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Parental Characteristics, Family Structure and Occupational Attainment in Britain

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Abstract

This article uses multivariate logistic regression analyses of the 2005 General Household Survey to assess the impact of parents' occupational and educational characteristics on occupational attainment in Britain, focusing specifically upon the salariat. Differences in outcomes according to family structure are then examined, controlling for such parental characteristics. The results indicate that both parents' characteristics are relevant, and that their effects interact. A smaller chance of a salariat occupation is evident for those who lived in a lone-mother family, lone-father family, or biological-mother stepfamily as a young teenager, reflecting different features of these family types, but consistently reflecting lower educational attainment. Both number of co-resident siblings and parental worklessness affect the odds of having a salariat occupation, this being relevant to family-type comparisons.

Keywords

class mobility; family size; lone-parent families; siblings; social mobility; stepfamilies.

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Introduction

This article examines the dependence of attaining a salariat occupation upon parental occupational and educational characteristics, and the extent to which family structure has an effect on this outcome stretching beyond these characteristics. Analyses of intergenerational class mobility in Britain (e.g. Goldthorpe and Mills, 2008; Kuha and Goldthorpe, 2010) still tend to focus on father's occupation or the 'dominant' parental occupation (Erikson, 1984); however, internationally, authors examining occupational outcomes have highlighted the desirability of taking into account both parents' characteristics (Lampard, 2007a; Schoon, 2008; Beller, 2009; Marks, 2009). Furthermore, studies of occupational outcomes in Britain have examined family-type effects surprisingly rarely, given the prominence of family 'disruption' in numerous educational attainment studies, often using longitudinal data (Kiernan, 1992, 1996, 1997; Ní Bhrolcháin *et al*., 2000; Ely *et al*., 2000; Ermisch and Francesconi, 2001; Ermisch *et al*., 2004; Scott, 2004; Lampard, 2007b; Cusworth, 2009). Beller (2004) noted the scarcity, internationally, of research on social mobility variations according to family type (although see Biblarz and Raftery, 1999); more specifically, she noted an absence of attempts within mobility analyses to integrate both parents plus family type into conceptualizations of family of origin (Beller, 2003).

In Britain, educational attainment studies routinely include both parents' educational levels; internationally, parental education also increasingly features in analyses of occupational outcomes (Korupp *et al*., 2002; Lampard, 2007a; Beller, 2009; Marks, 2009), with mother's education sometimes providing 'class-related' information when maternal occupations are not considered. While some authors (e.g. Biblarz and Raftery, 1999) adopt a 'dominance' approach to parental education, parents' educational levels are more often included separately, even if the 'dominance' approach is applied to parents' occupations (e.g. Gayle *et al*., 2002).

While both parental occupational and educational characteristics have been shown to affect educational and occupational outcomes, there is less clarity regarding what such measures represent conceptually (Lampard, 2007a: 2.10). Schoon (2008: 74) views them as indicators of 'family social position', citing the view that they can act as indicators of socio-economic resources and cultural characteristics (Gershuny, 2000). From a Bourdieusian perspective, they can be viewed as indicators of various forms of capital, and their impact on occupational outcomes interpreted as reflecting the intergenerational transmission of capital.[1] Bourdieu's suggestion that educational systems are increasingly important within this process provides theoretical support for a greater emphasis on the role of parental education in cross-generational examinations of 'social position' (Bourdieu, 1997: 55).

Educational attainment studies have also stretched beyond an emphasis on parental occupational class by focusing on maternal employment and parental or household worklessness (Scott, 2004; Ermisch *et al.*, 2004; Cusworth, 2009). Maternal paid employment impacts upon child outcomes in a complex way: negative effects have been identified, perhaps reflecting reduced maternal involvement, but also positive effects; children's ages and the full-time/part-time distinction appear crucial (Scott, 2004; Cusworth, 2009: 30–1). However, any negative effects may be less marked, or non-existent, in lone-mother families (Kiernan, 1996; Scott, 2004). Cusworth (2009) notes that maternal *non*-employment underpins lone-mother families' higher rate of household worklessness, interpreting negative effects of workless households as reflecting both socio-economic differences and also role model and cultural capital shortfalls (2009: 190), reflecting Bourdieu's influence on her work (2009: 24–7).

Some studies using longitudinal data to document family structure effects focus purely on the impact of experiencing a 'non-intact' family, perhaps qualified by the child's initial age (Ermisch *et al*., 2004), sometimes resulting in a lack of emphasis on stepfamilies (Ermisch and Francesconi, 2001). Authors focusing on the impact of family type *per se*, often with reference to a particular age (e.g. 16: Kiernan, 1996), stress the need to distinguish between female and male lone-parents and step-parents (e.g. Dronkers, 1994); some also distinguish between family change resulting from separation and from death (Kiernan, 1992; Ní Bhrolcháin *et al*., 2000). Sample sizes sometimes lead to lone-father or stepmother families being excluded or not examined separately (Kiernan, 1992; Ní Bhrolcháin *et al*., 2000; Scott, 2004; Cusworth, 2009), despite the conceptual importance of between-family-type variations in 'distance from mother' (Biblarz and Raftery, 1999: 348).

In UK studies examining the impact of family structure on educational attainment, in which the focus is on parental separation or any experience of a lone-parent family, or in which all ‘non-intact’ families are aggregated, controlling for factors like parental education and financial problems often leaves statistically significant net effects (Kiernan, 1997; Ermisch and Francesconi, 2001; Scott, 2004; Ermisch *et al*., 2004). However, in studies focusing on family type at a particular age, the net effects for lone-parent families tend to be small and non-significant (Kiernan, 1992; Ely *et al.*, 2000; Ní Bhrolcháin *et al*., 2000; Cusworth, 2009: 181–6), although net effects for sub-groups are sometimes significant (e.g. Kiernan, 1996). Conversely, such studies typically find significant net stepfamily effects, sometimes contingent upon sex or the stepfamily’s origins (Kiernan, 1992; Ely *et al.*, 2000; Ní Bhrolcháin *et al*., 2000; Lampard, 2007b; Cusworth, 2009: 181; see also Scott 2004). Small samples frequently undermine examinations of lone-father family effects. In the US, Biblarz and Raftery (1999) found the lone-mother family effect on educational and occupational outcomes was removed by controls, but found significant net stepfamily and lone-father effects. Beller (2009) notes that controlling for parental class when comparing lone-parent and two-parent families poses problems, given the lack of information about non-resident parents, and their distinctive status; Scott (2004) handled the parallel problem for parental education by only using information about the parent common to the family types being compared.

