The Scandinavian
Journal of Economics

Scand. J. of Economics 119(1), 5–33, 2017
DOI: 10.1111/sjoe.12205

The Evolution of Social Mobility: Norway during the Twentieth Century*

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Abstract

We document trends in social mobility in Norway using intergenerational income elasticities, the associations between the income percentiles of fathers and sons, and brother correlations. The results of all approaches suggest that social mobility increased substantially between cohorts born in the early 1930s and the early 1940s. Father–son associations remained stable for cohorts born after World War II, while brother correlations continued to decline. The relationship between father and son income percentile ranks is highly non-linear for early cohorts, but it approaches linearity over time. We discuss increasing educational attainment among low- and middle-income families as a possible mechanism underlying these trends.

Keywords: Intergenerational mobility
JEL classification: *D*31; *J*31; *J*62

I. Introduction

The debate on the consequences of income inequality has drawn attention to cross-country differences in social mobility. A large body of research has shown that countries that are known for redistributive welfare state institutions and low cross-sectional income inequality, such as the Nordic countries, have a much lower degree of intergenerational income persistence than, for example, the US or UK.[1] These cross-country differences have

*We thank two anonymous referees and seminar participants at EALE, FEA, HECER, NHH, SOLE, and the Conference on Social Mobility held at the University of Chicago for insightful comments and suggestions, and the Academy of Finland and the Norwegian Research Council for funding.

[1] See Black and Devereux (2011) and Corak (2013) for recent surveys.

led to speculation about their potential causes and implications. Yet, it is difficult to draw conclusions from a pattern present at a single point in time. As a response, recent research has shifted towards a complementary approach of documenting within-country changes in social mobility.

In this paper, we examine the evolution of social mobility in Norway for children born between the early 1930s and the mid-1970s using newly digitalized data and alternative measurement approaches. These birth cohorts are of particular interest because they cover a period in which the Norwegian economy underwent dramatic structural change and much of the Norwegian welfare state was built. The last few cohorts included in our sample were born into one of the world's richest countries, with extensive redistributive institutions and a high level of intergenerational mobility. In contrast, our earliest birth cohorts grew up in a relatively poor and unequal country. We show that they also experienced less social mobility than did subsequent birth cohorts.

We contribute to the earlier body of literature across several dimensions. First, we use high-quality register data augmented with military records from the early 1950s and newly digitalized municipal tax records from 1948. These data allow us to present precise estimates, even for those cohorts born before World War II (WWII). Moreover, we use three different measurement approaches – intergenerational income elasticities, associations between the income percentile ranks of fathers and sons, and brother correlations – in order to assess the robustness of patterns over time. We also examine non-linearities in father–son associations and, in particular, evidentiary changes in these across birth cohorts. Finally, we document the changes in the association between educational attainment and family background.

Our paper adds to the growing body of literature on historical trends in intergenerational mobility. Previous work examining Nordic countries includes Pekkala and Lucas (2007), who examine trends in intergenerational income elasticity in Finland, and Björklund *et al.* (2009), who investigate the evolution of brother income correlations in Sweden. Both of these studies present evidence on the increasing mobility between cohorts born in the 1930s and 1950s, and stable or decreasing social mobility for later birth cohorts. Modalsli (2017) documents a substantial increase in intergenerational occupational mobility in Norway between 1865 and 2011. In contrast, Lindahl *et al.* (2015) focus on the descendants of a single generation of schoolchildren in one Swedish city and find no evidence of changes in intergenerational income mobility. Clark (2012) examines the persistence of surnames among elite occupations and argues that rates of social mobility in Sweden have remained roughly stable since the pre-industrial era. Finally, and in line with our results, Bratberg *et al.* (2005) find that intergenerational income elasticities among post-WWII cohorts in Norway remained stable.[2]

Our main findings are as follows. All three approaches suggest that social mobility increased between the cohorts born in the early 1930s and the early 1940s. For cohorts born after WWII, our findings are more mixed, with father–son income associations remaining stable, while brother correlations continue to decline. A closer examination of the joint father–son income percentile distribution reveals a fairly complex evolution. Downward mobility among sons of the highest-earning fathers became more prevalent over time, while upward mobility from the 25th percentile of the fathers' income distribution steadily increased. The prospects of sons of the lowest-earning fathers at first improved and then deteriorated. We observe no changes for the sons of fathers between the 50th and 75th percentiles of the income distribution.

Guided by theoretical work starting with Becker and Tomes (1979) and extended by, among others, Solon (2004), Hassler *et al.* (2007), and Ichino *et al.* (2011), we augment our analysis by documenting trends in returns to education and association between family background and educational attainment. We show that among those cohorts for whom social mobility increased, educational attainment increased rapidly among the sons of fathers below the 80th percentile of the fathers' income distribution. At the same time, the returns to education decreased. This pattern is consistent with a hypothesis that the major educational reforms initiated in the 1930s substantially improved educational opportunities, with the exception of sons of the highest-earning fathers (who were already being highly educated). The resulting increase in the supply of educated workers may then have decreased the returns to education. While our analysis is purely descriptive, these stylized facts allow us to build a consistent narrative. We leave a more rigorous testing of this narrative to future research.

The rest of the paper is structured as follows. In Section II, we briefly discuss the changes in the institutional context in Norway during our period of study. In Section III, we review the estimation methods used for assessing social mobility. In Section IV, we explain how we combine information from Norwegian censuses, military records, tax registers, and municipality-level tax records to construct our data. We present the main results on changes in intergenerational mobility over time in Section V, and we discuss the role of education in Section VI. We conclude in Section VII.

[2] Studies examining trends in social mobility outside of the Nordic countries include Aaronson and Mazumder (2008), Lee and Solon (2009), Chetty *et al.* (2014), Olivetti *et al.* (2013), and Olivetti and Paserman (2015) in the US; Blanden *et al.* (2011) and Nicoletti and Ermisch (2008) in the UK; and Lefranc and Trannoy (2005) in France. Long and Ferrie (2013) provides a comparison of historical changes in mobility in the US and UK.

II. Institutional Context

In 1930, Norway was a poor and relatively unequal country compared with current standards, with a GDP per capita of around $4,000 (in 2002 US dollars), a top one-percent income share of about 13 percent, and a population with an average of seven years of education. However, the standard of living in Norway at the time was comparable to that of Sweden and the UK as measured by GDP per capita, average years of education, average height, life expectancy, and infant mortality.[3]

During the next few decades, Norway underwent a dramatic transformation. The economy industrialized and grew rapidly from the mid-1930s onwards. During this period – and particularly after WWII – Norway, like other Nordic countries, introduced extensive welfare institutions that provided public services and insurance to everyone either for free or at a highly subsidized price. An important political shift took place in 1935, when the Labor Party came to power and began to extend the old age pension, disability pension, sickness leave, and unemployment benefits to cover the entire country and all industries.[4]

One of the first initiatives of the new Labor government was to reform the education system with the aim of providing similar educational opportunities for all Norwegians. The background of this reform was the large regional differences in the supply of education. For example, the actual amount of teaching provided per year varied between 42 weeks in the cities to as low as 12 weeks in some rural municipalities. This reform was rolled out over the next decade and led to a major increase in public spending on education. Another major educational reform took place in the 1960s with the extension of the mandatory period of education from seven to nine years. Furthermore, the high school, regional college, and university sectors expanded, particularly from the early 1970s onwards.

