Mothers and Mobility: a Re-examination of (Trends in) Intergenerational Mobility in the UK*

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Abstract

Existing research on intergenerational class mobility typically uses either fathers' occupation, or the occupation of the 'class dominant' parent, as an indicator of 'class origin', ignoring heterogeneity in mothers' and fathers' class effects on children's mobility outcomes. Such an elision therefore misses how class reproduction may be shaped by gender in important ways. In this paper, I bring new empirical evidence to the role of mothers in intergenerational mobility in Britain for birth cohorts born in the latter half of the 20th century. I find a significant independent effect of mothers' class on individuals' class destinations, and evidence that fathers' and mothers' class positions show a stronger effect on same-gender children. Further, mothers and fathers' class effects do not show similar trends across cohorts; while fathers' effects have weakened for women but stayed constant for men, mothers' effects have weakened for both men and women. My findings highlight the importance of taking a gender-sensitive approach to the study of intergenerational mobility patterns and trends.

Keywords: social fluidity, log-linear models, class origin, mothers effect

Introduction

How important are mothers for individuals' class mobility chances? Do mothers and fathers have greater effects on children of the same gender? Do mothers' and fathers' effects show similar trends over time? While social scientists have long established patterns of and trends in the association

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between an individual's social origin and destination position (Glass, 1954; Blau and Duncan, 1967; Goldthorpe et al., 1980, ch. 8), the majority of this work focuses on only one parental class position at a time, leaving these and related questions unanswered.

In its early days, sociological research on class mobility, concerned specifically with the relationship between individuals' destination class position and the class position of their family when they grew up (their 'class origin'),[1] treated one's father's occupation as an indicator of this origin position (Erikson and Goldthorpe, 1992; Breen, 2004b; Goldthorpe and Jackson, 2007; Bukodi et al., 2015). This father-only practice was undoubtedly in part a matter of practical constraint, as very few 20th century mobility surveys contained information on respondents' mothers' occupation during childhood. Further, given that a large proportion of respondents in turn-of-the-century mobility research were born to mothers who would either not have been working or working in (lower-ranking) part-time employment (e.g. Crompton, 2006; Esping-Andersen, 2009; Bukodi et al., 2012), it seems reasonable to have concluded that respondents' fathers' occupation would be 'at worst an adequate, but in most cases by far the better, index of the class from which they originated.' (Breen, 2004a, p. 8)

Subsequently, social changes including the dramatic increase in women's labor market participation, job attachment, and continuity and financial contribution to families (Bianchi et al., 2006; Blau et al., 2006; DiPrete and Buchmann, 2013) have increasingly motivated the incorporation of women's occupation in analyses of mobility through what analysts label as Erikson's (1984) 'dominance principle' to measure class origin. This approach indexes social origin by the social class of the spouse with highest or dominant class position (Buscha and Sturgis, 2018; Li and Devine, 2011; Devine and Li, 2013). Such an approach clearly marks an improvement on male-centric earlier practice, which simply a priori assumed away the importance of female resources. Nevertheless, beyond methodological limitations to measurements of social origin based on the dominance criterion (Thaning and Hällsten, 2020; Ballarino et al., 2021),[2] a single-parental measure of class origin

[1] In this paper I focus on relative rates of mobility – which measure the *relative* chances of attaining a particular destination position net of intergenerational changes in the marginal distribution of these positions (Erikson and Goldthorpe, 1992).

[2] In particular, it has been shown that the dominance approach underperforms relative to joint measures of class origin in cross-sectional mobility analyses as predicts of child outcomes in terms of variance explained or model fit (Thaning and Hällsten, 2020; Ballarino et al., 2021). Indeed, in most cases the dominance principle will result in information on mothers', rather than fathers', class being discarded since mothers are less frequently found in a higher class position than their spouses. Further such a measurement strategy creates problems for over-time mobility analyses: if the correlation between the dominant parental class and non-dominant class changes over time as class homogamy increases or decreases, then father-only or dominance estimates of changes in mobility will misestimate the true mobility rate, since they will be partly a function of changes in measurement error (namely, in processes of marital sorting) rather

captures only some of the channels through which intergenerational transmission occurs, and thus does not allow researchers to isolate the effect of mothers' class position from that of fathers'. In other words, it is a methodological convenience if one wants to construct a single measure of class origin, but not adequate for answering research questions aimed at parsing out parental-specific effects on class outcomes, or at investigating the mobility effects of the class 'non-dominant' partner. Such an approach also means that research examines mobility from either fathers' or mothers' class to son's class, and from either fathers' or mothers' class to daughters' class, such that overall mobility estimates will conflate resource with gender dynamics. Yet, a research and measurement strategy that enables an examination of parent-specific effects on children's class outcomes would shed important insight into the intergenerational transmission of advantage and its change over time.

Mounting empirical evidence underscores the importance of mothers for children's class mobility chances. In an important contribution, Beller (2009) demonstrates that, in a cross-sectional snapshot of US relative mobility, mothers do contribute to mobility chances, and that mothers' class shows different trends over time from fathers' - a conclusion which prompts the question of whether this pattern holds more generally. Only two subsequent studies have pursued similar questions within the class mobility context. Mood (2017) finds independent effects of mother and fathers' income and social class on children's earnings in Sweden, while in a comprehensive study of parental socio-economic transmission by family type in Finland, Erola and Jalovaara (2017) find strong and independent effects of mothers' and fathers' socio-economic status on children's. However, these two studies' limitation to cross-sectional analyses does not enable us to assess the extent to which mothers' and fathers' effects on class outcomes show similar patterns of change or persistence over time; further, neither of these studies directly test gender-specific parent-child associations, testing whether the same-gender parent is the more influential for class reproduction.

This study extends prior research in two ways. First, it proposes a gender-sensitive theoretical model of intergenerational class transmission. My proposed model is grounded in classic status attainment theory (e.g. Warren and Hauser, 1997), but extends existing work by explicitly incorporating (a) both mothers and fathers, (b) potential heterogeneity in parental effects on children's class outcomes by parent-child gender pairings. The second contribution is to provide new empirical evidence from Britain to the debate on the role of mothers in social mobility. I analyze data from

than of an actual changing association between parental and child class (see Tach, 2014).

the UK Household Longitudinal Study (UKHLS) and the British Household Panel Study (BHPS), and fit a number of log-linear and related models which enable me to analyze the mothers effect in social mobility among individuals in Britain for two cohorts born over the course of the 20^{th} century. My empirical analysis consists of three steps. First, I ask whether individuals' class positions are shaped by their mothers' class over and above their fathers'. Since this indeed turns out to be the case, I next ask both what the substantive magnitude of this effect is for children of different genders,[3], as well as whether mothers' and fathers' class effects on children's class outcomes show different trends over time (thatis, across cohorts). In the following sections, I first briefly discuss the theoretical debate on conceptualizing and measuring origin, before outlining a framework to guide an examination of parents-specific class effects on children's outcomes. I next discuss my data and analytical strategy, and present the results in the following section.

Mothers and mobility

To elucidate the substantive benefits from a model of social reproduction attentive to parent and child gender, in this section I propose a gender-sensitive theoretical model of intergenerational *class* transmission. The basic motivation for analyzing mobility within a class structural context are the insights such an analysis gives into the transmission of a broad notion of economic advantage across generations that other measures would not fully address. Social class, conceptualized in terms of social relations in labor markets (Goldthorpe, 1987, 2007a), serves as a particularly appropriate indicator of economic status in each generation, since it is not only associated with level of current earnings but, further, with earnings security and stability as well as longer-term earnings prospects (Goldthorpe and McKnight, 2006; Bukodi and Goldthorpe, 2018, ch. 1). Measuring the association between parental class and child class therefore gives an indication of the degree of openness of a society with regard to a broad range of economic outcomes, especially as compared with other measures of status attainment and transmission such as current income or occupation.

Consider first Panel A in Figure 1, which is a stylized model of intergenerational reproduction adapted from Breen and Jonsson (2007) and which shows several ways in which parental class will be associated with the children's class of destination. The bivariate association between class of origin and class destination is a function of a set of "front-door" and "back-door" pathways

[3]Following in the descriptive tradition of mobility research, I use the term 'effect' for convenience without any pretense to causal identification.

which capture distinct channels of transmission of parental resources to children. How strong the intergenerational class association is will depend on the strength of each of these "front-door" and "back-door" pathways. To orient the following discussion around particular mechanisms for which we have empirical evidence, I define "assets" more tightly than Breen and Jonsson (2007) as "education", which for our purposes includes strict attainment in addition to cognitive capacities including ability, motivation and genetic factors.

According to this model, parental class position affects children's class position via two "front-door" paths: a direct path (parental class position $\rightarrow$ child class position) and an indirect path (parental class position $\rightarrow$ child education $\rightarrow$ child class position).[4] The direct path is the result of all the processes that link class of origin and child's class net of child's educational attainment - for instance, parental network resources, socialization processes that mould occupational preferences as well as 'bridging funds' to facilitate relocation or time to search for a preferred job (Friedman and Laurison, 2019). The indirect path captures the educational assets children gain from parental class position and returns to education in the labor market. Empirically, it has been shown that in the UK approximately half of the total effect of parental on children's earnings is through educational attainment (Kuha and Goldthorpe, 2010; Breen and Karlson, 2014). In particular, in the UK it has been estimated that approximately half of the association between class origins and destinations is mediated via educational attainment**.** Theoretically, the mediating role of education in class mobility is often explained by appeal to Goldthorpe's (Breen and Goldthorpe, 1997; Goldthorpe, 2007a) model of intergenerational mobility, which theorizes that higher-class parents' investment in their children's educational attainment is driven fundamentally by risk aversion or the need to avoid downward intergenerational mobility; iniquitously distributed resources are mobilized towards these mobility goals. As such, advantaged parents, with most to lose from downward mobility, mobilize their superior economic, cultural, and social resources to help their children succeed in the educational system.

Further, according to this model, parental class position will be associated with children's class position via a set of "back-door" pathways: parental education determines a parent's class position and are themselves consequential for both children's educational attainment and their ulti-

[4]The extent to which the possession of assets such as education is associated with particular class positions will also affect the degree of mobility, and will be in part a function of various labor market processes, including the scarcity of particular assets such as certain educational qualifications - what Breen and Jonsson (2007) label 'class returns'. I do not consider this association in this piece as it is the 'transmission' component of Breen and Jonsson's (2007) that is more directly relevant to an assessment of parent-specific effects on class outcomes.

mate class attainment (such as through socialization, genetic inheritance, and direct transfers of resources). Parental education can thus be said to confound the total causal effect of parental class position on child class position. In other words, cumulatively, parental education and class position characterize the bundle of socio-economic resources affecting children's mobility chances via two sets of pathways; estimates of the intergenerational class association will necessarily 'pick up' the effects of parental education on children's class destination, even if our primary focus is on intergenerational transmission of economic status.[5]

Gender should play a central role in our understanding of intergenerational transmission, yet mobility research overwhelmingly focuses on men and has paid little attention to mothers and women. Developing a theoretical model of mobility that is sensitive to gender dynamics enables us to consider how patterns and trends in class reproduction may be shaped by gender in important ways. Consider next Panel B in Figure 1, which adapts the above model in order to distinguish between two parents as well as the gender of the offspring (the box in the bottom right-hand side of the DAG denotes a conditioning set of child gender). Compared with the conventional model of intergenerational transmission, this framework accommodates parent- and gender-specific dynamics, since each arrow emanating from parental education and class may differ depending on (a) the gender of the parent, (b) the gender of the child. My theoretical model supposes that both mothers' and fathers' class position are associated with their children's mobility outcomes through parent-specific instantiations of the various pathways considered in Panel A in Figure 1. In order to understand these pathways will combine to shape class mobility between fathers and children and mothers and children, in the following section, I explore each of the constitutive pathways in Panel B in Figure 1 to outline a gender-sensitive framework of social class reproduction. Considering these mechanisms independently thus helps inform hypotheses for the relative effect size of mothers and fathers on intergenerational class transmission, as well as potential gender-specific effects of maternal class on child outcomes.[6]

[5] Of course, a concern with intergenerational economic transmission (that is, where both partners are employed) foregrounds the role of parental investments of economic capital or transfer of occupation-specific skills, rather than a more general transmission of assets through direct care (such as the development of cognitive capacity through cultural and genetic endowments), though of course asset transmission in part occurs independently of class resources (a highly educated mother who is unemployed for instance).

[6] One disadvantage to an approach which examines fathers' and mothers' effects separately is that many disadvantaged subjects born into (mother-only) single parental households are excluded from the analysis. This is increasingly so, since in recent years the proportion of single mother families has notably increased (see Appendix C) In this respect, it is important to add a cautionary note on the mother and father joint measurement of social origins. The risk of a mechanic adoption of this standard to measurement social origins is that only intact families are considered. In Appendix C, I replicate the main analyses in this paper on the population of individuals living in mother-only households in

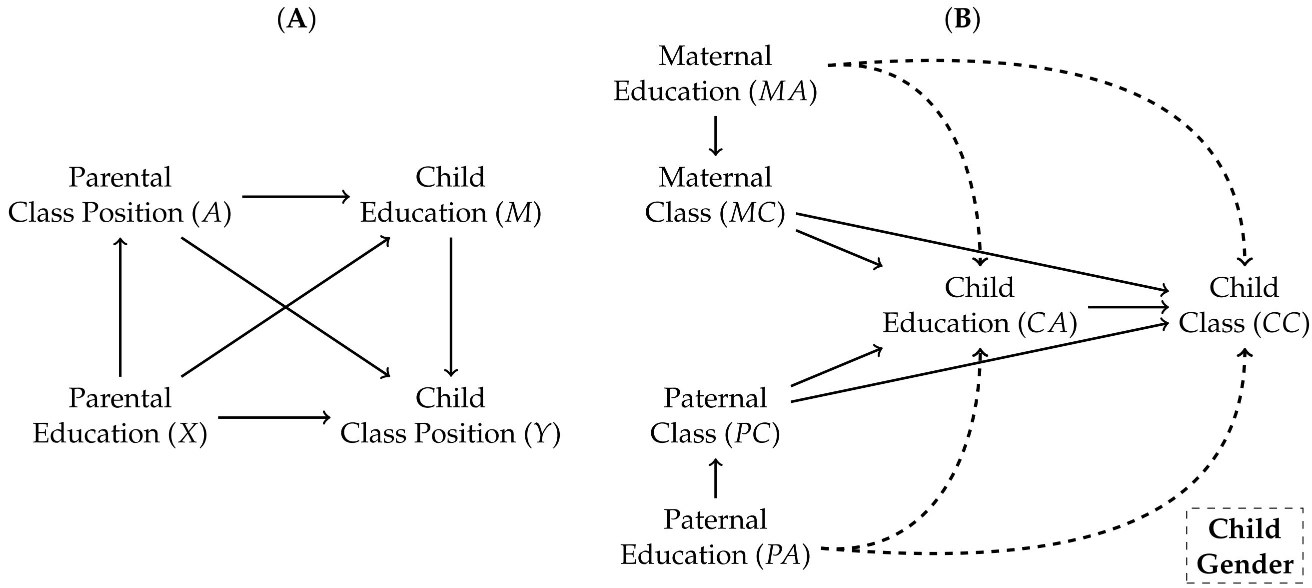

Figure 1: Mechanisms of intergenerational transmission from parental to child class position.

The relative size of mothers' class effects in a cross-sectional perspective

My model of intergenerational reproduction suggests that the "front-door" effects of both mothers' and fathers' class on children's class operate indirectly, through children's educational development, and directly, net of children's assets. First, what do we expect the sizes of the indirect pathways (paternal class position → child educatios → child class and maternal class position → child education → child class) to be? In the case of a gender-neutral model of reproduction (see Panel A in Figure 1), this pathway has received much attention both empirically and theoretically.

This empirical work on the mediating role of education, however, has been conducted exclusively with reference to the effect of the class position of the respondent's father (or head of the family), and not mothers, and as such gives us little ground for making hypotheses the maternal class position → child education → child class pathway. Further, Goldthorpe's original theory is developed without reference to either mothers' or fathers' class, although we could imagine that class resources from both parents are equally important in shaping both primary and secondary

childhood.

effects. In this case, we can understand the theory in more or less gender-neutral terms, and we might expect mothers' and fathers' indirect class effects to operate with equal magnitude. However, in considering the importance of the maternal class position → child education → child class pathway, I suggest focusing on two particular characteristics of mothers that may distinguish them from fathers in similar class positions. First, women (and mothers) are disproportionately found working part-time in a particular occupation. Thus, even if a mother is nominally in the same class as their husband, the mother may have fewer resources on account of her lower number of hours (Bukodi et al., 2012; Scott et al., 2010, 2012). Many social surveys do not distinguish between parents' occupations as part- or full-time, and thus do not enable us to draw an empirical distinction in this regard. If this is indeed the case, then net of paternal resources, maternal resources might therefore be less consequential for children's mobility outcomes as mothers have fewer resources to invest in their children's educational development. This is an important aspect of intergenerational transmission that goes unaddressed when researchers examine only father to child or dominant parental class to child associations, without taking into account the role of gender in both the parental and filial generations. This therefore leads to my first hypothesis:

H1a: Due to a larger proportion of part-time workers among mothers, mothers' class positions may show a weaker effect than fathers'.