Caution is advisable when attributing causal explanations to family disruption effects (Ní Bhrolcháin *et al*., 2000). However, whether viewing them as explanatory mechanisms or controls, authors often consider factors within two broad categories: family economic resources and parental behaviour/involvement. Between-family-type differences in educational outcomes have been attributed wholly to financial resources/disadvantages (Kiernan, 1997), or to parenting or family processes as well (Ely *et al*., 2000; Cusworth, 2009: 75–6). Lawson and Mace (2009) demonstrate that, when mothers are alone, their average ‘parental investment’ is higher, but that in families containing ‘unrelated’ father figures, average investment is lower (for both parents).[2] However, Chan and Koo (2011) suggest parenting styles which affect educational outcomes negatively are disproportionately common among *both* lone-parent families *and* stepfamilies.

Another family structure feature US studies consistently have shown affects attainment is number of siblings (Downey, 1995). While often included as a control, this is rarely a *focus* of UK studies, Iacovou (2001) excepted. Her findings echo van Eijck and de Graaf (1995: 282), who suggest that only children constitute the sole exception to a generally negative effect of increasing family size in Western societies. This may be stronger for boys and in lone-mother families, although the evidence is mixed (see Dronkers, 1994; Downey, 1995; Iacovou, 2001; Gayle *et al*., 2002; Jæger, 2008).

Iacovou (2001) shows much of the sibling effect reflects parental characteristics and behaviour (see also Downey, 1995; Jæger, 2008). Theoretical discussions often foreground 'Resource Dilution Theory', which suggests that number of siblings impacts upon receipt of various parental resources, confirmed for the US by Downey (1995); in Britain, Lawson and Mace (2009) found negative effects of higher family size on parental investment. Number of siblings is crucial for analyses of family-type effects; if lone-mother families contain fewer children, this 'advantage' will weaken any negative lone-mother family effects in studies not controlling for siblings (Biblarz and Raftery, 1999: 328).

In comparing occupational attainment between family types, this article takes account of siblings and also controls, in extensive, methodologically-appropriate ways, for parental occupational, employment and educational characteristics. In addition to drawing upon and complementing the literature on educational attainment in Britain, it adopts and develops approaches to examining 'family of origin' impacts on occupational outcomes advocated and applied in US analyses (Beller 2003, 2009; Biblarz and Raftery 1999). The 2005 General Household Survey (GHS) (ONS, 2007a, 2007b) provides the data analysed. Its sample size is sufficient for 'non-intact' families to be disaggregated, enabling conceptually-meaningful family-type comparisons (Dronkers, 1994), and its parental education measures are more detailed than those used in earlier studies (e.g. Lampard, 2007a); this is crucial, given parental education's importance as a control within family-type comparisons.[3]

Data and measures

In 2005, the GHS collected father's occupation data for the first time since 1992, and also data on mother's occupation and both parents' educational qualifications.[4] Respondents aged 25–65 completed the 'Social Mobility' interview section, hence this range is used here. Occupational attainment is measured using current/last occupation, coded into National Statistics Socio-economic Classification (NS-SEC) analytic classes (ONS, 2005), to which parents' (or step-parents') occupations, when the respondents were young teenagers[5], were also allocated. Respondents also reported how frequently their families had financial problems at that time.

For simplicity, this article uses a binary occupational outcome. The distinction between the 'service class' and other occupations is conceptually relevant (Goldthorpe, 2000: 248–50), gives balanced categories, and is used extensively (e.g. Erikson and Jonsson, 1998; Lampard, 2007a). However, the *term* 'service class' seems inappropriate, given that the Goldthorpe schema is not used; here we use the term 'salariat', as used in various other studies (e.g. Goldthorpe and Mills, 2008; Kuha and Goldthorpe, 2010), sometimes specifically to label NS-SEC Classes 1–2.[6]

While reducing the number of parameters to be estimated and presented, using a binary outcome ignores variations in the impact of parental characteristics and family structure, for example relating to the professional/managerial distinction, which influences occupational attainment processes (Savage *et al.*, 2005; Bühlmann, 2010). However, salariat occupations are consistently advantageous, a key characteristic for this article's purposes.[7]

Parental education measures based on highest qualifications were available (see Table 1).[8] The GHS also provided extensive information about respondents' educational qualifications. An amended highest qualification variable is used here (see Table 1), with categories aggregated to remove unnecessary detail and adjustments to the hierarchical positions of 'vocational' qualifications. Viewed as a determinant of salariat attainment, this measure's construct validity is evident from marked, monotonic variation in attainment across its categories.

The term ‘two-biological-parent families’ (see Biblarz and Raftery, 1999) is used here, in preference to ‘two-parent non-stepfamilies’ (see Lampard, 2007b), although question wording means that the relevant category may include adoptive couples.[9] To allow for non-linear effects (Downey, 1995: 749), a categorical siblings measure is used. Since this relates specifically to co-resident siblings when the respondents were young teenagers, it is partly determined by birth order and spacing; these influence the period spent sharing resources (Jæger, 2008), and can affect educational outcomes (van Eijck and de Graaf, 1995; Iacovou, 2001).