However, the transformation to a fully developed welfare state was not immediate, and partly relied on local initiatives by municipalities or private initiatives by philanthropic societies. For example, school breakfast programs were initiated by some municipalities from the mid-1930s onwards (Bütikofer *et al.*, 2016). Other examples are the well-child visit centers for mothers and new-born children, which were introduced by a philanthropic society in the 1930s and taken over by the state only in the early 1970s (Bütikofer *et al.*, 2016). By the mid-1960s, a fully developed social security system was in place. Family policies, such as maternity leave and

[3] The figures in this section are from Grytten (2004, 2014), Aaberge and Atkinson (2010), Statistics Norway (1995), and the Clio Infra web site (https://www.clio-infra.eu/datasets/).

[4] Partial versions of these programs had been introduced earlier in some municipalities and for some occupations/industries. WWII and the German invasion of Norway interrupted the implementation of these reforms, but they recommenced afterwards.

subsidized day care, were launched in the mid-1970s and implemented gradually (Carneiro *et al.*, 2015).

In short, the cohorts born in the 1970s grew up in a very different country than those born in the 1930s. By 1990, Norway had become one of the richest and most equal countries in the world, with a GDP per capita exceeding \$20,000 (in 2002 US dollars) and a top income share of 4 percent. While 40 percent of the population still had only mandatory schooling, 15 percent had a university degree.[5]

III. Measurement

Estimation of intergenerational mobility has a long history, and the econometric and measurement issues have been discussed extensively in numerous surveys.[6] We examine several measures of mobility – intergenerational income elasticity, rank–rank slopes, expected percentile ranks, and sibling income correlation – and focus on their changes over time. These measures provide alternative and complementary perspectives on intergenerational income persistence. In this section, we briefly discuss the estimation and interpretation of each measurement approach.

Intergenerational Income Elasticity

The most common measure of social mobility is the intergenerational income elasticity, typically measured by estimating the following regression

$$\ln Y_i = \alpha + \beta \ln X_i + \epsilon_i, \tag{1}$$

where Y_i is a measure of the son's income and X_i is a measure of the father's income.

The intergenerational elasticity, β, can change over time for several reasons. First, it can reflect changes in both the intergenerational correlation between fathers' and sons' income and changes in cross-sectional income inequality. To see this, note that

$$\beta = \frac{\sigma_y}{\sigma_x}\rho, \tag{2}$$

where ρ is the intergenerational income correlation, and σ_y and σ_x are the standard deviations of sons' and fathers' log income, respectively. Thus,

[5] As we discuss in Section VI, the last birth cohort we examine (i.e., those who were teenagers in 1990) ended up having much higher educational attainment than the 1990 population.

[6] See, for example, Solon (1999), Björklund and Jäntti (2009), Black and Devereux (2011), and Jäntti and Jenkins (2015).

a decrease in β can follow from either a decrease in intergenerational correlation or a decrease in cross-sectional income inequality.

An important practical challenge in interpreting and estimating intergenerational income elasticity is that the association between fathers' and sons' log income tends to be highly non-linear (Bratsberg *et al.*, 2007; Chetty *et al.*, 2014). As a consequence, the estimates for β can be highly sensitive to whether the tails of the fathers' income distribution are included in the estimations. As shown in Figure A1 in the Online Appendix, strong non-linearities are also present in our data. Thus, changes in the tails of the fathers' income distribution can have a disproportionate influence on the changes in β. To examine whether this affects our conclusions, we report estimates using our full data and a restricted sample, where we omit the top and the bottom deciles of the fathers' income distribution. We also investigate non-linearities in detail in the context of the income percentile ranks.

Rank–Rank Slope and Expected Percentile Ranks

Recent work on social mobility has shifted away from intergenerational income elasticities and towards the association between fathers' and sons' income percentile ranks. This approach has several advantages (Chetty *et al.*, 2014). First, intergenerational income elasticity estimates tend to suffer from important attenuation and life-cycle biases, particularly when one has to use snapshots of income data to construct proxies for lifetime income (Haider and Solon, 2006; Böhlmark and Lindquist, 2006; Bhuller *et al.*, 2014). In contrast, Nybom and Stuhler (2015) show that estimates for income percentile ranks are not sensitive to the age when measuring sons' income as long as it is measured during their mid-30s to late-40s. Second, percentile ranks provide a natural way to deal with zero incomes, which create an important measurement challenge when measuring income in logarithms (Solon, 1992).

We begin examining the association between fathers' and sons' income percentile ranks by estimating the regression

$$P_i = \alpha + \beta R_i + \epsilon_i, \tag{3}$$

where P_i is the son's percentile rank in the income distribution of his birth cohort and R_i is the percentile rank of his father. In regression (3), α corresponds to the expected income percentile of a son of the poorest father and the rank–rank slope β measures the difference in the expected percentile of the offspring of the poorest and the richest fathers. Thus, β is a measure of relative income mobility.

An alternative approach is to examine the expected percentile rank of a son with a father at the rth percentile. For example, Chetty *et al.* (2014)

use the expected percentile rank of children whose parents are at the 25th percentile, $\hat{\alpha} + 0.25\hat{\beta}$, as a measure of absolute upward mobility. We extend their approach in two ways. First, we report the expected percentile rank of sons over the entire distribution of the fathers' income distribution, and we focus on the changes in these expected percentile ranks over time. Second, we use local linear estimators to take into account the fact that the association between fathers' and sons' income percentile ranks is not linear in our data. This analysis allows us to pin down where in the fathers' income distribution the changes in mobility took place, and whether the importance of particular parts of the parental income distribution changed over time.

Brother Correlations

An alternative way to measure social mobility is to examine brother correlations instead of father–son associations. An advantage of this approach is that we do not need to observe parental income in order to calculate the income correlation between brothers. A conceptual advantage is that because brothers share a growth environment in a more general sense, we can interpret brother correlations as a broader measure of the importance of childhood conditions than intergenerational associations. Thus, comparison of trends in intergenerational associations and sibling correlations might be informative about the changes in the importance of those factors shared by brothers, such as school quality or changes in the importance of residential neighborhood, but not fully captured by their father's income.

