

Educational Transitions and Educational Inequality: A Multiple Pathways Sequential Logit Model Analysis of Finnish Birth Cohorts 1960–1985

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Abstract

We developed a multiple pathways sequential logit model for analysing social background inequality in completed education and applied it to analyse educational inequality in Finland (birth cohorts 1960–1985). Our model builds on the sequential logit model for educational transitions, originally presented by Robert D. Mare and later extended by Maarten Buis, which disaggregates inequality in completed education into the weighted sum of inequalities in the transitions leading to it. Although the educational transitions framework is popular among educational stratification researchers, its applications have almost exclusively focused on analysing inequalities in separate educational transitions. Buis presented a unifying model of inequalities in educational transitions and completed education, which gives a substantive interpretation to the weights that link them. We applied this to an educational system in which the same educational outcomes can be reached through multiple pathways. Our analysis of Finnish register data shows that intergenerational educational persistence increased, particularly among women. The main reasons are increased inequality in academic upper-secondary (gymnasium) completion and gymnasium expansion that increased the weight of this transition as well as of the transition to university. We discuss the integration of structural and allocative mechanisms in educational stratification research.

Introduction

A large body of research into educational stratification has analysed social background effects on educational transitions. This research has been motivated by Mare's (1981) disaggregation of social background inequalities in final attained education into the weighted effects of

social background on the transitions that lead to it. Social background effects on educational transitions—typically analysed through a sequence of logistic regressions—have often been argued to reveal the 'true' (Mare, 1981: p. 75) or 'pure' (Breen and Jonsson, 2005: p. 225) stratification mechanisms that allocate students

along the educational pathway. This research has produced important findings on the (in)stability of educational stratification over different stages of the educational career (e.g. Mare, 1980, 2011; Cameron and Heckman, 1998) and over the process of educational expansion (Mare, 1981; Shavit and Blossfeld, 1993; Breen and Jonsson, 2005; Hout and DiPrete, 2006).

Although analyses of inequalities in completed education remained a ‘legitimate enterprise’ (Breen and Jonsson, 2005: p. 225), regressions of completed education on social background were considered confounded by the marginal distributions of education—captured by the weights in Mare’s disaggregation—and of less sociological interest than analyses of inequalities in educational transitions. Yet social background inequalities in final educational attainment are, akin to measures of absolute intergenerational income mobility (Chetty *et al.*, 2017), salient features of intergenerational inequality that are close to the lived experience of generations.

The focus on educational transitions reduced analytical interest in the structures of educational systems that produce the marginal distributions of educational attainment (cf. Ballarino and Schadee, 2010). Changes in the volume of education—such as the general upgrading of educational attainment over time—have been strongly driven by economic and political forces (e.g. Meyer *et al.*, 1992; Schofer and Meyer, 2005; Ansell, 2008). By governing the chances of pursuing different educations, the structure of the educational system shapes the unequal chances of educational attainment.

Analyses of educational inequalities that have integrated (changes in) the volume of education and the stratified allocation of students to educational levels have remained rare (exceptions include Breen *et al.*, 2009; Ballarino *et al.*, 2009, 2014; Ballarino and Schadee, 2010). Recently, a tool for such analyses was developed by Buis (2017), who disaggregated the weights in Mare’s (1981) sequential logit model into three components—the share of a cohort at risk of an educational transition, the differentiating capacity of each transition, and the educational gains of making the transition—that give a substantive interpretation to the structural features of educational systems that shape inequality in educational attainment.

We extended the Mare–Buis sequential logit model into a multiple pathways sequential logit model that corresponds to tracked educational systems with multiple pathways to the same completed level of education. We applied the model to analyse trends in inequality in completed education over 26 birth cohorts (born 1960–1985) in Finland. Finnish educational expansion halted and even reversed during the past decade (Statistics

Finland, 2011). Although Finland is considered a country with comparative educational equality (Pfeffer, 2008), we show below that educational inequality—measured as the association between highest parental and filial education—has increased. We asked whether increasing educational inequality has resulted from increasing inequalities in educational transitions or from changes in the volume of education. Our analysis shows how analysis of structural features in educational systems can be integrated with research on inequalities in educational transitions to understand changing educational inequality.

Educational Transitions and Educational Inequality

An individual’s final level of educational attainment results from a sequence of transitions made along the educational pathway. Mare (1980, 1981) used this intuition to show how the linear regression of the highest completed education on family background can be disaggregated into the weighted sum of the effects of family background on the logged odds of making each transition along the educational pathway. Because the weights are a function of the proportions of students passing each transition, linear regressions of completed education on social background are affected by the marginal distributions of education.

Mare argued that the distribution of education across a population and, given this distribution, its allocation to different socio-demographic groups are ‘conceptually independent’ (Mare, 1981: p. 73). He considered the latter as the more intriguing quest for stratification research (also, Mare, 1993: pp. 352–353). The process of allocation of education could be modelled through a sequence of logistic regression models on making each transition that lead to the final educational outcome.

A number of studies built on Mare’s model to analyse whether inequality in educational opportunity has changed with educational expansion. Many analyses concluded that inequality in educational opportunities has remained generally stable (Raftery and Hout, 1993; Shavit and Blossfeld, 1993; Gerber, 2000; Whelan and Layte, 2002), whereas other results using the Mare model or other methods contested this conclusion (e.g. Breen and Jonsson, 2005; Ballarino *et al.*, 2009; Breen *et al.*, 2009; Ballarino and Schadee, 2010; Blossfeld *et al.*, 2015), often because of declining inequalities in making early transitions in the educational career. There need not be a straightforward relationship between changes in the volume of education and intergenerational educational inequality (Breen and Jonsson, 2005).

Although educational expansion widens access to education, the principles that allocate it by socioeconomic background need not change (Mare, 1981).

The Mare model also led to theoretical reorientations. An underlying assumption in educational transitions analyses is that individuals accumulate formal schooling through a sequence of separate and often irreversible decisions along the educational career (Mare, 1981, 1993). Although this view has been criticized as ‘myopically’ assuming that decisions are not made as a part of a long-term strategy for reaching educational goals (Cameron and Heckman, 1998), it became an important starting point in many sociological studies on educational inequality (cf., Breen and Jonsson, 2005).

Much of the theorizing on inequalities in educational transitions concerns socially stratified educational decision-making, net of the differences in performance. This has been generally conceptualized as reflecting variation in levels of education aspired by different social groups (Breen and Goldthorpe, 1997; Davies *et al.*, 2002; Pfeffer, 2008), as well as variation in the expected benefits, (direct and opportunity) costs, and assessments of the probability of success of educational choices (Raftery and Hout, 1993; Erikson and Jonsson, 1996; Breen and Goldthorpe, 1997; Davies *et al.*, 2002; Becker, 2003). Educational expansion or educational reforms can affect the inequality of educational opportunity if they affect the perceived costs, benefits, and the probability of success of education (Erikson and Jonsson, 1996; Breen and Goldthorpe, 1997; Breen and Jonsson, 2005; Pfeffer, 2008; Ballarino *et al.*, 2009). Other theoretical arguments maintained that educational expansion and reforms may not affect educational inequality, for instance if the supply of education does not surpass its demand by the more advantaged socioeconomic groups (Raftery and Hout, 1993).

Methodologically, the Mare model has been extended to consider multiple—rather than binary (continuation vs drop out)—choices that allow for more realistic analyses of educational transitions in systems with parallel tracks, common to many European countries (e.g. Breen and Jonsson, 2000; Hillmert and Jacob, 2010; Karlson, 2011). Another line of methodological work has considered whether the common finding that logit regression coefficients of social background effects wane towards higher-order transitions (Mare, 1980) results from actual changes in the effects or from bias due to selective attrition and the rescaling of logit regression coefficients along the educational pathway (e.g. Mare, 1980, 1993; Cameron and Heckman, 1998; Buis, 2011; Karlson, 2011).

The Multiple Pathways Sequential Logit Model

The perception of the distribution and allocation of education as two ‘conceptually independent’ dimensions of educational stratification (Mare, 1981: p. 73) reinforced a separation between macro-sociological analyses of educational systems and their expansion and micro-sociological ones of inequality in educational opportunity, which often focused on educational transitions (Ballarino and Schadee, 2010). Yet the association between social background and completed education is the product of allocative mechanisms as well as structural features of educational systems that govern the flows of cohorts through them. Identifying the relative importance of each provides a more comprehensive understanding of the drivers of social background inequality in educational attainment.

A tool for doing so was developed by Buis (2017), who extended Mare’s (1981) sequential logit model for educational transitions. In this model, a linear regression of completed education on social background is the weighted sum of the logit coefficients of social background on making each educational transition.

$$\frac{\partial E(L_i)}{\partial x_j} = \sum_{k=1}^K \lambda_k w_{ki}. \quad (1)$$

The marginal effect of a social background measure x_j on the expected level of educational attainment $E(L_i)$ is the sum of the weighted (w_i) logit coefficients (λ) of passing each educational transition k (Mare, 1981). The weights are a function of the conditional probabilities of passing each transition. Buis (2017) presented them as the product of three components that represent different features of the educational system at each transition point: the population at risk of making the transition, the differentiating capacity of the transition, and the expected gain in educational attainment following a successful transition. Expressing each of these components as α , β , and γ , respectively, the weight at transition k can be written as

$$w_{ki} = \alpha_{ki} \times \beta_{ki} \times \gamma_{ki}, \quad (2)$$

where

$$\alpha_{ki} = \prod_{k=1}^{K-1} \hat{p}_{ki}, \quad (3)$$

$$\beta_{ki} = \hat{p}_{ki}(1 - \hat{p}_{ki}), \quad (4)$$

and

$$\gamma_{ki} = E(L_{i|k=m}) - E(L_{i|k \neq m}). \quad (5)$$

\hat{p}_{ki} is the predicted probability of passing educational transition k . Equation (3) shows that the population at risk of making transition k (α_{ki}) is the (predicted) probability of passing the previous transitions ($k - n$). This follows from the funnel-like structure of educational systems: making an educational transition (say, to university) is conditional on successfully completing one (e.g. upper-secondary education) or several earlier transitions. The fraction of each birth cohort at risk of passing the first transition is 1, and it becomes progressively smaller at higher-order transitions. Because earlier transitions affect a larger share of each cohort than later transitions, earlier transitions contribute more to inequality in completed education through this weight component. Similarly, expansion of one educational level puts a higher fraction of a cohort at risk of the following transition and, thus, increases the contribution of the latter to inequality in completed education.

