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Intergenerational Mobility Trends and the Changing Role of Female Labor *

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Abstract

We present new evidence on the existence and drivers of trends in intergenerational income mobility using administrative income data from Scandinavia along with survey data from the United States. Harmonizing the data from Sweden, Denmark and Norway, we first find that intergenerational rank associations in income have increased uniformly across Scandinavia for cohorts of children born between 1951 and 1979. These trends are robust to a large set of empirical specifications that are common in the associated literature. However, splitting the trends by gender, we find that father-son mobility has been stable in all three countries, while correlations involving females display substantial trends. Similar patterns are confirmed in the US data, albeit with slightly different timing. Utilizing information about individual occupation, education and income in the Scandinavian data, we find that intergenerational mobility in latent economic status has remained relatively constant for all gender combinations. This suggests that a gradual reduction in gender-specific labor market segregation, increased female labor force participation and increased female access to higher education has strengthened the signal value that maternal income carries about productivity passed on to children. Based on these results, we argue that the observed decline in intergenerational mobility in Scandinavia is consistent with a socially desirable development where female skills are increasingly valued at the labor market, and that the same is likely to be true also in the US.

Keywords— Intergenerational Mobility, Labor Force Participation   JEL— J62, J21

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1 Introduction

A central question in the social sciences is how the childhood family environment shapes economic fortune in adulthood. If the family environment plays an important role in determining outcomes in adulthood, a common interpretation is that children are not born with equal opportunities in life. In the early economic literature that focused on inferring influence from the family environment through studies of the observational relationship between the earnings of parents and their children, it was highlighted that labor market conditions — determining for instance the return on human capital investments — may play an important role in shaping the persistence of economic outcomes across generations (Becker and Tomes, 1979; Solon, 1999). Accordingly, variation in labor market conditions may be an important determinant of variation in estimates of intergenerational mobility across space and potentially also time (Corak, 2013). While spatial variation in intergenerational mobility is well documented (see e.g. Solon (2002), Chetty et al. (2014a) and Bratberg et al. (2017) for an overview), far less is known about the intertemporal aspect (see Lee and Solon (2009a), Olivetti and Paserman (2015), Chetty et al. (2014b) and Song et al. (2020) for notable exceptions).

Over the past 50 years, labor market conditions for particularly women have changed substantially in all Western economies (Goldin, 2014). Women today are more likely to participate in market work (Olivetti and Petrongolo, 2016) and occupational segregation of men and women has decreased (Blau, Brummund and Liu, 2013). However, often due to constrains on the quality of linked survey data, it has proven difficult for researchers to estimate correlations between a child's income and that of her mother and father separately (Chadwick and Solon, 2002; Björklund, Jäntti and Lindquist, 2009; Blanden et al., 2004). For this reason, the extent to which the secular trend in female labor force participation has affected measures of intergenerational mobility has largely been left unexplored.

In this paper, we address this data issue by turning to the three Scandinavian countries. The high quality of Scandinavian administrative data allows us to follow how the changing patterns in female labor supply may have affected earnings at the individual level for both men and women. Scandinavia provides an ideal setting for understanding how the changing role of women at the labor market can affect intergenerational mobility, as the development toward gender equality precedes that in other countries (Kleven, Landais and Søgaard, 2019). First, we document trends in intergenerational earnings mobility in Sweden, Denmark, and Norway for cohorts of children born between 1951 and 1979 leveraging administrative earnings data from 1968 up until 2017. By applying a unified approach to long panels of full-population administrative data for three different countries, we can investigate the extent to which intergenerational mobility follows similar trends across countries that have been subject to different political and demographic developments, and we can ensure that any differences in findings are not related to the choice of data period or income definition.

Our results reveal a substantial decline in intergenerational mobility across Scandinavia that remains robust across a large set of common empirical specifications. In particular, we show that the results are largely unchanged when studying intergenerational correlations in log earnings rather than within-cohort earnings ranks, and when considering intergenerational correlations in gross or net-of-tax income rather than earnings. This suggests that the observed mobility trends were not driven by simultaneous rank-distorting changes in taxes or transfers across Scandinavia.

Second, after having documented that mobility has followed similar declining patterns across Scandinavia, we then turn our attention towards understanding how changes in female labor market conditions and access to education have affected estimates of intergenerational mobility over time. When breaking mobility trends down by the gender of parents and children, it is evident that earnings of children have become increasingly correlated with maternal earnings over time, while the correlation with paternal earnings has remained close to constant. In the earliest cohorts in our analysis, child earnings — in particular earnings of sons — were virtually uncorrelated with earnings of mothers while exhibiting a clear and economically significant correlation with earnings of fathers. Over time, these parent-specific mobility estimates between children and their separate mothers and fathers have all converged to similar levels. Conducting a similar analysis on Panel Study of Income Dynamics (PSID) survey data from the US, we find similar patterns, albeit with a slightly lagged timing. This suggests that the observed patterns exist outside the Scandinavian countries.

Third, for purposes of decomposing the observed trends in the data, we build a simple model of gender-specific mobility and latent productivity that rationalizes the empirical patterns that we observe in the data. Inspired by some of the building blocks in the model by Becker and Tomes (1979), we assume that income is determined by an inheritable component, say skills or productivity, and a non-inheritable, idiosyncratic determinant. Doing so, we quantify the extent to which the observational trend in intergenerational mobility can be attributed to determinants associated with assortative mating (correlations in parental skills), gender-neutral skills transmission, gender-specific skills transmission and gender-specific return on skill. Calibrating our simple model to country-specific aggregate data, we show that the observational downwards trend in intergenerational mobility is largely compatible with a trend of increasing return on inheritable skills among women relative to men and that this phenomenon explains up to a five rank point increase in the intergenerational rank association in all three countries for cohorts of children born from 1962 to 1979. Most of this trend is driven by mothers rather than daughters. To build intuition for this rise in gender-specific return on skills and the associated implications for mobility, we can think of an early period where a woman with a significant cognitive endowment is more likely to become a secretary than an equally skilled man with similar preferences who sorts into university and obtains a job that requires an academic degree. In this case, the female skills are arguably less well reflected by her earnings, which effectively attenuates the association between her earnings and that of her children. If this segregation becomes smaller over time, the observational relationship between maternal earnings and child earnings will increase. Bridging the model with this simple example, the decomposition suggests that the observed trends in income mobility could simply be an artifact of changes in how women participate in the labor market.

In the final part of the paper, we corroborate this decomposition empirically by showing that gender-specific intergenerational correlations in latent economic status — measured by combining own income, years of education, and occupation using the proxy variable method developed by Lubotsky and Wittenberg (2006) — remained constant over time, or are only weakly increasing. Mobility also remains at a constant level when correlating sons with their maternal uncles, as another way to proxy for maternal skills. Hence, our evidence suggests that the observed trends in intergenerational income mobility can be interpreted as a result of income rank correlations between children and parents — and in particular mothers — becoming gradually less attenuated by frictions caused by gender-specific segregation in the labor market. In other words, our results suggest that intergenerational mobility in income did in fact decline consistently in Scandinavia across cohorts born between 1951 and 1979, but they also suggest that this was almost solely driven by female earnings becoming more reflective of their actual skills. In other words, the return on latent productivity of women has converged towards that of men. Hence, female skills have increasingly become valued in the labor market in the same way as those of males and thus, the observed development in intergenerational earnings correlations can potentially be thought of as a natural implication of a socially desirable development rather than a sign of actually declining equality of opportunity.

With this paper, we make three contributions to the understanding of time variation in intergenerational earnings mobility. The first contribution is related to a series of recent empirical studies from Western economies which indicate that intergenerational mobility may have been declining in the past few decades, in turn suggesting that income inequality to a higher degree persists between generations. The results, however, are not conclusive, and the estimated trends show quantitatively large variation across the existing literature. In particular, Connolly, Haeck and Laliberté (2020), Harding and Munk (2019) and Markussen and Røed (2020) all find that intergenerational mobility has declined rapidly for cohorts of children born between 1960 and 1980 in the US, Canada, Denmark and Norway, respectively. Another set of recent studies, Pekkarinen, Salvanes and Sarvimäki (2017), Song et al. (2020) and Brandén and Nybom (2019) are only capable of detecting weakly declining — or even stable — trends in a similar set of countries. Davis and Mazumder (2020) find declining mobility in the US for children born between 1950 and 1960, while Chetty et al. (2014c) find no change in rank associations between children born in 1971 and later cohorts. In this paper, we provide clear evidence of a uniform decline in intergenerational mobility across Scandinavia for cohorts born between 1951 and 1979. In addition, we show that this trend persists across a range of common empirical specifications in the literature, and that the trends that have been observed in the existing literature are not simply a result of certain empirical specifications or country-specific policies. We also provide suggestive evidence of a similar pattern in the US from panels of linked survey data. To our knowledge, we are the first to estimate and compare trends in relative mobility across multiple countries, thereby providing suggestive evidence of a general phenomenon in Western economies.

The second contribution lies in explicitly documenting substantial gender-variation in mobility trends and showing that gender-specific mobility trends are surprisingly similar across a range of Western economies. A noteworthy strand in the mobility literature has previously highlighted that cross-sectional estimates of intergenerational mobility may differ substantially by gender due to different opportunities for men and women in the labor market (Corak, 2013; Lee and Solon, 2009a). With this paper, we show that mobility has seemingly been stable for father-son relations during the last few decades, while it has been declining considerably whenever female earnings are taken into account — a pattern that, to our knowledge, has only been documented in a Swedish setting by Engzell and Mood (2021) and Brandén and Nybom (2019). These findings suggest that not only do mobility levels vary by gender, but secular changes in gender-specific earnings determinants have also caused trends to differ substantially, in turn causing levels to converge. These patterns are present across all countries in our analysis, suggesting that one explanation why the recent literature has been reaching different conclusions in regards to the existence of mobility trends is choices in regards to dealing with female earnings.

The third and final contribution of this paper is that we provide an explanation for the observed pattern of declining mobility that is compatible with the gender-specific trends that we observe in Denmark, Sweden, and Norway. In recent studies, various explanations for downward trends in mobility have been proposed, none of which are consolidated across countries and specifications. One dominant explanation put forward by Davis and

Mazumder (2020) is that the return on education has increased. Given that education and human capital are significant channels for the transmission of income across generations, this has led to a decline in mobility. A similar explanation put forward by Connolly, Haeck and Laliberté (2020) is that the degree to which women obtain secondary education has increased. Observing that conditional on parental income, income in the child generation is 'boosted' by a higher level of education among parents, the authors conclude that this upward trend in mothers' level of education must have led to a decline in social mobility. However, the underlying mechanism of this relationship remains unclear. Finally, Harding and Munk (2019) suggest other explanations, such as changes in family structure including marital status, assortative mating, and childbearing among women. While the importance of changes in educational attainment has thus already been discussed in the context of mobility trends by Davis and Mazumder (2020) and Connolly, Haeck and Laliberté (2020), our paper is the first to explicitly show a connection to meritocracy and valuation of female skills in the labor market.[1]

The remainder of the paper is structured as follows. Section 2 provides a brief overview of the key features of the Scandinavian welfare states. Section 5 lays out a theoretical framework for the connection between intergenerational rank associations and increasing female labor force attachment, while section 3 presents our data sources and the common methodology used to estimate intergenerational income mobility. In Section 4, we present our main results and discuss the changing role of women in the labor market before we conclude with Section 7.

2 Institutional Context

Denmark, Norway, and Sweden share similar traits in terms of economic development, political culture, and institutions. The welfare state in all three countries is of universal character which means access to social security benefits, health care, subsidized childcare, and tuition-free higher education (Baldacchinoel and Wivel, 2020) for the whole population. In order to finance the provision of these public goods, marginal tax rates at the top of the income distribution, as well as the average tax burden, are substantially higher in Scandinavia than in other developed countries (Kleven, 2014). Employees are to a large degree organized in unions and wages are often collectively bargained (Pareliussen et al., 2018). Historically, all three countries have also been characterized by low levels of inequality and high levels of income mobility, in comparison to other Western countries (Søgaard,

[1] This hypothesis is also put forward, but not further investigated by Engzell and Mood (2021).

2018; Bratberg et al., 2017).

During the second half of the 20th century the role of women in society, and in the labor market, in particular, experienced a "grand convergence" towards the position of men (Goldin, 2014). Contributing to this development were the individualization of the tax system (Selin, 2014), the introduction and expansion of paid paternity leave (Ruhm, 1998), and the expansion of compulsory and higher education (Meghir and Palme, 2005; Black, Devereux and Salvanes, 2005). As a result, female labor force participation increased from the early 1950s and is currently higher in Scandinavia than in most other Western economies.[2] Over the same period occupational segregation strongly decreased, indicating that women increasingly entered occupations that were previously male-dominated. In Figure 1 we provide some descriptive evidence on the development of female labor in the countries under study.

In Panels A and B we show how labor force participation rates of women converged to the male level.[3] In the 1950s, participation rates of mothers were less than half the rate of fathers. This gap had closed almost entirely for mothers of children born in the 1970s and is even less pronounced when we compare sons and daughters of a given birth year. Even though the extensive margin labor supply gap narrowed considerably, women still work substantially more in part-time positions than men (Blau and Kahn, 2017). Panels C and D of Figure 1 show the development of occupational segregation across birth cohorts, capturing the extent to which men and women work in the same occupations. The segregation index is calculated as the difference in the share of all women and men in the labor force who work in a given occupation, summed over all observed occupations. To make comparisons of trends easier, we normalize the index to the base year 1962, allowing for an interpretation of occupational segregation relative to the 1962 level.[4] Evidently, occupational segregation has seen a substantial and persistent decline over time, similar to development documented by Blau, Brummund and Liu (2013) and Blau and Kahn (2017). In contrast to the development of female labor force participation, the decline in occupational segregation is to a larger extent present in the child generation, rather than the parent generation.

[2]See Appendix Figure 4 for a comparison of labor force participation rates across Scandinavia and the United States or Figure 5 for the development of labor force participation as defined in our samples.

[3]The labor force participation rate for men and women is based on the income definitions we use in our later analysis and always relates to the birth year of the child. A person is considered "in the labor force" in a given year if they have annual earnings exceeding the equivalent of 10,000 USD (2017).

[4]The occupational segregation index is defined by three-digit occupation codes for Norway and Sweden and one-digit codes for Denmark due to data limitations.