We follow the estimation approach in Björklund *et al.* (2009) and regress the log income of each brother i in family j at time t, Y_{ijt}, on year and age dummies Z_{ijt}

$$Y_{ijt} = \gamma Z_{ijt} + \epsilon_{ijt}. \tag{4}$$

We model the error term ϵ_{ijt} to consist of a permanent family component shared by all brothers in family j, a_j, a permanent component that is specific to individual i, b_{ij}, and an error term that picks up deviations from lifetime income, v_{ijt}, so that $\epsilon_{ijt} = a_j + b_{ij} + v_{ijt}$. The brother correlation ρ_{Y_i,Y_k} is then

$$\rho_{Y_i,Y_k} = \frac{\sigma_a^2}{\sigma_a^2 + \sigma_b^2}. \tag{5}$$

To estimate the brother correlation, we need to estimate both variances σ_a^2 and σ_b^2. Björklund *et al.* (2009) show that it is important to take into account the persistence in the transitory term v_{ijt} in this estimation. We use their generalized method of moments (GMM) approach under the assumption that the transitory term follows an AR(1) process, that is,

$v_{ijt} = \lambda v_{ijt-1} + u_{ijt}$, where u_{ijt} is a mean zero, constant variance random shock to current income.[7]

IV. Data

Documenting social mobility imposes two requirements on the data. First, we require reasonable proxies of lifetime income for both parents and their children. This means that we need to observe income at an age when the association between annual and lifetime income is reasonably strong. In addition, we need to link family members together in order to obtain information on individuals' own income as well as the income of their fathers and, in part of the analysis, their brothers. These criteria determine our estimation sample, which contains information on individuals born between 1932 and 1974, and their fathers and brothers. We also examine the cohort born in 1974–1979 in our analysis for educational attainment.

We derive most of our data from several longitudinal databases maintained by Statistics Norway, including information on the demographic and socioeconomic characteristics of the Norwegian population. We augment these data with census data from 1960 and military records from the early 1950s (see below for details). All data sources include personal identifiers and thus we can link them together, as well as link children with their parents and brothers. The information for family links is from the Norwegian population register, established in the early 1960s using information collected from the 1960 national and local censuses. For men born after 1950, we can virtually identify all mothers (and thus, all brothers). Also, in the case of the cohorts born in the mid-1930s, we identify most of the fathers and mothers, while for the cohorts born in the early 1930s we are able identify them for more than a third of the cohort.

Sons' and Brothers' Income

Our measure of sons' income is from the tax register, which records annual (pre-tax) income for the period 1967–2010. Our income measure is the sum of labor income (from wages and self-employment) and work-related cash transfers (such as unemployment and short-term sickness benefits). We measure income at age 35 for all birth cohorts, because the oldest sons included in our analysis were born in 1932 (and are thus 35 years of age

[7] To ensure our results are as comparable as possible to those of Björklund *et al.* (2009), we use the same birth cohorts and focus on brothers born within seven calendar years of each other. We differ from their specification in measuring income at age 35–44 (instead of 30–38), given data restrictions. Furthermore, we conduct inference using block bootstrapping with brother pairs as the resampling unit.

in 1967 when we first observe their income). This measure allows us to observe sons' income for the cohorts born between 1932 and 1974. We also examine sons' educational attainment, which we measure using information from the education register.

Income at age 35 provides us with a reasonable proxy of lifetime income (Böhlmark and Lindquist, 2006; Bhuller *et al.*, 2014). By this age, most men have completed their education and have entered the labor market. As shown in Tables A1 and A2 in the Online Appendix, the intergenerational income elasticities and the rank–rank slope estimates are slightly larger when we measure sons' income at ages 30–34, 35–39, and 40–44 rather than at age 35. However, the differences are small and do not significantly alter our conclusions regarding the trend in social mobility. We find nearly all sons in the register, but 7–9 percent have zero income at the age of 35. We include these observations in the analysis using percentile ranks, but omit them from the log specifications when estimating the intergenerational income elasticity.

Fathers' Income

We use two complementary approaches for measuring the income of fathers. First, we directly observe annual income from the tax register for those fathers still of working age in 1967 and for whom we can establish father–son links from the population register. We meet these conditions for almost everyone in the later birth cohorts. However, for earlier cohorts, we face the challenge that many of the fathers are toward the end of their working careers or already retired in 1967. Thus, our primary measure of a father's income is his average income at age 55–64, when sons are, on average, 29.6 years old. Importantly, while measuring fathers' income at quite an old age may lead to some measurement error, the resulting attenuation bias is likely similar for all birth cohorts.

Despite this, the share of sons for whom we directly observe a father's income declines as we move towards earlier birth cohorts. This could distort our conclusions if the subpopulation for which we observe a father's income differs from the full population in terms of intergenerational mobility. To examine this possibility, we construct an alternative measure of fathers' income using military records. In Norway, military service is mandatory for all young men of normal health. In the cohorts born between the 1930s and 1950s, roughly 75 percent of men served in the military (Rossow and Amundsen, 1986). Importantly, the military recorded information on the occupation of the father for each conscript (but not the father's identification number). We have access to the full draft records for men born in the period 1932–1933. For other cohorts, we observe the father's occupation from the 1960 census, provided we observe the father–son link from the population register.

We use the information on father's occupation and son's resident municipality to impute income for the father. This draws on Statistics Norway (1950), which reports information on average salaries by occupation across 735 Norwegian municipalities using 1948 tax records. As the military records provide us with information on the father's occupation in 20 categories, we can use this information to impute the father's income using over 10,000 income values from the tax records.[8] These sources allow us to construct imputed fathers' income for almost 80 percent of men born in the period 1932–1933. The match is lower for the late 1930s and the 1940s cohorts, but increases to 95 percent for the cohorts born in the early 1950s.

The strength of our two proxies for fathers' lifetime income is that their limitations differ greatly. The tax register provides accurate information on income, but these are from the later stages of the father's career.[9] In contrast, the matching of sons to fathers is not perfect for the early 1930s cohorts. However, the quality of the imputed income measure is likely to improve as we move towards the earlier cohorts. The reason is that the imputation uses the 1948 tax records and thus the occupation–municipality-level averages are a better proxy for the father's true income if the father was in his prime working age in around 1948.

Table A3 in the Online Appendix presents a closer examination of the relationship between the two measures by reporting the estimates obtained by regressing the observed income for fathers on imputed income for those fathers for whom we observe both measures. The results suggest a strong correlation between the cohorts born in the 1930s, a slightly less strong correlation for the 1940s birth cohorts, and an even less strong correlation for the 1950–1954 birth cohort. This pattern is consistent with the hypothesis that measurement error in imputed income becomes more severe as we move towards later birth cohorts. Consequently, we expect attenuation bias to increase over time in an analysis based on imputed income.

[8] This is a major improvement on earlier studies such as Pekkala and Lucas (2007), which have relied on simple occupational averages to proxy for fathers' incomes in the earliest cohorts.

[9] In the Online Appendix, we show that estimates using fathers' income rank are not sensitive to the age at which fathers' income is measured (Table A1), but the intergenerational income elasticity estimates tend to be substantially larger when fathers' income is measured at a younger age (Table A2). Importantly, these differences do not affect our conclusions regarding the trends in social mobility.