GHS2005 interviewed 11,955 respondents aged 25–65. For some, necessary data were unavailable, hence 722 (6.0%) were omitted[10], leaving a sample of 11,233. Table 1 shows key variables’ frequency distributions. Within this final sample, insufficient information regarding the father’s occupation was provided in 913 cases (8.1%), the same applying to mothers in 562 cases (5.0%). These cases were retained to avoid distorting the sample, although retaining them reduced the explanatory power of parental class, since potentially diverse occupations are lumped together.[11]

All the logistic regression analyses control for gender differences in salariat attainment; age and the age-sex interaction were also included, to control for gendered age and cohort effects (see Bühlmann 2010). Using age categories allowed for non-linearity; five-year bands significantly increased model fit relative to less detail.

[Tables 1–2]

Findings

Parents' occupational and educational characteristics and salariat attainment

The pseudo-r^2 values and model fit (-2 Log Likelihood) values for Models 1–4 (Table 2) indicate that the impacts on attaining a salariat occupation of parental class and parental education only partially overlap, both making substantial independent contributions. However, the results for Models 5–8 indicate that, for both parental class and parental education, the vast majority of the impact operates indirectly via educational attainment, although there are small, statistically significant, direct effects. Removing the class or educational level of *either* parent from Model 4 significantly reduces the model fit.[12]

More than two-fifths of the overall improvement in model fit relating directly or indirectly to parental characteristics can be viewed as corresponding to a path from parental education to salariat attainment via both parental class and educational attainment. Furthermore, nearly three-quarters of the improvement both relates to parental education and also operates via educational attainment.[13] The impact of parental educational and occupational characteristics on attaining a salariat occupation can thus very largely be viewed as reflecting 'educational mobility', or, more precisely, educational inheritance (Lampard, 2007a: 1.2). The relatively small difference in fit between Models 5 and 8 also highlights the crucial role of educational attainment in transmitting 'social position', consistent with the idea that the transmission process operates primarily via qualifications acting as 'institutionalized cultural capital' (Bourdieu, 1997: 47-55).

Thus, like Lampard (2007a), this article indicates a key role for parental education in the intergenerational transmission of social position. However, the small difference in fit between Models 6 and 8 (see Table 2) indicates a relatively small effect specific to parental *class*; both this effect and the overall impact of parental class on salariat attainment are markedly less substantial than in an earlier study (Lampard, 2007a: Table 2). This may partly reflect the detailed educational measures used here, but the results may also under-estimate the explanatory relevance of parental class, other studies having shown a substantial direct impact of class origin on class destination, albeit varying according to the class in question and sex (Kuha and Goldthorpe, 2010).

Part of the discrepancy with Lampard (2007a) reflects the sub-sample with parental occupational information that could not be coded.[14] Some of the remainder may reflect a weaker intergenerational relationship for NS-SEC than for Goldthorpe class.[15] Disaggregation into NS-SEC sub-categories (Rose *et al*., 2005: 51) only removes part of the discrepancy, suggesting the *detailed* allocation of occupations within NS-SEC may adversely affect class mobility analyses.[16]

Extensive checks were made for interaction effects. The first identified involves father's class and child's sex. Specifically, Class 4 fathers were found to have a significantly greater positive impact on salariat attainment for women than for men, possibly reflecting more sons inheriting their fathers' Class 4 positions (see Erikson and Jonsson, 1998; Lampard, 2007a; Breen *et al*., 2010). Additionally, Class 1 fathers' positive effect appeared *less* strong for women (at a borderline significance level: p=0.063).

Two further interaction effects are more specifically relevant. Including separate variables for parents' classes (or educational levels) assumes additive effects rather than any interaction (Korupp *et al*., 2002; Beller, 2004), contrary to the 'dominance' approach to family class measurement (Erikson, 1984). Variables were thus added to the model to check whether the impact of parental class (or education) depended more on the 'higher' class (or educational level). These interaction terms were statistically significant, and could be represented parsimoniously by focusing on individuals with both parents in Classes 1–3 and individuals whose parents both had (any) qualifications. For the former, the second parent's class had a weaker effect; however, its additional impact was still substantial[17], hence the 'dominance' approach would have been inappropriate (see also Lampard 2007a: 4.18). While the additional impact of a second parent with qualifications was weakened to a markedly greater extent by the interaction term, including it alongside both parental education variables nevertheless provided a substantially better model fit than the 'dominance' approach.

Thus, the preferred model here for the impact of parental characteristics on salariat attainment is Model 9 (see Tables 2 and 4), which includes interactions corresponding to the effects of father's class and child's sex, the parents' classes, and the parents' educational levels.

[Tables 3–4]

Family type and salariat attainment: Two-parent families

Table 3 shows the effects of family type on salariat attainment when family type is added to a model controlling only for sex, age and their interaction. Having lived as a young teenager in any other context than a two-biological-parent family is associated with lower odds of salariat attainment. With one exception, the comparisons are statistically significant; the exception involves biological-father/stepmother families, a small category with an effect similar in magnitude to other stepfamilies and lone-parent families. The reductions in odds for lone-parent families and stepfamilies are in the range 19%–33%, with more substantial reductions (>42%) for 312 individuals (2.8%) not living with either biological parent, who are henceforth omitted.

From now on, this section focuses upon a 'two-parent sub-sample': individuals who lived as young teenagers either with both biological parents or with one biological parent and their partner/spouse. In theory, equivalent parental information is available for these individuals, although the *impacts* of step-parents' and biological parents' characteristics are not necessarily equivalent (Beller, 2004: 9). To examine the effect of family type controlling for parental classes and educational levels, Model T1 (i.e. Model 9, plus a variable distinguishing between the three family types) was applied to the two-parent sub-sample.[18] In this model, the impact corresponding to biological-mother/stepfather families diminished substantially (compared to Table 3), becoming statistically non-significant (see Table 5). The effect for biological-father stepfamilies increased in magnitude, remaining non-significant.

A step-by-step assessment of the consequences of controlling for parental classes and educational levels showed the reduction in the biological-mother stepfamily effect was attributable to (step-)father’s education, and, more specifically, to the markedly higher proportion of individuals living in such families as a young teenager who were in the ‘Don’t know’ category. In other words, much of the initial negative effect appears to relate to children’s lack of knowledge about their mothers’ partners’ qualifications, perhaps indicating those partners’ lack of involvement with or investment in the children’s education, or perhaps their lack of involvement more generally, although these suggestions are quite speculative.