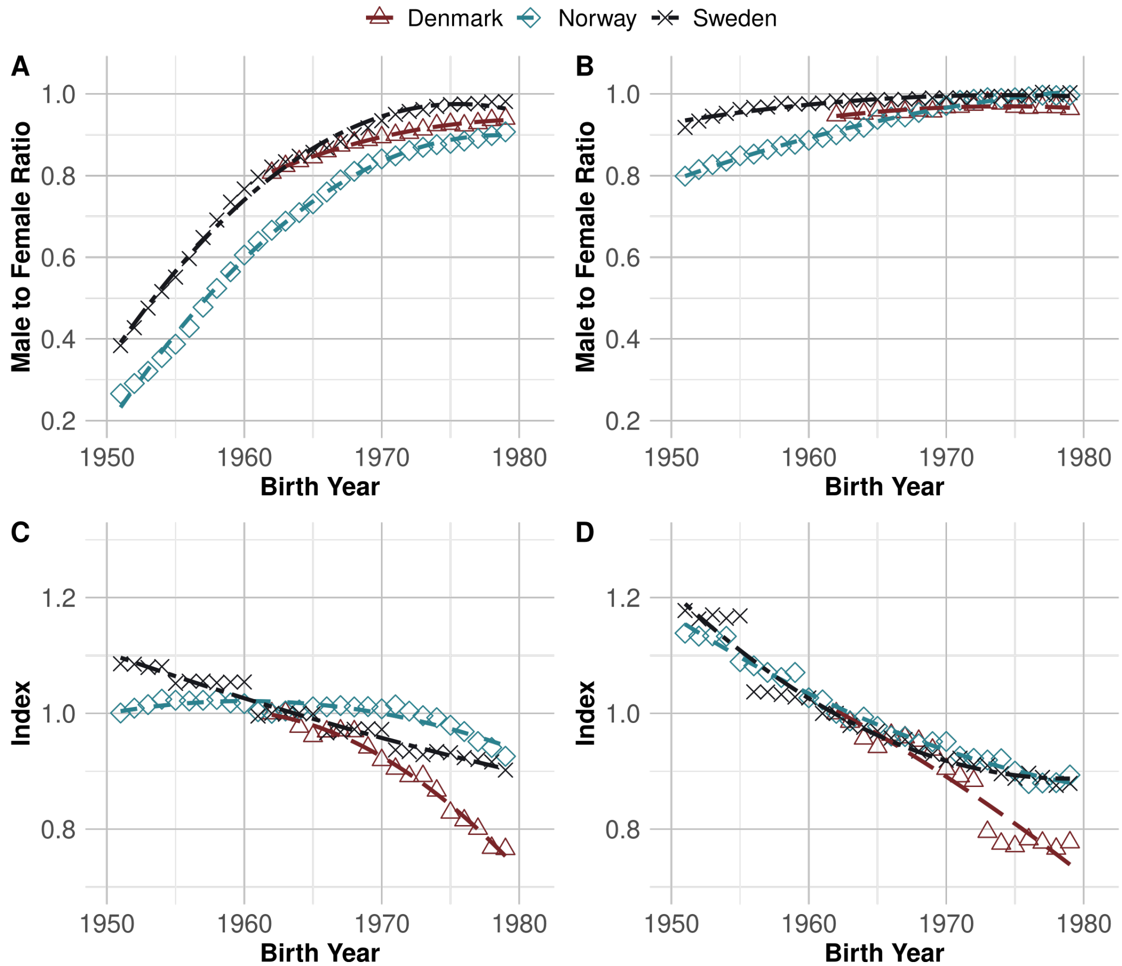

Figure 1: Labor Market Developments.

Note: Panel A and B depict female-to-male ratios of labor force participation in our main samples. Panel A shows participation ratios for parents by birth year of the child. Panel B shows participation ratios for children by birth year. Panel C and D depict an index for labor market segregation separately for parents and children respectively. The index is normalized to the base year 1962.

3 Data and Methodology

3.1 Data

For our main analysis, we rely on register data from Denmark, Norway, and Sweden that cover the whole population of each country from 1968 for Norway and Sweden and from 1980 for Denmark, and up until 2017. The data consist of linked administrative records that provide a variety of information, including birth year, educational attainment, earnings and other income measures, family status, and various demographic variables. Individuals can be linked to their parents. This allows us to create three unique data sets containing all child-parent pairs in a given time frame, with relevant individual income measures. For more details about the registers used, see Appendix A.

Our Scandinavian estimation sample consists of all children born between 1951 (1962 for Denmark) and 1979, who (i) have a valid personal identifier, and (ii) have at least one parent with a valid identifier. As this means that we remove a significant share of immigrants from our samples — in particular in early years — we remove all individuals who are immigrants or are children of immigrants. Sample sizes per birth year are approximately 70,000 child-parent pairs in Denmark, 60,000 pairs in Norway, and 100,000 pairs in Sweden, with variation over time. The results involving US data are based on the Panel Study of Income Dynamics (PSID). The PSID is a nationally representative survey that covers information on employment, income, occupation, education, and family links, starting from 1968. The PSID follows families and individuals across time and has a relatively low attrition rate. With this data, we create a sample of child-parent pairs and measure rank-rank correlations for the US in a comparable, yet more limited, fashion than our analysis on the main Scandinavian samples. In total, the US PSID sample contains about 5,000 child-parent pairs. See e.g. (Lee and Solon, 2009a; Vosters, 2018) for previous applications of the PSID to intergenerational mobility estimation.

The main income specifications are chosen to meet recent standards in the literature.[5] Child income is defined as three-year averages of annual labor income.[6] See Appendix Table 6 for an overview of the earnings components, and how these compare across countries. This is measured at ages 35-37, which balances the needs for a measure of permanent income rank with the needs for measuring child incomes relatively early in order to maximize the number of cohorts that can be included in the analysis (Nybom and Stuhler, 2016; Bhuller, Mogstad and Salvanes, 2017).

Parental income is defined in our main specification as the average of maternal and paternal individual labor earnings, measured as three-year averages of annual income around age 18 of the child. In general, this means measuring the parents' income around age 45, which is considered a good proxy for lifetime income in the literature (Nybom and Stuhler, 2016). In our Appendix, we provide robustness checks to different income definitions for child and parent income variables, such as estimating trends in total factor (gross) income or net-of-tax income. We also evaluate sensitivity to the exact age at which we measure child income.[7]

[5]See e.g. Chetty et al. (2014a) and Lee and Solon (2009b).

[6]Averages are calculated including zeroes. Individuals with one or more missing observations in the years averaged over are dropped from the sample.

[7]Additional results, available upon request, show robustness of the trend to controlling for parental year of birth to address concerns that trends in parental age at childbirth may be driving our main results.

3.2 Empirical Method

In order to measure the intergenerational income persistence, we transform observed income into cohort-specific ranks, as in Dahl and DeLeire (2008) and Chetty et al. (2014a). Using ranks, rather than levels or logs, offers certain advantages in this context. First, estimated rank correlations have proven to be less prone to life-cycle bias than other measures (Haider and Solon, 2006; Bhuller, Mogstad and Salvanes, 2017), and in addition, the use of ranks enables the inclusion of zero incomes. However, in order to ensure that our results are not driven by the rank transformation, we also present mobility trends in intergenerational income elasticities (IGE) in the Appendix.

Trends in intergenerational income mobility are estimated with the following regression, separately by birth cohort and country:

$$\text{Rank}_{it}^{\text{C}} = \alpha_t + \beta_t \text{Rank}_{it}^{\text{P}} + \varepsilon_{it} \text{ with } t = 1951, ..., 1979, \tag{1}$$

where $\text{Rank}_{it}^{\text{C}}$ is the percentile rank of child i's average income at age 35-37 within the distribution of all children born in year t. When we analyze sons and daughters separately, we calculate their income rank separately by gender. $\text{Rank}_{it}^{\text{P}}$ is the percentile rank of the same child's parents' income within the distribution of all parents with children in birth cohort t, averaged over ages 17-19 of the child. The coefficient β_t captures the average cohort-specific parent-child correlation in income ranks, sometimes referred to as the intergenerational rank association (IRA). Lower values of β_t are interpreted as lower rank-associations in income, and thus higher levels of intergenerational mobility.

Intuitively, one can think of the IRA as the correlation in inheritable skills and values that are transmitted across generations. These are attenuated by earnings determinants that cannot be passed on to children, which reduce the signal value of parental income. Such "noise" may stem from individual-specific idiosyncratic shocks to the parental earnings process or time-specific characteristics of the labor market. In particular, changes in the IRA over time are not necessarily driven by the way that skills are transmitted, but rather by the importance of earnings determinants that cannot be passed on to children. In the context of analyzing how changing female labor market participation may have affected the intergenerational association in income, this is a relevant consideration.

3.2.1 Intergenerational Correlation in Latent Economic Status

Income correlations between mothers and their children are complicated by the fact that female labor earnings are a poor measure of their earnings potential during most of our studied time frame. Estimates of the model in Equation 1 for maternal income will not capture the intergenerational relationship between maternal and child labor market skills, which is the main interest in this paper. To fix ideas, denote the underlying relationship of interest as:

$$x_{it}^{*C} = \alpha_t + x_{it}^{*P} + \varepsilon_{it},$$

where x_{it}^{*} is a person's true economic status, unobserved by the researcher. In our setting, it is reasonable to assume that lifetime earnings alone are a good proxy for economic status among sons and fathers, but less so for mothers and daughters. We follow recent work by Vosters and Nybom (2017); Vosters (2018) and Adermon, Lindahl and Palme (2021) and apply the Lubotsky-Wittenberg (Lubotsky and Wittenberg (2006), from now on "LW") method in the intergenerational mobility context. In essence, this method amounts to using a set of proxy variables for latent economic status and weighting these together in an optimal way, given some outcome variable (in our case, child income percentile ranks). The estimation procedure has been shown to minimize attenuation bias among its class of estimators (Lubotsky and Wittenberg (2006), p.552), and only requires the assumptions that 1) the proxies do not have independent effects on the outcome, and 2) that the proxy variables are independent factors. In sum, the proxy variables should be thought of as together representing a single missing variable; economic status.

The proxy variables for parental economic status that we use are income ranks, years of education and occupation: these are denoted $x_j, \quad j \in 1, .., k$. The LW estimator is constructed as follows:

$$\beta_{LW} = \sum_{j=1}^{k} \rho_j b_j, \tag{2}$$

where $\rho_j = \frac{cov(Rank_{it}^{\mathrm{C}}, x_{jit})}{cov(Rank_{it}^{\mathrm{C}}, Rank_{it}^{\mathrm{P}})}$, and the $\mathrm{b_j}$'s are OLS coefficients from a multiple regression of child income rank on the set of proxy variables.

This method has previously been used to estimate mother-child intergenerational income elasticities for Swedish birth cohorts 1951-1961 in Vosters and Nybom (2017). Our application uses the same set of proxy variables and the same methodology, with two exceptions.

First, we calculate year-specific LW estimates, in order to study the time trend in latent economic status mobility. We also extend the analysis to later-born cohorts, which necessitates measuring parental income somewhat earlier in life than in Vosters and Nybom (2017). Second, we make use of the explicit index construction mentioned in Lubotsky and Wittenberg (2006) (p.554):

$$x_{it}^{\rho,P} = \frac{1}{\beta_{LW}} \sum_{j=1}^{k} x_j b_j. \tag{3}$$

We calculate index values for each mother-son or father-son pair using the logarithm of child and parental labor income and then create percentile ranks from these. Finally, we regress the child income ranks on these parental index ranks, for mothers and fathers separately. In order to keep the interpretation as close as possible to that in our main analysis, we assign individuals with zero labor income a token low level of log earnings.[8]

The method described so far addresses the problem of unrepresentative maternal earnings. If trends in intergenerational rank correlations in latent economic status between mothers and sons resemble those found between fathers and sons, it stands to reason that mother-son trends in earnings correlations are attributable to increased economic opportunities of women, and subsequently less attenuation bias in mobility estimates. In order to understand whether daughter-father correlations are subject to the same issue (and bias in estimation), we repeat the above procedure for daughters. Since the LW method deals with measurement error in the right-hand-side (independent) variable, this requires "flipping" the intergenerational model (Eq. 1), and estimating rank associations between fathers and their daughters. This has only minor impacts on the year-specific IRA estimates and does not alter the trend. Apart from this first step, the analysis proceeds in an identical manner as for son-parent estimates.

4 Results

In this section, we first present the trend in child-parent rank associations for Scandinavia. Then we analyze rank associations when we split the sample into sons, daughter, mothers and fathers, and compare our Scandinavian results to suggestive US estimates.

[8]Sensitivity checks show that the exact level of earnings assigned does not alter the conclusions from this analysis. Results are available on request.

4.1 Trends in Intergenerational Mobility

In Figure 2, we present estimates for country-specific trends in intergenerational rank associations in individual labor income. Each point in the graph represents a slope parameter for a cohort-specific regression of Equation (1) with linear trends estimated separately for 1951-1961 and 1962-1979. We provide fitted lines separately to facilitate comparisons between Denmark, Norway, and Sweden for the cohorts where all countries have available data.[9]

From Figure 2, it appears that intergenerational mobility, measured using the IRA, has declined in all three countries, with the fastest rate of decline in Denmark. There, the rank association in income increased by 7.3 rank points (39 %) from 1962 to 1979 — equivalent to an average annual increase of 0.5 rank points. While smaller than in Denmark, the trends in Norway and Sweden are by no means negligible. From 1962 to 1979, the rank association in income increased by 6.4 and 3.4 rank points (38 vs. 19 %) in Norway and Sweden, respectively, yielding annual increases of 0.4 and 0.2. Over the entire range of birth cohorts, from 1951 to 1979, the total change in IRA for Norway is 7.8 rank points (50 %) and 4.6 rank points for Sweden (28 %). Comparing this to...

One may wonder what it actually means, in economic terms, that the rank association in income increased by up to 0.5 rank points per year in Scandinavia. Abstracting from nonlinearities in the relationship between parent and child income ranks, a straightforward interpretation is the following: for two children born by parents in the bottom versus the top percentile, the difference in the conditional expectation of their income ranks as adults increased by 0.5 each year — amounting to as much as five rank points over a decade. Taking the Norwegian results as an example, another interpretation of the observed trends is that in the earliest observed birth cohort, a ten rank points difference in parental income corresponded to an average difference in income ranks of 1.6 between their children. In contrast, the same difference was 2.3 rank points for children born in the latest cohort. While still indicating relatively high levels of mobility by international standards, such changes over relatively short periods of time are by all means economically substantial.

In order to ensure that the trends are robust and reflect structural changes in the economy (as opposed to being something that purely exists within a narrow set of specifications), we document similar trends for a large set of different specifications in Appendix B. Most importantly, we show that the trends remain largely similar when measured in net-of-tax-

[9] In addition to providing graphical illustrations of the trends in the IRA, Appendix Table 10 provides an overview of IRA coefficients for different specifications and tests whether trends are statistically different across countries.