V. Results

This section presents our main results. We begin with intergenerational rank–rank slopes, and compare these to traditional intergenerational income elasticities as well as to the brother correlations. The last subsection examines the non-linearities in the association between fathers' and sons' income percentile ranks.

Rank–Rank Slope

Table 1 reports estimates from regressing sons' percentile rank at age 35 in the income distribution of their birth cohort on their fathers' percentile rank in their income distribution (i.e., relative to other fathers with children in the same birth cohort). Each estimate pair comes from a separate regression, which differs in the birth cohort used in the estimation (columns) and the way we approximate fathers' lifetime income (panels). In Panel A, we use fathers' average annual income at age 55–64. The results show that the intergenerational rank correlation decreased from 0.28 for the cohort born in 1932–1933 to 0.20 in the cohort born in 1940–1944. This corresponds to an almost 30 percent decrease in the rank–rank slope and is highly statistically significant. For the cohorts born after WWII, the intergenerational rank correlation remains remarkably stable. Panel B provides similar estimates for a restricted sample, where we excluded those in the bottom and top deciles of the fathers' income distribution from the sample. The results are very similar to those obtained from the full sample.

Panel C of Table 1 reports the estimates using imputed fathers' income. For the 1932–1933 birth cohort, the sample size increases eightfold in comparison with that when using the fathers' observed income. However, the estimates for the rank–rank slope are very similar to those reported in Panels A and B. Furthermore, we again find a clear decline in the intergenerational rank–rank slope for the cohorts born between the early 1930s and the early 1940s. In relative terms, the decline, from 0.25 to 0.16, is larger than when using actual income. After the cohort born in 1945, the rank–rank slopes continue to decline, but at a much slower rate. These later declines are likely to reflect increasing attenuation bias as the 1948 average occupation–municipality earnings become an increasingly poorer proxy for fathers' true income (see Section IV, *Fathers' Income*, for a discussion).

Intergenerational Income Elasticity

Table 2 provides estimates of the intergenerational income elasticity using an identical approach to the rank–rank slopes above. In Panel A, we use

Table 1. *Rank–rank regressions*

	Sons' birth cohort								
	1932–1933	1935–1939	1940–1944	1945–1949	1950–1954	1955–1959	1960–1964	1965–1969	1970–1974
Panel A: Fathers' average percentile rank at age 55–64									
Fathers' inc. perc.	0.280	0.252	0.198	0.192	0.190	0.196	0.195	0.194	0.192
	(0.015)	(0.006)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)
Constant	0.322	0.380	0.419	0.416	0.411	0.407	0.410	0.405	0.406
	(0.009)	(0.003)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Obs.	3,909	32,114	78,526	130,079	140,411	150,457	152,663	165,436	156,723
R^2	0.08	0.06	0.04	0.04	0.04	0.04	0.04	0.04	0.04
Panel B: Fathers' average percentile rank at age 55–64, excluding bottom and top decile									
Fathers' inc. perc.	0.272	0.217	0.173	0.174	0.173	0.182	0.180	0.176	0.163
	(0.021)	(0.007)	(0.005)	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.004)
Constant	0.314	0.386	0.421	0.415	0.411	0.406	0.410	0.411	0.420
	(0.011)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Obs.	3,125	25,688	62,803	104,037	112,312	120,379	121,213	129,554	123,326
R^2	0.05	0.03	0.02	0.02	0.02	0.02	0.02	0.02	0.02
Panel C: Fathers' imputed income rank									
Fathers' earn. perc.	0.251	0.214	0.165	0.148	0.133				
	(0.005)	(0.004)	(0.003)	(0.003)	(0.003)				
Constant	0.377	0.383	0.426	0.432	0.437				
	(0.003)	(0.002)	(0.002)	(0.001)	(0.001)				
Obs.	31,568	55,075	97,074	142,569	145,105				
R^2	0.07	0.05	0.03	0.02	0.02				

Notes: Estimates and robust standard errors (in parentheses) from regressing sons' income percentile at age 35 in their birth cohort on fathers' income percentile.

Table 2. *Intergenerational income elasticities*

	Sons' birth cohort								
	1932–1933	1935–1939	1940–1944	1945–1949	1950–1954	1955–1959	1960–1964	1965–1969	1970–1974
Panel A: Fathers' average income at age 55–64, full									
Fathers' log inc.	0.121	0.109	0.069	0.064	0.067	0.068	0.064	0.062	0.060
	(0.011)	(0.004)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Obs.	3,529	29,634	72,165	117,057	123,032	126,959	126,706	136,156	130,623
R^2	0.05	0.03	0.02	0.01	0.01	0.01	0.01	0.01	0.01
Panel B: Fathers' average income at age 55–64, excluding bottom and top decile									
Fathers' log inc.	0.209	0.193	0.136	0.127	0.149	0.158	0.147	0.133	0.123
	(0.020)	(0.008)	(0.005)	(0.004)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
Obs.	2,826	23,797	57,907	93,995	98,733	102,074	101,834	109,376	104,934
R^2	0.04	0.03	0.02	0.01	0.01	0.01	0.01	0.01	0.01
Panel C: Fathers' imputed income									
Fathers' log inc.	0.229	0.228	0.156	0.143	0.142				
	(0.006)	(0.005)	(0.004)	(0.003)	(0.004)				
Obs.	29,879	51,581	91,497	133,976	135,891				
R^2	0.05	0.04	0.02	0.01	0.01				

Notes: Estimates and robust standard errors (in parentheses) from regressing sons' log income at age 35 in their birth cohort on fathers' log income.

fathers' actual income at age 55–64 to proxy for lifetime income. Once again, we observe a clear decline in the intergenerational persistence between cohorts born in the early 1930s and early 1940s, after which the intergenerational income elasticity remains roughly constant.

Panel B of Table 2 illustrates the importance of the tails of the fathers' income distribution for the estimation of intergenerational income elasticities. The estimates reported in Panel B are from otherwise identical regressions as those reported in Panel A, but we now omit observations from the top and bottom deciles of the fathers' income distribution. Consequently, the estimated elasticities increase by 69–129 percent.

However, while the levels of the elasticity estimates are highly sensitive to including/excluding the tails of the fathers' income distribution, the trends presented in Panels A and B are similar. Between the 1932–1933 and 1940–1944 birth cohorts, the elasticity falls by 43 percent in the full sample and by 33 percent in the restricted sample, and remains roughly constant afterwards. The exception is the elasticity, which increases slightly in the trimmed sample for cohorts born in the late 1940s and late 1950s, and then declines back to the level of the 1940s.

Panel C of Table 2 confirms similar patterns when using fathers' imputed income to proxy for lifetime income. The estimates are very similar to those using the trimmed sample of fathers' observed income. The elasticity estimates decline by 32 percent between the birth cohorts born in the early 1930s and the early 1940s, and remain roughly stable for the remaining birth cohorts.