The differentiating capacity of transition k (β_{ki}) is calculated as the variance of the probability of passing that transition [$\hat{p}_{ki}(1 - \hat{p}_{ki})$] (4). The differentiating capacity is at its lowest when nearly everyone or no-one passes the transition; in such a case, social background affects the outcome only for few people and the transition has limited importance for population-level inequality. Differentiating capacity peaks (at 0.25) when half of the risk population passes it. This implies a reversed U-shape relationship between the probability of passing the transition and its contribution to educational inequality through this weight component. Likewise, the expansion of a particular educational level has a non-linear effect on educational inequality.

Finally, passing an educational transition implies an increase in attained education, but it can also open doors for further education. The expected educational gain of passing transition k , γ_{ki} , summarizes this and is calculated as the difference in the expected final outcome of those who pass the transition and those who do not (or pass some other transition when multiple options m are available). The expected gain of passing transition k can increase with an expansion of education at subsequent levels, $k + n$; on the other hand, it can decrease following an expansion of levels reached by passing alternative educational transitions.

The Mare–Buis model can be used to locate the educational transitions that matter most for educational inequality at the population level. Not all transitions are equal (Buis, 2017); for example equalization at transitions with a large weight component (such as early ones that affect the whole cohort) can reduce educational inequality in completed education even when inequality at other transitions remains stable or even increases (Breen

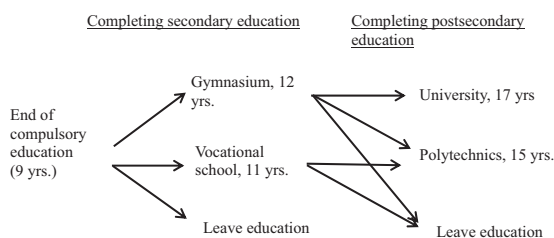


Figure 1. Compressed model of educational pathways in Finland

et al., 2009; Ballarino *et al.*, 2009, 2014; Ballarino and Schadee, 2010; Buis, 2017). Adding the demographic component into models of educational inequality can insure against erroneous conclusions based on an analysis of separate transitions alone. Furthermore, an analysis of the weight components can direct towards a closer analysis of educational systems' organizational features that shape progression along the educational pathway (cf. Sørensen, 1983; Kerckhoff, 1993; Blau, 1994).

A limitation of Buis's model (2017) and its empirical applications (Ballarino *et al.*, 2014) is that it assumed that each final level of education can only be reached through a single pathway.¹ This applies poorly to tracked educational systems, in which the same educational outcome can be reached through multiple pathways.

Finland has such a tracked education system as summarized in Figure 1 and discussed in detail in the next section. After compulsory school, students face a choice between academic upper-secondary education (gymnasium), vocational upper-secondary education, or dropping out. After gymnasium, they can continue to university or post-secondary vocational education (polytechnics), whereas vocational school graduates can continue to polytechnics. This model has some simplifications. First, we only consider transitions between completed levels of education (thus, transition k refers to completing educational level k). Second, we do not consider transitions from vocational schools to universities, as they remain rare with approximately 1–2 per cent of vocational school graduates doing so. Third, although combined academic and vocational upper-secondary education degrees are possible, we group these among the gymnasium graduates.

To apply the Mare–Buis sequential logit model to such tracked educational systems, we present what we call the multiple pathways sequential logit (MPSL) model. Starting from Buis's (2017: p. 657) decomposition of the marginal effect of social background on completed education into the sequential logit model for

single pathways, and suppressing the individual subscript i , the MPSL model applied to the Finnish case is

$$\begin{aligned}
 \frac{\partial E(L)}{\partial x_j} = & \left\{ 1 \times \hat{p}_{GYM}(1 - \hat{p}_{GYM}) \times \left[l_{GYM}(1 - \hat{p}_{POLY|GYM} - \hat{p}_{UNI}) \right. \right. \\
 & + l_{UNI}\hat{p}_{UNI} + l_{POLY}\hat{p}_{POLY|GYM} - \frac{l_{POLY}\hat{p}_{VOC}\hat{p}_{POLY|VOC}}{1 - \hat{p}_{GYM}} \\
 & \left. \left. - \frac{l_{VOC}\hat{p}_{VOC}(1 - \hat{p}_{POLY|VOC})}{1 - \hat{p}_{GYM}} - l_{COMP}(1 - \frac{\hat{p}_{VOC}}{1 - \hat{p}_{GYM}}) \right] \right\} \lambda_{GYM} \\
 & + \left\{ 1 \times \hat{p}_{VOC}(1 - \hat{p}_{VOC}) \times \left[l_{VOC}(1 - \hat{p}_{POLY|VOC}) \right. \right. \\
 & + l_{POLY}\hat{p}_{POLY|VOC} - \frac{l_{UNI}\hat{p}_{UNI}\hat{p}_{GYM}}{1 - \hat{p}_{VOC}} - \frac{l_{POLY}\hat{p}_{POLY|GYM}\hat{p}_{GYM}}{1 - \hat{p}_{VOC}} \\
 & \left. \left. - \frac{l_{GYM}(1 - \hat{p}_{POLY|GYM} - \hat{p}_{UNI})\hat{p}_{GYM}}{1 - \hat{p}_{VOC}} \right. \right. \\
 & \left. \left. - l_{COMP} \left(1 - \frac{\hat{p}_{GYM}}{1 - \hat{p}_{VOC}} \right) \right] \right\} \lambda_{VOC} \\
 & + \left\{ \hat{p}_{GYM} \times \hat{p}_{UNI}(1 - \hat{p}_{UNI}) \right. \\
 & \left. \times \left[l_{UNI} - l_{GYM} - \frac{(l_{POLY} - l_{GYM})\hat{p}_{POLY|GYM}}{1 - \hat{p}_{UNI}} \right] \right\} \lambda_{UNI} \\
 & + \left\{ \hat{p}_{GYM} \times \hat{p}_{POLY|GYM}(1 - \hat{p}_{POLY|GYM}) \right. \\
 & \left. \times \left[l_{POLY} - l_{GYM} - \frac{(l_{UNI} - l_{GYM})\hat{p}_{UNI}}{1 - \hat{p}_{POLY|GYM}} \right] \right\} \lambda_{POLY|GYM} \\
 & + \left\{ \hat{p}_{VOC} \times \hat{p}_{POLY|VOC}(1 - \hat{p}_{POLY|VOC}) \right. \\
 & \left. \times \left[l_{POLY} - l_{VOC} \right] \right\} \lambda_{POLY|VOC}, \tag{6}
 \end{aligned}$$

where l_k is the (pseudo) years to complete each educational level k , \hat{p}_k is the (predicted) probability of passing the transition to educational level k , and λ_k is the effect² of social background on transition k .

The derivation of the model is presented in Appendix A.

The main differences to the model presented by Buis (2017: p. 657) are the following. First, we allow multiple choices at each educational branching point. Indeed, the only transition with a binary choice is faced by vocational school graduates. The effect of social background on this transition is estimated using binary logit regression, whereas the other transitions are analysed using multinomial logit models (e.g. Breen and Jonsson, 2000). Second, the model allows multiple pathways to reach the same outcome. In our case, polytechnics can be reached through gymnasium or through vocational school.

The first weight component gives the population at risk of passing transition k [see (3)], and the second weight component (4) is the differentiating capacity of

transition k . The remaining part of the weight, within the square brackets, expresses the expected gain in educational attainment of passing transition k . For example, the expected gain of completing gymnasium consists of its immediate gain (the additional years of schooling), plus the expected educational outcomes that this transition opens up (i.e. to a university or polytechnics, weighted by the probability of completing this schooling after gymnasium), minus its opportunity costs. The opportunity costs are the expected outcomes from choosing the alternatives (attaining a polytechnical college after vocational school, completing vocational school and not continuing, or not continuing beyond compulsory education).

The model presented above relies on similar assumptions as Buis's (2017: p. 657) single pathway model. First, it requires that social background is measured using a continuous variable, such as years of parental education or a socioeconomic index. Although a similar model can be developed for discrete measures of social background, constraining assumptions limit its use (Buis, 2017: pp. 659–660). The models are furthermore non-linear. The non-linearity of the model means that the marginal effect of social background on final education can vary depending on the value of the social background variable. For instance, the marginal effect of adding 1 year of parental education to the lowest level of parental education can be different than the marginal effect at the average level of parental education.

Second, the effects of social background on making the educational transitions—parameters λ_k —are estimated using binary and multinomial logit models. These estimate the effect of social background on the transition on the logit scale and unlike linear probability models, for instance, separate the effect of social background on the transition from its distribution (Mare, 1981). They are thus useful for identifying the patterns of social background differences in educational transitions that are insensitive to their marginal distributions (Mare, 2011; Xie, 2011; Buis, 2017: pp. 654–655).

Education in Finland

The Finnish educational system is less differentiated and more standardized than in many other Western countries (e.g. Pfeffer, 2008). Education is free of charge and supported by student benefits as well as state-guaranteed loans.

Finnish educational expansion occurred relatively late together with a number of educational reforms. The 1970s comprehensive school reform replaced the old two-track compulsory school system with a uniform and

standardized 9-year comprehensive school (Pekkarinen *et al.*, 2009). Currently, approximately 95 per cent of each birth cohort continue immediately to upper-secondary education in gymnasiums or vocational schools and the former is more popular than the latter among female students, while the opposite is true for male students (Statistics Finland, 2010).³ Acceptance to different tracks is generally based on compulsory school grades. Upper-secondary education expansion was particularly strong from the 1960s onwards and the share of each birth cohort with the lowest (compulsory) level of education declined (Figure 2; Statistics Finland, 2011), despite a subsequent stabilization.

Gymnasium certificates give eligibility to academic and vocational post-secondary education. Gymnasium is clearly the most common pathway to the post-secondary level and to universities in particular. Admission to post-secondary education is based on upper-secondary school grades and, often, entrance examinations.