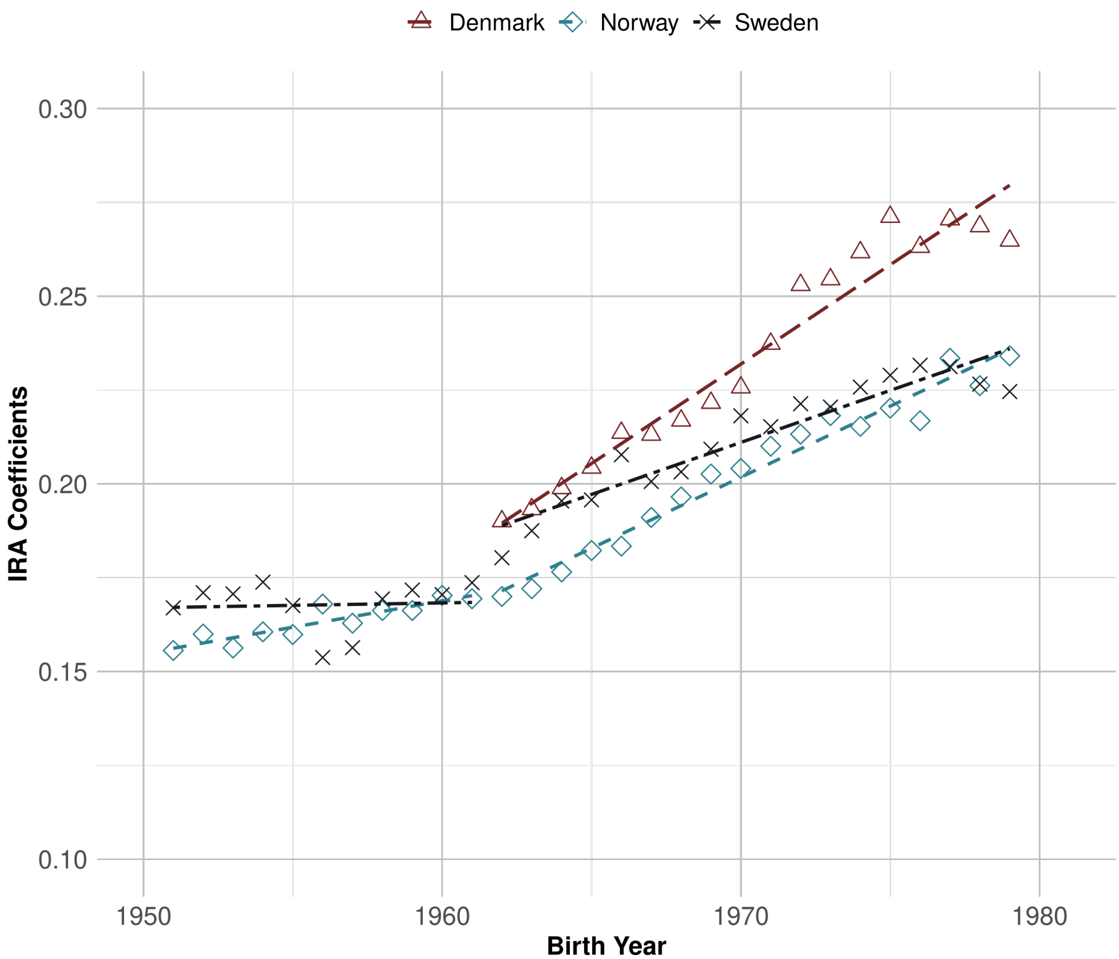

Figure 2: Trends in Intergenerational Mobility in Individual Labor Income.

Note: The figure plots the coefficients for the intergenerational rank association in individual labor income for Sweden, Denmark and Norway over the period from 1951 (1962) to 1979. Each panel shows fitted trend lines separately for the period 1951 to 1962 and 1962 to 1979.

and gross income (Figure 7), and when measuring child income at various ages (Figure 8).[10]

In Figure 9, we restrict the sample to parent-child pairs with labor earnings surpassing 10,000 USD (2017). In other words, we calculate rank associations for the subset of the population that is fully active in the labor market. In general, the mobility trends persist and are similar in magnitude in this specification. However, some cross-country differences are also revealed. Rank associations in Denmark and Norway are markedly lower when excluding non-participating workers from our samples, indicating that intergenerational correlations in labor market participation contribute greatly to intergenerational persistence

[10]We also tested a specification where we rank parental income within both child cohort and their own cohort in order to account for potential changes in life-cycle behavior. The trends remain stable, but the results are not presented in the current version of the paper.

in income — or at least that children of non-participating parents do disproportionately bad in the labor market themselves. In Sweden, on the other hand, the level of mobility largely remains the same after excluding non-participating parents from the estimation sample (panel B), and even increases slightly when excluding both non-participating parents and children (panel C).

4.2 Mobility Trends by Gender of the Child and Parent

Figure 3 presents estimates of country-specific IRA coefficients for pairs consisting of, in turn, sons and fathers (panel A), sons and mothers (panel B), daughters and fathers (panel C), and daughters and mothers (panel D). Each point is again obtained by separately estimating Equation (1) for the respective combination of child and parent.

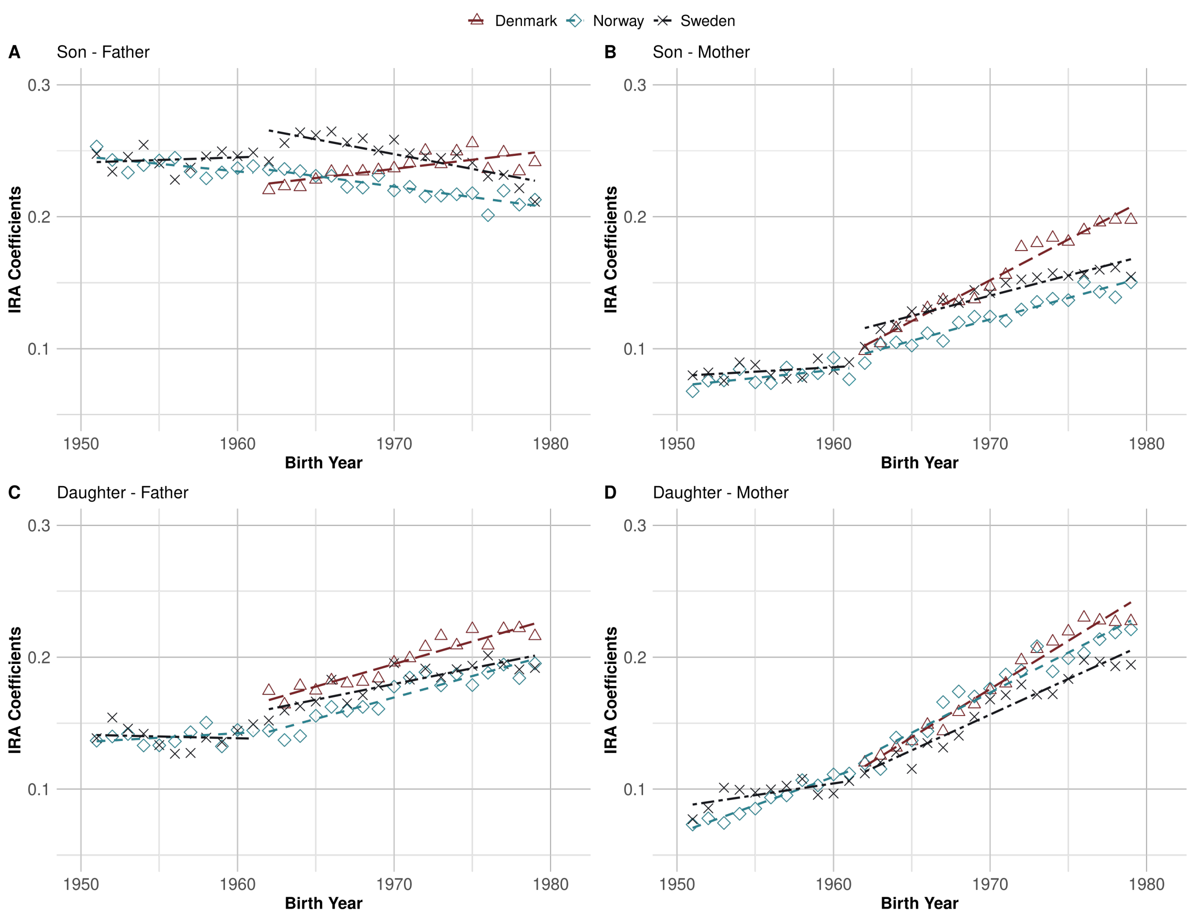

Figure 3: Trends in Intergenerational Mobility by Gender of Parent and Child.

Note: The four panels plot the coefficients for the intergenerational rank association in individual income for Denmark, Sweden and Norway over the period from 1951 (1962) to 1979. Each panel provides estimates separately by gender of the parent and child. Each marker indicates the coefficient of a separate regression and each line indicates fitted trend lines separately for the period 1951 to 1962 and 1962 to 1979.

The four sets of graphs make clear that — at least from 1962 and onward — the trends in IRA for all combinations of child and parent are similar in Sweden, Denmark, and Norway. Estimates for birth cohorts 1951-1979 are strikingly similar in Norway and Sweden. The trend in the IRA between those two countries is statistically indistinguishable for all combinations and years except for the trend in the mother-daughter IRAs after 1961.[11] Across all panels, however, there are also several distinct differences. Most importantly, we see that the rank association between fathers and sons is generally decreasing (Sweden, Norway) or displays a much flatter trend over time, compared to all other graphs that display a clear upwards trend after 1962 (Denmark). The strongest trends in IRAs are found among mother and daughter correlations, closely followed by mother-son correlations. Father-daughter correlations depict slightly weaker trends.

In order to rule out that the mobility patterns that we observe in Scandinavia are just local phenomena, we compute comparable mobility estimates for the US for cohorts born between 1947 and 1983. Results from this exercise are presented in Table 1.[12] Similar to Scandinavia, US mobility trends are steepest for pairs involving mothers and — in particular — daughters. One interesting difference between gender-specific trends in the US and Scandinavia lies in the fact that the upwards trend in mother-son correlations in earnings ranks is not statistically significant in the US. However, the slightly later development toward female labor market equality in the US seems to account for this pattern[13]. Father-son rank associations appear to be relatively constant in the US, suggesting a comparable development as that observed in Scandinavia.[14]

To the extent that father-son correlations represent a credible measure of equality of opportunity that is stable over time, it is hard to argue that an actual decline in opportunity has taken place over time in neither Scandinavia nor the US. Thinking of transmission of skills and values as something passive, this suggests that neither determinants of male income ranks nor the value of skills that are passed on across generations have changed notably over time. Instead, since all combinations of parent-child income that do yield upwards trends in IRAs (panels B-D) involve women, a close-at-hand explanation lies in

[11] In the Appendix Table 10 we provide results for several hypothesis tests regarding the trends and also report slope coefficients for different specifications.

[12] In Appendix Table 7, we provide similar estimates with alternative sample specifications and weighting procedures. In Table 9, we document the cohort-specific number of parent-child pairs used to compute these trends.

[13] The validity of this explanation is confirmed in Table 8. Here, we estimate child incomes around age 30 rather than 36, allowing us to compute gender-specific rank-correlations for cohorts of children born in 1953 to 1989 rather than 1947 to 1983. Looking at this set of children born slightly later, we find that rank-correlations that include mothers exhibit a clear and significant upwards trend.

[14] Recent evidence by Song et al. (2020) supports relatively stable father-son trends for the relevant cohorts in our samples. Moreover, the IRA estimates provided in Song et al. (2020) are similar in magnitude for cohorts between 1950 and 1980.

that women's increasing integration into the labor force has changed the way that incomes are correlated across generations. The fact that the son-mother and daughter-mother IRA is generally not larger than 0.1 among children born before the 1970s suggests that maternal income ranks did not reflect maternal skills well. If work-sharing within the household was more pronounced among parents of these cohorts, and if mothers, irrespective of skills, were more likely to be the ones staying home, this could explain the initial low levels of rank correlations for mothers. As more mothers started participating in the labor force, their income would likely better reflect their skills — skills that they would eventually also pass on to their children.[15] Higher participation and earnings over time among women may also be the key driver of the trend in the daughter-father rank association. The difference in maternal trends between the US and Scandinavia would also be in line with such an explanation, as developments concerning decreases in occupational segregation and increases in female labor force participation started later in the United States and therefore likely impacted mothers only for later born cohorts, while having a potentially larger impact trough changing equality in the labor markets for daughters.

Table 1: IRA Coefficients and Trends (United States)

	Parents	Father		Mother	
	Child	Son	Daughter	Son	Daughter
Pooled IRA	0.317***	0.336***	0.195***	0.097***	0.137***
	(0.017)	(.022)	(0.031)	(0.025)	(0.029)
Trend × 100	0.603***	-0.240	0.980***	0.136	1.047***
	(0.149)	(0.205)	(0.277)	(0.253)	(0.292)
N	5,392	2,272	1,637	2,477	2,205

Note: The table presents estimates of the IRA and linear trends in the IRA separately for different child-parent combinations. Due to the small sample sizes, trends have been estimated directly on the underlying micro data by regressing cohort-specific child ranks on cohort-specific parent ranks interacted with a linear time trend. The trend coefficients and corresponding standard errors have been multiplied by 100 in order to avoid too many digits after the separator. Estimates are based on the full sample of individuals in the PSID born between 1947 and 1983 using PSID sample weights. Standard errors are in parentheses. P-values indicated by * < 0.1, ** < 0.05, *** < 0.01.

[15] The weak link between maternal income and skills for the earliest birth cohorts is also suggested by patterns of assortative mating. Appendix figure 12 shows that maternal income and paternal income become more correlated over time, suggesting that mothers' income becomes more predictive of their true social status. An alternative explanation for the pattern in figure 12 would involve rapid and strong changes in underlying mating patterns which appear to be unlikely given recent research by e.g. Bratsberg et al. (2018).

One last feature of Figure 3 and Table 1 is that incomes appear to be more strongly related for parent-child pairs within gender (i.e., son-father and daughter-mother) than across gender (i.e., son-mother and daughter-father). In fact, while the association in income ranks is generally higher among sons and fathers than among any other combination of genders, the daughter-mother correlation reaches almost the same level towards the end of the considered period in Scandinavia. For the US, we only provide a pooled IRA coefficient due to the small sample. Nevertheless, the pattern that within-gender correlations are stronger than cross-gender correlations and that father-child correlations exceed mother-child correlations is also found in the US sample. This finding could have several reasons, such as intergenerational occupational mobility being lower within- than across gender, and the general tendency of men and women to sort into different occupations (see e.g., Blau and Kahn (2017) for a review on this latter point). Altonji and Dunn (2000) also find within-gender correlations in work hour preferences between parents and children and a recent working paper by Galassi, Koll and Mayr (2021) highlights how employment correlates between mothers and their children, especially so for daughters.

5 Decomposition by Earnings Determinants

In the previous section, we documented that the intergenerational rank association in earnings has increased rapidly in Scandinavia. However, we also showed that there is considerable variation in gender-specific trends across the three countries in our samples. In particular, we described how the intergenerational rank association in earnings between fathers and sons has been fairly stable across Sweden, Denmark, and Norway, while all rank associations involving women have trended upwards. At this point, the exact mechanism driving this upward trend in intergenerational income correlations is unknown. We cannot a priori distinguish a trend in the extent to which skills are transmitted across generations from a trend in the extent to which inheritable skills are valued in the labor market. However, we can use the gender-specific variation in mobility trends along with correlations in parental earnings to quantify the extent to which the trend in the overall intergenerational rank association in earnings is driven by females or higher correlations in skills or latent productivity among parents. In this section, we build a simple model that exactly allows us to quantify the importance of these channels through a simple decomposition exercise.