Brother Correlations

As the measurement of fathers' lifetime income is incomplete in the intergenerational regressions reported earlier, it is useful to compare our intergenerational results with the estimates of brother income correlations. In Table 3, we report estimates for the components of the income variance along with estimates of the autocorrelation in the transitory shock and the overall brother correlation. The main estimates concern the family component σ_a^2 and the individual component σ_b^2. To ease comparison, we followed the cohort and income definitions in Björklund *et al.* (2009) as closely as possible (see footnote 7).

The second and third columns of Table 3 report a clear declining trend in the family component, which falls by a third between the cohorts born in the 1930s and those born in the 1940s. Therefore, brother correlation in income decreases, which provides further evidence on the decreasing importance of family background. The estimated levels and trends of brother correlations during this period are very similar to those reported by Björklund *et al.* (2009) for Sweden, suggesting that similar mechanisms

Table 3. *Brother correlations*

	Variance component				
Birth cohort	Family, σ_a^2	Individual, σ_b^2	Error, σ_v^2	Autocorrelation λ	Sibling correlation, ρ
1932–1938	0.070	0.081	0.126	0.583	0.463
	(0.005)	(0.004)	(0.004)	(0.032)	(0.024)
1935–1941	0.059	0.078	0.121	0.547	0.430
	(0.004)	(0.004)	(0.004)	(0.027)	(0.024)
1938–1944	0.050	0.076	0.122	0.567	0.397
	(0.003)	(0.004)	(0.004)	(0.026)	(0.024)
1941–1947	0.044	0.073	0.131	0.635	0.378
	(0.003)	(0.004)	(0.004)	(0.026)	(0.025)
1944–1950	0.049	0.078	0.137	0.637	0.383
	(0.003)	(0.005)	(0.004)	(0.029)	(0.029)
1947–1953	0.053	0.090	0.154	0.637	0.370
	(0.003)	(0.006)	(0.005)	(0.033)	(0.033)
1950–1956	0.052	0.096	0.166	0.669	0.350
	(0.004)	(0.006)	(0.005)	(0.031)	(0.029)
1953–1959	0.054	0.106	0.168	0.650	0.337
	(0.004)	(0.007)	(0.006)	(0.036)	(0.032)
1956–1962	0.054	0.114	0.163	0.652	0.322
	(0.004)	(0.007)	(0.005)	(0.034)	(0.028)
1959–1965	0.054	0.122	0.154	0.648	0.306
	(0.004)	(0.006)	(0.005)	(0.037)	(0.027)
1962–1968	0.056	0.122	0.146	0.646	0.315
	(0.005)	(0.004)	(0.004)	(0.024)	(0.017)

Notes: Point estimates and block bootstrapped standard errors (in parentheses) using 1,000 replications. See Section IV for details.

are likely to be behind the respective changes in social mobility in Sweden and Norway.

Interestingly, the brother correlation in income continues to decline also among the cohorts for whom father–son rank correlations and intergenerational income elasticities remain stable. These divergent patterns suggest that the importance of the factors shared by brothers, but not related to fathers' income, continued to lose their importance. This could be because of a decrease in residential income segregation, the quality of education becoming more similar across schools, or other factors relating to social class not picked up by fathers' income. However, a full examination of these potential mechanisms is beyond the scope of this study.

Figure 1 summarizes our results thus far by plotting together the rank–rank slopes, intergenerational elasticities, and brother correlations from our preferred specifications. It illustrates that while the alternative specifications yield different persistence estimates, all of the estimation approaches suggest that social mobility increased between the cohorts born in the early 1930s and the early 1940s. For the cohorts born after WWII,

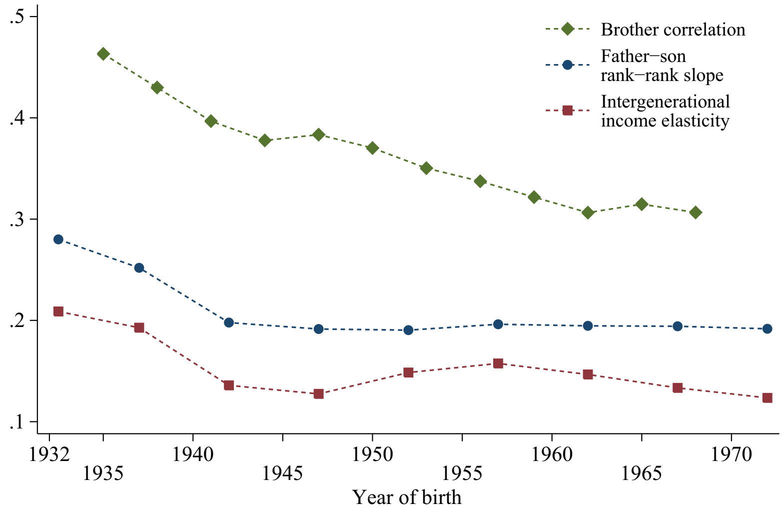

Fig. 1. Trends in social mobility

Notes: This figure presents the point estimates for the rank–rank slopes from regressing sons' income percentile on fathers' income percentile, intergenerational income elasticities from regressing sons' log income on fathers' log income, and brothers' income correlations estimated using the GMM approach in Björklund *et al.* (2009). Each estimate is from a separate regression. In the intergenerational regressions, sons' income is at age 35 and fathers' income at age 55–64 using pre-tax annual income. Intergenerational income elasticities are estimated using a sample that omits the top and bottom deciles of the fathers' income distribution. For brother correlations, we use pre-tax annual income at age 35–44 and include only brothers born within seven calendar years of each other.

the father–son associations remain stable, while the brother correlations continue to decline. The stability of father–son associations for cohorts born after WWII is in line with earlier results for the US (Aaronson and Mazumder, 2008; Lee and Solon, 2009; Chetty *et al.*, 2014) and Norway (Bratberg *et al.*, 2005). For the pre-WWII birth cohorts, the brother correlations are similar to earlier results for Sweden (see above). However, the continuing decline of brother correlations in Norway among the post-WWII birth cohorts differs from the results of Björklund *et al.* (2009), which suggest a slight increase in brother correlations in Sweden starting from the cohort born in the mid-1950s.

Trends across the Parental Income Distribution

The estimates discussed above are consistent with various patterns of mobility. For example, increases in the upward mobility of sons from low- or middle-income families or, alternatively, increased downward mobility from the top of the fathers' income distribution could drive the decline in income persistence between the cohorts born in the early 1930s and 1940s.