Taking account of how many siblings respondents lived with as young teenagers indicated that three or more siblings, and especially five or more, reduced the odds of salariat attainment significantly. Including this effect in the model substantially reduced the negative effect for biological-father stepfamilies, which tend to contain more siblings. A significant interaction was found between the sibling effect and sex, the negative effect of three-plus siblings being less marked for women, echoing other studies (Jæger 2008: 221; see also Iacovou 2001: 44).[19]

If added to Model 1 (see Table 2), the frequency of financial problems had a significant negative impact upon salariat attainment, but there was minimal evidence of an impact net of other family and parental characteristics.[20] Thus the sibling effect identified, controlling for parents’ educational and occupational characteristics, may reflect ‘resource dilution’ but does not operate via (perceived) frequency of financial problems.

Adding educational attainment showed that the effect of a large number of siblings only partially operates via this route (see Models T2–T3); conversely, adding educational attainment removed the remaining negative effect for biological-mother stepfamilies.

[Table 5]

Family type and salariat attainment: Lone-parent families

Comparing lone-parent and two-parent families poses practical and conceptual problems regarding controlling for parental characteristics (Beller, 2004), since the number and sex of the (resident) parent(s) for whom data are available vary with family type. To circumvent such problems, we first focus solely on families containing ‘mothers’ (including stepmothers/female partners), and then on families containing ‘fathers’, including in the models in each analysis only the class and education variables corresponding to the sex of parent present in all families.

An alternative would be to construct a single parental class variable and a single parental education variable using all the parental information. However, household or couple-based measures can blur the distinction between the effects of class or education and of family type. In an earlier study, controlling for the class of the ‘Household Reference Person’ (HRP) accounted for some of the difference in educational attainment between lone-mother families and ‘married non-stepfamilies’ (Lampard, 2007b: 45–6). However, a problem with this finding is that occupational segregation by sex can induce a greater class difference between the HRPs of different family types than is obtained by comparing mothers’ occupations (2007b: 44). Thus the relative class position of lone-mother families may be depressed artefactually, leading to some of the difference in attainment between family types being attributed to class rather than family type. Similarly, the ‘dominance’ approach (Erikson, 1984) leads to greater differences between the class or educational profiles of different family types than examining parents of a specified sex does, since the ‘second’ parents’ characteristics raise the classes or educational levels of (some) two-parent families, while those of lone-parent families are unaffected.[21]

Of course, a parent’s departure may genuinely ‘change’ a family’s class position. Note, then, that the approach taken here controls for the *broad* positioning of families in the class structure, but retains within the family-type effect any indirect effect of family disruption on salariat attainment operating via its impact upon family class position.

To contextualise the following sex-specific lone-parent analyses, note that adding educational attainment to the 'baseline' model, containing family type and age, sex and their interaction, results in odds ratios of very close to one for the comparisons of each form of lone-parent family with two-biological-parent families. The overall differences in salariat attainment between lone-parent and two-biological-parent families thus appear to reflect educational attainment (see also Table 6: Models M3/F3).

The remainder of this section consists of narrative accounts of the sequential addition of various independent variables within the two sex-specific analyses, starting in each case with the 'baseline' model and ending with one containing all the independent variables, except educational attainment (Models M2/F2: see Table 6). Consequently, quoted parameter estimates sometimes come from 'intermediate' models, not included in the tables. The concluding paragraph highlights a key difference between the lone-mother and lone-father analyses.

Restricting attention to families containing a mother (or female partner), adding mother's class and mother's education to the baseline model only slightly reduced the difference between lone-mother and two-biological-parent families, compared to Table 3 (the odds ratio rising to 0.82, $p=0.007$). However, the difference might reflect lone-mother families experiencing more financial problems. Further controlling for the frequency of household financial problems experienced as a young teenager significantly improved the model fit.[22] The difference in salariat attainment between lone-mother and two-biological-parent families was substantially reduced, leaving an odds ratio of 0.88, still substantively interesting but statistically non-significant ($p>0.05$).

A check for interaction effects identified a substantially and significantly greater negative impact of non-employment for lone-mothers, echoing other studies (Kiernan, 1996; Biblarz and Raftery, 1999), and suggesting that controlling for parental worklessness might be pertinent. Indeed, it significantly improved the model fit, with no interaction evident between the parental worklessness and family-type effects. The inclusion of parental worklessness rendered the difference in salariat attainment between lone-mother and two-biological-parent families negligible (an odds ratio of

0.97). There was surprisingly little overlap between the financial problems and parental worklessness effects; parental worklessness appearing to affect salariat attainment irrespective of financial problems.

However, when number of siblings was added to the model, the financial problems measure's effect decreased, suggesting that large families do not simply induce financial problems. The larger difference in attainment between lone-mother and two-biological-parent families in Model M1 (see Table 6), compared to Table 3, reflects the inclusion of number of siblings, since fewer lone-mother families included three-plus siblings, highlighting that sibship size is a crucial control (Biblarz and Raftery, 1999: 337–44). No interaction between the sibling effect and family type was apparent. Unlike the financial problems and parental worklessness effects, the siblings effect remained statistically significant net of educational attainment (see Table 6: Models M2–M3).

[Table 6]

We now compare lone-father and two-parent families, restricting the analysis to families containing a father (or male partner). The results partly mirror those for lone-mother families: adding father's class and father's education to the baseline model only accounted for a small proportion of the difference between lone-father and two-biological-parent families in Table 3 (the odds ratio rising slightly to 0.71, p=0.030). However, the impact of controlling for other factors on the difference between lone-parent and two-biological-parent families varies between the lone-mother and lone-father analyses. In the latter, when the financial problems measure was added, only the comparison between frequent problems and no problems was significant ($p<0.05$); consequently the difference in salariat attainment between lone-father and two-biological-parent families was only slightly reduced, and stayed significant ($p<0.05$). Adding parental worklessness did not improve the model fit significantly.[23] When number of siblings was added, the financial problems measure became statistical non-significant, and the difference between the odds of salariat attainment for lone-father and two-biological-parent families increased, since fewer lone-father families included three-plus siblings. Overall, the other factors in Model F2 (see Table 6) explained little of the difference between lone-father and two-biological-parent families.