5.1 Model Setup and Calibration

In our simple framework, individual gender-specific earnings at any time t, y_{it}^k, are determined by two factors; inheritable skills, x_{it}^k, and a non-inheritable determinant ε_{it}^k. This generalizes to all fathers, mothers, sons, and daughters, i.e. all $k \in \{F, M, S, D\}$. Interpreting the setup in the context of a highly simplified version of the frameworks formulated by Becker and Tomes (1979) and Solon (2004), we can think of x_{it}^k as representing an aggregate measure of earnings determinants that can be transmitted across generations such as skills, values, and connections, while ε_{it}^k represents the value of all other income determinants that are uncorrelated to skills that can be transmitted across generations (it may be instructive — yet slightly naïve — to think of this as sheer 'luck').

We assume that inheritable skills in the parental generation follow a bivariate Gaussian distribution on the following form:

$$\begin{pmatrix} x_{it}^F \\ x_{it}^M \end{pmatrix} = \mathcal{N}\left(0, \Sigma_t\right), \; \Sigma_t = \begin{pmatrix} 1 & \\ \frac{\psi_t}{\sqrt{\psi_t^2 + (1-\psi_t)^2}} & 1 \end{pmatrix}$$

where Σ_t denotes the cohort-specific covariance matrix that summarizes the joint mean-zero distribution of parental skills. Standardizing the variance of skills to one, ψ_t is a coefficient that summarizes cohort-specific correlations in parental skills, i.e. it is a simple measure of assortative mating in the model.

We assume that skills are simply transmitted passively from the parental generation to the child generation on the following form:

$$x_{it}^k = \begin{cases} \left(\kappa_t \left[\alpha_t x_{it}^F + (1-\alpha_t)\, x_{it}^M\right] + (1-\kappa_t)\, u_{it}\right)/\Gamma_t, & \text{for } k = S \\ \left(\kappa_t \left[\alpha_t x_{it}^M + (1-\alpha_t)\, x_{it}^F\right] + (1-\kappa_t)\, u_{it}\right)/\Gamma_t, & \text{for } k = D \end{cases}$$

Here, κ_t is a measure of correlation in inheritable skills — or the rate at which skills are transmitted — across generations within a given cohort of children, and α_t is a coefficient that allows the transmission of skills within gender to be stronger than skills across gender. Finally, Γ_t is a trivial scaling coefficient that ensures that the distribution of skills is standard normal.

Individual income is a monotone transformation of a linear index composed of inheritable and non-inheritable determinants:

$$y_{it}^{k}=\hat{F}_{t}^{k}\left(\phi_{t}^{k}x_{it}^{k}+\left(1-\phi_{t}^{k}\right)\varepsilon_{it}^{k}\right),\quad \text{for } k\in\{F,M,S,D\}$$

Here, ϕ_t^F, ϕ_t^M, ϕ_t^S and ϕ_t^D reflect the relative importance of inheritable skills in the income process for fathers, mothers, sons and daughters, respectively. Making the simple assumption that the distribution of non-inheritable determinants can be summarized by a standard normal distribution, $\varepsilon_{it}^m \sim \mathcal{N}(0,1)$, the individual earnings index is standard normal[16].

When measuring gender-specific intergenerational mobility in income ranks, the functional form of the monotone transformation function, $\hat{F}_t^k(\cdot)$, is essentially unimportant; as long as it is monotone in the earnings index, any rank transformation of the earnings index will yield the same result as a rank transformation of earnings. However, in order to find both a pooled measure of child income ranks and a measure of joint parental earnings, such functional form can no longer be disregarded without also disregarding potentially non-negligible differences in gender-specific earnings distributions. Fortunately, we can obtain the functional forms directly from the data. Exploiting the assumed monotone relationship between the earnings index and earnings, we match index ranks to the earnings distribution observed in the data. This allows us to compute pooled earnings ranks in the child generation as well as a measure of joint parental earnings, y_{it}^P, that takes the true earnings distribution into account:

$$y_{it}^{P}=\hat{F}_{t}^{F}\left(\phi_{t}^{F}x_{it}^{F}+\left(1-\phi_{t}^{F}\right)\varepsilon_{it}^{F}\right)+\hat{F}_{t}^{M}\left(\phi_{t}^{M}x_{it}^{M}+\left(1-\phi_{t}^{M}\right)\varepsilon_{it}^{M}\right)$$

Here, $\hat{F}_t^F(\cdot)$ and $\hat{F}_t^M(\cdot)$ are year-specific estimates of the functions that map the earnings index to the earnings distribution observed in the data.

For each country and cohort, we are currently calibrating a vector of seven decomposition parameters, $\left[\begin{array}{ccccccc}\psi_t & \kappa_t & \alpha_t & \phi_t^F & \phi_t^M & \phi_t^S & \phi_t^D\end{array}\right]'$, from only five equations. In order to avoid overidentification, we make two adjustments. First, we set $\phi_t^F = \phi_t^S$ such that the skill importance in earnings for mothers and daughters, ϕ_t^M and ϕ_t^D, must be interpreted relative to that of fathers and sons respectively — i.e. a generation-specific gender bias in the importance of skills for determination of earnings. Secondly, we set both ϕ_t^F and ϕ_t^S equal to 1, thereby effectively pinning down the level around which κ_t trends over time[17]. Finally, the vector of decomposition parameters that are now left for us to calibrate

[16] Through simulations, it can easily be verified that composing the individual income index of two sets of Gaussian components, one inheritable and one non-inheritable, replicates the aggregate functional relationship between parental and child income ranks remarkably well.

[17] The more skills are reflected in earnings, the less skills need to be transmitted across generations in

across countries and years is given by: $\begin{bmatrix} \psi_t & \kappa_t & \alpha_t & 1 & \phi_t^M & 1 & \phi_t^D \end{bmatrix}'$. The calibration procedure is explained in appendix section B.1. Here, we also document the quality of the calibration exercise for each set of country-year combinations of parameters.

5.2 Decomposition

By calibrating the model, we are eventually interested in understanding how country-specific changes in intergenerational mobility can be decomposed into changes in for instance the rate at which inheritable skills are valued among mothers and daughters relative to fathers and sons respectively, and the extent to which parents are mating on skills. Before doing so, we first investigate how the parameters associated with these channels have changed over time in our calibration exercise. Parameters for selected years are displayed in table 2.[18]

Table 2: Decomposition Parameters

	1952			1962			1979		
	SE	DK	NO	SE	DK	NO	SE	DK	NO
ψ_t	0.121	-	0.136	0.272	0.194	0.163	0.235	0.189	0.162
κ_t	0.296	-	0.290	0.255	0.261	0.265	0.255	0.291	0.273
α_t	0.586	-	0.614	0.639	0.587	0.636	0.562	0.553	0.574
ϕ_t^M	0.317	-	0.288	0.398	0.396	0.398	0.636	0.704	0.705
ψ_t^D	0.532	-	0.555	0.620	0.761	0.718	0.982	0.982	0.984

Note: The table presents calibrated decomposition parameters for Sweden, Denmark, and Norway in three selected years. The coefficients have been obtained by matching a simulated version of the aforementioned model to empirical gender-specific IRA-coefficients as well as the relation between father and mother income.

Several noteworthy features of our calibration exercise stand out. First, the decomposition parameters generally evolve very similar across countries. This observation adds credibility to the decomposition approach. In particular, the parameters associated with skill-importance in earnings and mothers, ϕ_t^M and ϕ_t^D, have indeed increased at a very similar

order to obtain a given correlation in earnings over time. Fixing the importance of skills for earnings among males therefore effectively pins down the skill transmission rate across time for a given intergenerational correlation in earnings.

[18]The full set of parameters is available upon request.

pace across all three countries. This, in turn, suggests that female earnings may have become more reflective of inheritable skills in both the parent and child generations. Second, the parameter associated with assortative mating is surprisingly constant across all three countries (at least from the early 1960s and onward) in spite of strongly increasing associations in maternal and paternal earnings over time. This discrepancy may indeed be an implication of the fact that maternal earnings have become more reflective of maternal inheritable skills, thereby mechanically increasing the observational correlation in father and mother earnings for a given correlation in skills. Third, within-gender correlations in skill do in fact seem to be stronger than cross-gender correlations in skills — α_t is approximately 0.6 across all countries but slowly declining from the early 1960s and onward. Finally, the coefficient associated with non-gendered skill-transmission is slowly downwards trending in both Sweden and Norway, while exhibiting a weak but robust upwards trend in Denmark.

While the trends in decomposition parameters are generally similar across countries, the direction and extent to which their changes may affect the intergenerational rank association in earnings between parents and children is a priory unclear. In order to decompose changes in this main parameter into effects associated with changes in the modeling parameters, we computer 'counterfactual' income associations holding one parameter fixed over time, while allowing the aggregate gender-specific income distributions that were obtained from the data to vary over time.

We do this by first defining $\tilde{\beta}_t$ as the rank association between joint parental and child earnings obtained from the calibrated set of parameters in the model stated above subject to a simulated set of data such that $\tilde{\beta}_t \equiv \beta\left(\psi_t, \kappa_t, \alpha_t\phi_t^M, \phi_t^D\right)$. Then we define $\tilde{\beta}_{t,\underline{t}}^b$ in a similar fashion, but we fix parameter $b_t \in \left(\psi_t, \kappa_t, \alpha_t\phi_t^M, \phi_t^D\right)$ to the calibrated value in period $\underline{t}$ such that for instance $\tilde{\beta}_t^{\psi_{\underline{t}}} \equiv \beta\left(\psi_{\underline{t}}, \kappa_t, \alpha_t\phi_t^M, \phi_t^D\right)$. Finally, the part of the trend in $\tilde{\beta}_t$ that can be attributed to parameter b is simply the difference in trend between $\tilde{\beta}_t$ and $\tilde{\beta}_t^{b_{\underline{t}}}$, while the part of the actual trend in β_t that can jointly be attributed other factors than decomposition parameters and changes in the aggregate gender-specific income distributions is the difference in trend between β_t and $\tilde{\beta}_t$. The results from this exercise are documented in table 3.

Table 3: Decomposition by Parameters

	1952-1961			1962-1979		
	SE	DK	NO	SE	DK	NO
Trend in β_t	0.013	-	0.140	0.277	0.530	0.379
Trend in $\tilde{\beta}_t$	0.067	-	0.151	0.250	0.515	0.338
Due to ψ_t	0.186	-	0.004	-0.052	-0.011	-0.001
Due to κ_t	-0.342	-	-0.165	-0.019	0.246	0.036
Due to α_t	0.005	-	0.005	0.002	0.001	0.001
Due to ϕ_t^M	0.048	-	0.114	0.142	0.214	0.164
Due to ψ_t^D	0.029	-	0.139	0.134	0.062	0.062

Note: The table presents trends in observational IRA coefficients, β_t, in the three countries as well as trends in IRA coefficients obtained from the calibrated models in the three countries, $\tilde{\beta}_t$. The contribution from each parameter is computed as the difference in $\tilde{\beta}_t$ that is obtained from holding one calibrated parameter fixed at a time. The sum of contributions from each parameter need not sum to the trend in $\tilde{\beta}_t$ as part of the trend will be driven by changes in the scale of gender-specific income distributions which is not modeled.

As the observational rank associations in earnings did not exhibit any clear, joint upwards trend for cohorts born between 1952 and 1961 across Sweden and Norway, there is not much to be explained by the decomposition parameters. However, there are certain noteworthy patterns in this period. In particular, the parameter associated with non-gendered skills transmission, κ_t, contributes negatively to the IRA over time, while the opposite is the case for the parameters associated with the extent to which female earnings are reflective of parental skills, ψ_t^M, and ψ_t^D. This suggests that skills transmission may in fact have declined over time, thereby pushing mobility up, but this effect could have been mitigated by changes in the extent to which females have inheritable skills valued in the labor market

From 1962 to 1979, IRA's are, in turn, increasing uniformly across Scandinavia, and this pattern seems to be captured well by the simple decomposition model. While both ψ_t and α_t generally seem to be unimportant contributors to mobility trends in the given period, κ_t seems more important — at least in Denmark, where close to half of the observed trend in mobility. In both Sweden and Norway, however, the importance of κ_t is negligible. Finally, changes in the extent to which female earnings — and particularly maternal earnings — are reflective of inheritable skills seem to be important drivers of downwards trends in mobility across Denmark, Norway, and Sweden. These effects jointly contribute to a yearly increase in the earnings IRA of between 0.23 and 0.28 rank points in all three countries, amounting to a total increase in the IRA of up to 5 rank points over the period. In the next section, we show that it is indeed plausible to interpret this phenomenon as female earnings becoming more reflective of inheritable skills over time.

6 Mobility Trends in Latent Economic Status

In this section, we present estimates of trends in intergenerational mobility in latent economic status in Scandinavia and compare these with trends in the intergenerational rank association in labor earnings. First, we show trends in the IRA and Lubotsky-Wittenberg (LW) coefficients for sons and their fathers and mothers, allowing us to isolate to what extent increased labor force attachment among mothers drives the observed trends. Second, we provide estimates of the trend in daughter-father IRA and LW coefficients, allowing us to investigate the extent to which changes in occupational segregation among the child generation influence trends in intergenerational mobility.[19] Table 4 provides estimates of the trend in IRA and LW estimates for the years 1962 to 1979, separately by country. We also report the difference between the trend estimates, which tests whether trends in intergenerational mobility are statistically distinguishable between the IRA and LW approaches. For a visual representation of the trends and corresponding estimates see Appendix Figure 13.

The son-father trends obtained from the LW method correspond well to the son-father IRA trends as suggested by Panel A in Table 4. Even though there are small differences between the estimated trends across all countries these differences are not statistically distinguishable from zero, indicating that son-father trends for the IRA and LW coefficients are similar. For Norway and Sweden IRA and LW, trends are negative indicating a development towards increased mobility, while Denmark's decline in mobility is supported by both the IRA and LW methods. Our interpretation of this similarity in estimated trends is that especially paternal, but also sons income ranks, provide a reasonably stable measure of socioeconomic status and therefore do not exhibit a large difference in trends compared to the LW trend.

[19] For the daughter-father correlation, we exchanged the dependent and independent variables of equation 1, such that we can account for the latent economic status of the daughter. See Section 3.2.1 for more details.