University expansion began in the 1960s and continued until the 1990s, after which it stabilized (Figure 2; Kivinen *et al.*, 2007). The main recent change in post-secondary education concerned the replacement of vocational colleges with vocationally oriented polytechnics (or ‘universities of applied sciences’). This reform standardized vocational post-secondary education into 4-year degrees and began at an experimental basis in 1991 and was fully implemented in 1996–2000. This reform lengthened some programmes and increased their theoretical content and was partly financed by the

polytechnics by reducing student intake (Hämäläinen and Uusitalo, 2008). Women have been clearly more likely to attain post-secondary education since the early-1960s cohorts (Figure 2).

Finnish educational expansion halted and partly reversed starting from the cohorts born in the mid-1970s (Figure 2). The share of each cohort completing no more than upper-secondary education by age 30 increased from the 1980 cohort, whereas the expansion of tertiary education ended among men from the 1970s cohorts on, and reversed among women from the cohorts born in the late-1970s. Thereafter, post-secondary degrees have been more common among (the highest educated) parents than their sons. The reasons for the educational contraction are not clear and debated.⁴ The deep economic recession in the beginning of the 1990s, which overlapped with the ages at which the educational contraction birth cohorts made their educational decisions, may have contributed to the trend.

Earlier studies of trends in intergenerational educational inequality in Finland have been somewhat contradictory, possibly due to methodological differences. In his 20-country comparison, Pfeffer (2008) found that educational inequality declined in Finland to an internationally low level from 1930s to the early-1970s cohorts. Kivinen *et al.*'s (2007, 2012) results of declining odds ratios between parental and filial university education drew similar conclusions. On the other hand, Karhunen and Uusitalo (2017) found—based on

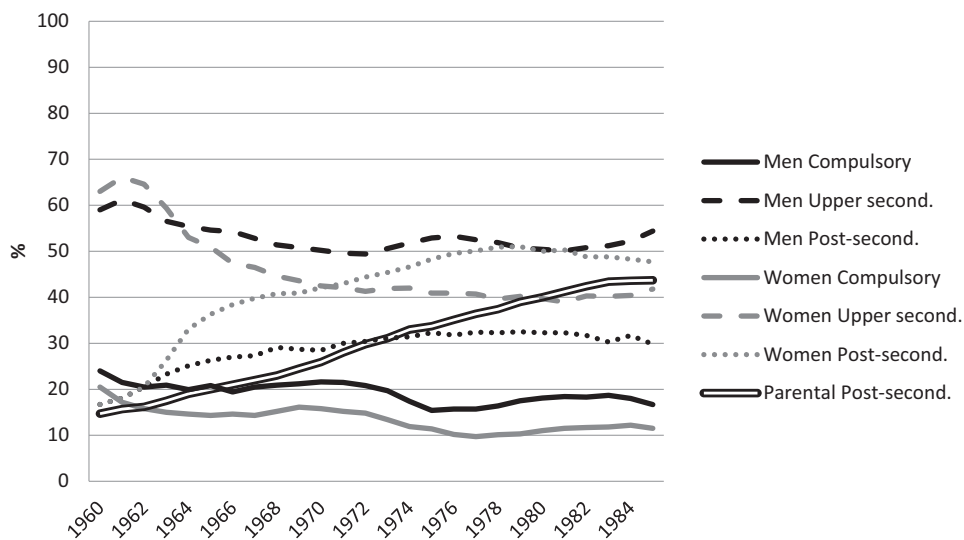


Figure 2. Highest completed educational degree among Finnish-born birth cohorts by the age of 30 years, and post-secondary education of their parents

Spearman correlations and Kendall's tau estimates—that educational inequality remained relatively stable for cohorts born in the 1950s and early-1960s and possibly increased somewhat for cohorts born after the mid-1960s. Other relevant findings include a slight recent increase in class inheritance over time (Erola, 2009), the equalizing effect of the 1970s comprehensive school reform on intergenerational income mobility (Pekkarinen *et al.*, 2009), and the rather constant trend in intergenerational income persistence for the birth cohorts schooled after the reform (Sirniö *et al.*, 2017).

Data and Analysis

We used total-population register-based longitudinal data provided by Statistics Finland. The data include annually updated information between 1987 and 2015 and are complemented with information from the 1975 and 1985 censuses.

We studied individuals born between 1960 and 1985. Only those born in Finland with at least one parent born in Finland were included (88.7 per cent of total population), because information on own and parental education is not reliable if completed outside Finland. We also excluded individuals who were institutionalized at age 30 (1.3 per cent) and those without information on parental education (2.3 per cent). Altogether, our data included 1,664,550 persons.

We analysed the association between parents' and their children's educational attainment. Information on education was obtained from the Register of Completed Education and Degrees, which includes information on all educational qualifications attained in Finland after compulsory schooling. The index person's highest level of education was measured when (s)he was 30 years old. This age restriction was done to include the youngest birth cohorts, which experienced educational contraction, in the analysis. Below, we discuss how this may affect our findings.

We focused on five levels of educational attainment. These educational levels were assigned pseudo-years of education corresponding to the formal years it takes to complete a given education. Those with compulsory schooling only (up to 9 years of education) have no registered non-compulsory education. After compulsory schooling, students face the choice of transitioning to gymnasium (12 years), vocational school (11 years), or dropping out. Gymnasium is the main pathway to university (17 years), and both gymnasium and vocational school graduates can continue to polytechnical colleges (15 years).

Parental education was indexed by the highest education of the two parents in two-parent families, and by the education of the co-residential parent in single-parent ones, and was measured—using the same values as for children's education—when the index persons were 15 years old.⁵

We analysed trends in educational inequality separately for men and women. Our analysis proceeded in three steps. First, we described trends in the intergenerational association in education using two measures: Spearman's rank correlation coefficient summarizes the strength and direction of the association between parental and filial education as ordinal variables (Karhunen and Uusitalo, 2017), and *unidiff* layer scores (Erikson and Goldthorpe, 1992; Xie, 1992) measure the association when educational is measured categorically. Both measures are robust to changes in the marginal distribution of education.

Second, we estimated the parameters for the multiple pathways sequential logit model. This involved estimating the binary logistic regression models for the transition from vocational school to polytechnical colleges, multinomial logistic regression models for transitions from compulsory to upper-secondary education and from gymnasium to post-secondary education, and the predicted probabilities (at average levels of parental education) of making each transition. These parameters were estimated separately for each transition and each birth cohort.

The resulting parameter estimates were plugged into (6) to decompose the marginal effect of an extra year of parental education at the average level of parental education. We do not present standard errors or confidence intervals. Our data cover the full population we are interested in and, thus, there is no sampling error. Even with full population data, standard errors can convey the information of uncertainty in the process that produces the observed data. However, because many of parts of the MPSL model are complex functions of parameter estimates, the estimators of the standard errors of these functions are unknown, at least to us.

Finally, we used the estimates from the above model to decompose the change in inequality in completed education (linear regressions of filial on parental schooling) into its constituent parts. Based on (1), the change in intergenerational educational inequality is the sum of the changes in the contributions of each transition. We can, furthermore, decompose the change in the contribution of each transition (here denoted by T_k) into the changes in its constituent parts following the decomposition method presented by Das Gupta (1993: pp. 9–10) and derived and explained in more detail in Appendix B. As a result, the change in the association between

parental and filial educations can be decomposed into the change in the effects of parental education on each transition, as well as the changes in the respective weight components:

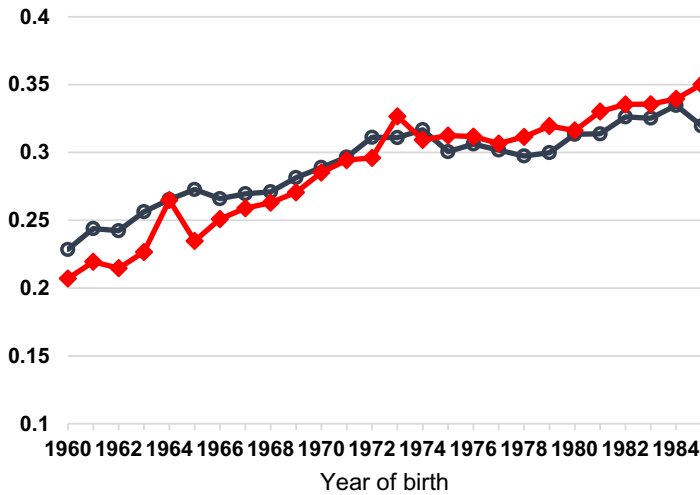
$$\begin{aligned} \Delta \left(\frac{\partial E(L)}{\partial x_j} \right) &= \sum_{k=1}^K \Delta T_k \\ &= \sum_{k=1}^K \Delta \alpha_k + \sum_{k=1}^K \Delta \beta_k + \sum_{k=1}^K \Delta \gamma_k \\ &\quad + \sum_{k=1}^K \Delta \lambda_k. \end{aligned} \tag{7}$$

Results

Descriptive Analysis: Spearman and Unidiff

Figure 3A and B describes the trends in educational inequality for women and for men using Spearman’s rank-order correlations and unidiff coefficients, respectively. The association between parental and filial educations has strengthened from the 1960 to 1985 cohorts according to both measures, the increase being stronger among women (Figure 3).