Table 6 facilitates comparisons between the lone-mother and lone-father family findings. Models M1 and F1 include the age, sex and age-sex interaction controls, together with family type and number of siblings. The odds ratios comparing lone-parent and two-biological-parent families are 0.79 (p=0.001) for lone-mother and 0.64 (p=0.004) for lone-father families. However, the corresponding values from Models M2 and F2, which also include mothers' or fathers' characteristics, the financial problems measure and parental worklessness, are 0.93 (p=0.360) and 0.70 (p=0.025). Thus, factors accounting for most of the difference between the odds of salariat attainment for lone-*mother* and two-biological-parent families account for little of the (larger) difference between lone-*father* and two-biological-parent families.

Conclusion

This article demonstrates that someone's likelihood of obtaining a salariat occupation depends on the occupational and educational characteristics of *both* their parents, with the effects of the two parents' characteristics interacting. These effects largely operate via educational attainment; viewed in combination with the important role of parental education, this suggests that class mobility is best viewed as one facet of a broader intergenerational transmission of social position.

The lower likelihood of salariat attainment for various forms of 'non-intact' family is also shown here to vary in origin, since the impact of the inclusion of some specific independent variables on this deficit varies according to family type. Among lone-mother families, the findings suggest that the shortfall reflects financial problems and non-employment (see Kiernan, 1996), whereas among biological-mother stepfamilies and lone-father families one might speculate that it relates to reduced parental involvement (see Lawson and Mace, 2009) and mother absence (see Biblarz and Raftery, 1999) respectively. However, for all these three family types, lower educational attainment is the vehicle of their negative impact; something about them thus seems either to reduce the amount of capital that can be converted into institutionalized cultural/educational capital (Bourdieu, 1997), or to make the process of ‘transforming’ it less effective. Nevertheless, broad generalisations about the negative impact of ‘non-intact’ families seem unjustified; such an effect may not exist, for example, among working-lone-mother families with no financial problems. Concerns about the impact of family structure *per se* may be more justifiable with reference to ‘large’ families; these appear related to marked reductions in salariat attainment (see Downey, 1995), particularly for men, even when educational attainment and the other factors considered here are taken into account.

In addition to extending the previously limited research evidence relating to family-type effects on occupational outcomes in Britain, this article also paints a more detailed picture of differences between family types than most of the recent literature on educational attainment in Britain, resonating with the US findings of Biblarz and Raftery (1999), but controlling for parental characteristics in a more satisfactory, sophisticated way. However, social change can reduce the salience of results based on past behaviour (Kiernan, 1996)[24]; fortunately, unlike Biblarz and Raftery (1999: 348–9), we found little evidence of any changes in *effects* over time[25].

The findings’ implications for policy debates are highly dependent on one’s causal assumptions (Ní Bhrolcháin *et al*., 2000: 67–8). More specifically, the findings may seem to resonate with a widely-held positive view of lone mothers taking up paid work (Biblarz and Raftery, 1999: 353; Cusworth, 2009: 194–5). However, employment may not, in itself, enable lone mothers to act as effective role models or to increase their children’s educational capital, and may instead reduce levels of parental contact or involvement.[26]

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Endnotes

[1] Bourdieu (1997: 51–4) suggests the capital transmitted is not necessarily reducible to an occupational position during this process, providing an argument for examining educational and occupational outcomes in parallel (Biblarz and Raftery, 1999), or in a 'composite' way (Schoon, 2008).

[2] For related ideas, see Biblarz and Raftery (1999: 356).

[3] Longitudinal studies have some advantages compared to the cross-sectional GHS2005, but may not, as in this instance, provide the desired sample size or measures (e.g. Ermisch and Francesconi, 2001: 137).

[4] This reflects the UK's obligation to collect data for a cross-European survey (EU-SILC).

[5] GHS definition: 'Between the ages of 12 and 16'.

[6] Similar occupational aggregations are sometimes labelled 'middle class' or 'professional' (e.g. Cabinet Office, 2009; McKay, 2010).

[7] A sensitivity analysis was carried out to assess the implications of using a binary, salariat-focused outcome for the salience of this article's findings to occupational attainment more generally, given the relevance of heterogeneity within the outcome's categories. Neither binary logistic regressions focusing on different outcomes within and outside the salariat nor an ordinal logistic regression analysis splitting the seven NS-SEC classes into five levels (by combining Classes 3–5) indicated that any of the findings are misleading or unduly specific to salariat attainment.

[8] Some detail was empirically irrelevant: doctorates were aggregated with other degrees; CSEs/YT certificates with other, basic qualifications.

[9] It is unclear which category/categories include(s) same-sex couples.

[10] Reasons: data needed for NS-SEC not collected (92); never worked/long-term unemployed (331); unclassifiable occupations (178); unavailable qualifications data (5); 'Social Mobility' section not completed (28); family type not reported (24); qualifications information or occupational information for a parent/step-parent not provided (9 and 55).

[11] In the multivariate analyses, the modal number of siblings category (1–2) includes 5 cases not reporting a number.

[12] So does aggregating, within any of these four variables, the categories containing parents with classifiable occupations or qualifications ($p<0.001$ for each variable).

[13] More precise proportions are (((2-1)-(4-3))-((6-5)-(8-7)))/(4-1) = 0.426 and ((2-1)-(6-5))/(4-1) = 0.747; the numbers within the LHS of these equations represent the Model Chi-square values for the corresponding models (see Table 2). Lampard (2007a) further documents this approach.

[14] Also, given the substantial proportion of unclassified cases, GHS2005 may have less reliable coding than the BHPS.