The second characteristic relates to social selectivity. Given the barriers that women and especially mothers face in gaining employment, it is plausible that women found in a given occupation are more highly selected on attributes such as 'drive', motivation and ability than their male counterparts. If these attributes mean that these women are especially effective with their deployment of resources to bolster their children's attainment, this could cause the mother-child class association to be as strong (if mothers are disproportionately found in part-time work) or even stronger than that of the father-child. This is especially so in the UK; when compared with other advanced industrial countries where the independent role of mothers in class mobility has received some attention, such as the US, rates of maternal employment in the UK have never quite reached levels evidenced elsewhere. For instance, for cohorts born in the early 1980s, maternal employment in the US was near 80%, whereas in the UK, it was just under 70%. This contrast has consistently been the case: for the earliest birth cohorts for which we have data (individuals born in the 1920s), mothers' employment was approximately 10% lower in the UK than that in the US (see Figure B1 in Appendix B). A lower degree of female labor market participation in the UK may indicate a higher degree of selectivity among those women who are engaged in paid work. The conjecture that mothers' effect might be as strong as the fathers' effect is bolstered by literature that suggests that the maternal backdoor path via education might be stronger. Maternal influences on children's educational outcomes have received by far greatest attention in the recent literature examining parent specific social reproduction, and a long line of quantitative educational inequality research suggests that mothers' resources are equally, if not more, important than fathers' for children's educational attainment. In early studies of the US, as well as of the Netherlands and West Germany, Kalmijn (1994) and Korupp and Ganzeboom (2002a) found that maternal education has strong and independent effects on children's schooling attainment; more recent studies have found that the effect of mothers' education is usually greater or equal to that of fathers' (Marks, 2008; Buis, 2013; Jerrim and Micklewright, 2011). This recent quantitative work which highlights the importance of mothers' resources on educational outcomes complements an older ethnographic literature exploring the cultural and social mechanisms of resource transmission operating in the home environment. It is often reported that the mother, rather than father, plays the crucial role in investing in school-based activities, carefully monitoring the children's homework, and actively engaging with teachers (Reay, 1998; Lareau, 2003). I therefore propose the following alternative hypothesis:

H1b: Due to their stronger selectivity into the workforce, the overall mother-child association is as strong or stronger than the father-child association.

Are parental effects gender-specific? Next, the direct effect of parental class on child class, net of parental education, can be understood as an endowment effect, where endowments are both economic and material resources as well as parental social networks that children can exhaust for their own good, and the role model that parents provide. It is plausible that these endowment effects are gender-specific. For instance, some research provides evidence in support of gender role theory, namely that mothers are an especially important role model of socio-economic and occupational achievement for daughters, and consequently daughters prefer an occupation closer to their mother's (Smith and Self, 1980; Korupp et al., 2002b; Rosenfeld, 1978; Payne and Abbott, 2005). It has also been noted that the propensity for class immobility between fathers and sons (or for men to follow their fathers in specific occupations) is stronger than that for mothers and daughters - the function of direct transfer of non-educational assets such as business ownership (Jonsson et al., 2009). Some research alludes to potential gender-specific effects in the transmission of education across generations. In particular, in an examination of intergenerational mobility in Ireland, Miller and Hayes (1990) find that the advantage of having a mother with higher educational attainment is more salient for daughters than for sons; more recently, Jerrim and Micklewright (2011) have found suggestive evidence across OECD countries that mothers' educational attainment has a greater effect on their daughters than on sons, although the differences are small. This then leads to my second hypothesis:

H2: Both fathers' and mothers' class positions show a stronger effect on same-gender children than on other-gender children.

Mothers and mobility over time

Do the effects of parental socio-economic resources on children's outcomes not only have different effect sizes but show different patterns of persistence or change over time? The few studies which directly address this question in the context of primarily educational mobility have produced rather contradictory results. While several studies have found that parental occupational effects on educational outcomes show similar trends over time (Korupp et al., 2002a; Buis, 2013), several studies have reached the opposite conclusion: joint measures of social origin lead to different conclusions about mobility trends (Beller, 2009; Ballarino et al., 2021).

Instead, for a theoretical basis for expectations about whether mothers' and fathers' class effects show similar trends over time, we can again turn to Breen and Jonsson's (2007) theoretical models of status attainment, as described above. As Breen and Jonsson write, changes in fluidity over time, particularly those resulting from changes in transmissibility of resources and aspects, are likely to be cohort- rather than period-driven since 'most transmission of assets occurs during childhood and early adulthood, manifested particularly in educational qualifications' (p. 1781). Further, the authors note that variation across time in the levels of mobility will therefore be a function of institutional arrangements that constrain or facilitate the transmission of parental assets such as education and class-specific resources to offspring. For example, educational expansion and reform across the 20^{th} century in a large number of European countries made educational opportunity less dependent on parental assets and class-specific resources, and weakened the association between parental resources and children's educational attainment (e.g. Goldthorpe, 2007b). In this regard, we might expect paternal and maternal class effects to display similar trends over time as the institutional drivers of change in mobility are gender neutral (for instance, educational opportunity weakens the association between parental education and children's education, irrespective of the individual's gender in either generation).

On the other hand, as I have already noted, female labor market participation differs from men's in the degree of selectivity. This particular feature of mothers' employment mean that men and women occupying a given class position may be qualitatively distinct in terms of unmeasured attributes and resources important for social reproduction. This then leads to the following observation: if the degree of women's social selectivity into the labor force has declined over time as women's work becomes increasingly normalized and policies enables women to pursue employment-oriented careers (Blau et al., 2006; Gornick and Meyers, 2003; Gangl and Ziefle, 2009; Waldfogel, 2001), it is possible that this declining selectivity has increased mother-child mobility over time. In earlier periods when it was the norm for married women with children not to work, mothers' positive selection into the labor market in respect of unmeasured positive attributes such as 'occupational drive' and high (intergenerational) aspirations may differentiate them from their later-cohort counterparts in similar occupational positions, who are less positively selected in this way. If women's labor market participation is selective, it is likely that the strength of this selection has weakened over the period as their labor force participation has increased. The UK provides an especially apt context in which to explore hypotheses related to changes in mothers' labor market selectivity. As shown in Figure 2, the proportion of individuals with mothers in employment in childhood in two-parent households increased steeply and almost linearly across birth cohorts between 1950 and 1990 - a trend which contrasts sharply with the high degree of fathers' employment over time (shown in the same Figure as for comparison).

The majority of research on intergenerational class mobility in the UK has focused on father-son and dominant-parent-child mobility, thus giving little in the way of expectation about trends in mother-child class mobility. Most recently, drawing on a novel dataset covering four birth cohorts, Bukodi et al. (2015) find stability in mobility patterns over the 20^{th} century in the UK for men - a degree of inter-temporal constancy in line with earlier research indicating little change in levels of social fluidity within the class structure Goldthorpe and Mills 2004, 2008; Goldthorpe and Jackson 2007. In the case of women, however, the authors find persuasive evidence of an increase in class fluidity, a trend the authors in a later study find to be confined who have worked part-time at some point (Bukodi et al., 2017a). Therefore, insofar as we expect mothers' class trends to be distinct from fathers' on account of their declining selection into the labor market over time, this contrast is likely to be strongest for men as opposed to women (where research using paternal or dominance measures already finds an increase in mobility across cohorts). I therefore test the following hypothesis:

H3: Due to decreasing labor force selectivity, mothers' class effects have declined over time.

To summarize, my hypotheses are oriented around the relationship between mothers' and children's class both cross-sectionally and over-time. Cross-sectionally, Hypothesis 1a presents a null hypothesis in which mothers effects not matter, while Hypothesis 1b presents an alternative hypothesis in which mothers matter as much as fathers, and Hypothesis 2 suggests that mothers have heterogeneous effects on children. Regarding trends over time, Hypothesis 3 suggests that mothers' effects over time may either increase (a) or decline (b) due to distinctive and varying characteristics of their workforce participation. In testing for these hypotheses, I improve upon existing studies by explicitly interrogating the relative importance of mothers in class reproduction (Hypotheses 1a and 1b), and further consider both heterogeneity in mothers' class effects on children's class outcomes by gender pairings as well as variation in mother and father class effects over time. In particular, while studies have produced the important finding that mothers and fathers' socio-economic status have independent effects on children's (see Beller [2009] for the US case, Mood [2017] for the Swedish case and Erola and Jalovaara [2017] for Finland), none of these studies explicitly address substantive questions about the relative size of effects and gender-specific heterogeneity.

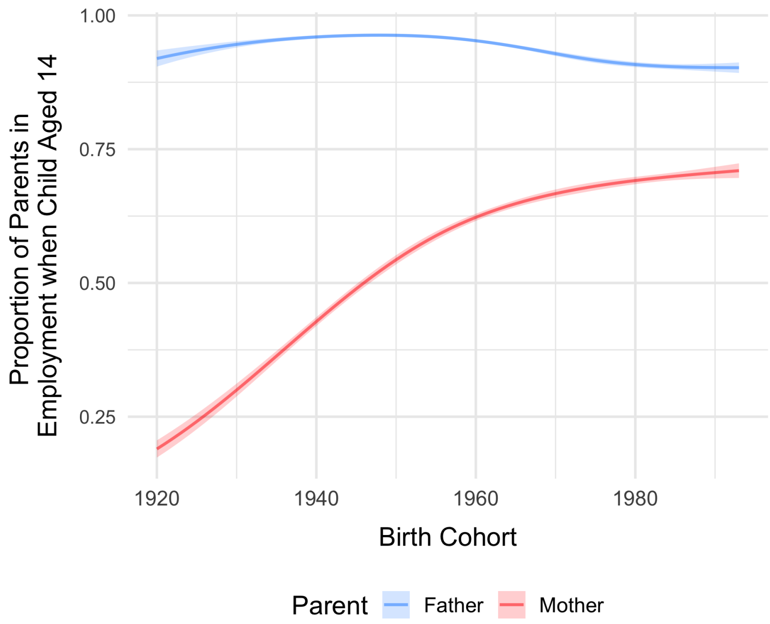

Figure 2: Fitted proportion of respondents living with a mother or father figure in employment, by year of birth, estimated using UKHLS 2009. The conditional means are fitted as a natural spline of year of birth with three degrees of freedom and adjusted by sampling weights. Ribbons represent 95 percent confidence intervals.

Empirical analysis

Data

I use data from two cross-sectional samples of the British population, which contain information on individuals' fathers' as well as mothers' occupations: I use the first wave of the UK Household Longitudinal Study (UKHLS, 'Understanding Society'), carried out between 2009 and 2011, for my cross-sectional analysis (Hypotheses 1 and 2), and combine this with data from the first wave of the British Household Panel Study (BHPS), carried out in 1991, for a two-point inter-temporal comparison (Hypothesis 3). UKHLS and BHPS are two nationally representative longitudinal surveys of the members of approximately 40,000 and 5,500 UK households, respectively. I restrict the sample to respondents who lived in two-parent households at age 14, and who have valid data on both their own class destination and two parents' occupations at this time (i.e. whose parents were both employed).[7][8] The focus on two-parent households is based on the finding that the socioeconomic status of non-residential mothers and stepmothers contribute little to a child's adult socioeconomic attainment (see Erola and Jalovaara, 2017), though in Appendices D and E I replicate my analyses on the full sample of respondents with maternal occupational information (thus including respondents with employed but non-resident mothers) and find that my conclusions are substantively unchanged. I assign respondents (but not parents) unemployed or out of the labor force at the time of the survey to a class position on the basis of their last employment. Since UKHLS contains data on Northern Ireland while BHPS does not, for my inter-temporal comparison I also restrict the sample to England, Scotland and Wales to maximize comparability. Regarding the specific measurement of respondents' and parents' class positions, I adopt the National Statistics Socioeconomic Classification (NS-SEC) and use the 'analytic', 7 class schema for both analyses (see Table 1).[9][10]

I conduct a cross-sectional analysis to address hypotheses (1-2) and a separate over-time analysis to address Hypothesis 3. For my cross-sectional analysis, I restrict the sample to individuals

[7] My primary concern with the transmission of economic status across generations, directly and indirectly via child education (such as parental investments in educational attainment) justifies focusing on employed parents. Specifically, I measure individuals as growing up in one of four household types at age 14: (a) living in a two parent heterosexual household (where parents are either (i) both the biological parents, (ii) both adoptive parents, (iii) a biological mother and co-resident father figure, or (iv) a biological father and co-resident mother figure); (b) living with both biological parents, (c) living in a single-parent maternal household, (d) other. Note that (a)/(b) and (c) and (d) are mutually exclusive, while (b) is a subset of (a). This latter category encompasses individuals who grew up in a father-only household, or individuals who have information on both parents' occupations but are missing information on household type. My main analyses all analyze individuals growing up in household type (a), although in the Appendices I explore the populations of (b) and (c) further.

[8] Only 0.4% of respondents whose mothers were in work at age 14 are missing information on maternal occupation in BHPS 1991 and 2.2% in UKHLS 2009. The corresponding figures for fathers are 1.2% and 1.9%. Note that the proportion of respondents missing information on mothers' work situation (i.e. in or out of employment when the respondent was age 14) is 3.4% in BHPS 1991 and 0.2% in UKHLS 2009. The corresponding figures for fathers are 3.5% and 0.5%.

[9] Although none of my tables contain zero margins, they contain a number of sampling zeros. I therefore add a small constant (0.001) to the cell counts for my over-time analysis.

[10] I encountered problems at two levels as regards recoding to NS-SEC. First, and regarding my cross-sectional analysis, in UKHLS the respondent's class is coded to the full analytic version of NS-SEC, but this is not the case for their father and mother, for whom parental employment status information is not available. In order to establish respondents' parents' NS-SEC, I therefore resort to the 'simplified derivation' approach and derive NS-SEC solely on the basis of parental occupational codes (Office for National Statistics, 2010, chs. 13, 16) – a procedure which yields only an approximation to the schema. For the cross-sectional analysis, I code parents' NS-SEC using SOC2010 codes since these yield slightly less sparse tables than does SOC90. Second, comparability across the samples poses difficulties for my over-time analysis: whereas UKHLS contains employment status information for respondents only, BHPS contains information on occupation and employment status for both respondents and their parents, enabling NS-SEC to be derived with great accuracy in all cases. Therefore, while I code BHPS respondents to the full analytic version of NS-SEC on the basis of their employment status and SOC90 three-digit occupational classification, my coding procedure for respondents' parents differs slightly across the two samples: I use SOC90 occupational codings and employment status data to allocate BHPS respondents' parents to the 7 class version of NS-SEC, but in the case of UKHLS respondents' parents, for whom no employment status information is available, I use the 'occupation only' derivation of NS-SEC (Office for National Statistics, 2005, ch. 13).

aged 25–64 at the time of the surveys. Since my hypothesis concerning changes in social fluidity over time is centered on cohort rather than by period effects (Breen and Jonsson, 2007), in order to examine cohort- rather than period-driven changes in fluidity, for my over-time analysis, I construct two synthetic birth cohorts from the BHPS and UKHLS surveys capturing, respectively, individuals born between 1950 and 1966 and those born between 1973 and 1984. This procedure then enables me to compare fluidity changes for two sets of cohorts who grew up in the second half of the 20^{th} century in Britain.[11] Further details on cohort construction are provided in Appendix E, where i also show that my conclusions about the changing effects of mothers' and fathers' classes are not sensitive to the exact construction of cohort groups from the BHPS and UKHLS.

Table 1: NS-SEC class schema, 'analytic' 7 class version.

