	(.053)	(.028)
Males					
Husbands of female cohort members					
log earnings of *her* parents	.009	.009	.006	.044	.015
	(.002)	(.004)	(.002)	(.017)	(.007)
Married male cohort members					
log earnings of *his* parents	.013	.014	.007	.035	.025
	(.002)	(.004)	(.001)	(.015)	(.008)
Single male cohort members					
log earnings of *his* parents	.049	053	.026	.139	.039
	(.007)	(.011)	(.005)	(.048)	(.023)
B: Model 2, child interaction					
Wives of male cohort members					
log earnings of *his* parents	.023	.020	.020	.004	.116
	(.006)	(.014)	(.008)	(.059)	(.035)
log earnings * child	.001	-.016	-.016	-.052	-.180
	(.007)	(.015)	(.009)	(.067)	(.042)
child below 15	.065	.001	.011	-.094	-.152
	(.007)	(.013)	(.007)	(.019)	(.022)
Married female cohort members					
log earnings of *her* parents	.025	.003	.018	.079	.009
	(.005)	(.012)	(.006)	(.044)	(.036)
log earnings * child	.001	.001	.004	-.098	-.004
	(.006)	(.013)	(.007)	(.052)	(.041)
child below 15	.032	.030	-.005	-.047	-.118
	(.006)	(.012)	(.005)	(.017)	(.019)
Single female cohort members					
log earnings of *her* parents	.061	.017	.035	.206	.081
	(.011)	(.013)	(.006)	(.072)	(.029)
log earnings * child	-.005	.011	.020	-.055	-.057
	(.017)	(.019)	(.011)	(.105)	(.044)
child below 15	.071	.029	.016	-.078	-.068
	(.013)	(.016)	(.008)	(.040)	(.032)

Note: Dependent variable is an employment indicator set to unity if employed (part or full-time) and 0 otherwise. Listed coefficients are marginal effects. Effect of indicator variable for dependent child at home is evaluated at mean log parental earnings. Regressions control for father’s age and its square.