[15] Being employment relations-based (Rose *et al.*, 2005), NS-SEC may be a less effective indicator of occupational features central to class reproduction; Goldthorpe and Mills (2008) found a discontinuity between mobility levels in NS-SEC-based analyses of GHS2005 and Goldthorpe schema-based analyses of older GHS data.

[16] Note the discrepancy partly relates to categories other than the NS-SEC analytic classes.

[17] The 'second' parent's main effect outweighs the negative interaction effect.

[18] The hypothesis that parental characteristics have weaker effects within stepfamilies (Biblarz and Raftery, 1999: 350–7) was not supported.

[19] The sibling effect/family type interaction was statistically non-significant.

[20] For consistency, parental worklessness was added to the model, as an effect *additional to* the (Model 9) non-working mother/father parameters.

[21] Applying the 'dominance' approach here, the lone-mother family effect switched to being significantly *positive* ($p<0.05$), reflecting the implicit 'rise' in class/educational levels for many two-parent families.

[22] The measure's effect was consistent across family types, suggesting it was proxying for fathers' occupational and/or educational characteristics within two-parent families, given its non-significance in the two-parent-family analysis.

[23] This reflects the broad similarity of *all* scenarios involving a non-working father.

[24] This article's other limitations include the absence of measures of parental involvement and of ability, aspirations, attitudes and agency (Scott, 2004; Schoon, 2008), and of longitudinal data facilitating event-history analyses and analyses accounting for unobserved family heterogeneity (e.g. Ermisch *et al.*, 2004).

[25] The sole exception to this was the (negative) lone-parent family effect appearing stronger for those under 30, possibly an age effect.

[26] Furthermore, policy-makers need to 'factor in' working hours, adequacy of pay, and children's ages.

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Table 1: Frequencies for the dependent and independent variables

Occupational class	**n = 11,233**	**Parent(s) workless**	**n = 10,921**
Salariat (NS-SEC 1–2)	4,502 (40.1%)	Yes	491 (4.5%)
Other class (NS-SEC 3–7)	6,731 (59.9%)	No (or unclear)	10,430 (95.5%)
Family type	**n = 11,233**	**Financial problems**	**n = 10,921**
Two-biological-parent	9,347 (83.2%)	Never	3,423 (31.3%)
Lone mother	983 (8.8%)	Rarely	1,884 (17.3%)
Lone father	203 (1.8%)	Occasionally	2,201 (20.2%)
Mother/stepfather	314 (2.8%)	Often	1,207 (11.1%)
Father/stepmother	74 (0.7%)	Most of time	1,160 (10.6%)
Other relatives	179 (1.6%)	Don't know	1,046 (9.6%)
Other household	59 (0.5%)	**Mother's highest qual.**	**n = 10,718**
Institution	74 (0.7%)	Degree	871 (8.1%)
No. of siblings	**n = 11,228**	'A' Level/ONC	306 (2.9%)
0	1,437 (12.8%)	BTEC or equivalent	354 (3.3%)
1–2	6,612 (58.9%)	'O' level	853 (8.0%)
3–4	2,171 (19.3%)	GCSE	76 (0.7%)
5+	1,008 (9.0%)	City & Guilds	146 (1.4%)
Mother's class	**n = 10,718**	Other	696 (6.5%)
Not classified	562 (5.2%)	None	5,606 (52.3%)
Not working	4,786 (44.7%)	Don't know	1,810 (16.9%)
NS-SEC 1	88 (0.8%)	**Father's highest qual.**	**n = 9,938**
NS-SEC 2	921 (8.6%)	Degree	1,073 (10.8%)
NS-SEC 3	948 (8.8%)	'A' Level/ONC	266 (2.7%)
NS-SEC 4	389 (3.6%)	BTEC or equivalent	1,362 (13.7%)
NS-SEC 5	299 (2.8%)	'O' level	399 (4.0%)
NS-SEC 6	1,392 (13.0%)	GCSE	42 (0.4%)
NS-SEC 7	1,286 (12.0%)	City & Guilds	312 (3.1%)
Don't know	47 (0.4%)	Other	583 (5.9%)
Father's class	**n = 9,938**	None	4,120 (41.5%)
Not classified	913 (9.2%)	Don't know	1,781 (17.9%)
Not working	289 (2.9%)	**Highest qualification**	**n = 11,233**
NS-SEC 1	1,022 (10.3%)	Higher degree	777 (6.9%)
NS-SEC 2	1,519 (15.3%)	First degree	1,494 (13.3%)
NS-SEC 3	469 (4.7%)	Teaching/Nursing	491 (4.4%)
NS-SEC 4	1,110 (11.2%)	Above 'A' level	777 (6.9%)
NS-SEC 5	1,599 (16.1%)	'A' level	1,058 (9.4%)
NS-SEC 6	1,054 (10.6%)	'O' level (or above)	2,650 (23.6%)
NS-SEC 7	1,914 (19.3%)	Other	1,845 (16.4%)
Don't know	49 (0.5%)	None	2,141 (19.1%)

Table 2: Goodness-of-fit of logistic regression models of salariat attainment: the impact of parents' occupational classes and educational levels (n=11,233)

Model	Terms included in the model	-2 Log Likelihood	Model Chi-square	d.f.	Pseudo r^2
Model 1	Base (Age, Sex, Age by Sex)	14,944.0	183.0	15	0.016
Model 2	+ PEDUC	14,011.5	1,115.5	33	0.095
Model 3	+ PCLASS	14,213.2	913.8	35	0.078
Model 4	+ PEDUC + PCLASS	13,791.8	1,335.2	51	0.112
Model 5	+ EDUC	11,371.5	3,755.4	22	0.284
Model 6	+ PEDUC + EDUC	11,299.3	3,827.7	40	0.289
Model 7	+ PCLASS + EDUC	11,300.6	3,826.4	42	0.289
Model 8	+ PEDUC + PCLASS + EDUC	11,248.3	3,878.7	58	0.292
Model 9	Model 4 + Interaction terms	13,750.4	1,376.6	55	0.115

Notes:

- PEDUC = Parents' qualification variables; PCLASS = Parents' class variables; EDUC = Respondent's qualifications; d.f. = Degrees of freedom.
- The Model 9 interactions are between the effects of father's class and child's sex, between the parents' classes, and between the parents' qualifications. (See Table 4 and the text for more details).
- $p<0.001$ for the Model Chi-square of each of the models.
- For a discussion of the pseudo-r^2 measure, see Cox, D.R. and Snell, E.J. (1989). *Analysis of binary data* (2nd edition). London: Chapman and Hall.