I Higher managerial, administrative and professional occupations
II Lower managerial, administrative and professional occupations
III Intermediate occupations
IV Small employers and own account workers
V Lower supervisory and technical occupations
VI Semi-routine occupations
VII Routine occupations

Analytical Strategy

To measure the association between parental and child class net of the confounding effects of heterogenous marginals (Goldthorpe, 2007a, ch.7), I use log-linear and log-multiplicative models as is standard practice in mobility research, applying these techniques to four-way tables. Through fitting these models, I test a series of smoothed representations against the data in order to describe the association between parents' and their children's class. Specifically, I adjudicate between models using the Pearson chi-squared statistic (X^2) and dissimilarity index (DI) which provides a measure of the proportion of cases misclassified; when models are nested I compare their fit using the

[11] Specifically, the employment rate of mothers increased by 6% for the two sets of cohorts I analyze. The range of ages of my respondents is therefore 25-36 for the first period, and 25-41 for the second. Restricting the BHPS sample even further to make the age distributions of each set of cohorts more comparable reproduces the patterns evident in the larger sample, though they are clearly chararacterized by lower statistical power due to the lower sample size (see Appendix E for more details). Despite the advantage of this approach to directly assess cohort-related changes in social fluidity, it necessarily reduces the sample size and further means that some of the individuals who feature in my cross-sectional analysis are in the event dropped for the over-time analysis. Including the full sample of respondents for my over-time analysis (so that it is the same set of UKHLS respondents who feature in the cross-sectional and over-time analysis) does not alter my results.

likelihood ratio statistic (i.e the difference in the deviances(G^2) between the two models) (Powers and Xie, 2008, ch. 4).

For my cross-sectional analysis, I construct four-way father x destination x mother x gender contingency tables to the pooled UKHLS sample of men and women. I begin by fitting the baseline model 1 of 'conditional independence' which states that individuals' class destinations are statistically independent of both their parents' occupations (i.e. that 'perfect mobility' prevails):

$$\log F_{ijkl} = \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_l^G + \lambda_{jl}^{CG} + \lambda_{il}^{FG} + \lambda_{kl}^{MG} + \lambda_{ik}^{FM} + \lambda_{ikl}^{FMG}, \quad (1)$$

where $\log F_{ijk}$ is the natural log of the expected frequency of the ijk^{th} cell, μ is a scale parameter, λ_i^F, λ_j^C, λ_k^M and λ_k^G and are the main effects for fathers', children's and mothers' NS-SEC, as well as children's gender, respectively; and the remaining terms refer to the two-way associations between fathers' NS-SEC and gender, gender and children's' NS-SEC, and fathers' NS-SEC and children's NS-SEC, as well as a three-way interaction. In models 2 and 3, I extend the baseline model by fitting an additional term to account for the separate associations of father-respondent (λ_{ij}^{FC}, model 2) and mother-respondent (λ_{kj}^{MC}, model 3): these posit, respectively, that there is no net mother- or father-respondent association once the other parents' occupational effects are accounted for. In models 4-5, I then test for joint-parental effects by adding to model 3 the net father-respondent association: model 4 constrains the mother and father effects to be equal, then in model 5 I relax this equality constraint to test if an unrestricted "heterogeneous effects" model improves the fit. A comparison of, on the one hand, models 2, 3 and 4 and, on the other, model 5, enables me to directly test Hypotheses 1a and 1b.

Next, I fit a further model 6 that tests for gender-specific effects of fathers' and mother's class positions. Specifically, model 6 adds two log-multiplicative layer effect, or uniform difference (UNIDIFF), parameters to model 1 (Erikson and Goldthorpe, 1992; Xie, 1992), and can be written as follows:

$$\log F_{ijkl} = \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_l^G + \lambda_{jl}^{CG} + \lambda_{il}^{FG} + \lambda_{kl}^{MG} + \lambda_{ik}^{FM} + \psi_{ij}^{FC}\phi_{l1}^G + \psi_{kj}^{MC}\phi_{l2}^G + \lambda_{ikl}^{FMG}, \quad (2)$$

where ψ_{ij}^{FC} and ψ_{kj}^{MC} represent the association between fathers' class and class destination and mothers' class and class destination, respectively, and ϕ_{l1}^G and ϕ_{l2}^G are the UNIDIFF coefficients and represent how the strength of each of these associations varies by gender. This model enables us to assess whether the father-child and mother-child associations vary uniformly across child gender, and thus to directly address Hypothesis 2 - that is, whether the father-child association is larger for sons than for daughters, and vice versa for the mother-child association. Identification of ϕ_k requires a constraint such as $\phi_1 = 1$; since my cross-tabulation assigns men (sons) to the first table, the ϕ_l^G parameters denote whether the parent-child association increases or decreases for women (daughters).

In this second stage of the analysis I address Hypothesis 3 , concerned with change in mobility between two cohorts in Britain. Unlike my first set of analyses, I now conduct analyses separately for men and women, which is standard procedure in research on mobility trends because of changing rates of selectivity of women's employment. When analyzing trends in intergenerational mobility over a period which saw a dramatic increase in the rate of female labor market participation, changes in the degree of selectivity of women's workforce participation could be erroneously interpreted as substantive changes in levels of mobility over time. Reporting results for men and women separately enables this caveat to be made about women's results.

Separately for men and women (that is, sons and daughters), I construct four-way father x destination x mother x time contingency tables, and model the separate father-child and mother-child associations through four models. Model 7 tests for constant father & mother effects over the two periods – akin to the model of 'constant social fluidity' (henceforth CSF) (Goldthorpe, 1987, ch. 3):

$$\log F_{ijkl} = \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_l^T + \lambda_{jl}^{CT} + \lambda_{il}^{FT} + \lambda_{kl}^{MT} + \lambda_{ik}^{FM} + \lambda_{ikl}^{FMT} + \lambda_{ij}^{FC} + \lambda_{kj}^{MC}. \qquad (3)$$

Finally, models 8-10 are variants of model 1, and specify combinations of separate UNIDIFF parameters for each parental effect in which only fathers' (model 8: father UNIDIFF & mother CSF), or only mothers' (model 9: father CSF & mother UNIDIFF) or both fathers' and mothers' occupational effects (model 10: father UNIDIFF & mother UNIDIFF), are allowed to change over time:

$$\log F_{ijkl} = \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_l^T + \lambda_{jl}^{CT} + \lambda_{il}^{FT} + \lambda_{kl}^{MT} + \lambda_{ik}^{FM} + \lambda_{kj}^{MC} + \lambda_{ikl}^{FMT} + \psi_{ij}^{FC}\phi_{l1}^T, \quad (4)$$

$$\log F_{ijkl} = \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_l^T + \lambda_{jl}^{CT} + \lambda_{il}^{FT} + \lambda_{kl}^{MT} + \lambda_{ik}^{FM} + \lambda_{ikl}^{FMT} + \lambda_{ij}^{FC} + \psi_{kj}^{MC}\phi_{l2}^T, \quad (5)$$

$$\log F_{ijkl} = \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_l^T + \lambda_{jl}^{CT} + \lambda_{il}^{FT} + \lambda_{kl}^{MT} + \lambda_{ik}^{FM} + \lambda_{ikl}^{FMT} + \psi_{ij}^{FC}\phi_{l1}^T + \psi_{kj}^{MC}\phi_{l2}^T, \quad (6)$$

where $\phi_{l_1}^T$ and $\phi_{l_2}^T$ are UNIDIFF parameters representing the relative strength of the father-respondent and mother-respondent associations in the second birth cohort as compared with the first. This set of models thus explicitly tests the possibility of distinct patterns of constancy or change of each parent's effect. For men and women in turn I first select a model for each of the father-only and joint-parental measures of class origin that best captures trends in relative mobility across the two survey periods and display the UNIDIFF parameters returned under this model to address my third hypothesis. All models are fitted using LEM (Vermunt, 1997) or the R package gnm (Turner and Firth, 2007), and for each UNIDIFF parameter of interest I display the associated 'quasi-standard error' (Firth, 2003).

Results

The relative magnitude of the mothers effect

Table 2 shows the results of fitting the cross-sectional log-linear and log-multiplicative models with both single-and joint-parental measures for the pooled sample. respectively. The baseline model (1) of conditional independence – which naively assumes that mothers' and fathers' NS-SEC are unassociated with respondents' class after adjusting for the marginal distributions of each – clearly provides an inadequate fit to the data. However, while models 2 and 3 clearly improve the fit to the data compared with the independence model in terms of G^2 and the Dissimilarity Index, neither of these models provide an adequate fit ($p = 0.00$). I thus consider next the models (4-5) employing joint-parental measures. Both of these models provide a substantially better fit to the data than the single-parental models. Because models 2 and 3 (father- and mother-only, respectively) and 5 (father & mother main effects) are nested, I can directly compare their goodness-of-fit using the likelihood ratio test: for 36 degrees of freedom (df), model 5 reduces the deviance of model 2 by 542.2 and of model 3 by 465.2 – both producing significant improvements in fit. These results therefore provide clear evidence that mothers' and fathers' class have independent effects on children's class destinations.

What is the relative strength of the mothers effect in social mobility? Hypotheses 1a and 1b can be initially addressed by comparing models 4 and 5. Model 4, which posits that mothers' and fathers' occupational effects on children's class outcomes are equal, provides a good fit to the data ($p = 0.38$). Further, the main effects model (which results from relaxing the assumption of equality in parental effects in model 4) fails to make any significant improvement in fit, achieving a reduction in G^2 of only 39.2 for the loss of 36 d.f. Moreover, comparing model 4 with model 5 in terms of (X^2) indicates that a significant proportion of the parent-child association ([2088.7-549.7]/[2088.7-511.6] = 97.6% is captured by equal parental effects. Comparing model 4 with model 5 shows that the added information by allowing parental effects to be of different sizes is not warranted given the loss of d.f., implying that, for both men and women, mothers' class position has a significant direct net effect on children's class outcomes that is roughly equal in strength to fathers'. Thus, Hypothesis 1a (a null effect) is unsupported by these analyses; instead, I find support for Hypothesis 1b, that mothers' and fathers' class effects are equal.

In order to test for gender-specific effects of fathers' and mother's class positions, I next examine UNIDIFF model 6. According to the likelihood ratio test, this UNIDIFF model out-performs the joint effects model 5, which implies that parental effects do in fact vary by child's gender. Figure 3 presents the parameter estimates from this model (with men as the reference category); the UNIDIFF coefficients indicates how fathers' class effects and mothers' class effects vary for daughters compared with men. The parameter estimates and their upper bounds for fathers' class effects of 0.765 (0.929, 0.630), and for mothers' class effect of 1.519 (1.853, 1.246), indicate that fathers' class effects are weaker for daughers than for sons, while mothers' class effects are in fact stronger for daughters than for sons.

In order to illustrate the substantive implications of the mothers effect, I re-specify model 5 as a multinomial logit model (DiPrete and Grusky, 1990; Breen, 1994) with respondent's occupation as the dependent variable (model 6). Figures 4 and 5 plot the conditional probabilities of individuals' attainment of NS-SEC Class 1 (that is, the higher salariat) by parental class combinations as estimated under the main effects multinomial model. The probabilities of individuals with different parental class combinations attaining particular class destinations indicate the extent to which mothers' and fathers' occupations are associated with inequalities in this respect – that is, how much differentiation in probability there is by mothers' occupation for a given level of fathers' occupation, and vice versa. A number of features are noticeable which reflect the already established finding of gender-specific heterogeneity in father and mother effects (Hypothesis 2). Comparing the extremes (that is, having a parent in Class 1 rather than 7), for sons, as shown in figure 4, having a mother in Class 1 rather than 7, controlling for fathers' class position, increases the probability of entering Class 1 by 15 ppts on average. Having a father in Class 1 rather than 7 increases this same probability by a larger amount (18 ppts) on average, net of mothers' class position. In the case of daughters, as is shown in Figure 5, having either a mother in Class 1 rather than 7, adjusting for the spouse's class position, increases the chances of reaching a higher salariat destination by a smaller amount on average compared with men (12 ppts); having a father in Class 1 rather than 7, net of mother's class position, increases probability of Class 1 attainment by a smaller amount (9 ppts). Based on the UNIDIFF parameters reported above as well as the predicted probabilities from the multinomial logit, it is clear that, for both men and women, parental class effects are more consequential for children of the same gender, supporting Hypothesis 2.

To further examine potential gender mother and father gender-specific effects, in Appendix A I fit a topological model which constrain the parent-child association parameters, such as the λ_{ij}^{FC} terms in the model above, to be a linear function of a limited set of binary characteristics (Hout, 1983, ch.4; Bukodi et al., 2017), and as such is able to characterize these associations in terms of a number of substantively interpretable effect parameters. The results show that while mothers' and fathers' effects are on the whole similar for both men and women, there are some important gender-specific differences in terms of parental inheritance effects.

Table 2: Fit statistics for log-linear models 1-6 fitted to 2009 mobility table. Men and women aged 25-64 ($N = 14,327$).

						Model comparison			
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
1 Independence	2088.7	2138.2	576	0.00	15.3				
2 Father only	1058.3	1079.8	540	0.00	10.0	1 v 2	1058.4	36.00	0.00
3 Mother only	1007.3	1002.8	540	0.00	9.0	1 v 3	1135.4	36.00	0.00
4 Equal effects	549.7	576.8	540	0.38	6.6				
5 Main effects	511.6	537.7	504	0.40	6.3	2 v 5	542.2	36.00	0.00
						3 v 5	465.2	36.00	0.00
						4 v 5	39.2	36.00	0.33
6 Father & Mother UNIDIFF	489.5	515.1	502	0.65	6.1	5 v 6	22.5	2.00	0.00

Note: (1) dissimilarity index showing percentage of cases misclassified.

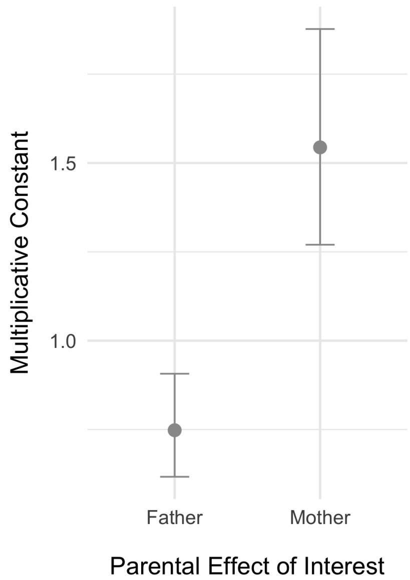

Figure 3: UNIDIFF parameter estimates with 95% confidence intervals scaled to $\phi^G_{\text{men}} = 1$: men and women aged 25-64 ($N = 14,327$). My cross-tabulation assigns men (sons) to the first table, such that ϕ^G_l parameters denote whether the parent-child association increases or decreases for women (daughters). In other words, for each parental effect, if the point estimate $\phi^G_l > 1$, then, relative to that parent's effect on sons, the parental effect is greater for daughers, and if the point estimate $\phi^G_l < 1$, then, relative to that parent's effect on sons, the parental effect is weaker for daughters.

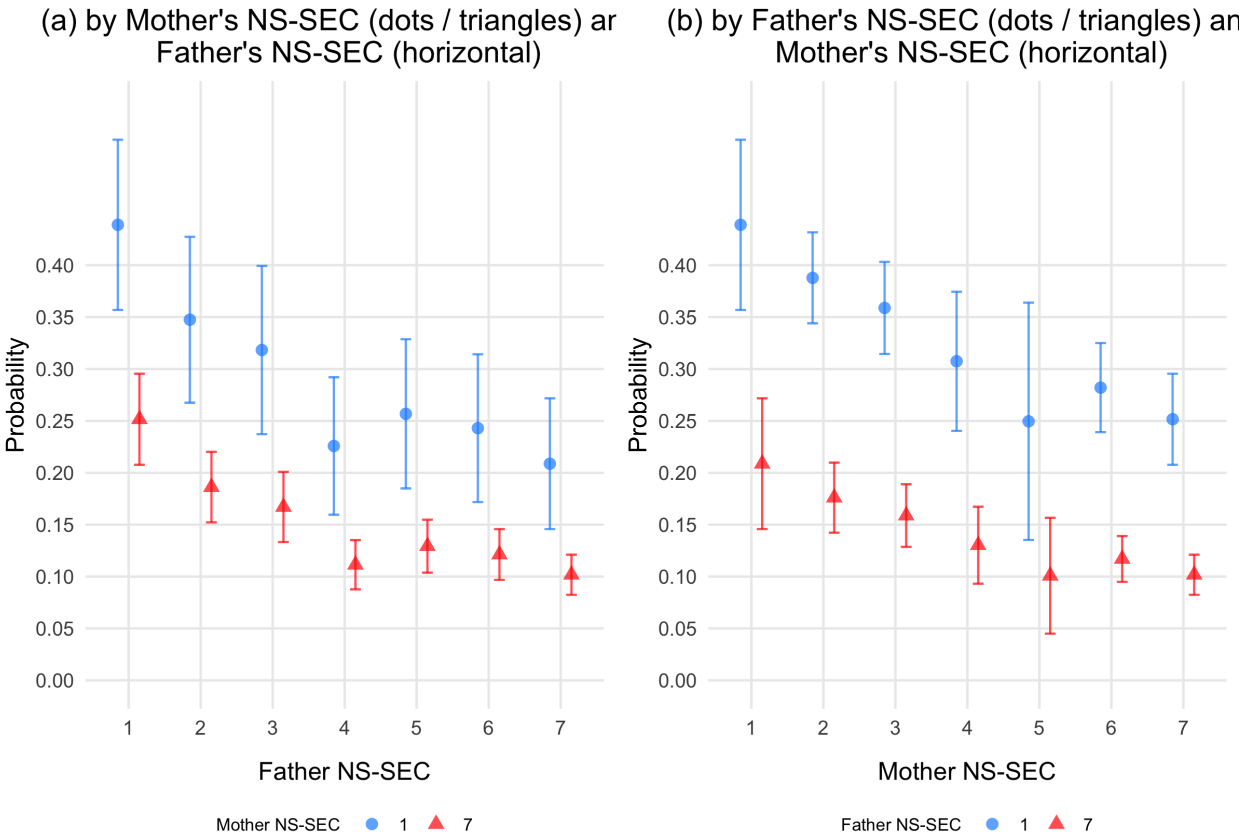

Figure 4: Estimated probabilities of attaining NS-SEC Class 1 destination by parental combinations: (a) father's and mother's NS-SEC and (b) mother's and father's NS-SEC: men aged 25-64.