A Spearman



B Unidiff

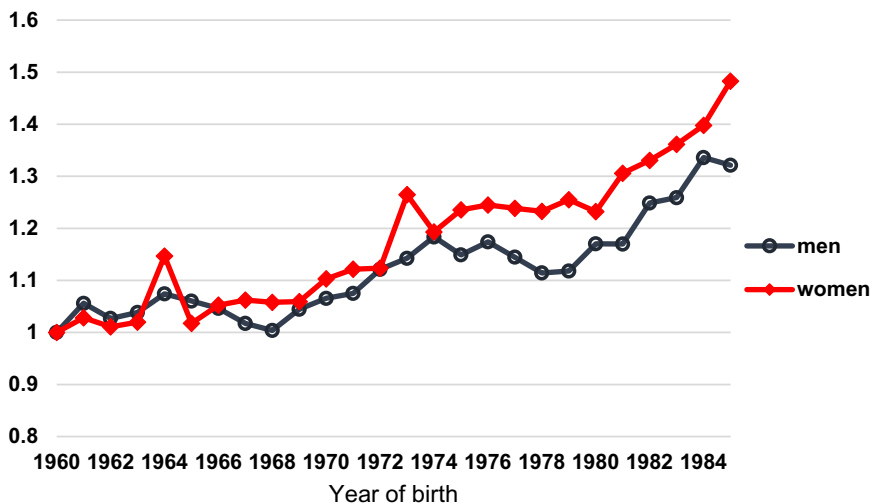


Figure 3. Spearman’s rank correlation and unidiff layer scores (ref = 1.00) from log-multiplicative uniform difference models describing the association between parental and personal education by birth cohort. (A) Spearman and (B) unidiff

The Spearman correlations increased from 0.21 (women) and 0.23 (men) in the 1960 cohort to around 0.3 in the early-1970s cohorts, and again from the late-1970s cohorts onwards reaching 0.35 for women and 0.32 for men in the 1985 cohort. The 1960 cohort was taken as the reference category for the unidiff models (1.00). Although the trend is somewhat choppier, one can detect a first larger increase from the late-1960s cohorts to the mid-1970s cohorts and another one from the 1980 cohort onwards. The unidiff coefficients for women and men, respectively, reached 1.48 and 1.32 in the 1985 cohort.

Which Transitions Matter?

Figure 4A (men) and B (women) presents the MPSL decomposition of the linear regression of the (pseudo)

years of completed education on parental education into the contributions of each educational transition. These were estimated at the average level of parental education. Despite being based on different scales, Figure 4 reproduces the overall finding of increased inequality in educational attainment and that the increase in educational persistence has been stronger among women.

The transition from compulsory education to gymnasium is, by far, the most important one for inequality in completed education. It alone accounts for 80–90 per cent of the inequality in educational outcomes among men and 70–80 per cent among women. The other transitions are less important. The transition from gymnasium to university is more important for women than for men, and the opposite is true for the transition from

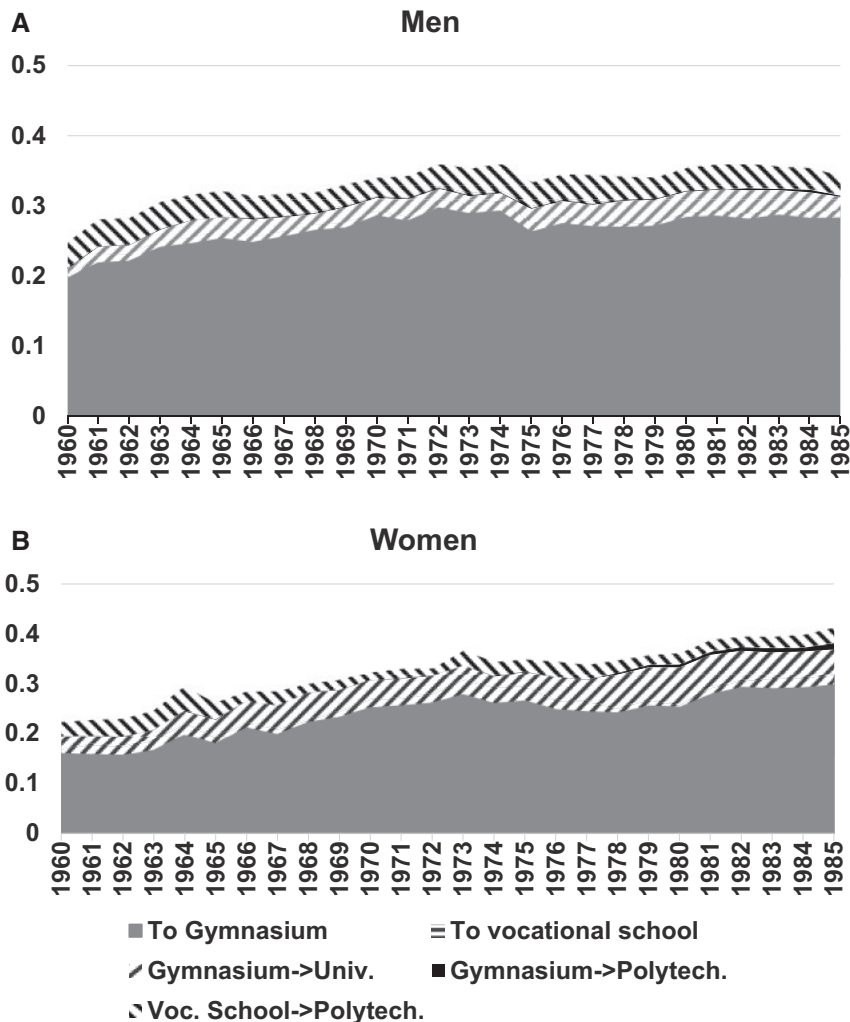


Figure 4. Contribution of educational transitions to the marginal effect of parental education on filial education. (A) Men and (B) women

vocational school to polytechnics. Neither transition contributes more than 20 per cent to the overall association. The importance of early educational transitions on intergenerational educational inequality has been reported in other countries by analyses using the sequential logit decomposition (Ballarino *et al.*, 2014; Buis, 2017) or other methods (Ballarino *et al.*, 2009; Breen *et al.*, 2009; Ballarino and Schadee, 2010). Our MPSL model decomposition shows that not all early transitions are equal, with one (to gymnasium) contributing much more than the other (to vocational school).

What makes the transition to gymnasium so important? To answer this question, we decomposed each transition's contribution to its constituent parts for selected birth cohorts. Figure 5A (men) and B (women) visualizes the importance of the structural factors (weights) and allocative mechanisms (binary and multinomial logit regression coefficients) at each transition. Table 1 presents the (binomial and multinomial) logit estimates as well as a further decomposition of the weights into the three components: the cohort share at risk of the transition, the differentiating effect of the transition, and the gain from making the transition. The last row (Prob) gives the fraction of the at-risk population that makes each transition.

The areas covered in each graph in Figure 5A and B highlight how the contribution of each transition is the product of the effect of parental education and its weight. Parental education has a strong effect on the transition to gymnasium, particularly for men, and this effect has become stronger over time. A closer look suggests that inequality in the transition to gymnasium increased when the probability of making this transition stalled or decreased (such as between the 1979 and 1985 cohorts) and decreased when the probability of the transition to gymnasium increased (for instance from the 1960 to the 1967 cohorts). The large weight of this transition amplifies its strong parental background effect. Each of the three weight components is large (Table 1). Entire birth cohorts are at risk of this transition, and a large share of each cohort (half of the female cohorts from 1967 onwards) completed it, strengthening its differentiating capacity. Gymnasium opens doors to further education, which increases the expected educational gain of a gymnasium degree.

The overwhelming importance of the transition to gymnasium contrasts with the next-to-nothing importance of the transition to vocational school. Indeed, the contribution is negative, meaning that this transition contributes to equalization in completed education. Parental education has a weak effect on the transition to vocational school—resonating with Breen's and

Jonsson's (2000) results for Sweden—and its expected educational gain is negative (Table 1). The negative expected gain reflects the low transition probabilities from vocational school to polytechnics, as well as the opportunity costs of choosing vocational school over gymnasium.

The transition to university explains up to 20 per cent of inequality in completed education for women and up to 10 per cent for men. Parental education has a strong effect on this transition, for women especially. For most cohorts—a clear exception being the 1985 male cohort—the weight of this transition is relatively large. All gymnasium graduates are at risk of making this transition, and gymnasium expansion increased this share. Graduating from university brings obvious educational gains, whereas the relatively modest—and for men especially, decreasing—share of gymnasium graduates who continue to university reduces its differentiating capacity. Altogether, the 'academic path'—the transition to gymnasium and later to university—explains almost all of the inequality in completed education in Finland.

Parental education has no effect on the transition from gymnasium to polytechnics; indeed, in many cohorts, the effect is even negative. Like us, Breen and Jonsson (2000) found that social background has a stronger effect on continuing to university than to vocational post-graduate education after gymnasium, but our result stands out by reporting no effect. Together, our results of no or very weak parental education effects on transitions to vocationally oriented educations suggest a persistent hierarchy in educational choices, in which individuals from higher social backgrounds choose academic tracks over vocational ones when possible. The weight of this transition has increased for men especially, tracking the increase in gymnasium graduates who make this transition (increasing its differentiating capacity) and the increasing gain of making this transition (shadowing the decrease in male gymnasium graduates continuing to university, which decreased the opportunity costs of continuing to polytechnics). However, the small parental education effect minimizes the overall contribution of this transition on inequality in completed education.

Parental education has a larger effect on continuing to polytechnics after vocational school, and for men, this is of similar size as for the transition from gymnasium to university. The probability of making this transition is lower than the one from gymnasium to polytechnics, and social background plays a stronger role for this less conventional pathway to vocational post-secondary education (cf. Breen and Jonsson, 2000).