Table 4: Comparison of Trends 1962 - 1979

	Denmark	Norway	Sweden
Panel A: Son - Father			
IRA	0.1385	-0.1598	-0.2243
	(0.0349)	(0.0222)	(0.0605)
LW	0.1504	-0.2062	-0.1898
	(0.0239)	(0.0350)	(0.0613)
Difference	-0.0118	0.0464	-0.0346
	(0.0423)	(0.0414)	(0.0861)
Panel B: Son - Mother			
IRA	0.6186	0.3244	0.3069
	(0.0256)	(0.0262)	(0.0408)
LW	0.2994	-0.1200	0.0175
	(0.0353)	(0.0273)	(0.0495)
Difference	0.3192	0.4444	0.2894
	(0.0436)	(0.0379)	(0.0642)
Panel C: Daughter - Father			
IRA	0.3416	0.3247	0.2388
	(0.0309)	(0.0318)	(0.0339)
LW	0.0658	-0.0385	-0.0897
	(0.0324)	(0.0314)	(0.0201)
Difference	0.2758	0.3632	0.3285
	(0.0447)	(0.0447)	(0.0394)

Note: IRA indicates linear trends estimated through all coefficients of the intergenerational rank association. LW specifies linear trends estimated through all coefficients obtained from applying the Lubotsky-Wittenberg method. The trend coefficients and corresponding standard errors have been multiplied by 100 in order to avoid too many digits after the separator. Difference indicates differences between LW and IRA trends and tests the null-hypothesis of equality in trends between the IRA and LW coefficients. Heteroskedasticity robust standard errors are in parentheses.

Panel B presents estimates of the son-mother IRA and LW trends. Noticeably, the trends in the IRA are significantly steeper for all three countries, with Denmark depicting the largest increase in IRA coefficients over the respective time period. In comparison the trends for the LW coefficients are less steep and in the case of Norway and Sweden even negative, again suggesting a development towards increased mobility in those two countries while providing evidence for a less pronounced decline in mobility for Denmark. Moreover, the difference between the trends of the IRA and LW coefficients is statistically meaningful and different from zero, with a very similar absolute difference across all three countries, which indicates that applying the LW method to capture latent economic status mitigates attenuation similarly across all countries. In addition, the estimated LW trends are a lot closer to the trends obtained in the son-father case. Evidently, when using mothers' years of education and occupations - rather than just labor earnings - to proxy for their latent economic status, the extent to which male children achieve similar economic success as their parents has remained relatively constant over time.

In Panel C of Table 4, we additionally present the comparison between trends in the LW and IRA coefficients for daughters and fathers. Similar to Panel B the trends in the IRA are significantly steeper than what the LW trends suggest. The differences between IRA and LW trends by country are almost identical across countries, suggesting that the use of additional proxy variables in the LW approach captures latent economic status in a similar fashion across all three countries. For Denmark, the adjusted trend still indicates that over time mobility in economic status decreases, however at a significantly lower rate, in Norway the relationship is stable, while in Sweden daughters experience a small increase in mobility over time.

In summary, Table 4 provides three important takeaways. First, in all three countries trends between sons and fathers are similar for the IRA and the LW approach, indicating that the IRA reasonably captures actual developments of intergenerational mobility in latent economic status. Second, trends in the son-mother and daughter-father IRA appear to overestimate declines in mobility and, third, differences in trends between the IRA and LW method are comparable across countries. In addition to the comparison of trends, the levels of the son-father, son-mother, and daughter-father LW coefficients are more similar to the IRA coefficients of son-father pairs which is what would be expected when accounting for attenuation in the coefficients and is also supported by findings in e.g. Vosters and Nybom (2017). Estimating rank associations in latent economic status by birth cohort shows that over time, father-daughter correlations have remained roughly constant at a level just below 0.3. The transmission of economic potential between parents and their female children, as well as their male children, has thus seen little change across birth cohorts from 1962 to 1979. That girls are not over time increasingly "invested in" by their parents might reflect the particular setting, with schooling relatively equally distributed among boys and girls already among individuals born in the 1950s. On the other hand, the fact that father-daughter correlations are as high as the father-son ones suggests that whatever skills relevant to economic success are transmitted between parents and their children, these are gender-neutral.

By estimating correlations in "latent economic status" rather than observed income, our goal is a measure that better approximates the transmission of income-generating skills between parents and their children. One could argue, however, that occupational and educational choices are so strongly correlated with realized income, that the approach adds little by way of intuition. This would also invalidate the primary assumption behind the LW approach — that of independence between the proxy variables. To corroborate the

LW results, we also estimate the intergenerational rank association in labor income between sons and their maternal uncles. Given a constant level of brother-sister correlation in earnings potential, this estimated trend captures changes in the importance of parental earnings potential for child outcomes.[20] Using observed skills of maternal uncles to proxy for unobserved female values is a strategy previously used by e.g. Grönqvist, Öckert and Vlachos (2017). Due to high data demands needed for parental generation sibling links the sample size used to estimate the IRAs is relatively low, particularly for the earliest birth cohorts. Appendix Figure 11, Panel A, presents the results, which reveals a constant level of rank associations over time. Panel B shows the original mother-son associations for comparison, and in Panels C-D, the same results are shown for daughters and maternal uncles. Daughter-uncle trends are substantially flatter than daughter-mother trends, indicating that a certain part of the mother-daughter trends is driven by the mothers. However, the remaining IRA trend shows that increased labor force attachment by daughters over time also contributes to the observed mobility trend.

To further clarify the intuition behind our central theme, Appendix Figure 12 provides evidence that maternal "skills" and income are virtually unrelated in the early period of our sample. The figure plots maternal income ranks by ventiles (5 percentile rank bins) of the paternal income distribution, for 1951, 1962, and 1979 samples, respectively. Mothers of the 1951 cohort evidently earned the same (low) level of income irrespective of their husband's earnings: the average rank hovers around 50 across the whole of the fathers' income distribution. In 1962 — and even more so in 1979 — however, maternal income rises almost monotonically in paternal income. Assuming a time-invariant pattern of assortative mating, this is evidence favoring our hypothesis of female incomes better reflecting underlying skills over time.[21].

7 Conclusion

In this paper, we have documented trends in intergenerational income mobility in Denmark, Norway, and Sweden, for children born in 1951 (1962) to 1979. Harmonizing data and definitions, we have shown that the intergenerational rank association between parents and children in individual income has increased significantly in all three countries.

[20]Using Swedish data, Björklund, Jäntti and Lindquist (2009) show that brother correlations in income remain similar for cohorts born between 1953 and 1968.

[21]Whether assortative mating in income and education has declined or inclined over time is a topic of recent research by e.g. Eika, Mogstad and Zafar (2019) and Bratsberg et al. (2018), with the latter suggesting that trends in assortative mating by social class have stayed considerably more constant than assortative mating by education.

These trends are robust to using different types of income measures, as well as to restricting the analysis to labor market active individuals. Splitting trends by gender of parents and children, son-father correlations exhibit the weakest trend in all three countries, whereas all correlations involving mothers and daughters increase over time. The strongest trend is found between mothers and daughters. To extrapolate our findings to countries outside of Scandinavia, we show that similar patterns can be found for US parent-child pairs from the PSID. In line with the Scandinavian results that are based on more detailed data of higher quality, we find a similar, but delayed, development in changes of the IRA in the US. Our results suggest that rising female labor supply and participation results in higher child-parent rank associations through better manifestation of maternal skills in income, such that the intergenerational correlation in "potential income", or latent economic status is revealed. In other words, the fact that maternal economic status was poorly reflected in maternal income among early cohorts of our sample caused rank associations between child income and joint parental income to be an attenuated measure of mobility of economic status or opportunity. Over time, as female labor supply and participation has increased, this attenuation has declined accordingly.

Our results clearly point to the importance of accounting for changes in female economic status when estimating trends in intergenerational mobility. The interpretation that higher rank associations in income or earnings between children and parents reflect a lower degree of social mobility or equality of opportunity is not always easily applicable when labor market conditions change substantially. In particular, our findings suggest that women's income over time is to a larger extent determined by their earnings potential, meaning that the traits and norms that women inherit from their parents are also better reflected in their income. While such a development must be seen as a necessary side-effect of increased gender equality in the labor market, it is a priori unclear whether such development should be seen as a reduction or advancement in equality of opportunity.

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Appendix

A Data Registers and Variable Definitions

A.1 Denmark

The Danish income registries start in 1980 and contain detailed information on the individual income composition of Danish adults. The registries are based on information from the Danish tax authorities and supplemented with information from other Danish authorities, including unemployment insurance funds and the municipalities.

The measure of labor income that is being used in this paper consists of wage payments (incl. perks, non-taxable wage payments, stock options, and more) and any net surplus from own, private company. Gross income is equal to labor income, transfers, property income, and any other non-classifiable income that the individual may have received throughout the year. Net-of-tax income is finally equivalent to gross income net of all taxes that have been paid to either the government, municipalities, or other public authorities. Individuals with no parents in the sample (generally people who moved to Denmark, whose parents have moved abroad, or whose parents do not live anymore) are naturally dropped from the sample.

When constructing household income measures, individuals are being linked to their spouses. In the Danish sample, a spouse is generally defined by marriage, registered partnership or simply from the fact that they are registered as a cohabiting couple. Matching individuals to spouses as well as parents is based on the population registries of Denmark.

A.2 Norway

For the Norwegian part of the analysis, we are able to include birth cohorts from 1951 onward. We combine information from the central population registry with information about income and earnings from the tax registry. Income data in Norway is available from 1967 to 2018. Labor income, which includes payments related to employment, including overtime pay, taxable sickness, parental leave, short-term disability, and rehabilitation benefits, is top-coded for a few years in the 1970s at the maximum amount for contributions to the national social security scheme (folketrygden). Gross income is the sum of labor income and taxable and non-taxable transfers and income from capital. Disposable income is defined as gross income minus taxes and is also sometimes referred to as net-of-tax income. The definitions mentioned change to some degree over time due to reforms of the benefit, insurance, and tax system. For the net-of-tax and gross income variable, the data series ends in 2014, which is why these income measures are then constructed from more detailed income data only available from 1993. Spouses are linked through their personal identifiers and include married couples as well as couples in civil unions.

The occupation data used for implementing the method proposed by Lubotsky and Wittenberg (2006) is pooled from matched employer-employee data (Registerbasert sysselsettingsstatistikk) available annually starting with the year 2000. In addition occupation data from the censuses 1960, 1970, and 1980 are added. To achieve a comparable classification of occupations we use the STYRK-08 one-digit code to group individuals into broad occupational groups (see Table 5). Individuals are assigned the occupation they have at age 36. In cases where this is not possible we use the closest applicable occupation we observe in the data. Due to the long break in occupational data between the 1980 census and the start of the employer-employee data, there might be some differences in the age at which we observe occupations for individuals that are also connected to the relevant birth year.

The educational data for the LW method is also pooled from different registries. Most individuals we observe are included in the national education database available from 1970. These data include variables for the highest achieved education of all individuals which we can link via personal identifiers. For individuals who are not included in the national education database, we try to obtain information about their educational attainment via census data from 1960, 1970, and 1980.

A.3 Sweden

The Swedish Income and Taxation registry starts in 1968 and holds official records of income for all individuals with any recorded income. In general, it contains all earned income from employment or businesses, capital income, taxable (mostly social insurances), and non-taxable transfers (social welfare, educational grants, child benefits, etc.). Identifiers for biological or adoptive parents are linked to the child identifier through the multigenerational register. Households are constructed by linking individuals (children, mothers, and fathers) to their spouses. This is available only for married couples (and those in registered partnerships) and thus excludes households formed by cohabiting partners.

Data on occupations are taken from two sources. First, the population censuses (Folk- och bostadsräkningarna) contain occupational codes corresponding to the ISCO-58 classification system. This information is available from 1960, and then every five years between 1970 and 1990 for the whole adult population. Individuals without an occupational code can be either classified as "undefined" or have a missing value. In our applications, both these are coded as missing. The census data are used to infer occupations for all parents in our Swedish sample, and we assign each parent an occupational code from the census closest in time to when the child is 18 years old (for example, a mother with a child born in 1951 will primarily be assigned an occupational code from the 1970 census, and occupations for fathers with children born in 1975 will be taken from the 1990 census). If no occupations is observed in this year, we search iteratively through the second and third closest waves, and so on. Parents who are missing an occupational code after this procedure, and who are at least 18 years old in 1960, are assigned occupations from that year's census. This mainly serves to capture occupations of women who are out of the labor force continuously after the birth of their first child; about 6.5 percent of the mother sample (3 percent of the fathers).

Occupational codes for the child generation are taken from the 1990 census for individuals born in the years 1951-1955, and from population register data for those born between 1956 and 1979. The population occupations register uses an adapted version of the ISCO-08 classifications, called SSYK 2012, and is available in our data for the years 2012-2017. As a result, the age at which occupations are observed among the child sample varies between 35 and 56, which might induce noise in between-birth cohort comparisons. On the other hand, this age span corresponds to prime working age, and occupational choice is relatively constant, especially given the broad classes we use in our analysis.

The highest attained level of education is observed in the 1970 census, and in the annual population registers that start in 1990. Each person is assigned the level of education that he or she displays in the year closest in time to when income is observed (age 36 for children; age 18 of the child for the parents). Years of education is then inferred from these categorical data (e.g. completing a three-year secondary education program is coded as twelve years of education, or eleven years if the person completed primary school when it was still only seven years in duration).

B Additional Figures and Tables

Code	Definition
Norway	
0	Armed forces and unspecified
1	Managers
2	Professionals
3	Technicians and associate professionals
4	Clerical support workers
5	Service and sales workers
6	Skilled agricultural, forestry and fishery workers
7	Craft and related trades workers
8	Plant and machine operators and assemblers
9	Elementary occupations
Sweden	
1	Professional work (arts and sciences)
2	Managerial work
3	Clerical Work
4	Wholesale, retail and commerce
5	Agriculture, forestry, hunting, and fishing
6	Mining and quarrying
7	Transportation and communication
8	Manufacturing
9	Services
10	Military/Armed Forces
Denmark	
0	Military work
1	Management work
2	Work that requires knowledge at the highest level in the area in question
3	Work that requires knowledge at intermediate level
4	Ordinary office and customer service work
5	Service and service work
6	Work in agriculture, forestry and fisheries
7	Craft and related trades workers
8	Operator and assembly work, transport work
9	Elementary occupations

Table 5: Occupation Classification by Country

Note: Occupational categories for Norway are assigned using the STYRK-08 classificaiton provided by SSB. For Sweden the classification follows SSYK-2012 similar to Vosters and Nybom (2017). For Denmark, we use the first integer from the Danish ISCO classication (link). In the Danish case, note that this variable is not available for all years in the data. For this reason, we generate it from a set of other available occupation related variables. Code is available upon request.