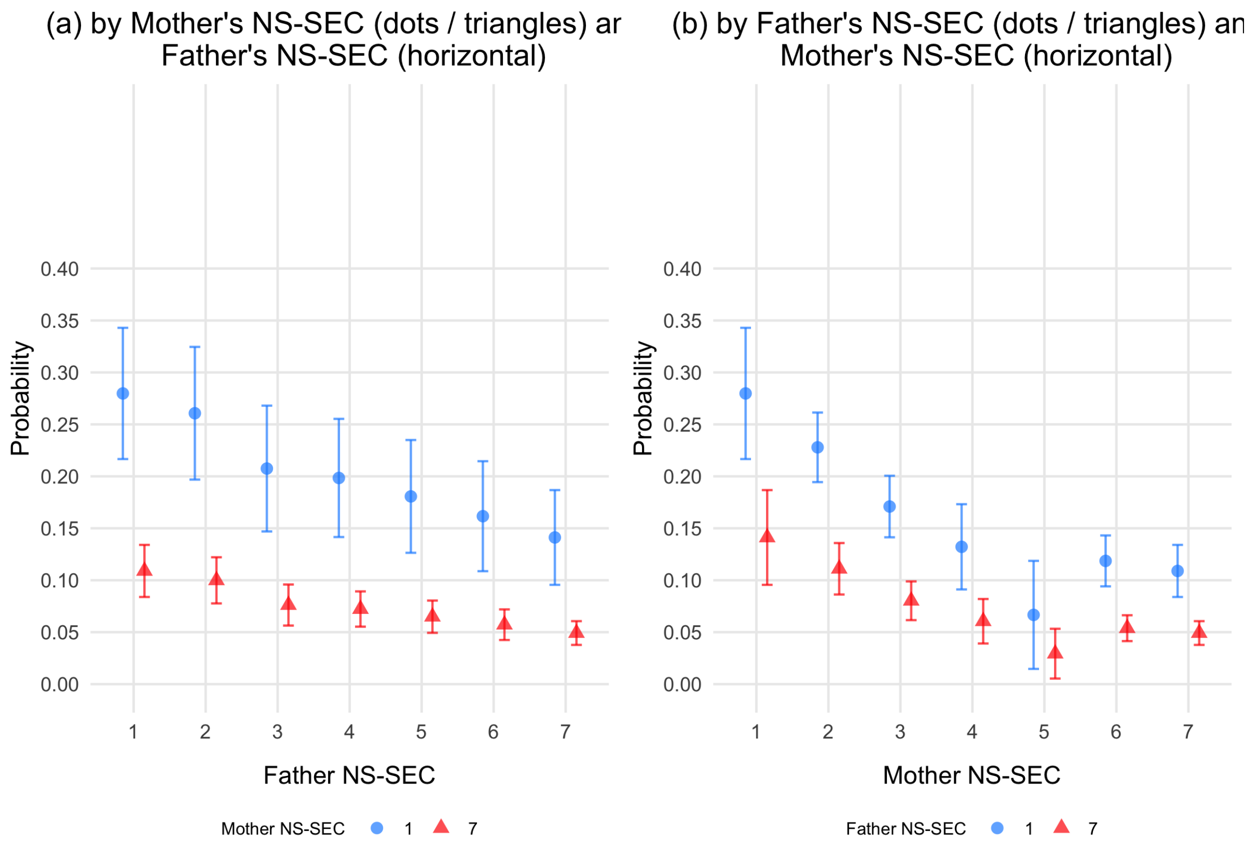

Figure 5: Estimated probabilities of attaining NS-SEC Class 1 destination by parental combinations: (a) father's and mother's NS-SEC and (b) mother's and father's NS-SEC: women aged 25-64.

Mothers and mobility trends

Goodness-of-fit statistics are presented in Table 3 and 6 for both men and women. First, for men, I sequentially explore different variants of the joint-parental model (7) characterized by equation 3. The father & mother constant effects model (7) fits the data reasonably well, and the mother constant effects & father UNIDIFF model (8) fails to achieve a significant improvement in fit. By contrast, the father constant effects & mother UNIDIFF model (9) does in fact perform significantly better ($p = .03$), and adding a UNIDIFF parameter for the father-son association to model (model 10) fails to improve the goodness-of-fit. I therefore consider model 9 (which suggests that only the mothers' effects on sons' class outcomes changed over time) to be my preferred joint-parent model for characterizing male mobility trends across the two male cohorts (1950-1966 and 1973-1984).

Figure 6 displays the UNIDIFF parameter returned under this model, which indicates that the effect of mothers' class on class outcomes declined significantly across the two cohorts: the ϕ parameter estimated suggests that all odds ratios defining the mother-son association decreased by a factor of 0.45. By way of comparison, the UNIDIFF parameter for the father-son association under a model that only allows the father-son association to change over time (model 8) is estimated at 1.105 - a substantively far smaller (and statistically insigificant) change comapred with that pertaining to the mother-son association.

For women, the mother constant effects & father UNIDIFF model (8) clearly improves over the father & mother constant effects model ($p = .00$), providing evidence of a change in fathers' occupational effects over time. There is also evidence in favor of a change in mothers' NS-SEC effects: the father constant effects & mother UNIDIFF model (9) performs significantly better than the father & mother constant effects model (3) ($p = .01$), and given that model 10 (containing separate UNIDIFF parameters for the father and mother effects) produces a significant improvement in fit over this model, I consider model 10 my preferred joint-parental model for characterizing female mobility trends, indicating that both parental effects changed across the two cohorts. Figure 6, which shows the ϕ parameters estimated under model 10 for women, reveals a similar rate of increase in both father-daughter and mother-daughter fluidity over the cohorts: the odds ratios defining the mother-daughter association decreased by a factor of 0.57, while fathers' class effect decreased by a factor of 0.45. Appendix E shows that these results are highly robust across alternative sample restriction criteria, household types and cutoffs for defining birth cohorts. These results therefore provide strong evidence in support of Hypothesis 3: mothers' class effects have indeed weakened over time.

Table 3: Fit statistics for log-linear and log-multiplicative models 7-10 fitted to mobility tables for two British cohorts (1950-1966 and 1973-1984). Men and women aged 25-64.

Model	X^2	G^2	df	$P(X^2)$	DI^1		Model comparison rG^2	rdf	p
Men, $N = 2,937$									
7 Father & mother CSF	475.9	459.6	504	0.81	14.1				
8 Father UNIDIFF & mother CSF	474.0	459.4	503	0.82	14.1	7 v 8	0.1	1.00	0.71
9 Father CSF & mother UNIDIFF	477.5	454.7	503	0.79	14.1	7 v 9	4.8	1.00	0.03
10 Father UNIDIFF & mother UNIDIFF	471.0	453.7	502	0.84	14.0	9 v 10	1.1	1.00	0.30
Women, $N = 2,157$									
7 Father & mother CSF	462.3	493.7	504	0.91	11.6				
8 Father UNIDIFF & mother CSF	462.6	483.7	503	0.90	11.7	7 v 8	10.0	1.00	0.00
9 Father CSF & mother UNIDIFF	460.2	487.6	503	0.91	11.7	7 v 9	6.1	1.00	0.01
10 Father UNIDIFF & mother UNIDIFF	460.1	481.1	502	0.91	11.7	9 v 10	6.5	1.00	0.01

Note: (1) dissimilarity index showing percentage of cases misclassified.

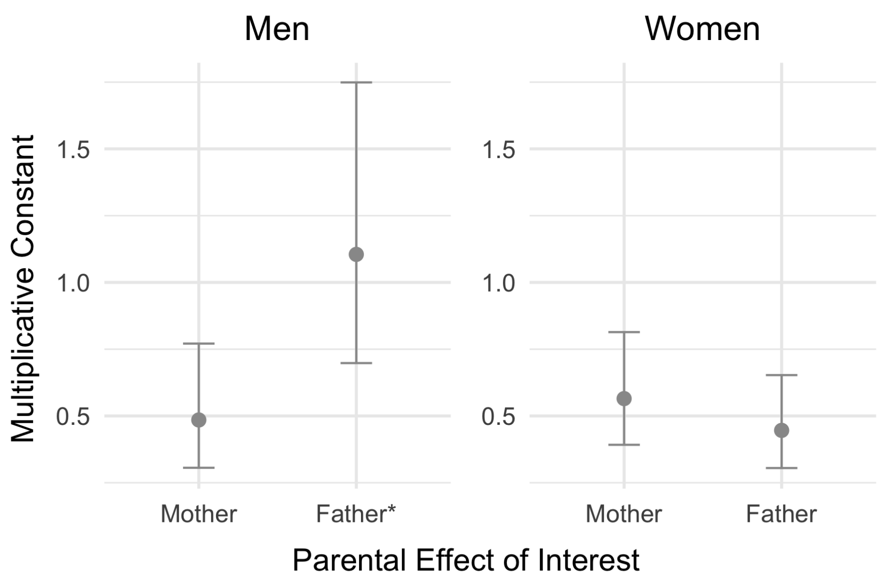

Figure 6: UNIDIFF parameter estimates with 95% confidence intervals scaled to $\phi_{1950\text{-}1966} = 1$: men and women aged 25-64. * denotes an a UNIDIFF parameter returned under a non-preferred model, as a baseline comparison.

Discussion

Motivated by the remarkable absence of empirical investigation of parental-specific effects on class outcomes, in this paper I have brought new empirical evidence to the question of mothers' role in social class reproduction. I have reported two primary results. Firstly, in a cross-sectional snapshot of mobility, net of fathers' NS-SEC, mothers' class contribute importantly to individuals' class mobility chances. Furthermore, whereas previous scholarship primarily focuses on the dichotomous question of whether mothers contribute to children's mobility changes, my analysis also reveals the striking result that, overall, this effect is of roughly equal magnitude to fathers, although a test for gender interaction shows that both fathers' and mothers' class positions show a stronger effect on same-gender children than on other-gender children. Secondly, in respect of mobility trends over time, my results show that, for both men and women, mothers' class effects have weakened across two British cohorts.

The takeaway from this empirical analysis is evident enough: when we consider the combined effects of mothers' and fathers' occupation in analyses of mobility, we reveal far wider inequalities in class attainment than when father-only measures are used and, in turn, how these wider inequalities change in strength over time. My findings are also of interest in relation to a set of important theoretical issues. In particular, they bring out that differing social processes could be at work in the production of mothers' and fathers' occupational effects on class outcomes as well as how these effects change over time. In particular, and regarding change over time, a plausible explanation for the declining influence of mothers on men's class attainment is the decreasing social selectivity of mothers between my two samples as dual-earning households became increasingly commonplace. Thus, the higher degree of immobility between mothers and sons in 1991 might reflect positive selection into the labor market in respect of unmeasured positive attributes such as 'occupational drive' and high (intergenerational) aspirations (cf. Breen and Goldthorpe, 1997) in a period when it was the norm for married women with children not to work. This may differentiate them from their 2009 counterparts in similar occupational positions. This account is, of course, rather speculative, yet it invites further theoretical and empirical investigation explicitly to parse out out the mechanisms underpinning distinct parental associational trends.

There are a number of substantive analyses that would shed further light on the trends I have explored here. As discussed in the main text, the intergenerational association between parental and child class is a function of multiple front-door and back-door pathways. The patterns and trends I have highlighted in this article are a result of the patterns or trends in all of these components, and these components may in fact show differing patterns of change of persistence over time. This implies at least two fruitful areas for research. First, future research could consider the distinct patterns of association between children's class position, on the one hand, and parental assets and parental class position, on the other, as well as the mechanisms underpinning such effects. Second, mediation-style analyses that disentangle the direct from the indirect effects of parental class and assets on children's class outcomes would further enable us to parse out which pathways are gender-specific, and which are not. For example, gender role theory, which contends that the same-gender parent is the more influential for social reproduction, often examines the gendered nature of occupational role modeling, and implies that the direct effect from parental class to child class may be gendered. By contrast, literature exploring maternal educational effects on children's educational outcomes does not emphasize these as gender-specific, which might suggest a gender-neutral indirect effect via children's educational attainment.

Second, in principle, the mechanisms behind the influence of mothers' class could be expected to apply irrespective of institutional context. Yet, as we have already seen, mothers' class effects are likely to be a function of mothers' selectivity into the labor force and extent of part-time work, which vary across countries. In the main text I note a number of differences between the UK and US with regard to the extent of, and change in, maternal employment over time. This pattern of cross-country differences in maternal economic activity following childbirth is consistent with research which shows US-UK variation in a range of economic behaviors following motherhood, including duration of work interruption and number of job changes, and probability of transitioning into part-time work. In particular, A key driver behind the higher rates of maternal employment in the US is its less market-oriented family public policy - namely, its maternity leave policies and public provision and subsidy of childcare - than those in the UK (Waldfogel, 2001; Gornick and Meyers, 2003; Lambert, 2008; Gangl and Ziefle, 2009). As a consequence of the lower proportion of women in the work force compared with lower countries overall, we might expect selectivity effects to be stronger than in the US. Indeed, in Appendix B I replicate my cross-sectional analyses in the contrasting context of the United States using the GSS and find, contrary to my results for the UK, that mothers' class effects are far smaller than fathers'. A useful extension of this study would therefore be to provide a more systematic cross-national comparison of mothers' effects and their changes over time, with a particular focus on how variation in relative effect size (compared with fathers' class effects) and in change over time correlates with variation in patterns of female labor market participation (in terms of both part-time work and selectivity). Addressing how maternal class effects vary across time and place promises to provide a fuller picture of the dynamics of class reproduction.

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A Gender-specific and Topological Analysis

In the main text, I show that, for both men and women, although a model allowing the mother-respondent and father-respondent associations to differ does not improve on a model constraining these associations to be equal (in terms of statistical significance), there is still some evidence of (perhaps small) parent effects that are particular to the gender of the child. To test this proposition more formally, I fit models 2-5 separately for men and women, and in addition fit a topological model which aims to parsimoniously capture the association between parent and child class in terms of a limited number of theoretically-derived parameters, each of which either applies or does not apply to a particular cell in the mother-respondent or father-respondent mobility tables.

For men, as shown in table A1, all of the models employing joint-parental measures (4 and 5) provide a substantially better fit to the data than the single-parental modelFor 36 degrees of freedom (df), model 5 reduces the deviance of model 2 by 167.8 and of model 3 by 308.6 – both significant improvements in fit. For women, as shown in table A2 ,from this test it is clear that the latter do make a significant improvement over the former in fit at the 0.00 level. Further, comparing models 4 and 5 shows that the added information by allowing parental effects to be of different sizes is not warranted given the loss of d.f., implying that, for both men and women, mothers' class position has a significant direct net effect on children's class outcomes that is roughly equal in strength to fathers'.