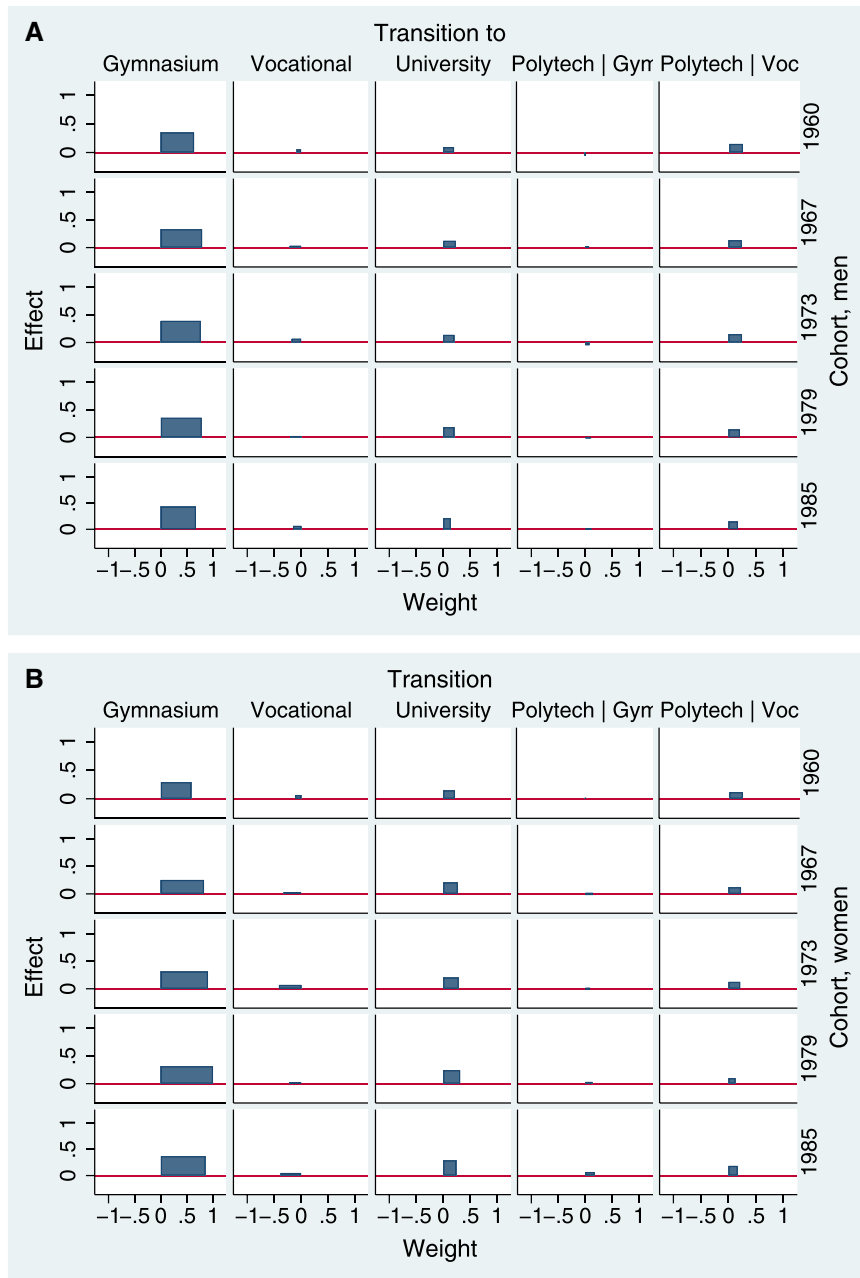


Figure 5. Decomposition of parental education into effects and weights

The weight of this transition is—for men particularly—similar to the weight of the transition to university, due to the clear gain from completing it as well as the relatively high share of each cohort at risk.⁶

Would we have reached the same conclusions had we focused only on the logit estimates, as done in most educational transitions research? This way we would have,

correctly, identified the transition to gymnasium as the most unequalizing one, and the transitions to vocational school and from gymnasium to polytechnics as the least consequential. However, we may also have given the transitions to university and from vocational school to polytechnics too much emphasis as contributors to educational inequality at the population level, given their

Table 1. Detailed decomposition of the contribution of parental education on each transition among selected cohorts

	Compulsory to upper secondary		Upper-secondary to post-secondary		
	Gymnasium	Vocational	Gymnasium to		Vocational to
			University	Polytechnic	Polytechnic
Men					
1960					
Effect	0.350	0.054	0.095	-0.060	0.151
Weight	0.638	-0.083	0.202	0.019	0.242
At risk	1.000	1.000	0.215	0.215	0.574
Differentiate	0.169	0.245	0.238	0.169	0.106
Gain	3.777	-0.341	3.944	0.522	4.000
Total contrib.	0.223	-0.005	0.019	-0.001	0.037
Prob	0.215	0.574	0.389	0.215	0.120
1967					
Effect	0.333	0.036	0.126	-0.019	0.134
Weight	0.792	-0.212	0.233	0.072	0.243
At risk	1.000	1.000	0.337	0.337	0.473
Differentiate	0.223	0.249	0.195	0.230	0.128
Gain	3.546	-0.852	3.568	0.928	4.000
Total contrib.	0.264	-0.008	0.029	-0.001	0.033
Prob	0.337	0.473	0.266	0.358	0.151
1973					
Effect	0.394	0.072	0.140	-0.053	0.152
Weight	0.767	-0.176	0.205	0.078	0.247
At risk	1.000	1.000	0.320	0.320	0.481
Differentiate	0.218	0.250	0.184	0.237	0.128
Gain	3.525	-0.705	3.478	1.028	4.000
Total contrib.	0.302	-0.013	0.029	-0.004	0.037
Prob	0.320	0.481	0.243	0.384	0.151
1979					
Effect	0.359	0.026	0.185	-0.013	0.144
Weight	0.783	-0.204	0.207	0.108	0.210
At risk	1.000	1.000	0.353	0.353	0.453
Differentiate	0.228	0.248	0.166	0.237	0.116
Gain	3.429	-0.823	3.538	1.293	4.000
Total contrib.	0.281	-0.005	0.038	-0.001	0.030
Prob	0.353	0.453	0.210	0.385	0.134
1985					
Effect	0.435	0.068	0.214	0.011	0.150
Weight	0.674	-0.151	0.141	0.116	0.173
At risk	1.000	1.000	0.291	0.291	0.513
Differentiate	0.206	0.250	0.135	0.239	0.084
Gain	3.268	-0.606	3.595	1.674	4.000
Total contrib.	0.293	-0.010	0.030	0.001	0.026
Prob.	0.291	0.513	0.161	0.393	0.093
Women					
1960					
Effect	0.290	0.066	0.147	0.017	0.116
Weight	0.584	-0.114	0.217	0.025	0.241
At risk	1.000	1.000	0.221	0.221	0.601
Differentiate	0.172	0.240	0.231	0.134	0.100
Gain	3.393	-0.475	4.249	0.851	4.000

(continued)

Table 1. (Continued)

	Compulsory to upper secondary		Upper-secondary to post-secondary		
	Gymnasium	Vocational	Gymnasium to		Vocational to
			University	Polytechnic	Polytechnic
Total contrib.	0.169	-0.008	0.032	0.000	0.028
Prob	0.221	0.601	0.361	0.160	0.113
1967					
Effect	0.253	0.028	0.208	-0.001	0.121
Weight	0.828	-0.323	0.270	0.146	0.229
At risk	1.000	1.000	0.514	0.514	0.360
Differentiate	0.250	0.230	0.161	0.248	0.159
Gain	3.314	-1.530	3.277	1.142	4.000
Total contrib.	0.209	-0.010	0.056	-0.000	0.028
Prob	0.514	0.360	0.201	0.459	0.198
1973					
Effect	0.314	0.071	0.204	-0.001	0.127
Weight	0.904	-0.414	0.286	0.089	0.216
At risk	1.000	1.000	0.508	0.508	0.365
Differentiate	0.250	0.232	0.181	0.250	0.148
Gain	3.615	-1.785	3.102	0.702	4.000
Total contrib.	0.284	-0.029	0.058	-0.000	0.027
Prob	0.508	0.365	0.238	0.482	0.181
1979					
Effect	0.315	0.029	0.245	0.033	0.100
Weight	1.007	-0.222	0.315	0.133	0.128
At risk	1.000	1.000	0.579	0.579	0.213
Differentiate	0.244	0.168	0.171	0.249	0.150
Gain	4.133	-1.323	3.179	0.918	4.000
Total contrib.	0.317	-0.006	0.077	0.004	0.013
Prob	0.579	0.213	0.219	0.474	0.184
1985					
Effect	0.371	0.051	0.292	0.069	0.187
Weight	0.856	-0.382	0.243	0.174	0.169
At risk	1.000	1.000	0.519	0.519	0.351
Differentiate	0.250	0.228	0.143	0.250	0.120
Gain	3.431	-1.678	3.266	1.343	4.000
Total contrib.	0.318	-0.019	0.071	0.012	0.032
Prob	0.519	0.351	0.173	0.478	0.140

relatively large logit coefficients (cf. Ballarino and Schadee, 2010). Our model—as well as its predecessor, Buis's (2017) model—emphasizes the transitions where the structural effects are large and disaggregates these structural effects to their constitutive components.

Finally, the estimated probabilities of the different transitions suggest sources of the recent educational contraction in Finland. The popularity of gymnasium decreased starting from the 1979 cohort. Although this was mainly offset by an increase in vocational school popularity—so that the cohort shares not completing upper-secondary education remained relatively stable at 20 per cent (men) and 11–13 per cent (women)—this

trend accompanied by the decrease in progression from gymnasium to graduation from post-secondary education contributed to educational contraction. Furthermore, fewer vocational school graduates graduated from polytechnical colleges; this decrease overlaps with the polytechnics reform that affected cohorts born from the late-1970s onwards. In other words, the pipe has leaked from several parts.

These results may partly reflect our choice of measuring educational attainment at age 30. Indeed, in additional analyses (not reported here), we found indications of decreasing completion rates from post-secondary education by age 30, particularly among men.

Therefore, extending the follow-up period with later data may partly correct the picture of contraction in completed education across cohorts. However, as we explain in more detail,⁷ this would likely lead to an even stronger increase in educational inequality.

Decomposing Change in Inequality of Educational Outcomes

Why did inequality in completed education increase in Finland? The previous analyses offer cues of the important factors, and Table 2 presents a numerical decomposition of the change into its sources between the 1960 and 1973 and 1960 and 1985 cohorts, respectively.

Changes in the transition to gymnasium stand out; for men, the total share attributable to changes in this transition adds up to above 100 per cent. The increased importance of this transition stemmed both from allocative and structural factors. The increased effect of parental education on completing gymnasium was particularly important among men. Moreover, the differentiating capacity of this transition increased as gymnasium expanded (in the 1960s), making it the single most important explanation for the growth in inequality in completed education. These trends were partly offset by the decreased expected gain to gymnasium among men—which overlaps with the decreased probability of continuing to university.