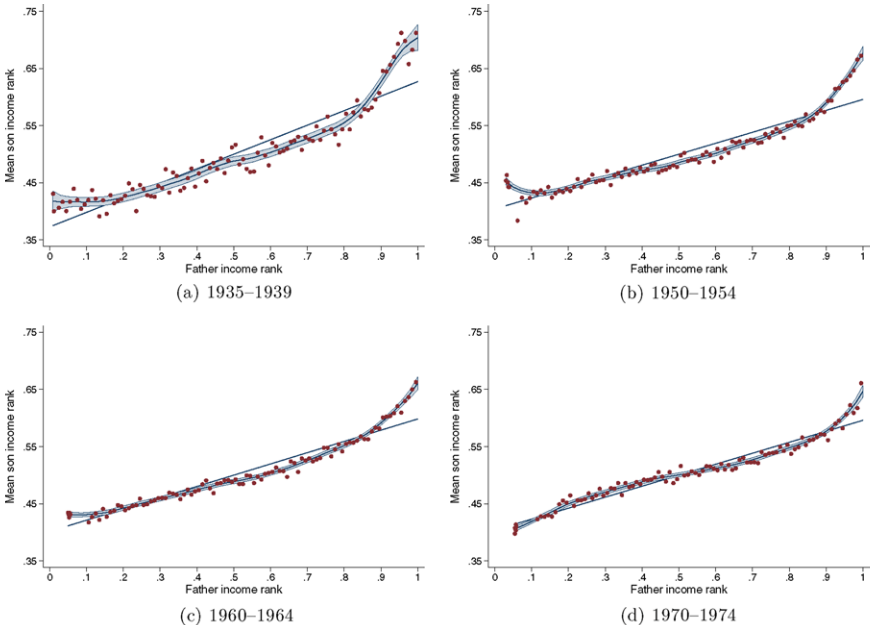

Fig. 2. Association between sons' and fathers' income percentile ranks

Notes: Sons' expected income percentile rank at age 35 as a function of fathers' income percentile rank at age 55–64. Each curve is estimated with a local linear regression using an edge (triangle) kernel and STATA's rule-of-thumb bandwidth selection routine. The shaded areas correspond to the 95 percent confidence intervals.

A shortcoming of the summary measures of mobility that Chetty *et al.* (2014) classified as "measures of relative mobility" is that they do not distinguish between these possibilities. Thus, next we focus on estimating absolute mobility measures over the fathers' income distribution.

In order to assess in which part of income the changes in mobility took place, Figure 2 presents the results for four birth cohorts by plotting sons' expected income percentile against fathers' income percentile.[10] We follow Chetty *et al.* (2014) and divide the horizontal axis into 100 percentile bins and plot the mean sons' income percentile for each bin. The figure also includes a linear fit corresponding to the rank–rank slope estimates reported in Table 1 and local linear estimates for the sons' expected income rank over the income distribution of the fathers. Table 4 reports the local linear estimates for some fathers' income percentiles for all birth cohorts included in our data.

[10] The corresponding figures for the remaining birth cohorts, and when using fathers' imputed income, are plotted in Figures A2 and A3 in the Online Appendix. We also report the transition matrices for fathers' and sons' income quintiles in Table A7.

Table 4. *Sons' expected income percentile by fathers' income percentile*

Fathers' percentile	Birth cohort 1932–1933	1935–1939	1940–1944	1945–1949	1950–1954	1955–1959	1960–1964	1965–1969	1970–1974
95th	0.672	0.676	0.647	0.634	0.629	0.626	0.621	0.610	0.603
	(0.013)	(0.005)	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
90th	0.629	0.627	0.600	0.594	0.590	0.592	0.591	0.584	0.577
	(0.010)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
75th	0.549	0.540	0.532	0.531	0.534	0.536	0.539	0.539	0.537
	(0.009)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
50th	0.490	0.488	0.482	0.486	0.485	0.488	0.489	0.492	0.500
	(0.010)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
25th	0.415	0.433	0.447	0.445	0.451	0.452	0.451	0.457	0.461
	(0.009)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
10th	0.407	0.416	0.434	0.434	0.433	0.429	0.431	0.424	0.419
	(0.010)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
5th	0.413	0.416	0.441	0.443	0.443	0.436	0.431	0.419	0.404
	(0.013)	(0.005)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)

Notes: Local linear estimates and standard errors (in parentheses). The estimates are from local linear regression using an edge (triangle) kernel and STATA's rule-of-thumb bandwidth selection routine where we regress sons' income percentile at age 35 on fathers' income percentile at age 55–64. The estimates for each column are from a separate regression.

Figure 2 and Table 4 present a complex picture of the evolution of the joint father–son income percentile distribution. The association between fathers' and sons' income percentile ranks is highly non-linear among the early cohorts, but approaches linearity over time. Nevertheless, changes in the rank–rank slope estimates (Table 1), and a comparison of the predicted percentile ranks at the bottom and the top of the fathers' income distribution (Table 4), lead to similar conclusions. Both suggest that the difference in average income ranks between sons coming from the top and the bottom of fathers' income distribution has fallen from roughly 30 to 20 percentiles. However, the expected income percentiles remain remarkably stable for sons whose fathers are between the 50th and 75th percentiles, while the expected income rank for sons of fathers at the 25th percentile steadily increases over time. Furthermore, upward mobility from the bottom of the fathers' income distribution increases among cohorts born before the early 1940s and then declines from the late 1950s birth cohort onwards. Finally, and most notably, the average income percentile of sons of the highest-income fathers declines steadily over time. For example, the expected percentile rank for sons of fathers at the 95th percentile declines from 67 for those born in the early 1930s to 60 for those born in the early 1970s.

To place our results into context, we compare them to the present-day US.[11] The expected income percentile of Norwegian men born in the 1932–1933 cohorts to fathers at the 95th income percentile, is very close to the expected percentile of Americans born in 1980–1982 in families at the 95th percentile of the parental income distribution (67 in Norway versus 66 in the US). However, the expected income percentile of Norwegians born in the 1930s to fathers at the 5th percentile, is already much higher than that in the present-day US (41 in Norway versus 34 in the US). It is also informative to contrast the changes over time in Norway to geographical variation in the US. According to the preferred measure used by Chetty *et al.* (2014) – the expected income percentile of children growing up in families at the 25th percentile – Norwegian men born in 1932–1933 experienced absolute upward mobility comparable to mid-ranking locations in the modern US, such as Denver or Buffalo. In contrast, the absolute upward mobility for Norwegian cohorts born in 1970–1974 is comparable to the most mobile locations in the US, such as Salt Lake City or Pittsburgh.

[11] The information for the US is from the Online Appendix for Chetty *et al.* (2014). It is important to keep in mind that our measures refer to the personal income of sons and their fathers, while Chetty *et al.* (2014) measure income at the family level.

VI. Education as a Potential Mechanism

While several alternative mechanisms might give rise to changes in social mobility, much of the discussion has focused on the role of human capital and changes in production technology. Theoretical work such as that of Becker and Tomes (1979), Solon (2004), Hassler *et al.* (2007), and Ichino *et al.* (2011) has shown that educational policies that decrease the cost of education for the offspring of disadvantaged families tend to increase social mobility.[12] However, changes in production technology that increase returns to skill can create incentives for poor families to invest in education, and lead to higher mobility. In this section, we present a set of stylized facts that examine these potential mechanisms. However, we stress that our analysis is purely descriptive and thus does not provide strong evidence on the causal impacts of educational reforms or changes in production processes.