Despite the high degree of similarity in fathers' and mothers' effect sizes, there is still a sizeable decrease in G^2 between the equal and main effects models, for both men and women, which would not be the case if the effect were truly of the same magnitude and type. What is more, the different G^2 and X^2 (with the same d.f.) reported in models 2 and 3 clearly demonstrate that depending on child's gender, either mother or father's class is more informative to understand the social mobility pattern: a significant reduction in X^2 is achieved by fitting a father- or mother-only model to sons and daughters, respectively, compared with the baseline model of independence. Further, comparing models 2 and 3 with 5 indicates that a more significant reduction in X^2 is achieved by adding the father-child association to a mother-only model for men than by adding the mother-child association to a father-only model for men; the reverse is true for women. These findings are consistent with those reported in the main text regarding mother and father effect heterogeneity by child gender.

How can this information be explained? One possible explanation is that the association between social origins and destinations consists of several parts. Following model 5, there is (1) the part that is proxied by either mother's or father's class. In addition, there is (2) a gender-specific part that results in the better of models using mothers rather than fathers to predict daughters' (Models 2-3 in Table A1) but not sons' (Models 2-3 in Table A2) position, and vice versa. Since model 5 does not outperform model 4, the gender-specific pattern appears to not add significantly to the resource part (which can be proxied by either parent).

In order to further explore potential gender mother and father gender-specific effects, I next fit a topological model which models the father-respondent and mother-respondent association more parsimoniously than the main effects model, and which is able to characterize these associations in terms of a number of substantively interpretable effect parameters (specifically, hierarchy, immobility and affinity effects). Specifically, the topological model I use is identical to that proposed in Bukodi and Goldthorpe (2017b; 2021) in that it uses eight "design matrices" to characterize three types of effects—hierarchy, inheritance, sector, and affinity— that promote or impede mobility between specific classes. The hierarchy, inheritance and affinity effects can be interpreted substantively as follows:

1. Hierarchy effects limit mobility, and result from the relative advantages or disadvantages of one's class of origin in terms of economic and cultural resources, as well as barriers to entry into class groupings in adulthood. Following Bukodi and Goldthorpe (2017b; 2021), I model four hierarchy effects which capture the five hierarchical levels into which the 7-class NS-SEC schema is commonly grouped (specifically, the classes in Table 1 are grouped such that classes 3, 4 and 5 are treated at the same hierarchical level). The first hierarchy effect (HI1), operates in cells parent-respondent mobility table for that implies the crossing of any one of the four hierarchical divisions; the second effect (HI2), in cells implying the crossing of two levels; the third effect (HI3), in cells implying the crossing of three levels, and the fourth (HI4), in cells implying the crossing of all four levels. These hierarchy effects are cumulative, such that the presence of a higher-level hierarchical effect implies the presence of its lower level effects.

2. Inheritance effects are those which promote immobility (i.e. which limit mobility), and reflect the general tendency for individuals to come to occupy the same social class as that in which

they grew up. The effects result from, among other factors, social networks, direct inheritance of property, and socialization processes. I use two inheritance effects. The first effect (IN1) captures the general propensity for immobility in all classes, and thus operates on the main diagonal of each parent-child mobility table. The second (IN2) applies only to two cells on the main diagonal (Class 1 and Class 4), and is intended to capture the greater tendency for immobility within these classes.

3. Finally, I model two affinity effects which promote mobility by capturing shared cultural orientations and behaviors within the white-collar world (Classes 1, 2 and 3) and blue-collar world (Classes 5, 6 and 7) respectively, and which facilitate intergenerational movement within these sets of classes. The first set of effects operates in all cells implying mobility between Classes 1-3 (AF1), and the second, in all cells implying mobility between Classes 5-7 (AF2).

Formally, we can write this model as

$$
\begin{aligned}
\log F_{ijk} &= \mu + \lambda_i^F + \lambda_j^C + \lambda_k^M + \lambda_{ik}^{FM} \\
&\quad + \sum_{i=1}^{4} \lambda^{MHI_i} + \sum_{i=1}^{2} \lambda^{MINH_i} + \sum_{i=1}^{2} \lambda^{MAF_i} + \sum_{i=1}^{4} \lambda^{FHI_i} + \sum_{i=1}^{2} \lambda^{FINH_i} + \sum_{i=1}^{2} \lambda^{FAF_i}.
\end{aligned} \tag{7}
$$

This model differs from the main effects model by replacing the unconstrained λ_{ij}^{FC} and λ_{kj}^{MC} parameters, respectively, with three distinct sets of topological effects. Tables A1-A2 show the results of fitting models 1-5, as well as this additional topological model, for men and women, respectively.

The relevant parameter estimates from this model are presented in Figure A1. To interpret the parameter estimates, we observe that the inheritance and hierarchy parameters, despite having opposite signs, in fact contribute in the same direction to increase the (log-)odds ratio defining the parent-child association, while the affinity effects work in the opposite direction, attenuating the odds ratio back towards zero.[12] Note first that father and mother status affinity effects are highly similar, with the effect size for each being slightly stronger in general for men than for women

[12]This can be seen via the fact that any local log-odds ratio for a 2 x 2 subtable can be expressed in terms of the model parameters as $\log \theta_{ij,i'j'} = \log F_{ij} + \log F_{i'j'} - \log F_{ij'} - \log F_{i'j} = \boldsymbol{\alpha}^\top \boldsymbol{\lambda}^{\text{INH}} - \boldsymbol{\beta}^\top (-\boldsymbol{\lambda}^{\text{HI}}) - \boldsymbol{\gamma}^\top \boldsymbol{\lambda}^{\text{AF}}$, where $\alpha \in [0, 1, 2]$ and $\beta, \gamma \in [0, 1]$, since inheritance effects, on the one hand, and the hierarchy and affinity effects, on the other, refer only to the diagonal and off-diagonal cells of the full table respectively.

in the sense of offsetting hierarchical barriers to mobility within white-collar and the blue- collar classes more for men than for women. This lower offsetting effect for women reproduces a key finding in a recent topological study of mobility in Britain using a different dataset (Bukodi and Goldthorpe, 2021, p. 8), which is ascribed to the lower levels of cross-class status homogeneity among women (the fact that women are more likely than men to be found in lower status groups even if occupying the same overall class position - e.g. of routine office and sales work within the salariat).

The important differences between parental parameters emerge with regard to inheritance and hierarchy effects. While the general tendency for class inheritance, represented by the IN1 effect, between mothers and their children is similar for men and women, the IN1 effect of fathers is stronger in the case of men than of women. Regarding the IN2 effects (representing immobility within Classes 1 and 4) the IN2 effect of fathers on men is larger than that of mothers, the latter also being insignificant. Inheritance effects overall are weaker and less significant for women, with the one exception of the maternal IN1 effect (0.235). Thus, while men appear more class immobile than women, there is still a notable parental-gender difference in immobility patterns overall. Parental effects are largely similar for the four hierarchy parameters, with the exception of HI3, in which case the relative size of the coefficients for fathers and mothers flip in the case of men and women: women have a lower propensity than men to experience long-range mobility (that is, from Classes 6 or 7 to 1 or 2, or vice versa). In other words, men seem to be more constrained in terms of upward mobility by their fathers than their mothers (and, similarly, are more protected from downward mobility by their fathers than their mothers), while the reverse is true for women. In short, the parameter estimates from the topological model are consistent with the finding of high similarity of mothers' and fathers' effects for both men and women above, though also reveal some notable gender-specific differences in terms of class immobility as well as long-range mobility.

Table A1: Fit statistics for mobility models fitted to 2009 mobility table. Men aged 25-64 (N=6,068).

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
1 Independence	905.8	938.8	288	0.00	15.6				
2 Father only	363.0	368.4	252	0.00	8.3	1 v 2	570.4	36.00	0.00
3 Mother only	511.4	509.2	252	0.00	9.9	1 v 3	429.6	36.00	0.00
4 Topological Effects	255.5	270.5	272	0.76	6.5				
5 Main effects	185.8	200.7	216	0.93	5.5	2 v 5	167.8	36.00	0.00
						3 v 5	308.6	36.00	0.00
						4 v 5	69.9	56.00	0.10

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table A2: Fit statistics for mobility models fitted to 2009 mobility table. Women aged 25-64 (N=8,259).

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
1 Independence	1182.8	1199.4	288	0.00	15.1				
2 Father only	655.6	666.2	252	0.00	10.6	1 v 2	533.2	36.00	0.00
3 Mother only	441.1	445.8	252	0.00	7.9	1 v 3	753.6	36.00	0.00
4 Topological Effects	341.6	354.1	272	0.00	7.1				
5 Main effects	233.5	245.9	216	0.20	5.6	2 v 5	420.3	36.00	0.00
						3 v 5	199.9	36.00	0.00
						4 v 5	108.2	56.00	0.00

Note: (1) dissimilarity index showing percentage of cases misclassified.

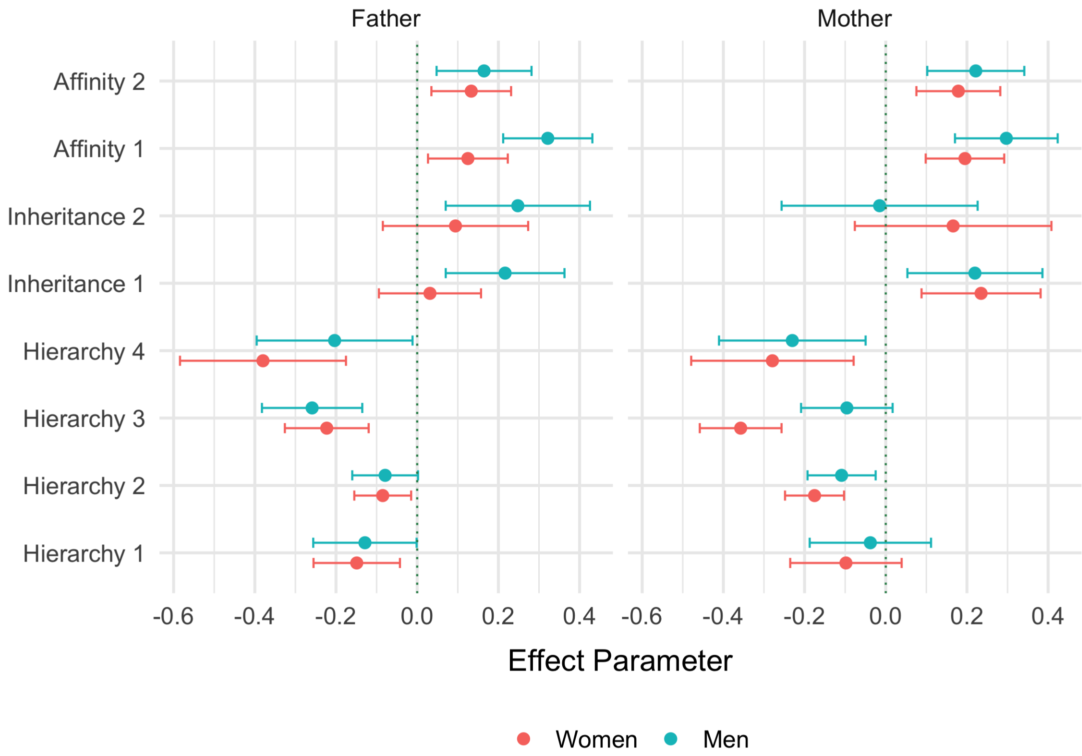

Figure A1: Effect parameters estimated under topological model.

B Replication of Main Analyses on US Data

In principle, the mechanisms behind the influence of mothers' class could be expected to apply irrespective of institutional context. Yet, in the main text I have highlighted some important contextual differences between the the UK and other countries especially the US. While my aim in this paper is not to provide a formal comparative analysis of the effect sizes of fathers' and mothers' class, I have noted some key contextual differences between the two countries. To draw out some particularities of the UK context compared with the US, where the issue of mothers' effects on children's class mobility outcomes has already been explored (Beller, 2009), I use the General Social Survey (GSS) to estimate trends in maternal employment across the 20th century in the US. Figure B1, I overlay fitted estimates (calculated using a spline function with 3 degrees of freedom) of mothers' employment rates in the US and UK as a function of year. As discussed in the main text, the general pattern of a steep rise in maternal employment in the UK is clearly consistent with other post-industrial countries including the US. Yet, as the Figure shows shows, there are two striking differences in maternal employment patterns in the US and UK. First, rates of maternal employment in the UK have never quite reached those evidence in the US: for cohorts born in the early 1980s, maternal employment in the US was near 80%, whereas in the UK, it was just under 70%. Second, for the earliest birth cohorts for which we have data (individuals born in the 1920s), mothers' employment was approximately 10% lower in the UK than in the US, and has since risen at a steeper rate.

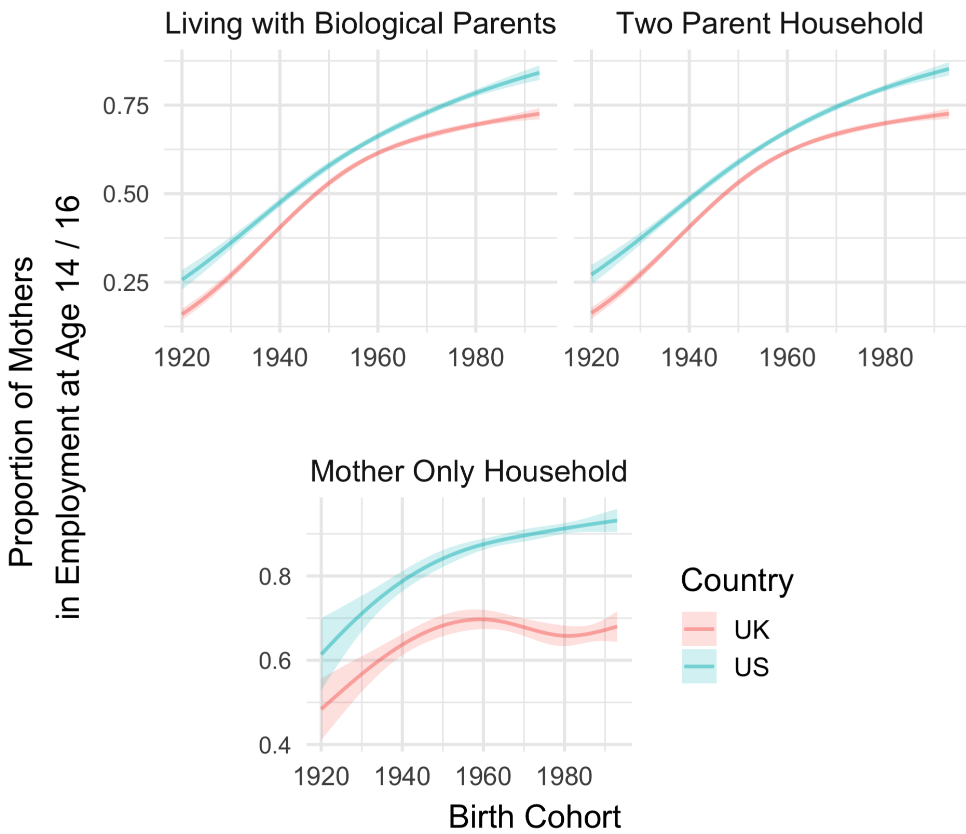

Figure B1: Fitted proportion of respondents living with a mother in employment by year of birth by country (UK and US), estimated per household type (respondents living with both biological parents, in a two parent household with a mother and father figure, or in a mother-only household), all at age 14 (UK) or age 16 (US). The conditional means are fitted as a natural spline of year of birth with three degrees of freedom on the UKHLS sample (UK) or the GSS 1972-2018 (US) and adjusted by sampling weights. Ribbons represent 95 percent confidence intervals.

Next, I use the 1994-2018 waves of the GSS to replicate my cross-sectional analyses for the US to outline some key similarities and differences between parental effects in the two countries. I restrict the GSS s ample to individuals born between 1944 and 1984 (to make the age distribution of the sample as similar to the UKHLS as possible). I apply the same sample restrictions as those applied to the UK data, as described in the main text. The key results can be summarized as follows. For men and women, from the reduction of X^2 following the addition of spousal class, it is clear that mothers and fathers independently affect children's class outcomes. However, model 4, which posits that mothers' and fathers' occupational effects on children's class outcomes are equal, provides for both men and women a poor fit to the data, while the main effects model (which results from relaxing the assumption of equality in parental effects in model 4) makes a significant improvement in fit in all cases. Therefore, unlike in the UK case, it appears that parental effects are not of equal size: fathers matter more for children's mobility classes.

Table B1: Fit statistics for mobility models fitted to 2009 mobility table. Men aged 25-64 (N=5,389).