Table 2. Decomposition of total change in linear regression of pseudo-years in education on parental education: 1960–1973 and 1960–1985

	Men				Women			
	1960–1973		1960–1985		1960–1973		1960–1985	
	Effect	Per cent	Effect	Per cent	Effect	Per cent	Effect	Per cent
Total change	0.072	100	0.067	100	0.118	100	0.191	100
Compulsory to gymnasium								
Effect	0.031	42.9	0.056	83.3	0.017	15.1	0.058	30.6
At risk	0	0	0	0	0	0	0	0
Differentiate	0.066	91.7	0.052	77.0	0.083	69.8	0.088	45.9
Gain	−0.018	−25.4	−0.038	−55.8	0.015	12.3	0.003	1.4
Total	0.079	109.3	0.070	104.4	0.115	97.3	0.148	77.9
Compulsory to vocational								
Effect	−0.002	−3.2	−0.002	2.4	−0.001	−1.1	0.004	2.0
At risk	0	0	0	0	0	0	0	0
Differentiate	−0.000	−0.2	−0.000	−0.4	0.001	0.5	0.001	0.4
Gain	−0.006	−7.6	−0.004	−5.8	−0.021	−17.8	−0.017	−8.8
Total	−0.008	−11.1	−0.006	−8.5	−0.022	−18.4	−0.012	−6.4
Gymnasium to university								
Effect	0.009	12.9	0.021	31.1	0.015	12.9	0.036	19.0
At risk	0.010	13.2	0.008	11.8	0.038	32.1	0.046	23.4
Differentiate	−0.006	−8.7	−0.015	−22.7	−0.012	−9.8	−0.027	−14.2
Gain	−0.003	−4.3	−0.002	−3.7	−0.015	−12.8	−0.015	−7.7
Total	0.009	13.2	0.011	16.5	0.026	22.4	0.039	20.5
Gymnasium to polytechnic								
Effect	0.000	0.4	0.004	6.4	−0.001	−0.8	0.004	2.3
At risk	−0.001	−1.3	−0.000	−0.4	0.000	0.3	0.003	1.6
Differentiate	−0.001	−1.1	−0.000	−0.5	0.000	0.2	0.002	1.2
Gain	−0.002	−2.3	−0.002	−2.5	−0.000	−0.1	0.002	0.9
Total	−0.003	−4.3	0.002	2.9	−0.001	−0.4	0.012	6.1
Vocational to polytechnic								
Effect	−0.006	−8.2	−0.000	−0.3	0.003	2.2	0.015	7.8
At risk	−0.006	−8.3	−0.003	−5.1	−0.014	−12.1	−0.017	−8.8
Differentiate	0.007	9.2	−0.007	−10.3	0.011	9.5	0.006	3.0
Gain	0	0	0	0	0	0	0	0
Total	−0.005	7.3	−0.011	−15.8	−0.000	−0.4	0.004	1.9

Gymnasium expansion increased educational inequality also by increasing the cohort share at the risk of continuing to university. This accounted for up to one third of the growth in women's educational inequality (1960–1973). The effect of parental education on continuing to university contributed to increased inequality among men, whereas the smaller shares of gymnasium graduates that continued to university reduced this transition's differentiating capacity. Overall, gymnasium expansion and increasing inequality in making the transitions along the academic track produced the increase in inequality in completed education in Finland.

Most of the other transitions had a negligible contribution to changes in educational inequality. Among women, the transition to vocational school decreased inequality in completed education between the 1960 to 1973 cohorts, because of the decreased expected gains of making this transition. Among men, the transition from vocational school to polytechnic colleges reduced educational inequality from the 1960 to 1985 cohorts, both because of a decrease in the cohort share completing vocational school (which decreased the population at risk of the transition) and a decrease in the share of vocational school graduates who continued to polytechnics (which reduced the differentiating capacity of this transition).

Discussion

Many educational stratification researchers have analysed the effects of social background on educational transitions. Linear regressions of completed education are affected by the marginal distributions of education, whereas logistic regressions of educational transitions are not. For this reason, the latter have been considered better measures of inequality in educational opportunities and given analytical primacy (Mare, 1980; 1981; 1993; Breen and Jonsson, 2005). As a consequence, analyses of social background inequalities in completed education have received less sociological interest. Likewise, the marginal distributions of education—shaped by demand for education and the structure of educational systems—have been considered confounders to be controlled away in statistical analysis rather than substantively interesting factors that can shape educational inequality.

In this study, we extended the sequential logit model, first introduced by Mare (1980, 1981) and later developed by Buis (2017), into a multiple pathways sequential logit model. Sequential logit models can be used to decompose the marginal effect of social background on completed education into the sum of weighted effects of

social background on (the logged odds of) the educational transitions that lead to completed education. Buis's (2017) contribution was to give the weights—that capture the marginal distribution of education—substantive interpretations. Our methodological contribution was to extend this model for tracked educational systems where the same completed education can be reached through multiple pathways. Our substantive contribution was to use this model to analyse the increase in the association between parental and filial education in Finland for cohorts born between 1960 and 1985.

Finland is generally considered a country of low and either decreasing (e.g. Kivinen *et al.*, 2007; Pfeffer, 2008) or stable educational inequality (Karhunen and Uusitalo, 2017). We analysed more recent birth cohorts and found that inequality in completed education increased from cohorts born in 1960s to those born in 1985, and more so for women. This result was robust to different measures (linear regressions, Spearman correlations, and unidiff layer score estimates). However, the early-1960s cohorts may have been among the educationally most mobile ones (Pfeffer, 2008; Karhunen and Uusitalo, 2017), also in international comparison (Hertz *et al.*, 2007; Pfeffer, 2008; Barone and Ruggera, 2018). The Spearman correlations strengthened from just above 0.2 up to a moderate 0.32–0.35. Educational inequality in Finland thus increased from a very low to a more moderate level, possibly coming close to those of other countries. Nevertheless, clear increases in intergenerational educational inheritance are remarkable enough after a long earlier trend of educational equalization both in Finland and internationally (Pfeffer, 2008; Breen *et al.*, 2009; Barone and Ruggera, 2018).

By far the most important source of the increase in educational inequality was the transition to gymnasium after compulsory school. Inequality in completing this transition by parental education increased from the early-1970s cohorts onwards. This can reflect multiple processes, such as unmet demand for gymnasium education when the shares of cohorts that made this transition stalled or decreased while levels of parental education continued to increase (cf. Raftery and Hout, 1993; Bernardi *et al.*, 2018), rural–urban migration trends, or changing selectivity at levels of parental education resulting from increased mobility in previous generations (Nyblom and Stuhler, 2013). The increased importance of the transition to gymnasium also reflected structural factors, particularly among women. Gymnasium expansion during the 1960s—which among women came to cover half of each cohort—meant that the transition to gymnasium became a stronger divider of cohorts' educational careers. Gymnasium expansion also increased the share

of each cohort that could continue to university, further increasing the importance of this transition for educational inequality at the population level. These allocative and structural mechanisms along the academic educational pathway were responsible for the increase in inequality in completed education. This increase would have been even larger without offsetting mechanisms, such as a decrease in the expected educational gain for men's gymnasium completion.

Sequential logit decompositions are among the methods that relate inequality in specific educational transitions to population-level educational inequality (cf. Ballarino *et al.*, 2009; Ballarino and Schadee, 2010; Buis, 2017). These methods allow an apt description of the reproduction process of educational inequality over the life course. They not only highlight inequalities at different educational branching points but also summarize the important structural features of educational systems that contribute to educational inequality. In this respect, they merge the tradition of research on inequalities in educational transitions with a structural-demographic analysis of the (changing) flows through the process of schooling attainment, and the processes that govern these flows and their consequences. Such analysis can help in highlighting how changes in the volume (such as expansion of contraction) of different levels of education shape inequality in completed education both through allocative and structural factors.

Appendix A: Derivation of the Multiple Pathways Sequential Logit Model

The multiple pathways sequential logit model extends the sequential logit model presented by Buis (2017: p. 657) and allows for multiple pathways to reach the same level of educational attainment. The derivations below have been adapted from the supplementary material to Buis (2017).

Suppressing the individual subscript i , expected educational attainment can be written as the weighted sum of the pseudo-years it takes to attain each educational level where the weights are the (predicted) probabilities of the transitions needed to reach that level:

$$\begin{aligned}
 E(L) = & (1 - \hat{p}_{\text{GYM}} - \hat{p}_{\text{VOC}})l_{\text{COMP}} \\
 & + \hat{p}_{\text{GYM}}(1 - \hat{p}_{\text{UNI}} - \hat{p}_{\text{POLY|GYM}})l_{\text{GYM}} \\
 & + \hat{p}_{\text{VOC}}(1 - \hat{p}_{\text{POLY|VOC}})l_{\text{VOC}} + \hat{p}_{\text{GYM}}\hat{p}_{\text{UNI}}l_{\text{UNI}} \\
 & + \hat{p}_{\text{GYM}}\hat{p}_{\text{POLY|GYM}}l_{\text{POLY}} + \hat{p}_{\text{VOC}}\hat{p}_{\text{POLY|VOC}}l_{\text{POLY}},
 \end{aligned}
 \tag{A.1}$$

where COMP, HS, VOC, POLY, and UNI denote compulsory schooling, gymnasium, vocational education,

polytechnics, and university, respectively. l is the pseudo-years of each educational level, and \hat{p} are the (predicted) probabilities of each transition. Taking the derivative of this by family background (x_j) and using the sum rule, we get the marginal effect of family background on completed education:

$$\begin{aligned}
 \frac{\partial E(L)}{\partial x_j} = & \frac{\partial(1 - \hat{p}_{\text{GYM}} - \hat{p}_{\text{VOC}})l_{\text{COMP}}}{\partial x_j} \\
 & + \frac{\partial \hat{p}_{\text{GYM}}(1 - \hat{p}_{\text{UNI}} - \hat{p}_{\text{POLY|GYM}})l_{\text{GYM}}}{\partial x_j} \\
 & + \frac{\partial \hat{p}_{\text{VOC}}(1 - \hat{p}_{\text{POLY|VOC}})l_{\text{VOC}}}{\partial x_j} \\
 & + \frac{\partial \hat{p}_{\text{GYM}}\hat{p}_{\text{UNI}}l_{\text{UNI}}}{\partial x_j} + \frac{\partial \hat{p}_{\text{GYM}}\hat{p}_{\text{POLY|GYM}}l_{\text{POLY}}}{\partial x_j} \\
 & + \frac{\partial \hat{p}_{\text{VOC}}\hat{p}_{\text{POLY|VOC}}l_{\text{POLY}}}{\partial x_j}.
 \end{aligned}
 \tag{A.2}$$