Table 6: Overview Income Definitions by Country

		Denmark	Norway	Sweden
1	Salary	taxable salary incl. fringe benefits, tax-free salary, anniversary and severance pay and value of stock options	all payments related to employment including overtime pay	all payments from employment
2	Net Profit	net profit from self-employment incl. profit of foreign company and net income as employed spouse	net income from self-employment and income from other businesses	net profit from self-employment, income from other businesses
3	Transfers	cash benefits, unemployment insurance benefits, sickness benefits, unemployment benefits, pensions, child allowance, and more	taxable sickness benefits, parental leave benefits, unemployment benefits, short-term disability payments, rehabilitation benefits	sickness benefit from employer (sjuklö), value of e.g. car, travel expenses (förmånsvärden)
		Earnings/Labor Income = Combination of 1+2+3		
4	Transfers	cash benefits, unemployment insurance benefits, sickness benefits, unemployment benefits, pensions, child allowance, and more	taxable transfers: benefits from the national insurance scheme (disability insurance, pensions, etc.) non-taxable: child benefits, housing allowance, scholarships, parental leave benefits, social assistance payments	taxable: social insurances (unemployment, parental leave etc), private pension income, stipends etc. non-taxable: pensions and annuities housing support child support social welfare alimony conscript support grants and loans for students
5	Property Income	capital and wealth income excl. calculated rental value of real estate	gross interest income, dividend income, return on life insurance, net realised capital gains (e.g. shares, house, land), other capital income (taxable rental income)	capital income (gross pre-2004, net post-2004) and after-tax rental income
		Gross Income = Combination of Earnings +4+5		
6	Other Income	other non-classifiable income		
7	Taxes	taxes on earnings, wealth taxes, property value tax, tax on share dividends/gains and more	taxes, maintenance paid, mandatory insurance premia	all taxes
8	Negative Transfers			repayments of study loans, paid alimony
		Disposable Income = Combination of Gross Income - 7+10		

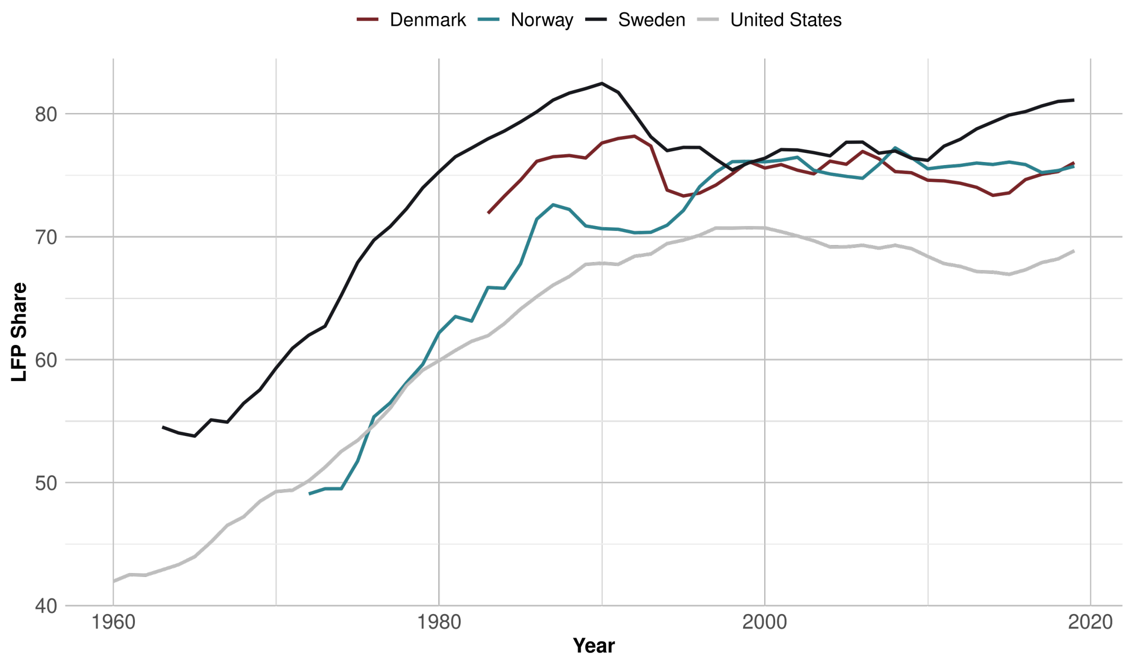

Figure 4: Labor Force Participation Rate.

Note: The figure depicts the labor force participation rates of women aged 15 to 64 for Denmark, Norway, Sweden and the United States. The data was obtained from the OECD (2021) and covers all years available for the respective countries.

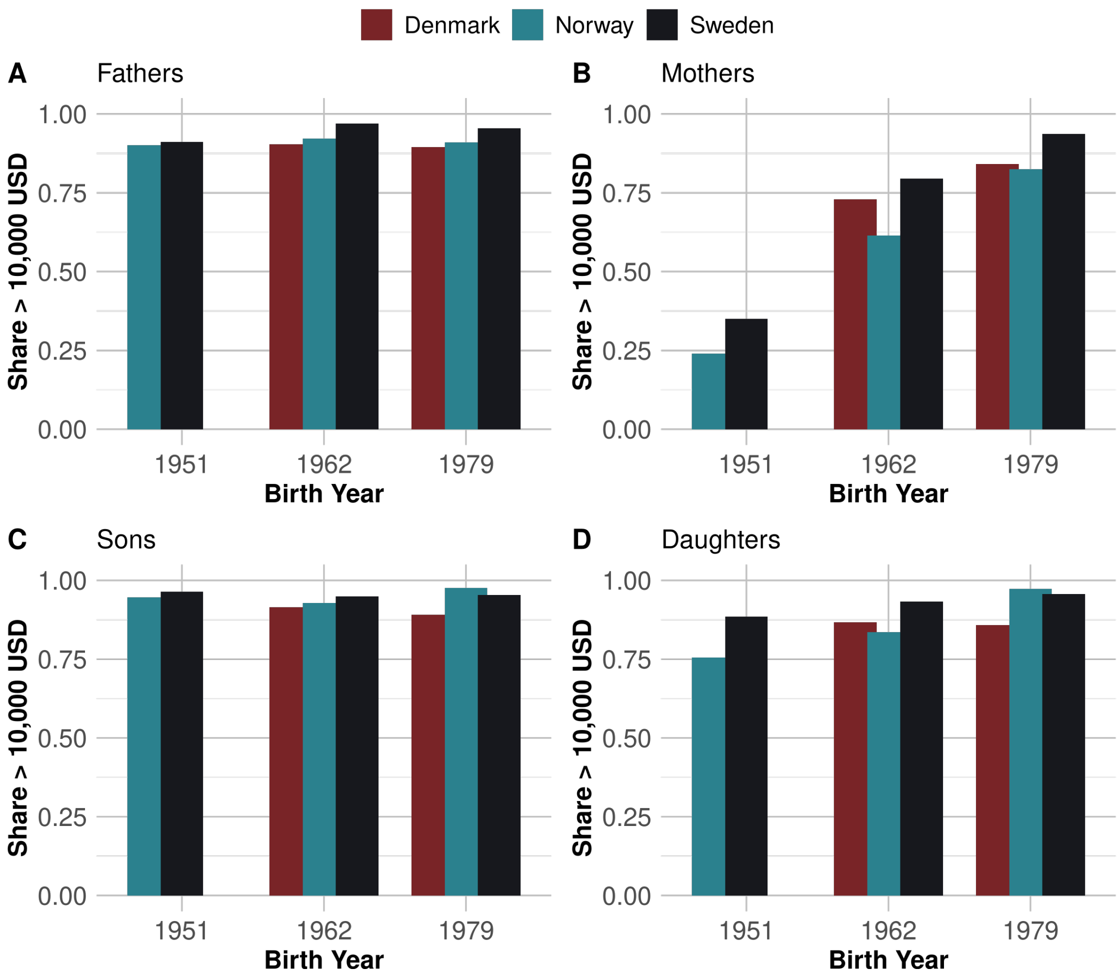

Figure 5: Labor Force Participation.

Note: Each panel depicts shares of individuals with labor income exceeding 10,000 USD (2017) in Sweden, Denmark and Norway for the years 1951, 1962 and 1979. Panel A provides information for fathers, panel B mothers, panel C sons and panel D daughters.

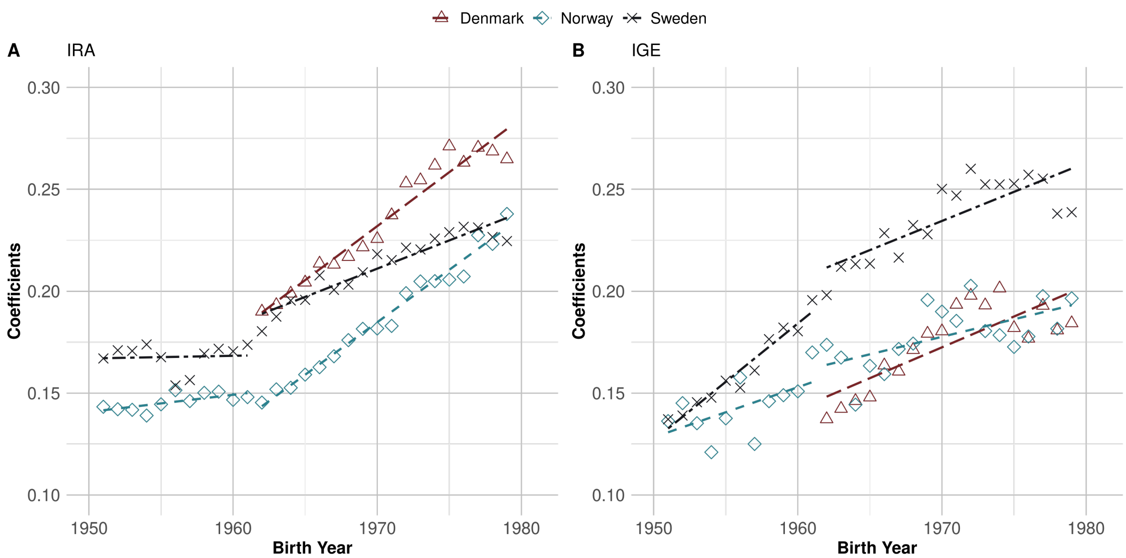

Figure 6: Estimates of IRA and IGE (Labor Income).

Note: Panel A depicts intergenerational rank associations between parents and children, estimated as in Equation (1), for each country. Panel B shows intergenerational income elasticities, i.e. correlations in log income between parent and child pairs (with zero incomes excluded from analysis). Parental income averaged over child ages 17-19, and child income averaged over ages 35-37 in all estimates.

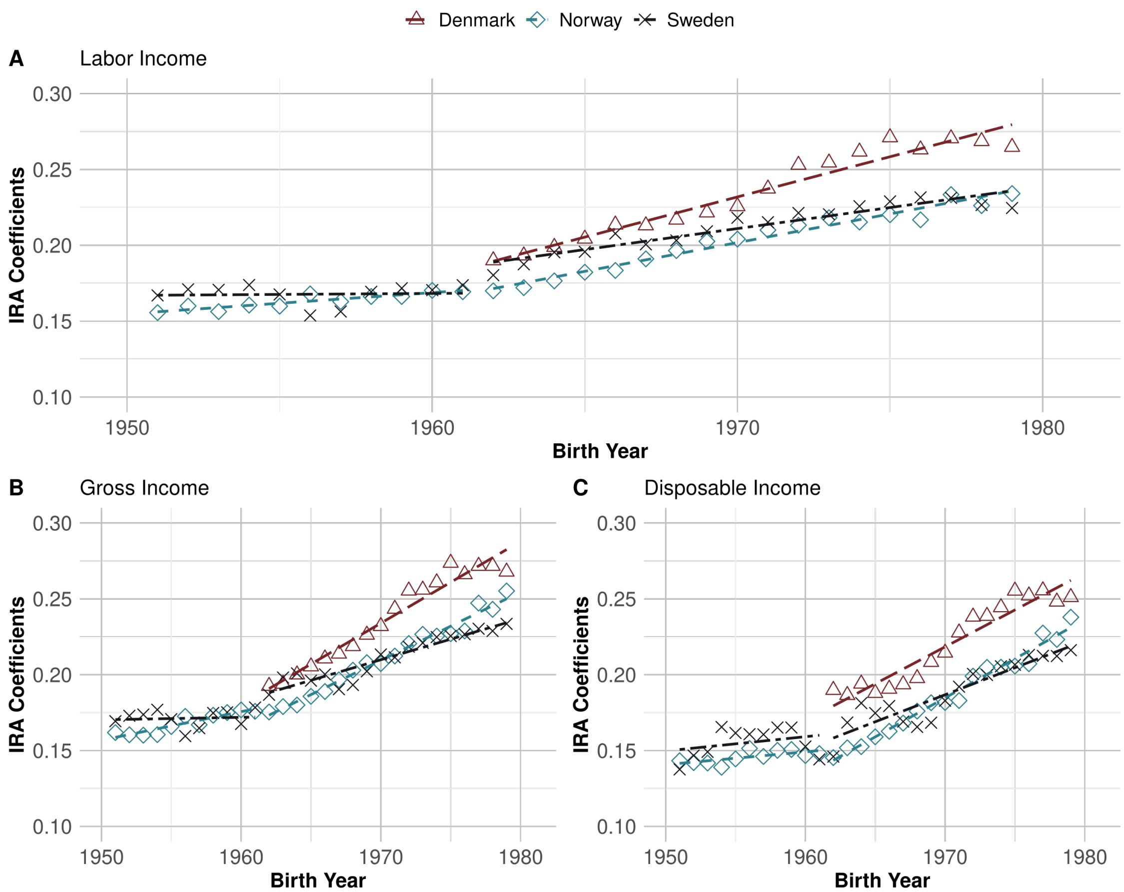

Figure 7: Estimates of IRA in Net-of-tax, Gross and Labor income.

Note: Each panel depicts intergenerational rank associations between parents and children, estimated as in Equation (1), for each country. Panel A shows estimates of the main specification: net-of-tax income. In panel B, total factor (gross) income is used, and panel C depicts labor earnings. Parental income averaged over child ages 17-19, and child income averaged over ages 35-37 in all estimates.

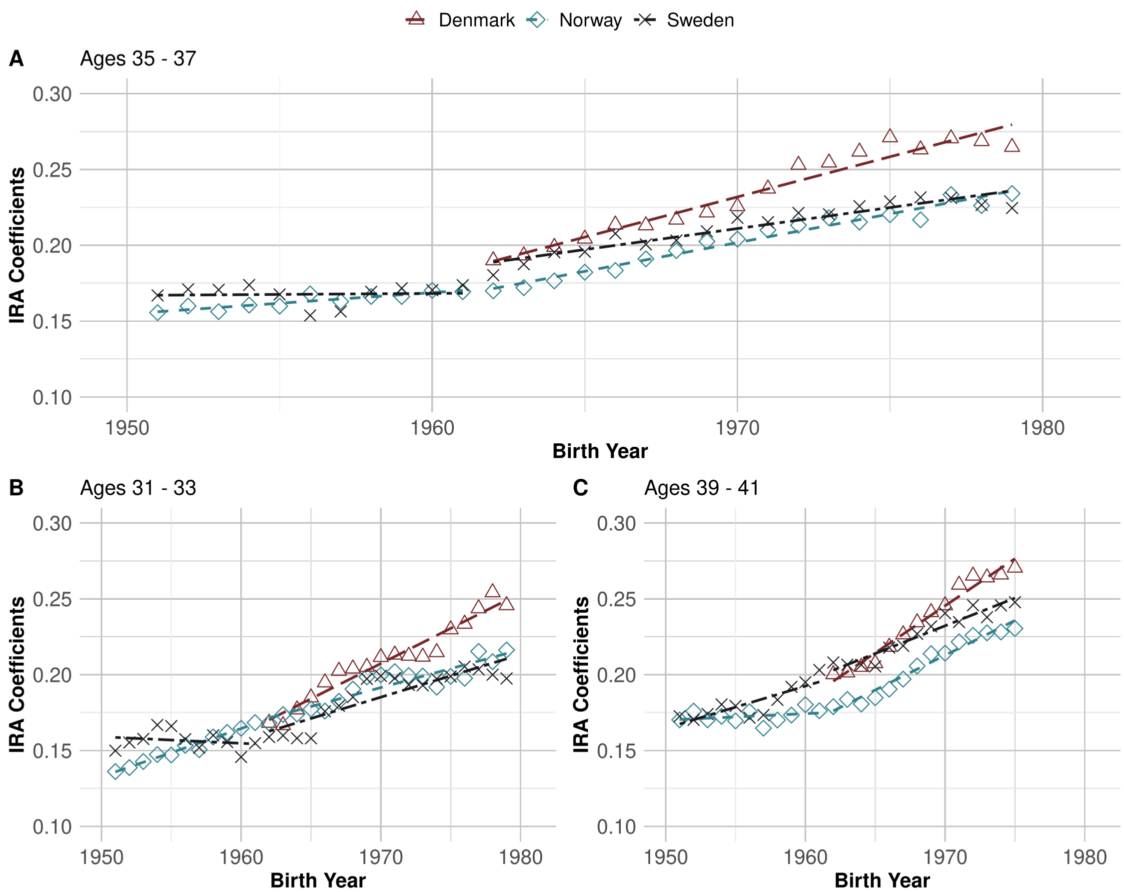

Figure 8: Estimates of IRA at Different Ages of the Child (Labor Income).