Table 3: **Results (odds ratios) corresponding to differences in salariat attainment between family types when family type is added to Model 1 (n=11,233)**

Family type		**(0.000)**
Two-biological-parent	1.00	
Lone mother	0.81	(0.003)
Lone father	0.67	(0.008)
Mother/stepfather	0.77	(0.033)
Father/stepmother	0.72	(0.182)
Other relatives	0.56	(0.001)
Other household	0.44	(0.007)
Institution	0.58	(0.036)

Notes:

- To save space, the parameter estimates for the control variables (age, sex and their interaction) are not presented in Tables 3 to 6, but are available from the author.
- In Tables 3 to 6, the figures in parentheses are *p*-values. A *p*-value in bold relates to the overall statistical significance of the variable in question.
- -2 Log Likelihood = 14,900.8; Pseudo-r^2 = 0.020; Model Chi-square = 226.2 (22 d.f.; p=0.000); Change in Model Chi-square = 43.2 (7 d.f.; p=0.000).

Table 4: Results (odds ratios) from the preferred model (Model 9) relating salariat attainment to parents' classes and parents' educational levels (n=11,233)

Mother's class		**(0.000)**	**Mother's h. qual.**		**(0.000)**
Not classified	1.66	(0.000)	Degree	2.65	(0.000)
Not working	1.19	(0.017)	'A' Level/ONC	2.42	(0.000)
NS-SEC 1	2.37	(0.001)	BTEC or equiv.	1.81	(0.000)
NS-SEC 2	1.61	(0.000)	'O' level	2.45	(0.000)
NS-SEC 3	1.96	(0.000)	GCSE	2.12	(0.003)
NS-SEC 4	1.60	(0.000)	City & Guilds	2.23	(0.000)
NS-SEC 5	1.40	(0.016)	Other	1.70	(0.000)
NS-SEC 6	1.22	(0.024)	No qualifications	1.00	
NS-SEC 7	1.00		Don't know	0.90	(0.160)
Don't know	1.02	(0.961)	No resident mother	1.06	(0.643)
Father's class		**(0.000)**	**Father's h. qual.**		**(0.000)**
Not classified	1.47	(0.001)	Degree	1.95	(0.000)
Not working	1.14	(0.365)	'A' Level/ONC	1.38	(0.028)
NS-SEC 1	2.99	(0.000)	BTEC or equiv.	1.37	(0.000)
NS-SEC 2	2.39	(0.000)	'O' level	1.53	(0.001)
NS-SEC 3	1.85	(0.000)	GCSE	0.70	(0.303)
NS-SEC 4	1.18	(0.143)	City & Guilds	1.26	(0.076)
NS-SEC 5	1.50	(0.000)	Other	1.64	(0.000)
NS-SEC 6	1.14	(0.139)	No qualifications	1.00	
NS-SEC 7	1.00		Don't know	0.70	(0.000)
Don't know	1.07	(0.862)	No resident father	1.31	(0.002)
2 parents in 1-3	0.72	**(0.003)**	**2 pars. with quals.**	0.73	**(0.001)**
Gender/class int.		**(0.000)**			
Woman/Father in 1	0.77	(0.063)			
Woman/Father in 4	1.69	(0.000)			

Notes:

- -2 Log Likelihood = 13,750.4; Pseudo-r^2 = 0.115; Model Chi-square = 1376.6 (55 d.f.; p=0.000).

Table 5: Results (odds ratios) from extensions of Model 9, comparing the salariat attainment of different types of two-parent families (n=9,735)

	Model T1		**Model T2**		**Model T3**	
Family type		**(0.156)**		**(0.325)**		**(0.462)**
Two-biological-parent	1.00		1.00		1.00	
Mother/Stepfather	0.87	(0.291)	0.89	(0.345)	1.00	(0.979)
Father/Stepmother	0.65	(0.104)	0.73	(0.239)	0.69	(0.214)
No. of siblings				**(0.000)**		**(0.003)**
0			1.00		1.00	
1–2			1.02	(0.842)	1.11	(0.196)
3–4			0.65	(0.000)	0.88	(0.244)
5+			0.47	(0.000)	0.74	(0.025)
Woman, 3+ sibs.			1.27	**(0.020)**	1.17	**(0.168)**
Financial problems				**(0.206)**		**(0.036)**
Never			1.00		1.00	
Rarely			1.00	(0.954)	1.00	(1.000)
Occasionally			1.10	(0.138)	1.14	(0.061)
Often			1.13	(0.121)	1.31	(0.004)
Most of time			0.99	(0.884)	1.17	(0.118)
Don't know			0.92	(0.302)	1.03	(0.723)
Workless parents			1.26	**(0.414)**	1.23	**(0.515)**
Highest quals.						**(0.000)**
Higher degree					1.00	
First degree					0.55	(0.000)
Teach./Nurs.					0.49	(0.000)
Above 'A' level					0.23	(0.000)
'A' level					0.13	(0.000)
'O' level (plus)					0.07	(0.000)
Other					0.04	(0.000)
None					0.02	(0.000)

Notes:

- To save space, the parameter estimates for the variables carried forward from Model 9 are not presented, but are available from the author.
- -2 Log Likelihood (Model T3) = 9,790.0; Pseudo-r^2 (Model T3) = 0.294; Changes in Model Chi-square: Model 9 to T1 = 3.8 (2 d.f.; p=0.150); T1 to T2 = 86.1 (10 d.f.; p=0.000); T2 to T3 = 2,094.3 (7 d.f.; p=0.000).