Using the product rule, and rearranging terms, the model becomes

$$\begin{aligned}
 \frac{\partial E(L)}{\partial x_j} = & \frac{\partial \hat{p}_{\text{GYM}}}{\partial x_j} \left[(1 - \hat{p}_{\text{UNI}} - \hat{p}_{\text{POLY|GYM}})l_{\text{GYM}} \right. \\
 & \left. + \hat{p}_{\text{UNI}}l_{\text{UNI}} + \hat{p}_{\text{POLY|GYM}}l_{\text{POLY}} - l_{\text{COMP}} \right] \\
 & + \frac{\partial \hat{p}_{\text{VOC}}}{\partial x_j} \left[(1 - \hat{p}_{\text{POLY|VOC}})l_{\text{VOC}} \right. \\
 & \left. + \hat{p}_{\text{POLY|VOC}}l_{\text{POLY}} - l_{\text{COMP}} \right] \\
 & + \frac{\partial \hat{p}_{\text{UNI}}}{\partial x_j} \left[(l_{\text{UNI}} - l_{\text{GYM}})\hat{p}_{\text{GYM}} \right] \\
 & + \frac{\partial \hat{p}_{\text{POLY|GYM}}}{\partial x_j} \left[(l_{\text{POLY}} - l_{\text{GYM}})\hat{p}_{\text{GYM}} \right] \\
 & + \frac{\partial \hat{p}_{\text{POLY|VOC}}}{\partial x_j} \left[(l_{\text{POLY}} - l_{\text{VOC}})\hat{p}_{\text{VOC}} \right].
 \end{aligned}
 \tag{A.3}$$

$\frac{\partial \hat{p}_k}{\partial x_j}$ show the marginal effects of family background on the probability of making a transition. Using the chain rule, the transition from vocational school to polytechnics is estimated using a binary logit model, and the corresponding effect is written as

$$\begin{aligned}
 \frac{\partial \hat{p}_{\text{POLY|VOC}}}{\partial x_j} = & \frac{\partial g^{-1}(\eta_{\text{POLY|VOC}})}{\partial \eta_{\text{POLY|VOC}}} \times \frac{\partial \eta_{\text{POLY|VOC}}}{\partial x_j} \\
 = & \lambda_{\text{POLY|VOC}} \hat{p}_{\text{POLY|VOC}} (1 - \hat{p}_{\text{POLY|VOC}}),
 \end{aligned}
 \tag{A.4}$$

where $\eta_{\text{POLY|VOC}}$ is the logged odds of the transition, $g^{-1}(\cdot)$ is the inverse link function of the logit model, and $\lambda_{\text{POLY|VOC}}$ is the logit coefficient. The other transitions

involve choosing between two available options (and dropping out) and are modelled using multinomial logit regressions. The corresponding derivations for transition k are

$$\frac{\partial \hat{p}_k}{\partial x_j} = \frac{\partial g^{-1}(\eta_k)}{\partial \eta_k} \frac{\partial \eta_k}{\partial x_j} = \lambda_k \hat{p}_k (1 - \hat{p}_k) - \hat{p}_k \lambda_{k \neq m} \hat{p}_{k \neq m}. \quad (\text{A.5})$$

Here, $\lambda_{k \neq m}$ and $\hat{p}_{k \neq m}$ are the multinomial logit coefficients for and probabilities of, respectively, the alternative educational transition (e.g. choosing vocational education instead of gymnasium).

Plugging in (A.4) and (A.5) and rearranging the terms, we arrive at (6):

$$\begin{aligned} \frac{\partial E(L)}{\partial x_j} = & \left\{ 1 \times \hat{p}_{\text{GYM}} (1 - \hat{p}_{\text{GYM}}) \times \left[l_{\text{GYM}} (1 - \hat{p}_{\text{POLY|GYM}} - \hat{p}_{\text{UNI}}) \right. \right. \\ & + l_{\text{UNI}} \hat{p}_{\text{UNI}} + l_{\text{POLY}} \hat{p}_{\text{POLY|GYM}} - \frac{l_{\text{POLY}} \hat{p}_{\text{VOC}} \hat{p}_{\text{POLY|VOC}}}{1 - \hat{p}_{\text{GYM}}} \\ & \left. \left. - \frac{l_{\text{VOC}} \hat{p}_{\text{VOC}} (1 - \hat{p}_{\text{POLY|VOC}})}{1 - \hat{p}_{\text{GYM}}} - l_{\text{COMP}} \left(1 - \frac{\hat{p}_{\text{VOC}}}{1 - \hat{p}_{\text{GYM}}} \right) \right] \right\} \lambda_{\text{GYM}} \\ & + \left\{ 1 \times \hat{p}_{\text{VOC}} (1 - \hat{p}_{\text{VOC}}) \times \left[l_{\text{VOC}} (1 - \hat{p}_{\text{POLY|VOC}}) \right. \right. \\ & + l_{\text{POLY}} \hat{p}_{\text{POLY|VOC}} - \frac{l_{\text{UNI}} \hat{p}_{\text{UNI}} \hat{p}_{\text{GYM}}}{1 - \hat{p}_{\text{VOC}}} - \frac{l_{\text{POLY}} \hat{p}_{\text{POLY|GYM}} \hat{p}_{\text{GYM}}}{1 - \hat{p}_{\text{VOC}}} \\ & \left. \left. - \frac{l_{\text{GYM}} (1 - \hat{p}_{\text{POLY|GYM}} - \hat{p}_{\text{UNI}}) \hat{p}_{\text{GYM}}}{1 - \hat{p}_{\text{VOC}}} \right. \right. \\ & \left. \left. - l_{\text{COMP}} \left(1 - \frac{\hat{p}_{\text{GYM}}}{1 - \hat{p}_{\text{VOC}}} \right) \right] \right\} \lambda_{\text{VOC}} \\ & + \left\{ \hat{p}_{\text{GYM}} \times \hat{p}_{\text{UNI}} (1 - \hat{p}_{\text{UNI}}) \right. \\ & \times \left[l_{\text{UNI}} - l_{\text{GYM}} - \frac{(l_{\text{POLY}} - l_{\text{GYM}}) \hat{p}_{\text{POLY|GYM}}}{1 - \hat{p}_{\text{UNI}}} \right] \left. \right\} \lambda_{\text{UNI}} \\ & + \left\{ \hat{p}_{\text{GYM}} \times \hat{p}_{\text{POLY|GYM}} (1 - \hat{p}_{\text{POLY|GYM}}) \right. \\ & \times \left[l_{\text{POLY}} - l_{\text{GYM}} - \frac{(l_{\text{UNI}} - l_{\text{GYM}}) \hat{p}_{\text{UNI}}}{1 - \hat{p}_{\text{POLY|GYM}}} \right] \left. \right\} \lambda_{\text{POLY|GYM}} \\ & + \left\{ \hat{p}_{\text{VOC}} \times \hat{p}_{\text{POLY|VOC}} (1 - \hat{p}_{\text{POLY|VOC}}) \right. \\ & \times \left[l_{\text{POLY}} - l_{\text{VOC}} \right] \left. \right\} \lambda_{\text{POLY|VOC}}. \end{aligned}$$

Appendix B: Decomposition of Inequality in Completed Education between Cohorts

The marginal effect of family background on completed education can be expressed as the sum of contributions of each transition (T_k) along the educational pathway:

$$\frac{\partial E(L)}{\partial x_j} = \sum_{k=1}^K T_k. \quad (\text{B.1})$$

T_k can further be expressed as the product of the family background effects on the logged odds of making each transition (λ), and the three weight components at risk (α), differentiation (β), and gain (γ):

$$T_k = \alpha_k \times \beta_k \times \gamma_k \times \lambda_k. \quad (\text{B.2})$$

To compare the components in cohort c with the reference cohort 1960, we can use the exact decomposition method for four factors presented by Das Gupta (1993: pp. 9–10). The values of the factors α , β , γ , and λ in cohort c are denoted by a, b, c, and d, and in the reference cohort 1960 by A, B, C, and D, respectively. The contribution of cohort change in λ to the total contribution of this transition (T_k) on educational inequality is:

$$\Delta \lambda_k = T_{kc}^\lambda - T_{k,1960}^\lambda = Q(d - D), \quad (\text{B.3})$$

where Q is a function of the other terms:

$$\begin{aligned} Q &= Q(a, b, c, A, B, C) \\ &= \frac{abc + ABC}{4} \\ &+ \frac{abC + aBc + Abc + ABC + AbC + aBC}{12}. \end{aligned} \quad (\text{B.4})$$

We can similarly estimate the contribution of the other factors. Because the transitions T_k have an additive contribution to overall educational inequality (B.1), and re-arranging terms, the change in educational inequality can be expressed as the change in each of the four factors:

$$\begin{aligned} \Delta \left(\frac{\partial E(L)}{\partial x_j} \right) &= \sum_{k=1}^K \Delta T_k \\ &= \sum_{k=1}^K \Delta \alpha_k + \sum_{k=1}^K \Delta \beta_k + \sum_{k=1}^K \Delta \gamma_k \\ &+ \sum_{k=1}^K \Delta \lambda_k, \end{aligned} \quad (\text{B.5})$$

which is (7).

Notes

- 1 However, Buis (2017) did in the supplementary material to the article discuss a model with transitions with multiple choices, which provided inspiration for our work.
- 2 For the sake of simplicity (cf. Buis, 2017), we refer to the logistic regression coefficients of parental education on the educational transitions as ‘effects’.
- 3 Students can also enrol to an optional preparatory 10th year of comprehensive school. A combined academic and vocational certificate is also possible.