Note: Each panel depicts intergenerational rank associations between parents and children, estimated as in Equation (1), for each country. Panel A shows estimates of the main specification: average income at child ages 35-37. In panel B, child income is measured at ages 31-33, and in panel C, it is measured at ages 39-41. Parental income averaged over child ages 17-19 in all estimations.

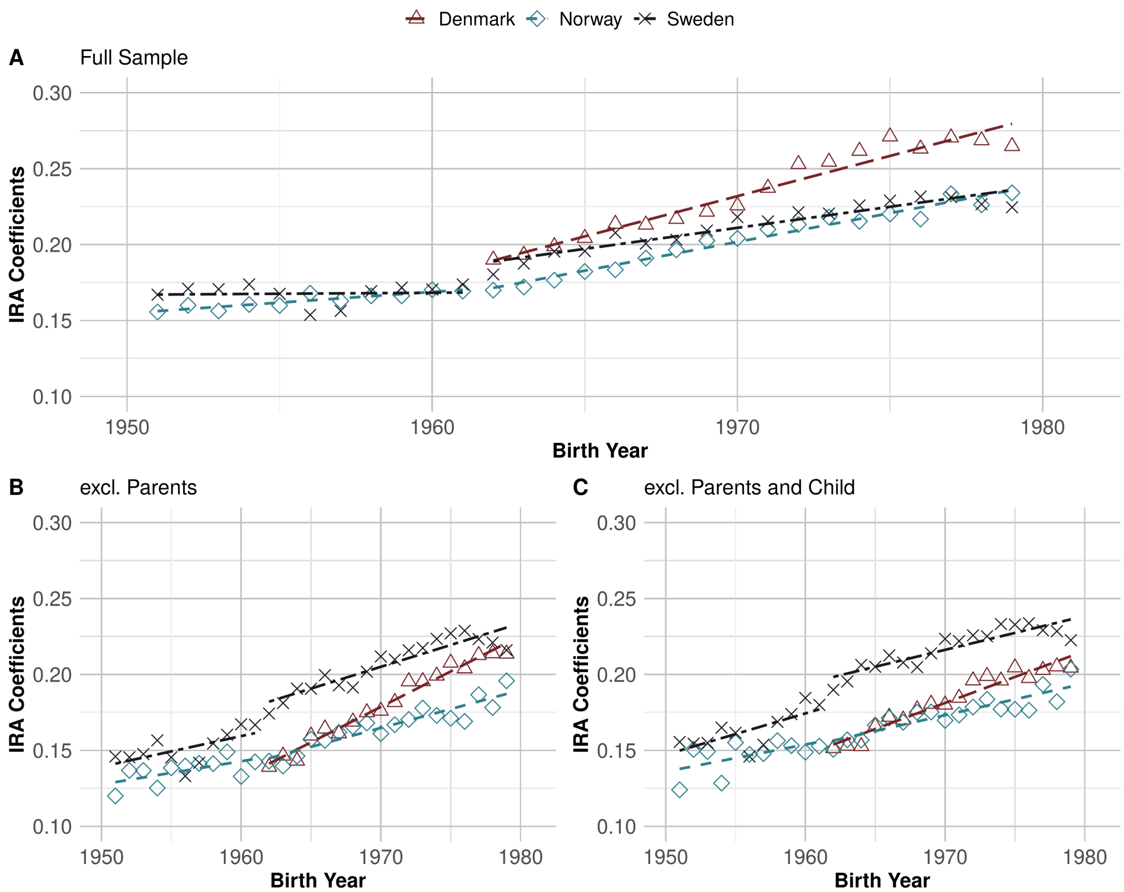

Figure 9: Estimates of IRA, Labor Force Participants Only (Labor Income).

Note: Panel A depicts intergenerational rank associations between parents and children, estimated as in Equation (1), for each country. Panel B shows equivalent estimates of IRA, when excluding child-parent pairs where either parent earns less than 10,000 USD (2017) in a given year. In panel C, we additionally exclude child-parent pairs where both the child and the parents have incomes below the 10,000 USD threshold. Parental income averaged over child ages 17-19, and child income averaged over ages 35-37 in all estimates.

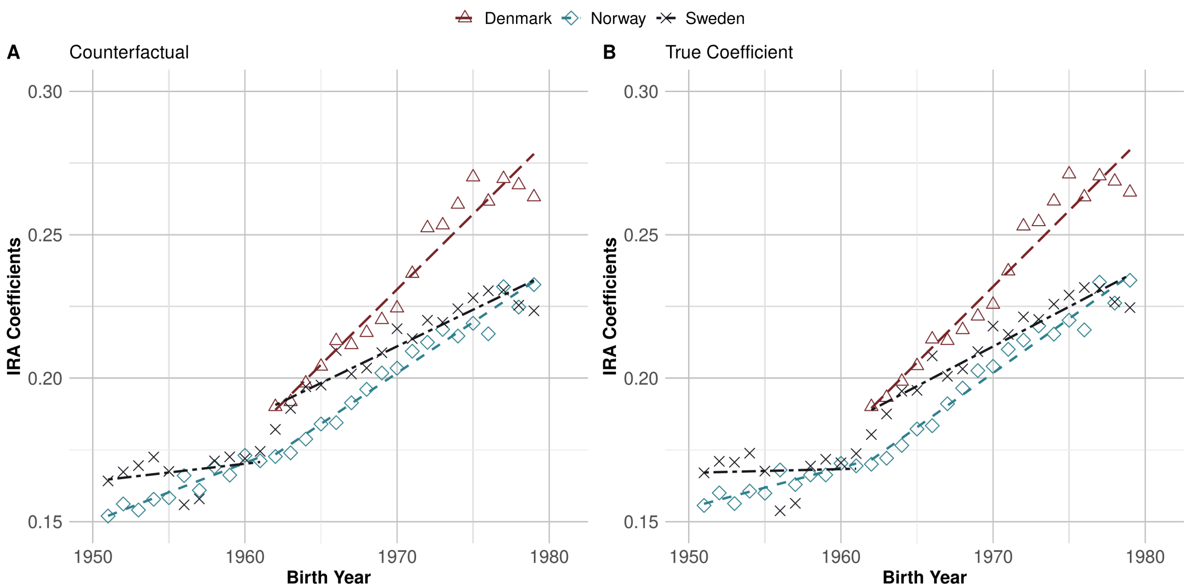

Figure 10: IRA Estimates Accounting for Participation Differences.

Note: The two panels depict IRA coefficients by year for the the counterfactual and the true relationship between child and parental income for each country. Panel A presents the plot for the counterfactual where maternal incomes are changed to the corresponding percentile income in 1979. Panel B shows the true coefficients estimated from the data.

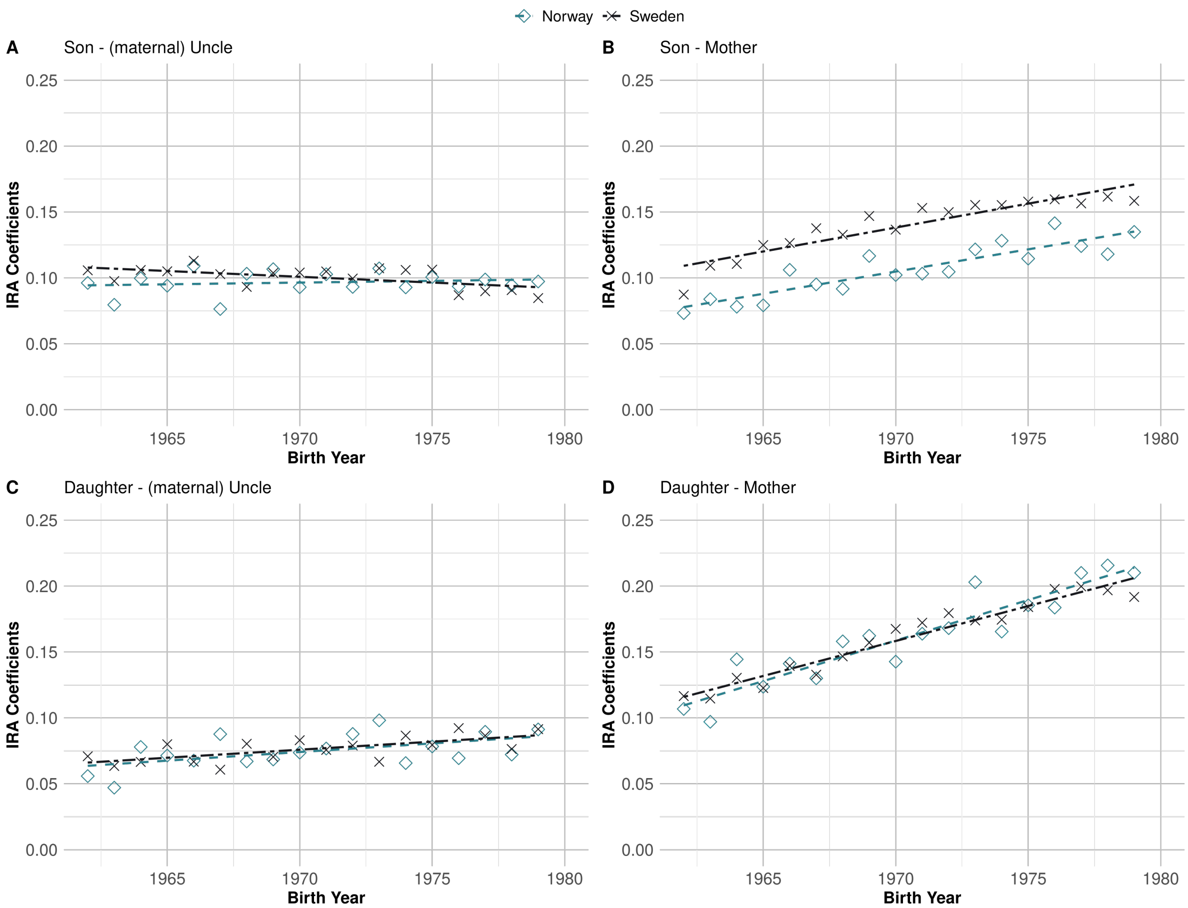

Figure 11: IRA Estimates Between Children and Their Maternal Uncles.

Note: The four panels depict IRA coefficients by year for the income association between sons (Panel A) and daughters (Panel C) and their mothers' brothers. i.e. maternal uncles. Panels B and D show the estimated IRA between sons and daughters and their mothers for the sample where maternal brothers are applicable. Estimates are birth-year specific. Each panel depicts these measures separately by country for the years 1951, 1962 and 1979.

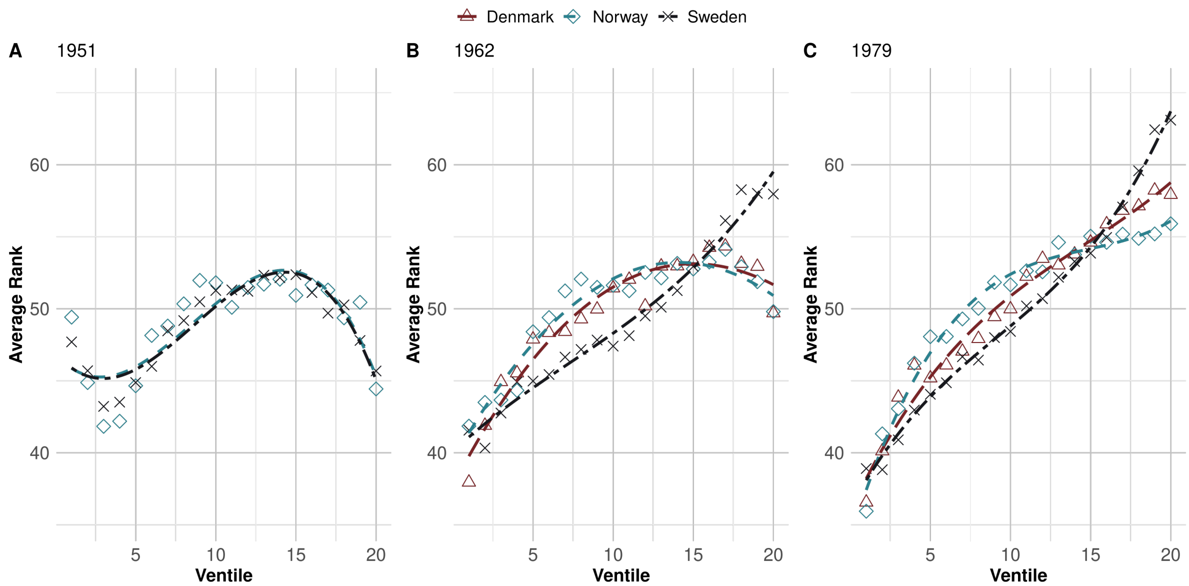

Figure 12: Average Maternal Ventile Rank by Paternal Ventile.

Note: The three panels show the average maternal income rank of mothers with children in the same birth cohort, by paternal (within parental pairs) income ventile. Each panel depicts these measures separately by country for the years 1951, 1962 and 1979. The fitted lines in panel A to B are estimated with local polynomial (third order) regressions.