						Model comparison			
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
1 Independence	654.1	648.6	175	0.00	14.1				
2 Father only	240.3	240.5	150	0.00	7.2	1 v 2	408.1	25.00	0.00
3 Mother only	421.4	416.7	150	0.00	11.2	1 v 3	231.9	25.00	0.00
4 Equal Effects	174.8	179.3	150	0.08	6.0				
5 Topological Effecs	224.7	230.4	159	0.00	6.8				
6 Main effects	120.8	126.7	125	0.59	4.4	2 v 6	113.8	25.00	0.00
						3 v 6	290.0	25.00	0.00
						4 v 6	52.6	25.00	0.00
						5 v 6	103.7	34.00	0.00
7 Father UNIDIFF	119.0	124.5	120	0.51	4.3	5 v 7	2.2	5.00	0.83
8 Mother UNIDIFF	116.9	121.4	120	0.56	4.3	5 v 8	5.3	5.00	0.38

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table B2: Fit statistics for mobility models fitted to 2009 mobility table. Women aged 25-64 (N=6,352).

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
1 Independence	582.2	553.7	175	0.00	11.1				
2 Father only	332.4	318.0	150	0.00	7.1	1 v 2	235.7	25.00	0.00
3 Mother only	271.8	272.2	150	0.00	7.6	1 v 3	281.5	25.00	0.00
4 Equal Effects	165.0	168.9	150	0.19	5.4				
5 Topological Effecs	200.8	203.4	159	0.01	5.8				
6 Main effects	120.4	128.0	125	0.60	4.4	2 v 6	190.0	25.00	0.00
						3 v 6	144.2	25.00	0.00
						4 v 6	40.8	25.00	0.02
						5 v 6	75.3	34.00	0.00
7 Father UNIDIFF	119.3	126.7	120	0.50	4.4	5 v 7	1.3	5.00	0.93
8 Mother UNIDIFF	109.4	116.1	120	0.75	4.2	5 v 8	11.9	5.00	0.04

Note: (1) dissimilarity index showing percentage of cases misclassified.

C Trends in and consequences of mothers' employment patterns

C.1 Trends in single parenthood

One disadvantage to an approach which examines fathers' and mothers' effects separately is that it necessarily excludes many disadvantaged subjects born into (mother-only) single parental households. This is increasingly so, since in recent years the proportion of single mother families has notably increased, as Figure C1 illustrates. In this respect, it is important to add a cautionary note on the mother and father joint measurement of social origins. The risk of a mechanic adoption of this standard to measurement social origins is that only intact families are considered.

At the same time, it is plausible that different single parent mothers exhibit different patterns and trends in mobility effects compared with mothers who live with a co-resident partner. If these patterns and effects operate in different directions to those among two-parent mothers, then full-sample estimates may fail to pick up patterns and trends specific to each of these groups. One particular source of difference between these two 'types' of mother can be understood in terms of differential rates of selectivity into employment. To illustrate this, Figure C2 plots the fitted proportion of respondents living with mother in employment over time for different household types, with the top two panels showing the proportion of working mothers in two-parent biological and two parental other households, and the bottom panel showing the proportion of working mothers in mother-only households. Two differences are striking. First, while approximately 50% of mothers in two-parent households were working for cohorts born in 1920 (that is, of mothers in 1934), the same figure is under 20% for individuals born in 1920 living in a two-parent household at age 14. Second, since for cohorts born in the early 1980s the proportion of working mothers stands at roughly the same figure (just under 70%) for all household types, the rate of change in maternal employment rates has been far steeper for two-parent households, compared with mother-only families. What could be the implications of these trends for trends in mobility rates among individuals from single-parent families? To the extent that the pronounced increase of and rate of increase in maternal employment has contributed to a weakening of positive selection among mothers in two-parent households, we might suppose that the degree of selectivity among mothers in single-parent households was never as high among early cohorts as it was among two-parent households: particularly before the expansion of welfare for single parent families, single-parent mothers would have undertaken employment out of material necessity. In consequence, the decline in social selectivity into employment among single-parents may have been less steep among later cohorts than among mothers living in two-parent households.

As a preliminary investigation of this account, in Tables C1 and C2 I present the results of fitting two mother-only models to respondents living in a mother-only household at age 14 for the 1991 and 2009 samples, with four different sample restrictions/specifications: 5 class categorization using the full sample of UKHLS 2009 respondents; 5 class categorization using only those respondents born strictly after BHPS 1991 respondents; 7 class categorization using the full sample of UKHLS 2009 respondents, and 7 class categorization using only those respondents born strictly after BHPS 1991 respondents. The two models I fit are as follows. First, I fit a main effects model for mothers' class position which states that the association between mothers and children is the same across the two time periods:

$$lnF_{ijkl} = \mu + \lambda_j^C + \lambda_i^M + \lambda_k^T + \lambda_{kj}^{TC} + \lambda_{ij}^{MC}$$

where F_{ijk} is the expected frequency of the ijk^{th} cell of the four-way table comprising child's (C) and mothers' (M) NS-SEC and time period (T). Second, a UNIDIFF model which tests for the possibility that the general pattern of association between mothers and children increases or decreases uniformly over time (thus replacing the λ_{ij}^{MC} term with $\psi_{ij}^{MC}\phi_k$).

As Tables C1 and C2 show, there is some indication that the effect of single parental mothers' class position on sons does not weaken as strongly as it does for mothers in two-parent households (as captured by the UNIDIFF point estimates). Nevertheless, the small sample sizes in each case been that our estimates are underpowered, and estimation uncertainty prevents us from reaching a definitive conclusion about trends in single mothers' fluidity patterns.

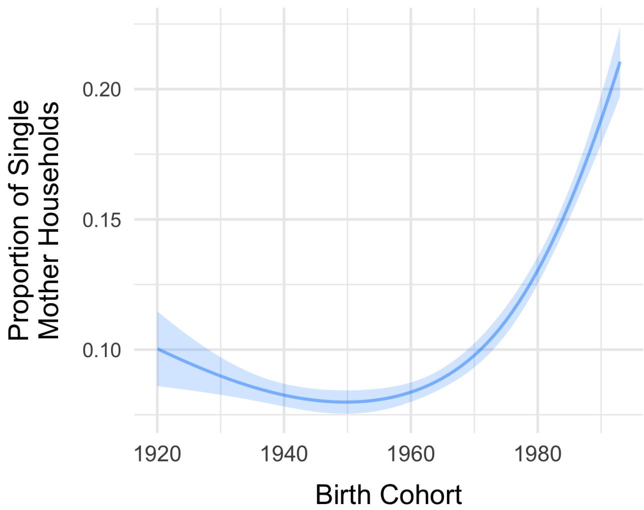

Figure C1: Fitted proportion of respondents living with in a single mother household, by year of birth, estimated using UKHLS 2009. The conditional means are fitted as a natural spline of year of birth with three degrees of freedom and adjusted by sampling weights. Ribbons represent 95 percent confidence intervals.

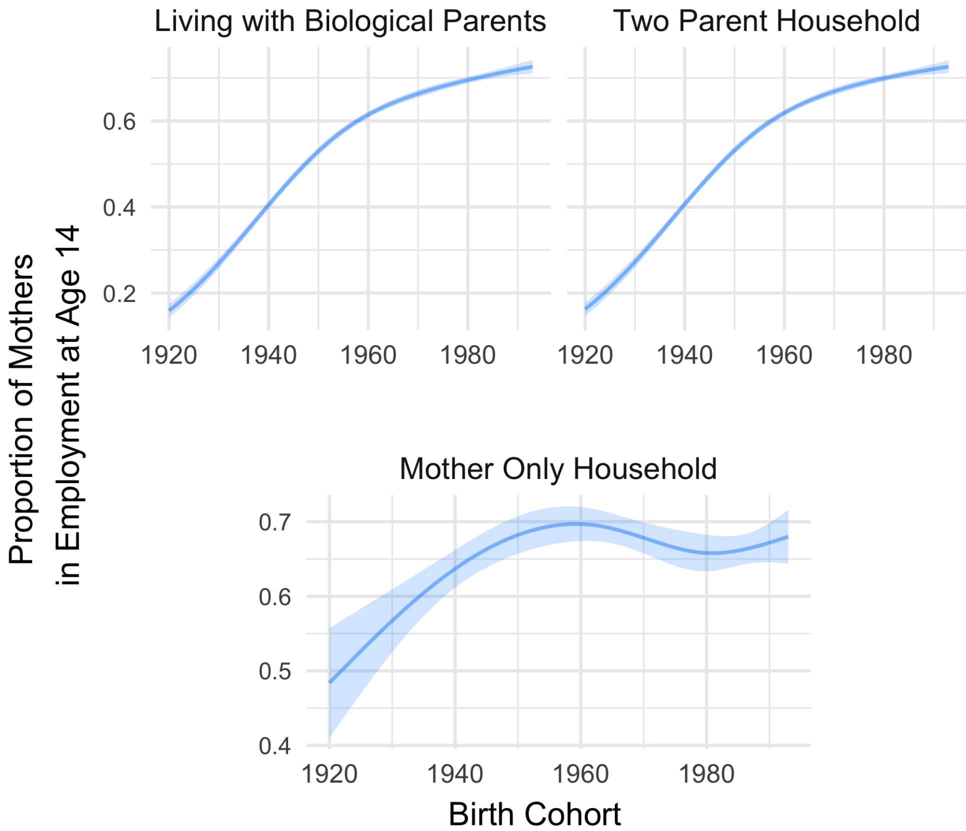

Figure C2: Fitted proportion of respondents living with a mother in employment by year of birth, estimated per household type (respondents living with both biological parents, in a two parent household with a mother and father figure, or in a mother-only household), all at age 14. The conditional means are fitted as a natural spline of year of birth with three degrees of freedom on the UKHLS sample and adjusted by sampling weights. Ribbons represent 95 percent confidence intervals.

Table C1: Fit statistics for log-linear and log-multiplicative mother-only models fitted to mobility tables 1991-2009 for respondents living in mother-only households at age 14. Men aged 25-64 ($N = 795$ for full sample and $N = 338$ for no birth overlap sample).

							Model comparison				
Sample	Model	X^2	G^2	df	$P(X^2)$	DI^1	ϕ (95% CI)	rG^2	rdf	p	
7 class (no overlap)	CSF	30.3	31.7	36	0.74	7.8					
7 class (no overlap)	UNIDIFF	33.3	31.7	35	0.55	8.0	0.880 (1.980, 0.390)	1 v 2	0.0	1.00	0.90
7 class (with overlap)	CSF	33.5	33.2	36	0.59	3.9					
7 class (with overlap)	UNIDIFF	31.4	33.2	35	0.64	3.7	1.120 (2.790, 0.450)	1 v 2	0.0	1.00	0.86
5 class (no overlap)	CSF	10.1	13.2	16	0.86	3.3					
5 class (no overlap)	UNIDIFF	14.8	11.4	15	0.47	4.9	0.410 (1.010, 0.170)	1 v 2	1.8	1.00	0.18
5 class (with overlap)	CSF	13.8	16.2	16	0.62	1.8					
5 class (with overlap)	UNIDIFF	13.6	16.2	15	0.55	1.8	1.060 (3.210, 0.350)	1 v 2	0.0	1.00	0.95

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table C2: Fit statistics for log-linear and log-multiplicative mother-only models fitted to mobility tables 1991-2009 for respondents living in mother-only households at age 14. Women aged 25-64 ($N = 1,161$ for full sample and $N = 508$ for no birth overlap sample).

							Model comparison				
Sample	Model	X^2	G^2	df	$P(X^2)$	DI^1	ϕ (95% CI)	rG^2	rdf	p	
7 class (no overlap)	CSF	22.8	28.3	36	0.96	5.3					
7 class (no overlap)	UNIDIFF	4518789493.0	20.6	35	0.00	6.1	0.190 (0.470, 0.080)	1 v 2	7.6	1.00	0.01
7 class (with overlap)	CSF	23.9	28.1	36	0.94	2.4					
7 class (with overlap)	UNIDIFF	24.4	27.4	35	0.91	2.7	0.680 (1.300, 0.360)	1 v 2	0.7	1.00	0.40
5 class (no overlap)	CSF	12.7	14.6	16	0.70	3.9					
5 class (no overlap)	UNIDIFF	10.7	10.9	15	0.78	3.6	0.360 (0.780, 0.160)	1 v 2	3.6	1.00	0.06
5 class (with overlap)	CSF	13.6	15.9	16	0.63	1.8					
5 class (with overlap)	UNIDIFF	13.3	14.5	15	0.58	1.9	0.590 (1.140, 0.310)	1 v 2	1.4	1.00	0.24

Note: (1) dissimilarity index showing percentage of cases misclassified.

D Results under Alternative Sample Restriction Criteria and Class Categorizations for Cross Sectional Analysis

In my main analyses, I use a 7-class categorization for parents' and respondents' class positions, and restrict to the sample to children living with any two resident parents. To assess the sensitivity of my results to these particular sample specifications, I conduct a series of parallel cross-sectional analyses under alternative restrictions and class categorizations. Specifically, I show results under three alternative sample selection criteria/class categorizations: (i) children living with their biological parents at age 14, using a 7-class categorization; (ii) children living with any two resident parents at age 14, using a 5-class categorization, and (iii) children living with their biological parents at age 14, using a 5-class categorization. The 5-class categorization is identical to the 7-class categorization shown in Table 1 in the main text, other than it merges Classes 1 and 2 and 6 and 7. The results are shown in Tables E1-E5 below, and the UNIDIFF estimates returned under model 6 in Tables We can see that, in all cases, the model allowing the mother and father effects to differ does not significantly improve over the model constraining these effects to be equal, and that, further, the father-only model fits better than the mother-only model for men, while the reverse is true for women. These alternative results strongly support the conclusions I draw in the main text.

Table E1: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 2009. Men and women aged 25-64 living with **both biological parents** aged 14, **7 class categorization**.

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
$N = 13,416$									
1 Independence	2088.7	2138.2	576	0.00	15.3				
2 Father only	1058.3	1079.8	540	0.00	10.0	1 v 2	1058.4	36.00	0.00
3 Mother only	1007.3	1002.8	540	0.00	9.0	1 v 3	1135.4	36.00	0.00
4 Equal effects	549.7	576.8	540	0.38	6.6				
5 Main effects	511.6	537.7	504	0.40	6.3	2 v 5	542.2	36.00	0.00
						3 v 5	465.2	36.00	0.00
						4 v 5	39.2	36.00	0.33
6 Father & Mother UNIDIFF	489.5	515.1	502	0.65	6.1	5 v 6	22.5	2.00	0.00

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table E2: UNIDIFF parameter estimates returned under model 6 in Table E1 with 95% confidence intervals scaled to $\phi_{\text{men}} = 1$: men and women aged 25-64.

Parental Effect	Point Estimate	Upper Bound	Lower Bound
Father	0.75	0.91	0.62
Mother	1.54	1.88	1.27

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

Table E3: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 2009. Men and women aged 25-64 living with **any two resident parents** aged 14, **5 class categorization**.

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
$N = 14,327$									
1 Independence	2088.7	2138.2	576	0.00	15.3				
2 Father only	1058.3	1079.8	540	0.00	10.0	1 v 2	1058.4	36.00	0.00
3 Mother only	1007.3	1002.8	540	0.00	9.0	1 v 3	1135.4	36.00	0.00
4 Equal effects	549.7	576.8	540	0.38	6.6				
5 Main effects	511.6	537.7	504	0.40	6.3	2 v 5	542.2	36.00	0.00
						3 v 5	465.2	36.00	0.00
						4 v 5	39.2	36.00	0.33
6 Father & Mother UNIDIFF	489.5	515.1	502	0.65	6.1	5 v 6	22.5	2.00	0.00

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table E4: UNIDIFF parameter estimates returned under model 6 in Table E3 with 95% confidence intervals scaled to $\phi_{\text{men}} = 1$: men and women aged 25-64.

Parental Effect	Point Estimate	Upper Bound	Lower Bound
Father	0.75	0.91	0.62
Mother	1.54	1.88	1.27

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

Table E5: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 2009. Men and women aged 25-64 living with **both biological parents** aged 14, **5 class categorization**.

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
$N = 13,416$									
1 Independence	2088.7	2138.2	576	0.00	15.3				
2 Father only	1058.3	1079.8	540	0.00	10.0	1 v 2	1058.4	36.00	0.00
3 Mother only	1007.3	1002.8	540	0.00	9.0	1 v 3	1135.4	36.00	0.00
4 Equal effects	549.7	576.8	540	0.38	6.6				
5 Main effects	511.6	537.7	504	0.40	6.3	2 v 5	542.2	36.00	0.00
						3 v 5	465.2	36.00	0.00
						4 v 5	39.2	36.00	0.33
6 Father & Mother UNIDIFF	489.5	515.1	502	0.65	6.1	5 v 6	22.5	2.00	0.00

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table E6: UNIDIFF parameter estimates returned under model 6 in Table E5 with 95% confidence intervals scaled to $\phi_{\text{men}} = 1$: men and women aged 25-64.