- 4 One question concerns how educational attainment is measured. The polytechnics reform increased the duration of vocational post-secondary education, which may have countered the decline in attained education. However, if one focuses on the attainment of post-secondary degrees—as we do—the resulting picture is one of education contraction. As a side note, the cohort pattern in educational contraction curiously overlaps with the decline in IQ scores of (male) conscripts from 1997 onwards (Koivunen, 2007).
- 5 Information on the year in which parents completed their highest degree was incomplete for the oldest cohorts. For individuals not having this information (up to 50 per cent of cohorts born before 1970), parental education was measured at an age in childhood that was closest to 15 years and not over 18 years.
- 6 Results from sequential logit models are sensitive to the coding of the independent and dependent variables. Giving polytechnics 16 pseudo-years (instead of 15) increased educational inequality, mainly because it increased the expected gain from Gymnasium. Giving vocational school 12 (instead of 11) pseudo-years decreased inequality because it decreased the expected gain from Gymnasium. Doing both led to a marginal decrease in inequality in all cohorts but similar results on the trends. Conclusions of the sources of educational inequality remained unchanged.
- 7 Higher rates of post-secondary education completion—which would happen if we followed the cohorts to later ages—would most directly affect educational inequality through the weights. Taking the transition to university graduation as the example, this would happen through increases in the gains from gymnasium and from university, and decreases in the gains from vocational school, and polytechnics after gymnasium (Equation 6). By a simple simulation, increasing men's gymnasium to university transition probabilities from the 1985 to the 1979 rates would, all else constant, increase the marginal effect of parents' on sons' education from 0.34 to 0.37. Extending the follow-up to later ages would probably reduce the social background effect on the transition from gymnasium to university, but it should decrease from 0.214 to 0.065 to offset the increase in inequality through the weights, which is unlikely.

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References

- Ansell, B. W. (2008). University challenges explaining institutional change in higher education. *World Politics*, 60, 189–230.
- Ballarino, G. *et al.* (2009). Persistent inequalities? Expansion of education and class inequality in Italy and Spain. *European Sociological Review*, 25, 123–138.
- Ballarino, G., Panichella, N. and Triventi, M. (2014). School expansion and uneven modernization. Comparing educational inequality in Northern and Southern Italy. *Research in Social Stratification and Mobility*, 36, 69–86.
- Ballarino, G. and Schadee, M. (2010). Allocation and distribution: a discussion of the educational transition model, with reference to the Italian case. *Research in Social Stratification and Mobility*, 28, 45–58.
- Barone, C. and Ruggera, L. (2018). Educational equalization stalled? Trends in inequality of educational opportunity between 1930 and 1980 in across 26 European nations. *European Societies*, 20, 1–25.
- Becker, R. (2003). Educational expansion and persistent inequalities of education: utilizing subjective expected utility theory to explain increasing participation rates in upper secondary school in the Federal Republic of Germany. *European Sociological Review*, 19, 1–24.
- Bernardi, F., Hertel, F. R. and Yastrebov, G. (2018). A U-turn in inequality in college attainment by parental education in the US? *Research in Social Stratification and Mobility*, 58, 33–43.
- Blau, P. M. (1994). *Structural Contexts of Opportunities*. Chicago: University of Chicago Press.
- Blossfeld, P. N., Blossfeld, G. J. and Blossfeld, H.-P. (2015). Educational expansion and inequalities in educational opportunity: long-term changes for East and West Germany. *European Sociological Review*, 31, 144–160.
- Breen, R. and Goldthorpe, J. H. (1997). Explaining educational differentials: towards a formal rational action theory. *Rationality and Society*, 9, 275–305.
- Breen, R. and Jonsson, J. O. (2000). Analyzing educational careers: a multinomial transitions model. *American Sociological Review*, 65, 754–772.
- Breen, R. and Jonsson, J. O. (2005). Inequality of opportunity in comparative perspective: recent research on educational attainment and social mobility. *Annual Review of Sociology*, 31, 223–243.
- Breen, R. *et al.* (2009). Nonpersistent inequality in educational attainment: evidence from eight European countries. *American Journal of Sociology*, 114, 1475–1521.

- Buis, M. L. (2011). The consequences of unobserved heterogeneity in a sequential logit model. *Research in Social Stratification and Mobility*, 29, 247–262.
- Buis, M. L. (2017). Not all transitions are equal: the relationship between effects on passing steps in a sequential process and effects on the final outcome. *Sociological Methods and Research*, 46, 649–680.
- Cameron, S. V. and Heckman, J. J. (1998). Life cycle schooling and dynamic selection bias: models and evidence for five cohorts of American males. *Journal of Political Economy*, 106, 262–333.
- Chetty, R. *et al.* (2017). The fading American dream: trends in absolute income mobility since 1940. *Science*, 356, 398–406.
- Das Gupta, P. D. (1993). *Standardization and Decomposition of Rates: A User's Manual*. Washington, DC: U.S. Bureau of the Census.
- Davies, R., Heinesen, E. and Holm, A. (2002). The relative risk aversion hypothesis of educational choice. *Journal of Population Economics*, 15, 683–713.
- Erikson, R. and Goldthorpe, J. H. (1992). *The Constant Flux: A Study of Class Mobility in Industrial Societies*. Oxford: Clarendon Press.
- Erikson, R. and Jonsson, J. O. (1996). *Can Education Be Equalized? The Swedish Case in Comparative Perspective*. Boulder, CO: Westview Press.
- Erola, J. (2009). Social mobility and education of Finnish cohorts born 1936–75: succeeding while failing in equality of opportunity? *Acta Sociologica*, 52, 307–327.
- Gerber, T. P. (2000). Educational stratification in contemporary Russia: stability and change in the face of economic and institutional crisis. *Sociology of Education*, 10, 219–246.
- Hämäläinen, U. and Uusitalo, R. (2008). Signalling or human capital: evidence from the Finnish polytechnic school reform. *Scandinavian Journal of Economics*, 110, 755–775.
- Hertz, T. *et al.* (2007). The inheritance of educational inequality: international comparisons and fifty-year trends. *The B.E. Journal of Economic Analysis and Policy*, 7.
- Hillmert, S. and Jacob, M. (2010). Selections and social selectivity on the academic track: a life-course analysis of educational attainment in Germany. *Research in Social Stratification and Mobility*, 28, 59–76.
- Hout, M. and DiPrete, T. A. (2006). What we have learned: RC28's contributions to knowledge about social stratification. *Research in Social Stratification and Mobility*, 24, 1–20.
- Karhunen, H. and Uusitalo, R. (2017). 50 vuotta koulutusmahdollisuuksien tasa-arvoa. *Yhteiskuntapolitiikka*, 82, 296–303.
- Karlson, K. B. (2011). Multiple paths in educational transitions: a multinomial transition model with unobserved heterogeneity. *Research in Social Stratification and Mobility*, 29, 323–341.
- Kerckhoff, A. (1993). *Diverging Pathways: Social Structure and Career Deflections*. Cambridge: Cambridge University Press.
- Kivinen, O., Hedman, J. and Kaipainen, P. (2007). From elite university to mass higher education: educational expansion, equality of opportunity and returns to university education. *Acta Sociologica*, 50, 231–247.
- Kivinen, O., Hedman, J. and Kaipainen, P. (2012). Koulutusmahdollisuuksien yhdenvertaisuus Suomessa. Eriarvoisuuden uudet ja vanhat muodot. *Yhteiskuntapolitiikka*, 77, 559–566.
- Koivunen, S. (2007). *Suomalaismiesten kognitiivisen kykyprofiilin muutokset 1988–2001: Flynnin efektiä suomalaisessa aineistossa?* MSc Thesis (University of Jyväskylä, Jyväskylä, Finland). Available at urn.fi/URN:NBN:fi:jyu-2007274
- Mare, R. D. (1980). Social background and school continuation decisions. *Journal of the American Statistical Association*, 75, 295–305.
- Mare, R. D. (1981). Change and stability in educational stratification. *American Sociological Review*, 46, 72–87.
- Mare, R. D. (1993). Educational stratification on observed and unobserved components of family background. In Shavit, Y. and Blossfeld, H.-P. (Eds.), *Persistent Inequality: Changing Educational Attainment in Thirteen Countries*. Boulder, CO: Westview Press, pp. 351–377.
- Mare, R. D. (2011). Introduction to symposium on unmeasured heterogeneity in school transition models. *Research in Social Stratification and Mobility*, 29, 239–245.
- Meyer, J. W., Ramirez, F. O. and Soysal, Y. N. (1992). World expansion of mass education, 1870–1980. *Sociology of Education*, 65, 128–149.
- Nybohm, M. and Stuhler, J. (2013). *Interpreting Trends in Intergenerational Income Mobility*. IZA Discussion Paper No. 7514. Bonn.
- Pekkarinen, T., Uusitalo, R. and Kerr, S. (2009). School tracking and intergenerational income mobility: evidence from the Finnish comprehensive school reform. *Journal of Public Economics*, 93, 965–973.
- Pfeffer, F. T. (2008). Persistent inequality in educational attainment and its institutional context. *European Sociological Review*, 24, 543–565.
- Raftery, A. E. and Hout, M. (1993). Maximally maintained inequality: expansion, reform, and opportunity in Irish education, 1921–75. *Sociology of Education*, 66, 41–62.
- Schofer, E. and Meyer, J. W. (2005). The worldwide expansion of higher education in the twentieth century. *American Sociological Review*, 70, 898–920.
- Shavit, Y. and Blossfeld, H.-P. (1993). *Persistent Inequality: Changing Educational Attainment in Thirteen Countries*. Boulder, CO: Westview Press.
- Sirniö, O., Kauppinen, T. M. and Martikainen, P. (2017). Cohort differences in intergenerational income transmission in Finland. *Acta Sociologica*, 60, 21–39.
- Sørensen, A. B. (1983). Processes of allocation to open and closed positions in social structure. *Zeitschrift für Soziologie*, 12, 203–224.
- Statistics Finland (2010). *Oppilaitostilastot 2009 [Educational Statistics 2009]*. Helsinki: Statistics Finland.
- Statistics Finland (2011). *Educational Structure of Population 2010*. Helsinki: Statistics Finland.

Whelan, C. T. and Layte, R. (2002). Late industrialization and the increased merit selection hypothesis. Ireland as a test case. *European Sociological Review*, 18, 35–50.

Xie, Y. (1992). The log-multiplicative layer effect model for comparing mobility tables. *American Sociological Review*, 57, 380–395.

Xie, Y. (2011). Values and limitations of statistical models. *Research in Social Stratification and Mobility*, 29, 343–349.

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