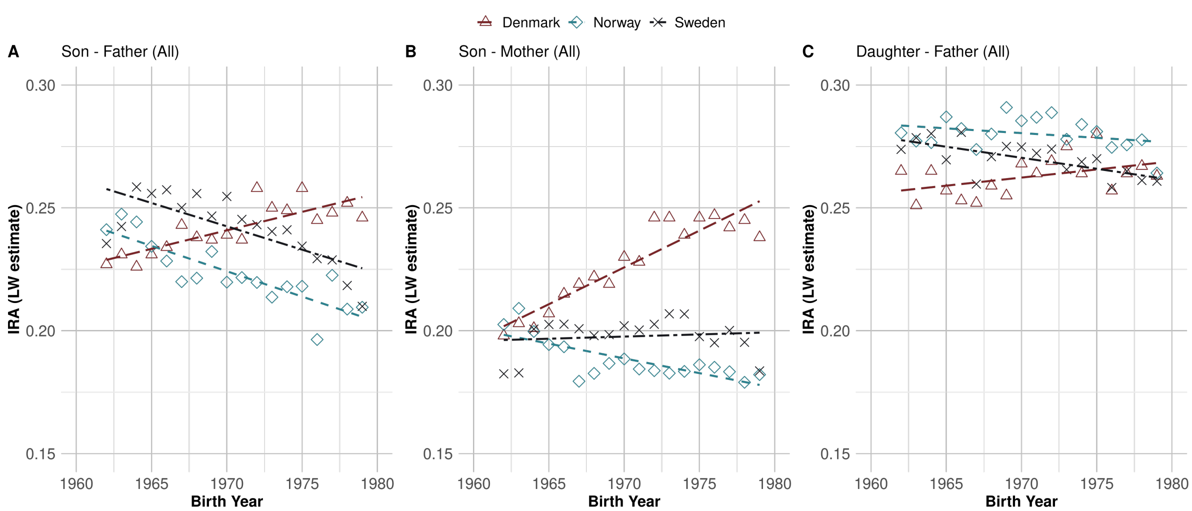

Figure 13: Trends in Intergenerational Mobility in Latent Economic Status.

Note: The three panels plot coefficients for intergenerational rank associations in latent economic status for Denmark, Sweden and Norway over the period from 1951 (1962) to 1979. Panel A shows son-father correlations, panel B son-mother correlations and panel C daughter-father correlations. Each marker indicates the coefficient of a separate regression and each line indicates fitted trend lines for the period 1962 to 1979.

Table 7: IRA Coefficients and Trends (United States)

	Parents	Father		Mother	
		Son	Daughter	Son	Daughter
Panel A:					
Pooled IRA	0.317***	0.336***	0.195***	0.097***	0.137***
	(0.017)	(.022)	(0.031)	(0.025)	(0.029)
Trend $\times$ 100	0.603***	-0.240	0.980***	0.136	1.047***
	(0.149)	(0.205)	(0.277)	(0.253)	(0.292)
N	5,392	2,272	1,637	2,477	2,205
Panel B:					
Pooled IRA	0.335***	0.360***	0.237***	0.107***	.152***
	(0.013)	(0.020)	(0.025)	(0.021)	(0.022)
Trend $\times$ 100	0.449***	-0.263*	0.728***	0.268	0.917***
	(0.118)	(0.178)	(0.229)	(0.202)	(0.213)
N	5,392	2,272	1,637	2,477	2,205
Panel C:					
Pooled IRA	0.294***	0.327***	0.192***	0.098***	0.126***
	(0.018)	(0.023)	(0.0353)	(0.026)	(0.032)
Trend $\times$ 100	0.433**	-0.393	1.156***	0.180	0.727
	(0.162)	(0.218)	(0.305)	(0.266)	(0.327)
N	2,927	1,583	904	1,497	1,001

Note: The table presents estimates of the IRA and linear trends in the IRA separately for different child-parent combinations. Due to the small sample sizes, trends have been estimated directly on the underlying micro data by regressing cohort-specific child ranks on cohort-specific parent ranks interacted with a linear time trend. The trend coefficients and standard errors have been multiplied by 100 in order to avoid too many digits after the separator. Panel A contains estimates for the full PSID sample using provided sample weights, Panel B uses the full sample without weights and Panel C includes estimates on the nationally representative SRC sample. Standard errors are in parentheses. P-values indicated by * $<$ 0.1, ** $<$ 0.05, *** $<$ 0.01.

Table 8: IRA Coefficients and Trends - Age 30 (United States)

	Parents	Father		Mother	
	Child	Son	Daughter	Son	Daughter
Panel A:					
Pooled IRA	0.327***	0.318***	0.222***	0.120***	0.151***
	(0.015)	(0.022)	(0.026)	(0.025)	(0.027)
Trend × 100	0.643***	0.133	0.661***	0.610**	0.571**
	(0.129)	(0.193)	(0.220)	(0.239)	(0.262)
N	6,652	2,664	2,109	2,685	2,611
Panel B:					
Pooled IRA	0.345***	0.341***	0.263***	0.148***	0.168***
	(0.012)	(0.018)	(0.021)	(0.019)	(0.020)
Trend × 100	0.457***	0.102	0.510***	0.429**	0.567***
	(0.101)	(0.59)	(0.183)	(0.176)	(0.181)
N	6,652	2,663	2,109	2,686	2,611
Panel C:					
Pooled IRA	0.303***	0.310***	0.225***	0.097***	0.133***
	(0.016)	(0.023)	(0.028)	(0.027)	(0.030)
Trend × 100	0.528***	0.020	0.586**	0.661**	0.352
	(0.146)	(0.210)	(0.245)	(0.261)	(0.307)
N	3,451	1,757	1,161	1,460	1,142

Note: The table presents estimates of the IRA and linear trends in the IRA separately for different child-parent combinations Children's income is measure at age 30. Due to the small sample sizes, trends have been estimated directly on the underlying micro data by regressing cohort-specific child ranks on cohort-specific parent ranks interacted with a linear time trend. The trend coefficients and standard errors have been multiplied by 100 in order to avoid too many digits after the separator. Panel A contains estimates for the full PSID sample using provided sample weights, Panel B uses the full sample without weights and Panel C includes estimates on the nationally representative SRC sample. Standard errors are in parentheses. P-values indicated by * < 0.1, ** < 0.05, *** < 0.01.

Table 9: Cohort-specific Parent-child Links, Main Spec. (United States)

	Parents	Father		Mother	
Birth year		Son	Daughter	Son	Daughter
1947	76	27	26	35	39
1948	107	38	39	48	56
1949	143	52	44	73	65
1950	171	57	72	68	99
1951	218	76	81	101	108
1952	193	70	78	88	102
1953	239	87	98	116	117
1954	235	78	96	101	128
1955	267	103	98	122	136
1956	263	86	106	107	147
1957	247	95	85	126	115
1958	220	75	95	95	119
1959	159	79	37	107	46
1960	177	96	34	121	54
1961	105	54	27	62	38
1962	117	50	37	63	53
1963	125	49	41	64	56
1964	100	47	30	56	43
1965	91	42	20	49	40
1966	88	39	17	52	34
1967	95	49	21	60	33
1968	66	32	17	39	26
1969	99	49	32	55	42
1970	87	40	20	50	35
1971	92	40	21	60	32
1972	111	49	22	63	43
1973	107	50	20	65	36
1974	117	51	25	64	48
1975	128	55	36	70	48
1976	132	58	37	63	53
1977	130	66	23	41	21
1978	138	66	34	27	30
1979	179	93	43	29	41

1980	142	65	36	33	26
1981	148	77	30	40	24
1982	120	58	27	26	30
1983	160	74	32	38	42
Total	5,392	2,272	1,637	2,477	2,205

Note: The table presents the number of cohort-specific parent-child links that were used to produce the main results from the PSID survey data.

Table 10: IRA Coefficients, Trends and Differences Across Countries and Time

	1951		1962			1979			Trend 1962-1979			Δ P-Value		
IRA Spec.	NO	SE	DK	NO	SE	DK	NO	SE	DK	NO	SE	DK-NO	DK-SE	NO-SE
All	0.156	0.167	0.190	0.170	0.180	0.265	0.234	0.225	0.530	0.379	0.277	0.065	0.004	0.176
									(0.035)	(0.018)	(0.033)			
Son-Parent	0.242	0.245	0.225	0.222	0.233	0.280	0.241	0.235	0.360	0.085	0.014	0.000	0.000	0.736
									(0.036)	(0.024)	(0.061)			
Daughter-Parent	0.146	0.158	0.197	0.173	0.169	0.276	0.262	0.240	0.592	0.552	0.428	0.363	0.067	0.020
									(0.035)	(0.037)	(0.038)			
Son-Father	0.253	0.248	0.220	0.236	0.242	0.241	0.213	0.211	0.139	-0.160	-0.224	0.000	0.000	0.378
									(0.035)	(0.022)	(0.060)			
Son-Mother	0.068	0.080	0.098	0.089	0.101	0.198	0.150	0.155	0.619	0.324	0.307	0.000	0.007	0.723
									(0.026)	(0.026)	(0.041)			
Daughter-Father	0.137	0.139	0.175	0.144	0.152	0.216	0.195	0.192	0.342	0.325	0.239	0.778	0.570	0.421
									(0.031)	(0.032)	(0.034)			
Daughter-Mother	0.073	0.077	0.120	0.119	0.112	0.227	0.221	0.194	0.731	0.607	0.541	0.439	0.181	0.107
									(0.036)	(0.037)	(0.034)			

Note: Columns (1)-(7) report the IRA coefficients of separated regressions in the years 1951, 1962 and 1979 separately for Denmark, Norway and Sweden. Columns (8)-(10) report the coefficient of the fitted regression lines of country specific regressions of the IRA coefficient on a linear trend for the years 1962 to 1979. The trend coefficients and corresponding standard errors have been multiplied by 100 in order to avoid too many digits after the separator. Columns (11)-(13) report rounded p-values for the null hypothesis that the slopes for the respective countries (see column header) are equal. Robust standard errors are reported in parentheses.

B.1 Appendix: Calibrating parameters in Model

Each set of country-year model parameters for trend decomposition are — loosely described — calibrated in the following steps:

1. If the year is the first year of observation for a given country, draw a random set of parameters. If the year is not the first year of observation, initialize the algorithm with the optimal set of parameters from the last year associated with the same country. These become the 'search parameters' until they are replaced.

2. Draw 400,000 parent-child pairs (the same in each year), and repeat the following procedure until there is a sufficiently close match between empirical rank associations and modelled rank associations:

   (a) Compute skills and incomes for all individuals (father, mother, son and daughter) using the set of 'search parameters' along with randomly drawn values for x_{it}^k, i_{it}^k and ε_{it}^k.

   (b) Compute associations in income ranks between (i) fathers and sons, (ii) fathers and daughters, and (iii) mothers and sons, and (iv) mothers and daughters, while (v) matching the relationship between mother and father income ranks.

   (c) Compare these associations to the empirical equivalents of the data for a given country and year. If the sum of squared distances between the rank associations from the data and their equivalents from the simulated data is smaller than the sum of squared distances obtained by the preferred parameters, the new set of parameters get to be preferred.

   (d) If the convergence criterion is not met, adjust the parameters using a customized variation of gradient descent. These now become the 'search parameters'.

In the following set of figures, we illustrate how the the implied empirical association between the two types of income in the (calibrated) simulated data compares to the empirical association between the same two incomes as observed in the data. These figures validate the quality of the calibration exercise.

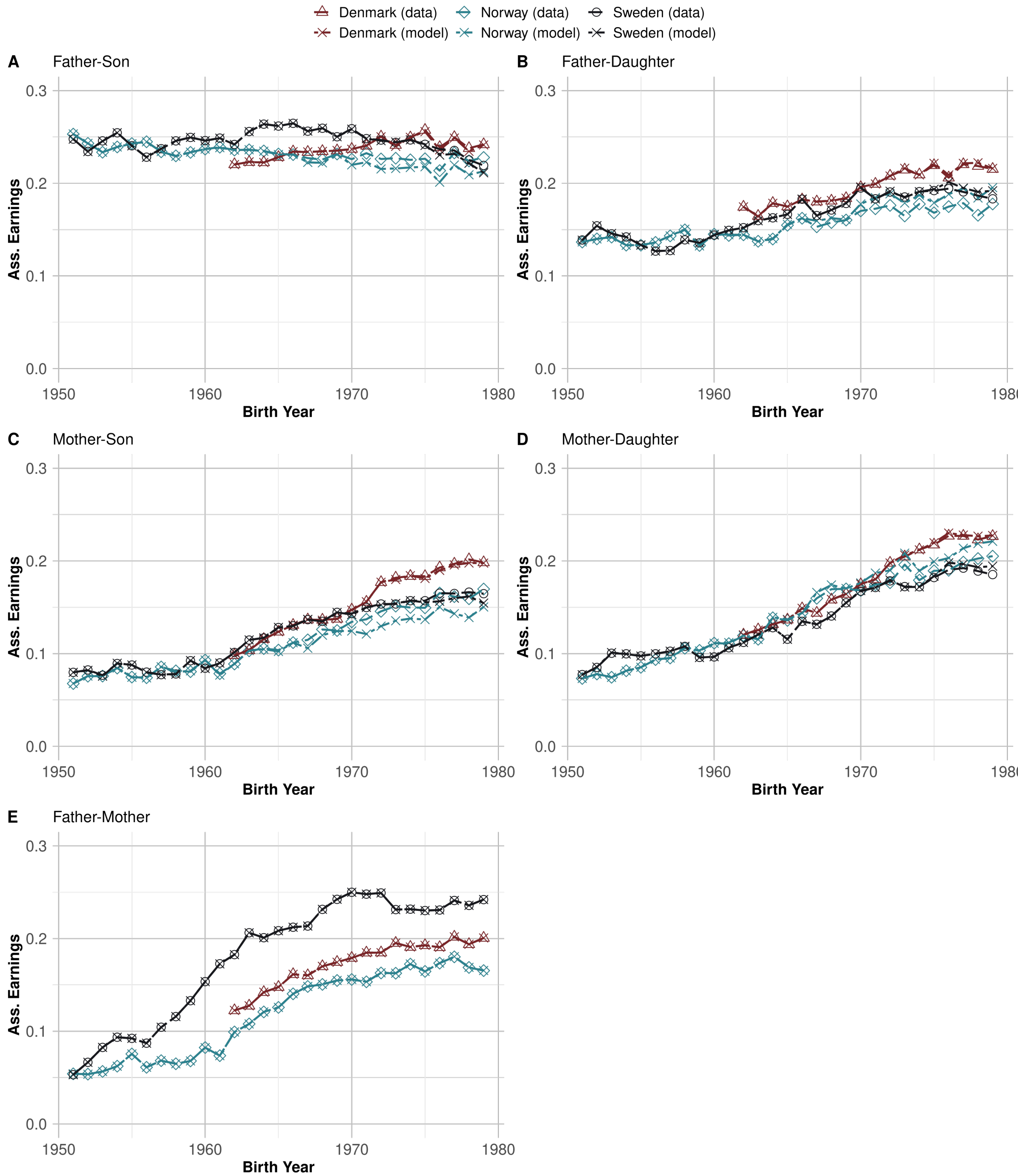

Figure 14: Validation of Calibration Exercise.

Note: Each panel displays the empirical association between two incomes as observed in the data as well as the implied empirical association between the same two types of income in the simulated data as calibrated in the decomposition model.