Parental Effect	Point Estimate	Upper Bound	Lower Bound
Father	0.75	0.91	0.62
Mother	1.54	1.88	1.27

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

E Results under Alternative Sample Restriction Criteria and Class Categorizations for Over Time Analysis

In my main over-time analyses, I use a 7-class categorization for parents' and respondents' class positions, and restrict to the sample to non-overlapping birth cohorts and to children living with any two resident parents. To assess the sensitivity of my results to these particular sample specifications, I have conducted a series of parallel over-time analyses under alternative restrictions and class categorizations. The results are shown in Tables F1- F6 below, which parallel the table and figure presented in Appendix **??**. We can see that the trend of decreasing maternal class effects on men and women, and decreasing paternal class effects on women only, are highly consistent across sample specifications (namely, regardless of whether we restrict parents to biological only, and whether we categorize social class using a 5-class collapse). As an illustration, Tables F1 and F2 report results for biological parents using a 7-class categorization. Here, for men, while the father UNIDIFF model clearly fails to improve in fit over the father CSF model, the father CSF & mother UNIDIFF performs significantly better than the father & mother constant effects at $p = .03$, and in turn adding an additional UNIDIFF parameter for the father-son association fails to improve on this model. The UNIDIFF parameter returned under the father CSF & mother UNIDIFF model clearly shows that the mother-son association decreased in strength over time. For women, as is indicated by the models in the main text, the UNIDIFF model significantly improves in fit for the father-only model, and the joint-parent measures also reject a pattern of constant fluidity for both mothers and fathers. The UNIDIFF parameters returned under the models for women suggest that both the father-daughter and mother-daughter relationships decreased over the cohorts.

Table F1: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 1991-2009 (no birth overlap). Men and women aged 25-64 living with **both biological parents** aged 14, **7 class categorization**.

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
Men, $N = 2,665$									
7 Father & mother CSF	501.0	458.4	504	0.53	14.6				
8 Father UNIDIFF & mother CSF	500.4	458.4	503	0.52	14.6	7 v 8	0.0	1.00	0.86
9 Father CSF & mother UNIDIFF	501.5	453.9	503	0.51	14.6	7 v 9	4.5	1.00	0.03
10 Father UNIDIFF & mother UNIDIFF	497.5	453.3	502	0.55	14.5	9 v 10	0.7	1.00	0.42
Women, $N = 2,010$									
7 Father & mother CSF	467.0	487.2	504	0.88	12.6				
8 Father UNIDIFF & mother CSF	474.7	478.3	503	0.81	12.5	7 v 8	8.9	1.00	0.00
9 Father CSF & mother UNIDIFF	465.4	480.9	503	0.88	12.6	7 v 9	6.2	1.00	0.01
10 Father UNIDIFF & mother UNIDIFF	471.3	475.2	502	0.83	12.5	9 v 10	5.7	1.00	0.02

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table F2: UNIDIFF parameter estimates from models 7-10 fitted to mobility tables 1991-2009 (no birth overlap) and scaled to $\phi_{1991} = 1$. Men and women aged 25-64 living with **both biological parents** aged 14, **7 class categorization**.

Parental Effect (Model Number)	Point Estimate	Upper Bound	Lower Bound
Men, $N = 2,665$			
Father (8)	0.47	0.75	0.29
Father (10)	0.50	0.79	0.31
Mother (9)	1.04	1.60	0.68
Mother (10)	1.25	2.06	0.76
Women, $N = 2,010$			
Father (8)	0.48	0.70	0.33
Father (10)	0.53	0.78	0.35
Mother (9)	0.54	0.80	0.37
Mother (10)	0.63	0.96	0.42

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

Table F3: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 1991-2009 (no birth overlap). Men and women aged 25-64 living with **any two resident parents** aged 14, **5 class categorization**.

						Model comparison			
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
Men, $N = 2,937$									
7 Father & mother CSF	178.3	168.0	160	0.15	7.3				
8 Father UNIDIFF & mother CSF	176.4	167.5	159	0.16	7.3	7 v 8	0.5	1.00	0.49
9 Father CSF & mother UNIDIFF	175.0	164.7	159	0.18	7.1	7 v 9	3.3	1.00	0.07
10 Father UNIDIFF & mother UNIDIFF	170.0	163.1	158	0.24	7.0	9 v 10	1.6	1.00	0.21
Women, $N = 2,157$									
7 Father & mother CSF	170.6	177.3	160	0.27	6.2				
8 Father UNIDIFF & mother CSF	160.1	170.2	159	0.46	5.9	7 v 8	7.1	1.00	0.01
9 Father CSF & mother UNIDIFF	167.2	173.9	159	0.31	6.1	7 v 9	3.4	1.00	0.07
10 Father UNIDIFF & mother UNIDIFF	159.1	169.1	158	0.46	6.0	9 v 10	4.8	1.00	0.03

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table F4: UNIDIFF parameter estimates from models 7-10 fitted to mobility tables 1991-2009 (no birth overlap) and scaled to $\phi_{1991} = 1$. Men and women aged 25-64 living with **any two resident parents** aged 14, **5 class categorization**.

Parental Effect (Model Number)	Point Estimate	Upper Bound	Lower Bound
Men, $N = 2,937$			
Father (8)	0.48	0.81	0.29
Father (10)	0.53	0.89	0.32
Mother (9)	1.13	1.73	0.74
Mother (10)	1.45	2.69	0.78
Women, $N = 2,157$			
Father (8)	0.50	0.76	0.33
Father (10)	0.53	0.83	0.34
Mother (9)	0.62	0.86	0.45
Mother (10)	0.77	1.21	0.49

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

Table F5: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 1991-2009 (no birth overlap). Men and women aged 25-64 living with **both biological parents** aged 14, **5 class categorization**.

Model	X^2	G^2	df	$P(X^2)$	DI^1		Model comparison rG^2	rdf	p
Men, $N = 2,665$									
7 Father & mother CSF	186.7	166.2	160	0.07	7.5				
8 Father UNIDIFF & mother CSF	185.5	165.6	159	0.07	7.6	7 v 8	0.6	1.00	0.42
9 Father CSF & mother UNIDIFF	184.4	163.3	159	0.08	7.2	7 v 9	2.9	1.00	0.09
10 Father UNIDIFF & mother UNIDIFF	180.3	161.5	158	0.11	7.2	9 v 10	1.8	1.00	0.18
Women, $N = 2,010$									
7 Father & mother CSF	174.3	170.0	160	0.21	6.3				
8 Father UNIDIFF & mother CSF	163.0	163.1	159	0.40	6.1	7 v 8	6.9	1.00	0.01
9 Father CSF & mother UNIDIFF	170.8	166.3	159	0.25	6.2	7 v 9	3.7	1.00	0.05
10 Father UNIDIFF & mother UNIDIFF	162.3	161.8	158	0.39	6.1	9 v 10	4.5	1.00	0.03

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table F6: UNIDIFF parameter estimates from models 7-10 fitted to mobility tables 1991-2009 (no birth overlap) and scaled to $\phi_{1991} = 1$. Men and women aged 25-64 living with **both biological parents** aged 14, **5 class categorization**.

Parental Effect (Model Number)	Point Estimate	Upper Bound	Lower Bound
Men, $N = 2,665$			
Father (8)	0.49	0.83	0.29
Father (10)	0.55	0.93	0.32
Mother (9)	1.16	1.75	0.77
Mother (10)	1.46	2.64	0.80
Women, $N = 2,010$			
Father (8)	0.51	0.77	0.34
Father (10)	0.55	0.85	0.36
Mother (9)	0.62	0.86	0.45
Mother (10)	0.74	1.19	0.46

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

Moreover, in my main analyses I reduce the UKHLS and BHPS samples to produce non-overlapping birth cohorts whose age distributions at the time of observation are similar. This is so that I am able to make a comparison of different cohorts, since the theory on which I base my hypotheses (see especially Breen and Jonsson, 2007) is based on a cohort, rather than period, effect.

The full and restricted sample distributions of age per wave is shown in Figure F1

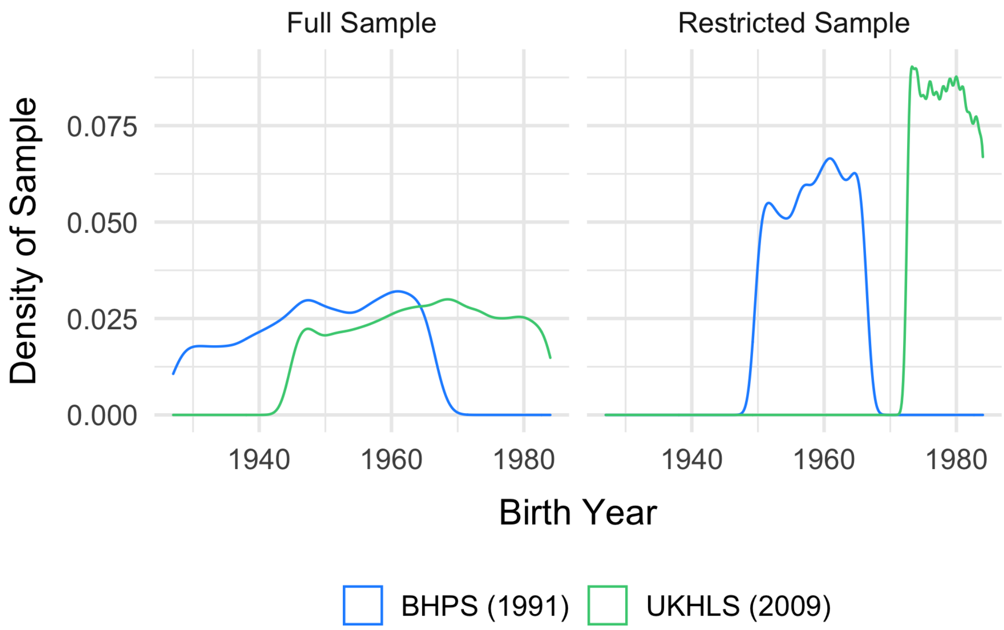

Figure F1: Distribution of birth year of BHPS and UKHLS respondents in the full sample of individuals aged 25-64 in each survey, and in the restricted sample of non-overlapping birth cohorts with similar age profiles, which I use in my main analyses.

Despite the advantage of this approach to directly assess cohort-related changes in social fluidity, it necessarily reduces the sample size by a considerable amount, and further means that some of the individuals who feature in my cross-sectional analysis are in the event dropped for the over-time analysis. To assess the sensitivity of my results to excluding these individuals, I have conducted a series of alternative analyses which include the full sample of in both the BHPS and UKHLS aged 25-64 at the time of the surveys. My results, which are shown in Tables F7-F10, are not sensitive to including the full sample of respondents.

Table F7: Fit statistics for log-linear and log-multiplicative models 7-10 fitted to mobility tables 1991-2009 (with birth overlap). Men and women aged 25-64 living with **any two resident parents** aged 14, **7 class categorization**.

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
Men, $N = 8,818$									
7 Father & mother CSF	493.7	515.9	504	0.62	7.5				
8 Father UNIDIFF & mother CSF	493.4	515.9	503	0.61	7.5	7 v 8	0.0	1.00	0.95
9 Father CSF & mother UNIDIFF	493.6	512.3	503	0.61	7.6	7 v 9	3.6	1.00	0.06
10 Father UNIDIFF & mother UNIDIFF	489.3	511.8	502	0.65	7.6	9 v 10	0.5	1.00	0.48
Women, $N = 6,499$									
7 Father & mother CSF	521.5	564.5	504	0.29	6.9				
8 Father UNIDIFF & mother CSF	522.4	554.2	503	0.27	6.9	7 v 8	10.2	1.00	0.00
9 Father CSF & mother UNIDIFF	516.3	558.8	503	0.33	6.9	7 v 9	5.7	1.00	0.02
10 Father UNIDIFF & mother UNIDIFF	517.7	552.2	502	0.30	6.9	9 v 10	6.6	1.00	0.01

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table F8: UNIDIFF parameter estimates from models 7-10 fitted to mobility tables 1991-2009 (with birth overlap) and scaled to $\phi_{1991} = 1$. Men and women aged 25-64 living with **any two both biological parents** aged 14, **7 class categorization**.

Parental Effect (Model Number)	Point Estimate	Upper Bound	Lower Bound
Men, $N = 8,218$			
Father (8)	0.60	0.89	0.41
Father (10)	0.63	0.92	0.43
Mother (9)	0.99	1.43	0.69
Mother (10)	1.13	1.74	0.74
Women, $N = 6,136$			
Father (8)	0.54	0.73	0.40
Father (10)	0.58	0.80	0.41
Mother (9)	0.66	0.85	0.52
Mother (10)	0.78	1.07	0.57

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

Table F9: Fit statistics for log-linear and log-multiplicative models 1-6 fitted to mobility tables 1991-2009 (with birth overlap). Men and women aged 25-64 living with **any two resident parents** aged 14, **5 class categorization**.

							Model comparison		
Model	X^2	G^2	df	$P(X^2)$	DI^1		rG^2	rdf	p
Men, $N = 8,818$									
7 Father & mother CSF	191.6	189.4	160	0.04	4.2				
8 Father UNIDIFF & mother CSF	191.0	189.3	159	0.04	4.2	7 v 8	0.1	1.00	0.76
9 Father CSF & mother UNIDIFF	190.6	187.4	159	0.04	4.1	7 v 9	2.0	1.00	0.16
10 Father UNIDIFF & mother UNIDIFF	188.7	186.7	158	0.05	4.1	9 v 10	0.6	1.00	0.42
Women, $N = 6,499$									
7 Father & mother CSF	212.0	213.3	160	0.00	3.9				
8 Father UNIDIFF & mother CSF	206.5	207.0	159	0.01	3.8	7 v 8	6.3	1.00	0.01
9 Father CSF & mother UNIDIFF	203.8	206.9	159	0.01	3.8	7 v 9	6.4	1.00	0.01
10 Father UNIDIFF & mother UNIDIFF	201.2	203.8	158	0.01	3.8	9 v 10	3.2	1.00	0.08

Note: (1) dissimilarity index showing percentage of cases misclassified.

Table F10: UNIDIFF parameter estimates from models 7-10 fitted to mobility tables 1991-2009 (with birth overlap) and scaled to $\phi_{1991} = 1$. Men and women aged 25-64 living with **any two biological parents** aged 14, **5 class categorization**.

Parental Effect (Model Number)	Point Estimate	Upper Bound	Lower Bound
Men, $N = 8,818$			
Father (8)	0.64	0.96	0.42
Father (10)	0.68	1.02	0.45
Mother (9)	1.04	1.44	0.75
Mother (10)	1.21	1.94	0.75
Women, $N = 6,499$			
Father (8)	0.58	0.81	0.41
Father (10)	0.64	0.86	0.48
Mother (9)	0.65	0.95	0.44
Mother (10)	0.70	0.97	0.51

Note: 95% upper & lower bounds calculated using quasi SEs (Firth, 2004).

As I describe in the main text, although the age distribution of each of my two samples (sets of cohorts) is similar, the BHPS sample is on average 2 years older than the 2009 sample (32.7 years vs 30.7 years), and features a greater number of older respondents (the age range of the BHPS sample is 25-41, whereas that of the UKHLS sample is 25-36). As an additional robustness check, I create a BHPS-based cohort with even more similar age profile to the UKHLS-based cohort, such that this cohort has an average age of 30.3 and age range of 25-36. The results under this alternative specification (not shown here) are highly consistent with those reported elsewhere. It should be noted that, although the substance of the results (in terms of parameters returned) are not sensitive to further alterations of this sort, further reducing the size of the BHPS sample leads to a significant number of sampling zeros, and thus underpowered results.

Finally, in order to assess the sensitivity of my results to the size of the birth cohorts chosen as well as the period of time between them, I have conducted a further set of over-time analyses based on a range of alternative cohort sizes, period spans and starting years. Tables F11-F14 below show my estimates of models testing, for each child gender and parent in turn: (i) a mother-father CSF model (Equation 3 in Table **??**) versus (ii) parental UNIDIFF model. The UNIDIFF model is represented by Equations 4 and 5 for fathers and mothers,, respectively, in Table **??**. We can see that the conclusions described in the main text regarding a weakening of maternal class effects for men are highly consistent with these supplementary analyses.