Education and Income

Table 5 summarizes the trends in educational attainment in Norway over our observation period. In the first column, the average years of education increased by 2.7 years or 27 percent between those cohorts born in the early 1930s and the late 1970s. These changes are partly because of the educational reforms discussed in the second section of the paper that made attendance in secondary education universal. In addition, in the second column, the share of birth cohorts obtaining a college degree increased dramatically alongside expansion of the Norwegian college and university sector.

The next two columns of Table 5 present estimates from regressing log income at age 35 on years of education (Column 3) or an indicator for having a college degree (Column 4).[13] Between the cohorts born in the early 1930s and the late 1940s, the association between log income and years of education decreases by 18 percent and returns to a college degree by 31 percent. This change is consistent with the hypothesis that the increased supply of educated workers decreased the returns to education. However, among cohorts born after 1950, the returns to education increased substantially, even though the supply of educated workers continued to increase. This pattern is consistent with the demand for educated workers

[12] See Björklund and Salvanes (2011) for an overview of empirical research on education and family background.

[13] In Figure A4 in the Online Appendix, we show that the relationship between income and years of education is roughly linear, and thus single regression coefficients provide a meaningful summarization of the association.

Table 5. *Trends in educational attainment*

			Association between log income at 35 and		Association between fathers' income rank and			
					years of educ.		tertiary deg.	
	Years of education (1)	Tertiary degree (2)	years of education (3)	tertiary degree (4)	Cons. (5)	Slope (6)	Cons. (7)	Slope (8)
1932–1933	10.0	0.13	0.062	0.41	8.4	2.3	0.00	0.18
1935–1939	10.5	0.17	0.065	0.41	9.0	3.2	0.01	0.35
1940–1944	11.1	0.22	0.052	0.30	9.7	3.1	0.06	0.36
1945–1949	11.4	0.24	0.051	0.28	10.1	2.9	0.08	0.36
1950–1954	11.8	0.27	0.060	0.29	10.5	2.7	0.09	0.36
1955–1959	12.0	0.26	0.075	0.35	10.8	2.4	0.09	0.34
1960–1964	12.0	0.26	0.074	0.31	10.9	2.2	0.09	0.35
1965–1969	12.3	0.29	0.073	0.31	11.2	2.2	0.12	0.35
1970–1974	12.6	0.34	0.074	0.29	11.6	2.1	0.16	0.35
1975–1979	12.6	0.36	0.080	0.31	11.7	2.0	0.19	0.37

Notes: Columns 1 and 2 report the average years of education and the share obtaining a tertiary degree for each birth cohort. Column 3 reports OLS point estimates from regressing log annual income at age 35 on years of education. Column 4 reports similar estimates when using an indicator variable for tertiary degree as a measure of education. Columns 5 and 6 report the estimates from regressing sons' years of education on fathers' income rank. Columns 7 and 8 report similar estimates for sons' tertiary degree.

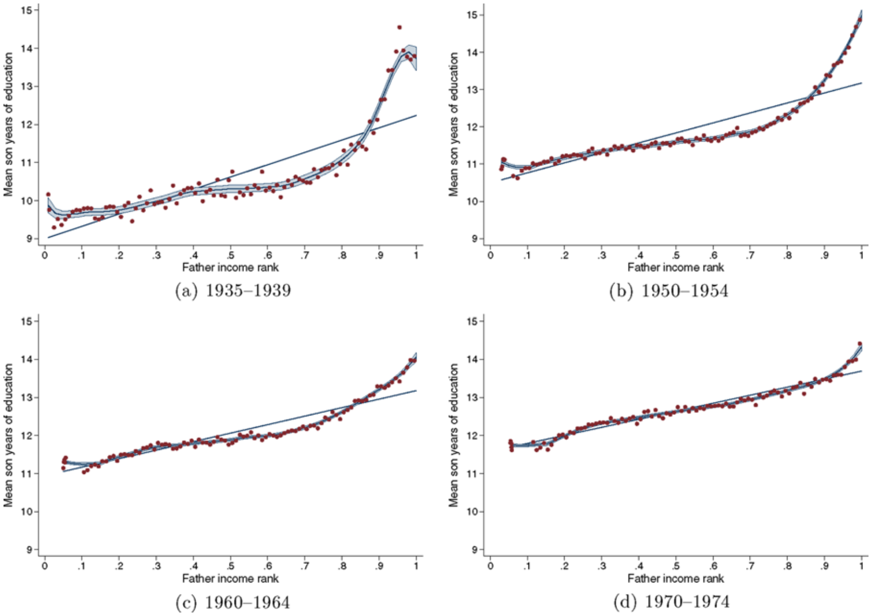

Fig. 3. Association between sons' years of education and fathers' income percentile rank
Notes: Sons' expected years of education as a function of fathers' income rank at age 55–64. Each curve is estimated with a local linear regression using an edge (triangle) kernel and STATA's rule-of-thumb bandwidth selection routine. The shaded areas correspond to the 95 percent confidence intervals.

increasing faster than its supply; see, for example, Goldin and Katz (2009) for a discussion.[14]

Education and Parental Background

We now turn to changes in the relationship between educational attainment and family background. Figure 3 and Table 6 present the results for years of education using an identical approach to that used in the final subsection of Section V for income percentile ranks. That is, we use local linear regressions to estimate the expected years of education across the fathers' income percentile.

Figure 3 reveals a highly convex relationship between parental background and years of education, particularly for the early birth cohorts. For the cohorts born between the 1930s and the 1950s, the relationship

[14] For brevity, we refer to the association between income and educational attainment as "returns to education". We recognize that this association might not measure a causal relationship, because unobserved factors are likely to affect educational choices. Furthermore, the nature of the selection process might change over time.

Table 6. *Sons' expected years of education by fathers' income percentile*

Fathers'	Birth cohort									
percentile	1932–1933	1935–1939	1940–1944	1945–1949	1950–1954	1955–1959	1960–1964	1965–1969	1970–1974	1975–1979
95th	11.7	13.6	13.9	13.9	14.0	13.8	13.5	13.6	13.8	13.8
	(0.16)	(0.06)	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
90th	10.6	12.5	12.9	13.1	13.2	13.3	13.2	13.3	13.5	13.5
	(0.12)	(0.06)	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
75th	9.8	10.8	11.4	11.7	12.0	12.2	12.4	12.8	13.1	13.2
	(0.11)	(0.05)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
50th	9.4	10.3	11.0	11.2	11.6	11.8	11.9	12.2	12.6	12.7
	(0.11)	(0.05)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
25th	9.0	9.9	10.6	10.9	11.3	11.5	11.6	11.8	12.2	12.3
	(0.11)	(0.05)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
10th	8.8	9.7	10.2	10.6	10.9	11.1	11.2	11.5	11.7	11.8
	(0.10)	(0.05)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
5th	8.9	9.6	10.2	10.5	11.0	11.2	11.3	11.5	11.8	11.7
	(0.12)	(0.05)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)

Notes: Local linear estimates and standard errors (in parentheses). The estimates are from a local linear regression using an edge (triangle) kernel and STATA's rule-of-thumb bandwidth selection routine where we regress sons' years of education on fathers' income percentile at age 55–64. The estimates for each column are from a separate regression.