<u>Table 6</u>: Results (odds ratios) from models comparing the salariat attainment of lone mother/lone father families and two-parent families (n=10,718 {mothers}; n=9,938 {fathers})

Parent/Model	Mother/**M1**		Mother/**M2**		Mother/**M3**		Father/**F1**		Father/**F2**		Father/**F3**	
Family type		**(0.003)**		**(0.222)**		**(0.730)**		**(0.006)**		**(0.104)**		**(0.704)**
Two-bio.-parent	1.00		1.00		1.00		1.00		1.00		1.00	
Lone parent	0.79	(0.001)	0.93	(0.363)	0.98	(0.861)	0.64	(0.004)	0.70	(0.024)	0.96	(0.807)
Moth./Stepfath.	0.78	(0.043)	0.81	(0.086)	0.96	(0.791)	0.78	(0.043)	0.93	(0.555)	1.04	(0.803)
Fath./Stepmoth.	0.85	(0.501)	0.78	(0.341)	0.72	(0.270)	0.85	(0.519)	0.79	(0.360)	0.72	(0.258)
No. of sibs.		**(0.000)**		**(0.000)**		**(0.000)**		**(0.000)**		**(0.000)**		**(0.000)**
0	1.00		1.00		1.00		1.00		1.00		1.00	
1–2	1.03	(0.678)	1.00	(0.960)	1.09	(0.241)	1.06	(0.423)	1.03	(0.658)	1.11	(0.206)
3–4	0.58	(0.000)	0.62	(0.000)	0.84	(0.103)	0.58	(0.000)	0.62	(0.000)	0.83	(0.098)
5+	0.38	(0.000)	0.44	(0.000)	0.69	(0.003)	0.39	(0.000)	0.44	(0.000)	0.68	(0.004)
Woman, 3+ sibs	1.23	**(0.023)**	1.27	**(0.012)**	1.23	**(0.057)**	1.23	**(0.034)**	1.30	**(0.008)**	1.21	**(0.092)**
Parent's class				**(0.000)**		**(0.012)**				**(0.000)**		**(0.000)**
Not classified			1.69	(0.000)	1.41	(0.007)			1.78	(0.000)	1.31	(0.008)
Not working			1.54	(0.000)	1.14	(0.106)			1.17	(0.479)	0.93	(0.762)
NS-SEC 1			2.52	(0.000)	1.68	(0.064)			2.74	(0.000)	1.49	(0.000)
NS-SEC 2			1.62	(0.000)	1.22	(0.090)			2.50	(0.000)	1.59	(0.000)
NS-SEC 3			1.93	(0.000)	1.35	(0.005)			1.86	(0.000)	1.37	(0.014)
NS-SEC 4			1.92	(0.000)	1.48	(0.007)			1.81	(0.000)	1.29	(0.008)
NS-SEC 5			1.47	(0.006)	1.42	(0.025)			1.56	(0.000)	1.33	(0.001)
NS-SEC 6			1.28	(0.005)	1.16	(0.132)			1.16	(0.098)	1.06	(0.575)
NS-SEC 7			1.00		1.00				1.00		1.00	
Don't know			0.94	(0.867)	0.57	(0.172)			1.00	(0.994)	0.79	(0.580)
Parent's educ.				**(0.000)**		**(0.000)**				**(0.000)**		**(0.000)**
Degree			3.23	(0.000)	1.21	(0.074)			2.35	(0.000)	1.07	(0.524)
'A' Level/ONC			2.82	(0.000)	1.12	(0.423)			1.61	(0.000)	0.82	(0.205)
BTEC or equiv.			1.78	(0.000)	1.10	(0.488)			1.37	(0.000)	1.06	(0.444)
'O' level			2.54	(0.000)	1.41	(0.000)			1.82	(0.000)	1.11	(0.412)
GCSE			1.90	(0.007)	1.33	(0.307)			0.75	(0.392)	0.47	(0.045)
City & Guilds			2.22	(0.000)	1.35	(0.131)			1.29	(0.037)	0.87	(0.331)
Other			1.76	(0.000)	1.00	(0.992)			1.62	(0.000)	1.10	(0.362)
No qualifications			1.00		1.00				1.00		1.00	
Don't know			0.79	(0.000)	0.85	(0.023)			0.66	(0.000)	0.74	(0.000)
Workless p(s)			0.72	**(0.008)**	0.88	**(0.355)**			1.15	**(0.610)**	1.16	**(0.643)**
Financ. probs				**(0.068)**		**(0.208)**				**(0.091)**		**(0.027)**
Never			1.00		1.00				1.00		1.00	
Rarely			0.97	(0.658)	1.04	(0.584)			0.98	(0.779)	1.00	(0.972)
Occasionally			0.96	(0.536)	1.08	(0.250)			1.09	(0.156)	1.16	(0.041)
Often			0.94	(0.407)	1.22	(0.019)			1.12	(0.142)	1.30	(0.004)
Most of time			0.85	(0.046)	1.15	(0.141)			0.96	(0.640)	1.17	(0.115)
Don't know			0.80	(0.006)	0.99	(0.910)			0.88	(0.119)	1.02	(0.822)
Highest quals.						**(0.000)**						**(0.000)**
Higher degree					1.00						1.00	
First degree					0.58	(0.000)					0.54	(0.000)
Teach./Nurs.					0.46	(0.000)					0.48	(0.000)
Above 'A' level					0.24	(0.000)					0.23	(0.000)
'A' level					0.13	(0.000)					0.13	(0.000)
'O' level (plus)					0.07	(0.000)					0.07	(0.000)
Other					0.04	(0.000)					0.04	(0.000)
None					0.02	(0.000)					0.02	(0.000)
-2 Log L'hood	14,085.8		13,389.0		10,817.6		13,063.0		12,303.1		10,031.0	
Pseudo-r^2	0.036		0.097		0.289		0.037		0.108		0.290	

[Blank]