Table F11: Fit statistics for log-linear and log-multiplicative models 3 and 4 (**mother-father CSF and mother UNIDIFF**) fitted to mobility tables 1991-2009 for various cohort specifications. **Men**.

									Model comparison			
Schema	Cohort	Model	X^2	G^2	df	$P(X^2)$	DI	ϕ (95% CI)	rG^2	rdf	p	
7 Class	1950-1964 & 1965-1979	CSF	455.1	450.8	504	0.94	11.2					
7 Class	1950-1964 & 1965-1979	UNIDIFF	450.9	446.8	503	0.95	11.4	0.480 (0.764, 0.302)	CSF v UNIDIFF	4.0	1.00	0.05
7 Class	1946-1960 & 1965-1979	CSF	479.5	465.5	504	0.78	11.2					
7 Class	1946-1960 & 1965-1979	UNIDIFF	480.1	461.3	503	0.76	11.4	0.467 (0.754, 0.290)	CSF v UNIDIFF	4.1	1.00	0.04
7 Class	1941-1955 & 1965-1979	CSF	507.1	463.0	504	0.45	11.1					
7 Class	1941-1955 & 1965-1979	UNIDIFF	503.1	452.9	503	0.49	11.2	0.308 (0.497, 0.191)	CSF v UNIDIFF	10.1	1.00	0.00
7 Class	1940-1959 & 1960-1979	CSF	486.3	473.7	504	0.71	9.7					
7 Class	1940-1959 & 1960-1979	UNIDIFF	487.0	468.2	503	0.69	9.8	0.473 (0.722, 0.310)	CSF v UNIDIFF	5.4	1.00	0.02
7 Class	1936-1955 & 1960-1979	CSF	493.9	461.3	504	0.62	9.5					
7 Class	1936-1955 & 1960-1979	UNIDIFF	488.1	453.8	503	0.68	9.5	0.387 (0.599, 0.250)	CSF v UNIDIFF	7.5	1.00	0.01
7 Class	1931-1950 & 1960-1979	CSF	421.9	396.4	504	1.00	8.9					
7 Class	1931-1950 & 1960-1979	UNIDIFF	417.6	391.6	503	1.00	8.9	0.422 (0.685, 0.260)	CSF v UNIDIFF	4.8	1.00	0.03
5 Class	1950-1964 & 1965-1979	CSF	154.3	160.5	160	0.61	5.7					
5 Class	1950-1964 & 1965-1979	UNIDIFF	151.8	158.9	159	0.65	5.7	0.617 (1.061, 0.359)	CSF v UNIDIFF	1.6	1.00	0.21
5 Class	1946-1960 & 1965-1979	CSF	163.6	173.3	160	0.41	6.1					
5 Class	1946-1960 & 1965-1979	UNIDIFF	163.7	169.3	159	0.38	5.8	0.498 (0.828, 0.299)	CSF v UNIDIFF	4.0	1.00	0.05
5 Class	1941-1955 & 1965-1979	CSF	165.3	173.9	160	0.37	5.9					
5 Class	1941-1955 & 1965-1979	UNIDIFF	176.7	164.8	159	0.16	5.7	0.333 (0.556, 0.200)	CSF v UNIDIFF	9.2	1.00	0.00
5 Class	1940-1959 & 1960-1979	CSF	169.1	175.7	160	0.29	5.1					
5 Class	1940-1959 & 1960-1979	UNIDIFF	176.8	171.2	159	0.16	5.1	0.505 (0.805, 0.317)	CSF v UNIDIFF	4.5	1.00	0.03
5 Class	1936-1955 & 1960-1979	CSF	165.9	175.6	160	0.36	5.0					
5 Class	1936-1955 & 1960-1979	UNIDIFF	173.6	169.7	159	0.20	4.9	0.414 (0.678, 0.253)	CSF v UNIDIFF	6.0	1.00	0.01
5 Class	1931-1950 & 1960-1979	CSF	143.0	151.9	160	0.83	4.6					
5 Class	1931-1950 & 1960-1979	UNIDIFF	146.6	148.8	159	0.75	4.6	0.488 (0.869, 0.273)	CSF v UNIDIFF	3.1	1.00	0.08

Note: (1) CSF and UNIDIFF models refer to models 3 and 5, respectively, in Table D1.

Table F12: Fit statistics for log-linear and log-multiplicative models 3 and 4 (**mother-father CSF and father UNIDIFF**) fitted to mobility tables 1991-2009 for various cohort specifications. **Men**.

Schema	Cohort	Model	X^2	G^2	df	$P(X^2)$	DI	ϕ (95% CI)	Model comparison rG^2	rdf	p	
7 Class	1950-1964 & 1965-1979	CSF	455.1	450.8	504	0.94	11.2					
7 Class	1950-1964 & 1965-1979	UNIDIFF	453.9	450.7	503	0.94	11.2	1.088 (1.780, 0.665)	CSF v UNIDIFF	0.1	1.00	0.76
7 Class	1946-1960 & 1965-1979	CSF	479.5	465.5	504	0.78	11.2					
7 Class	1946-1960 & 1965-1979	UNIDIFF	478.3	465.4	503	0.78	11.2	1.064 (1.810, 0.625)	CSF v UNIDIFF	0.0	1.00	0.85
7 Class	1941-1955 & 1965-1979	CSF	507.1	463.0	504	0.45	11.1					
7 Class	1941-1955 & 1965-1979	UNIDIFF	500.8	462.5	503	0.52	11.0	1.364 (2.944, 0.632)	CSF v UNIDIFF	0.5	1.00	0.49
7 Class	1940-1959 & 1960-1979	CSF	486.3	473.7	504	0.71	9.7					
7 Class	1940-1959 & 1960-1979	UNIDIFF	482.7	473.6	503	0.74	9.7	1.113 (1.877, 0.660)	CSF v UNIDIFF	0.1	1.00	0.73
7 Class	1936-1955 & 1960-1979	CSF	493.9	461.3	504	0.62	9.5					
7 Class	1936-1955 & 1960-1979	UNIDIFF	485.6	460.7	503	0.70	9.5	1.354 (2.738, 0.669)	CSF v UNIDIFF	0.7	1.00	0.42
7 Class	1931-1950 & 1960-1979	CSF	421.9	396.4	504	1.00	8.9					
7 Class	1931-1950 & 1960-1979	UNIDIFF	419.3	396.0	503	1.00	8.9	1.309 (3.058, 0.561)	CSF v UNIDIFF	0.4	1.00	0.55
5 Class	1950-1964 & 1965-1979	CSF	154.3	160.5	160	0.61	5.7					
5 Class	1950-1964 & 1965-1979	UNIDIFF	154.2	160.5	159	0.59	5.8	1.056 (1.774, 0.628)	CSF v UNIDIFF	0.0	1.00	0.84
5 Class	1946-1960 & 1965-1979	CSF	163.6	173.3	160	0.41	6.1					
5 Class	1946-1960 & 1965-1979	UNIDIFF	163.9	173.2	159	0.38	6.1	0.931 (1.557, 0.557)	CSF v UNIDIFF	0.1	1.00	0.81
5 Class	1941-1955 & 1965-1979	CSF	165.3	173.9	160	0.37	5.9					
5 Class	1941-1955 & 1965-1979	UNIDIFF	164.0	173.7	159	0.38	6.0	1.191 (2.503, 0.567)	CSF v UNIDIFF	0.2	1.00	0.65
5 Class	1940-1959 & 1960-1979	CSF	169.1	175.7	160	0.29	5.1					
5 Class	1940-1959 & 1960-1979	UNIDIFF	169.3	175.7	159	0.27	5.1	0.987 (1.624, 0.599)	CSF v UNIDIFF	0.0	1.00	0.96
5 Class	1936-1955 & 1960-1979	CSF	165.9	175.6	160	0.36	5.0					
5 Class	1936-1955 & 1960-1979	UNIDIFF	163.8	175.3	159	0.38	5.0	1.225 (2.453, 0.611)	CSF v UNIDIFF	0.3	1.00	0.56
5 Class	1931-1950 & 1960-1979	CSF	143.0	151.9	160	0.83	4.6					
5 Class	1931-1950 & 1960-1979	UNIDIFF	142.9	151.9	159	0.82	4.6	1.021 (2.116, 0.493)	CSF v UNIDIFF	0.0	1.00	0.96

Note: (1) CSF and UNIDIFF models refer to models 3 and 4, respectively, in Table D1.

Table F13: Fit statistics for log-linear and log-multiplicative models 3 and 4 (**mother-father CSF and mother UNIDIFF**) fitted to mobility tables 1991-2009 for various cohort specifications. **Women**.

									Model comparison			
Schema	Cohort	Model	X^2	G^2	df	$P(X^2)$	DI	ϕ (95% CI)	rG^2	rdf	p	
7 Class	1950-1964 & 1965-1979	CSF	455.1	450.8	504	0.94	11.2					
7 Class	1950-1964 & 1965-1979	UNIDIFF	450.9	446.8	503	0.95	11.4	0.480 (0.764, 0.302)	CSF v UNIDIFF	4.0	1.00	0.05
7 Class	1946-1960 & 1965-1979	CSF	479.5	465.5	504	0.78	11.2					
7 Class	1946-1960 & 1965-1979	UNIDIFF	480.1	461.3	503	0.76	11.4	0.467 (0.754, 0.290)	CSF v UNIDIFF	4.1	1.00	0.04
7 Class	1941-1955 & 1965-1979	CSF	507.1	463.0	504	0.45	11.1					
7 Class	1941-1955 & 1965-1979	UNIDIFF	503.1	452.9	503	0.49	11.2	0.308 (0.497, 0.191)	CSF v UNIDIFF	10.1	1.00	0.00
7 Class	1940-1959 & 1960-1979	CSF	486.3	473.7	504	0.71	9.7					
7 Class	1940-1959 & 1960-1979	UNIDIFF	487.0	468.2	503	0.69	9.8	0.473 (0.722, 0.310)	CSF v UNIDIFF	5.4	1.00	0.02
7 Class	1936-1955 & 1960-1979	CSF	493.9	461.3	504	0.62	9.5					
7 Class	1936-1955 & 1960-1979	UNIDIFF	488.1	453.8	503	0.68	9.5	0.387 (0.599, 0.250)	CSF v UNIDIFF	7.5	1.00	0.01
7 Class	1931-1950 & 1960-1979	CSF	421.9	396.4	504	1.00	8.9					
7 Class	1931-1950 & 1960-1979	UNIDIFF	417.6	391.6	503	1.00	8.9	0.422 (0.685, 0.260)	CSF v UNIDIFF	4.8	1.00	0.03
5 Class	1950-1964 & 1965-1979	CSF	154.3	160.5	160	0.61	5.7					
5 Class	1950-1964 & 1965-1979	UNIDIFF	151.8	158.9	159	0.65	5.7	0.617 (1.061, 0.359)	CSF v UNIDIFF	1.6	1.00	0.21
5 Class	1946-1960 & 1965-1979	CSF	163.6	173.3	160	0.41	6.1					
5 Class	1946-1960 & 1965-1979	UNIDIFF	163.7	169.3	159	0.38	5.8	0.498 (0.828, 0.299)	CSF v UNIDIFF	4.0	1.00	0.05
5 Class	1941-1955 & 1965-1979	CSF	165.3	173.9	160	0.37	5.9					
5 Class	1941-1955 & 1965-1979	UNIDIFF	176.7	164.8	159	0.16	5.7	0.333 (0.556, 0.200)	CSF v UNIDIFF	9.2	1.00	0.00
5 Class	1940-1959 & 1960-1979	CSF	169.1	175.7	160	0.29	5.1					
5 Class	1940-1959 & 1960-1979	UNIDIFF	176.8	171.2	159	0.16	5.1	0.505 (0.805, 0.317)	CSF v UNIDIFF	4.5	1.00	0.03
5 Class	1936-1955 & 1960-1979	CSF	165.9	175.6	160	0.36	5.0					
5 Class	1936-1955 & 1960-1979	UNIDIFF	173.6	169.7	159	0.20	4.9	0.414 (0.678, 0.253)	CSF v UNIDIFF	6.0	1.00	0.01
5 Class	1931-1950 & 1960-1979	CSF	143.0	151.9	160	0.83	4.6					
5 Class	1931-1950 & 1960-1979	UNIDIFF	146.6	148.8	159	0.75	4.6	0.488 (0.869, 0.273)	CSF v UNIDIFF	3.1	1.00	0.08

Note: (1) CSF and UNIDIFF models refer to models 3 and 5, respectively, in Table D1.

Table F14: Fit statistics for log-linear and log-multiplicative models 3 and 4 (**mother-father CSF and father UNIDIFF**) fitted to mobility tables 1991-2009 for various cohort specifications. **Women**.

									Model comparison			
Schema	Cohort	Model	X^2	G^2	df	$P(X^2)$	DI	ϕ (95% CI)	rG^2	rdf	p	
7 Class	1950-1964 & 1965-1979	CSF	455.1	450.8	504	0.94	11.2					
7 Class	1950-1964 & 1965-1979	UNIDIFF	453.9	450.7	503	0.94	11.2	1.088 (1.780, 0.665)	CSF v UNIDIFF	0.1	1.00	0.76
7 Class	1946-1960 & 1965-1979	CSF	479.5	465.5	504	0.78	11.2					
7 Class	1946-1960 & 1965-1979	UNIDIFF	478.3	465.4	503	0.78	11.2	1.064 (1.810, 0.625)	CSF v UNIDIFF	0.0	1.00	0.85
7 Class	1941-1955 & 1965-1979	CSF	507.1	463.0	504	0.45	11.1					
7 Class	1941-1955 & 1965-1979	UNIDIFF	500.8	462.5	503	0.52	11.0	1.364 (2.944, 0.632)	CSF v UNIDIFF	0.5	1.00	0.49
7 Class	1940-1959 & 1960-1979	CSF	486.3	473.7	504	0.71	9.7					
7 Class	1940-1959 & 1960-1979	UNIDIFF	482.7	473.6	503	0.74	9.7	1.113 (1.877, 0.660)	CSF v UNIDIFF	0.1	1.00	0.73
7 Class	1936-1955 & 1960-1979	CSF	493.9	461.3	504	0.62	9.5					
7 Class	1936-1955 & 1960-1979	UNIDIFF	485.6	460.7	503	0.70	9.5	1.354 (2.738, 0.669)	CSF v UNIDIFF	0.7	1.00	0.42
7 Class	1931-1950 & 1960-1979	CSF	421.9	396.4	504	1.00	8.9					
7 Class	1931-1950 & 1960-1979	UNIDIFF	419.3	396.0	503	1.00	8.9	1.309 (3.058, 0.561)	CSF v UNIDIFF	0.4	1.00	0.55
5 Class	1950-1964 & 1965-1979	CSF	154.3	160.5	160	0.61	5.7					
5 Class	1950-1964 & 1965-1979	UNIDIFF	154.2	160.5	159	0.59	5.8	1.056 (1.774, 0.628)	CSF v UNIDIFF	0.0	1.00	0.84
5 Class	1946-1960 & 1965-1979	CSF	163.6	173.3	160	0.41	6.1					
5 Class	1946-1960 & 1965-1979	UNIDIFF	163.9	173.2	159	0.38	6.1	0.931 (1.557, 0.557)	CSF v UNIDIFF	0.1	1.00	0.81
5 Class	1941-1955 & 1965-1979	CSF	165.3	173.9	160	0.37	5.9					
5 Class	1941-1955 & 1965-1979	UNIDIFF	164.0	173.7	159	0.38	6.0	1.191 (2.503, 0.567)	CSF v UNIDIFF	0.2	1.00	0.65
5 Class	1940-1959 & 1960-1979	CSF	169.1	175.7	160	0.29	5.1					
5 Class	1940-1959 & 1960-1979	UNIDIFF	169.3	175.7	159	0.27	5.1	0.987 (1.624, 0.599)	CSF v UNIDIFF	0.0	1.00	0.96
5 Class	1936-1955 & 1960-1979	CSF	165.9	175.6	160	0.36	5.0					
5 Class	1936-1955 & 1960-1979	UNIDIFF	163.8	175.3	159	0.38	5.0	1.225 (2.453, 0.611)	CSF v UNIDIFF	0.3	1.00	0.56
5 Class	1931-1950 & 1960-1979	CSF	143.0	151.9	160	0.83	4.6					
5 Class	1931-1950 & 1960-1979	UNIDIFF	142.9	151.9	159	0.82	4.6	1.021 (2.116, 0.493)	CSF v UNIDIFF	0.0	1.00	0.96

Note: (1) CSF and UNIDIFF models refer to models 3 and 4, respectively, in Table D1.