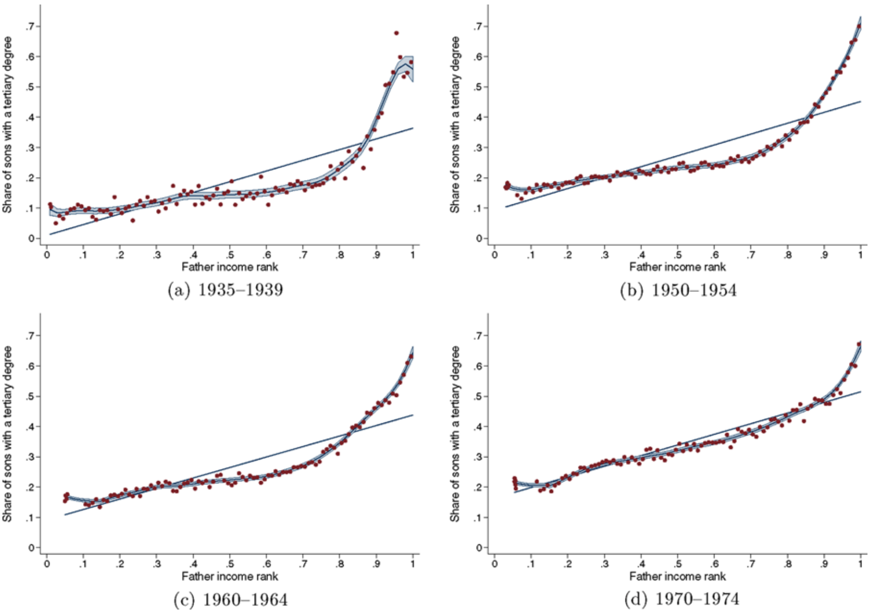

Fig. 4. Association between sons' likelihood of obtaining a college degree and fathers' income percentile rank

Notes: Sons' probability of holding a college degree as a function of fathers' income rank at age 55–64. Each curve is estimated with a local linear regression using an edge (triangle) kernel and STATA's rule-of-thumb bandwidth selection routine. The shaded areas correspond to the 95 percent confidence intervals.

is very steep above the 80th percentile rank and fairly flat below it. For the later birth cohorts, the relationship slowly becomes linear as the sons of the low- and middle-income fathers steadily increase their educational attainment, while the education of sons of high-income fathers remains remarkably stable. As a consequence, the gap between the expected years of education of sons born to fathers at the 95th and 5th percentiles decreases from three to two years between the cohorts born in the late 1930s and early 1970s (Table 6).[15]

Figure 4 and Table 7 repeat the analysis for the likelihood of the son obtaining a college degree. The pattern is qualitatively similar as for years

[15] For completeness, Columns 5 and 6 of Table 5 report estimates from regressing sons' years of education on fathers' observed income percentile. These estimates, and those reported in the first column of Table 6, indicate that men born in 1932–1933, for whom we observe fathers' income in the tax register, have low educational attainment. The most likely explanation is that for this cohort, we can observe fathers at age 55–64 in 1967 only if the father was quite young when the son was born. In the Online Appendix, we show that the expected years of education evolve smoothly over the early birth cohorts when we replicate Table 6 using fathers' imputed income.

Table 7. *Sons' likelihood of obtaining a tertiary degree by fathers' income percentile*

Fathers'	Birth cohort									
percentile	1932–1933	1935–1939	1940–1944	1945–1949	1950–1954	1955–1959	1960–1964	1965–1969	1970–1974	1975–1979
95th	0.31	0.54	0.57	0.57	0.57	0.54	0.53	0.52	0.55	0.59
	(0.02)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
90th	0.16	0.39	0.45	0.46	0.47	0.46	0.46	0.46	0.49	0.53
	(0.02)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
75th	0.08	0.19	0.26	0.28	0.30	0.29	0.30	0.36	0.40	0.46
	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
50th	0.08	0.14	0.20	0.22	0.23	0.22	0.22	0.26	0.32	0.35
	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
25th	0.05	0.11	0.16	0.18	0.19	0.18	0.19	0.22	0.26	0.30
	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
10th	0.04	0.09	0.13	0.15	0.16	0.15	0.16	0.18	0.21	0.23
	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
5th	0.04	0.09	0.13	0.15	0.17	0.17	0.17	0.18	0.22	0.22
	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Notes: Local linear estimates and standard errors (in parentheses). The estimates are from a local linear regression using an edge (triangle) kernel and STATA's rule-of-thumb bandwidth selection routine where we regress an indicator for sons holding a tertiary degree on fathers' income percentile at age 55–64. The estimates for each column are from a separate regression.

of education, but more pronounced across the fathers' income distribution. About a tenth of sons born in the 1930s into families below the 70th percentile in the fathers' income distribution had a college degree, while almost 70 percent of the sons of the highest income families did. Above the 80th percentile, the association between fathers' income rank and sons' likelihood of obtaining a college degree was very steep. The strong association at the top of the distribution remains over time, even though the pattern otherwise becomes more linear as the likelihood of obtaining a college degree increases among the sons of low-income and, particularly, middle-income fathers.

VII. Conclusions

In this paper, we have documented trends in social mobility among Norwegian men during the period when Norway transformed from a poor and relatively unequal country into one of the world's richest economies with extensive redistributive institutions. According to all of our measurement approaches, social mobility increased between the cohorts born in the early 1930s and the early 1940s. The increase in mobility coincides with equalization in educational attainment across the fathers' income distribution and a declining association between income and education. These patterns are consistent with a hypothesis that the expansion of the public provision of education simultaneously leveled educational opportunities and reduced the returns to education. However, it is important to recall that these results are purely descriptive. Thus, examining the causal impact of educational reforms affecting these birth cohorts might be a particularly promising avenue for future research.

The results for the post-WWII birth cohorts are more mixed. Father–son income correlations remained stable between the cohorts born in the late 1940s and the early 1970s, while brother correlations and the expected income ranks of sons of the highest and lowest earning fathers declined. At the same time, the returns to education increased and the educational attainment of children from low- and middle-income families increased rapidly. These patterns are consistent with a hypothesis that increasing returns to education would tend to reduce social mobility, while the continuing equalization of educational attainment would push towards higher mobility. A possible interpretation is that these forces largely offset each other during this period. However, we again stress that while the stylized facts are consistent with such an interpretation, there remains scope for future research that would put these hypotheses to a more rigorous test.

Supporting Information

The following supporting information can be found in the online version of this article at the publisher's web site.

Online Appendix

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