Title: Intergenerational social mobility predicts midlife well-being: prospective evidence from two large British cohorts

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Abstract

Rationale: It is often assumed that experiencing an upward shift in social position from one generation to the next will bring happiness, yet empirical evidence for this is limited.

Objective: We provide a large-scale test of the relationship between intergenerational mobility and midlife life satisfaction using data from two prospective UK studies (N = 20,948).

Method: Intergenerational mobility was modelled as a formative construct gauging the extent to which individuals moved up or down the social hierarchy compared to their parents, on a continuum ranging from high levels of downward mobility to high levels of upward mobility.

Results: An intergenerational increase in social mobility, captured by greater educational attainment, social status, and home size than one's parents was positively associated with life satisfaction at age 42 in both cohorts. Mediation analyses revealed that almost half of this relationship was explained by better self-reported health and fewer perceived financial difficulties amongst the upwardly mobile.

Conclusion: This study provides evidence that enhanced satisfaction with life may be a key outcome of intergenerational increases in social status.

Keywords: Intergenerational social mobility; life satisfaction; well-being; self-rated health

Introduction

Intergenerational social mobility refers to the extent to which a person's living standards and/or social position has improved or deteriorated relative to that of their parents. Social mobility between generations is considered a core societal goal in part because of the well-being gains thought to be experienced by the upwardly mobile. This assumption underpins concerns that rates of intergenerational mobility are slowing in some Western countries including the US (Aaronson & Mazumder, 2008) and has even motivated policies to increase mobility in others (UK Government, 2011). Yet, it remains uncertain whether moving up the social hierarchy actually matters for how people evaluate and feel about their lives, which is typically encapsulated under the term 'subjective well-being' (Diener, Suh, Lucas & Smith, 1999).

There are several reasons to suspect mobility should matter for well-being. Upward social mobility occurs when children grow up to become better educated, have greater access to material resources, and hold more prestigious jobs than their parents. In particular, education confers the knowledge and analytical skills needed to understand physical and psychological health information, the competence and resources required to achieve related goals, and a sense of control that can directly instil welfare benefits (Mirowsky & Ross, 1998; 2005; Mottus et al., 2014). Education has been shown to reduce exposure to blue-collar jobs which can impact health and well-being (Mazzonna, 2014) whilst also providing valuable skills needed to access high status jobs that pay well which has been shown to be associated with better satisfaction with life (e.g. Diener & Oishi, 2000; Kahneman & Deaton, 2010). Further, evidence from intervention studies and natural experiments suggests that decreases in poverty can lead to reductions in negative affect and stress (Haushofer & Fehr, 2014). Those who experience upward social mobility typically have greater access to material resources than they had as children and for this reason may perceive fewer financial concerns than others. Insofar as upwards intergenerational social mobility results from greater education, access to material resources and fewer financial concerns than one's parents, it may also generate improvements in subjective well-being.

A parallel route through which intergenerational mobility may relate to subjective well-being is through health. It is well known that the socioeconomic circumstances in which a person lives affects their health (Adler & Stewart, 2010). Whilst socioeconomic circumstances in childhood have been shown to be specifically important for establishing later health inequalities (Davey Smith, Hart, Blane & Hole, 1998) socioeconomic factors in adulthood remain important predictors of disease and health when early life factors are controlled for (Marmot, Shipley, Brunner & Hemingway, 2001) and factors in child- and adulthood have been shown to make independent contributions to later health (Hyde, Jakub, Melchior, Van Oort, & Weyers, 2006). Improving one's social circumstances relative to that of one's parents may offset some of the negative health consequences of coming from a background of disadvantage. It has been shown that individuals from lower socioeconomic status (SES) backgrounds who improve their social position have fewer chronic conditions such as diabetes, heart/lung disease, hypertension and cancer compared to those with stable or downward social trajectories (Luo & White, 2005). Elsewhere, upward mobility has been shown to be associated with lower levels of depressive symptoms (Gugushvili, Zhao & Bukodi, 2018) and self-reported physical health problems at age 32 (Cundiff, Boylan, Pardini & Matthews, 2017). We suggest that the positive health benefits of upward mobility are likely to be accompanied by improved subjective well-being.

Despite the various ways in which intergenerational social mobility may be linked to well-being, evidence demonstrating this association remains elusive. Whilst some reports show that well-being is higher among the upwardly mobile relative to the immobile (Chan, 2018; Zhao, Li, Heath & Shryane, 2017) and that upward (but not downward) mobility is associated with positive subjective well-being outcomes (Nikolaev & Burns, 2014), others observe no association between intergenerational mobility and subjective well-being (Hadjar & Samuel, 2015; Marshall & Firth, 1999; Zang & de Graaf, 2016). Prior research has been criticised for: (i) relying on bias-prone retrospective assessments of childhood socioeconomic status, (ii) surveying participants of varying ages who have experienced different societal economic circumstances and opportunities for social mobility throughout their lives, and (iii) failing to adjust for potentially confounding third variables (e.g. childhood cognitive ability, personality traits) that could affect both social mobility and well-being throughout adulthood (Iveson & Deary, 2017).

In a recent report, Iveson & Deary (2017) drew on longitudinal data from a Scottish cohort to address each of these issues. First, the authors employed a prospectively assessed measure of intergenerational social mobility (changes in occupational social class) that enabled the degree of social mobility each participant experienced to be operationalized. Second, they sampled a cohort of same-aged individuals, thus ensuring that any observed mobility-wellbeing associations reflected differences in well-being attributable to an individual's change in socioeconomic circumstances relative to their parents, rather than reflecting the impact of broader societal economic change (which was constant across the sample). Thirdly, they controlled for childhood intelligence: a key individual difference characteristic known to predict both social mobility and well-being in adulthood (Von Stumm, Gale, Batty, & Deary, 2009; Gale et al., 2012). Including controls of this kind is key to ensuring that associations cannot be explained by individual and psychological characteristics, which are likely to predict both better socioeconomic outcomes as well as higher satisfaction with life.

Iveson & Deary (2017) reported that the amount of change in occupational social class in adulthood relative to childhood (difference between own and father's occupational class) did not predict self-reported life satisfaction or general health in later life. This is arguably surprising given the multiple theoretical reasons to expect relationships of this kind as outlined above. We suggest that the absence of a relationship between intergenerational mobility and health and well-being may have arisen for two methodological reasons. First, life satisfaction and health were assessed in retirement (at age 78), when a number of factors may operate to diminish the benefits of intergenerational mobility for well-being. Participants were no longer actively engaged in the labour market at 78 and so the benefits of mobility in occupational class may no longer be tangible to them. If mobility has consequences for health then those who experience upward mobility may live longer and be overrepresented in an older sample making it more difficult to detect well-being associations with mobility. It is also possible that the prioritization of interpersonal relationships and emotional goals and improvements in the regulation and experience of well-being with age may mean mobility concerns are less impactful than at earlier points in the life-span (Carstensen & Mikels, 2005). Second, the authors relied on changes in a single indicator of social status (occupational class) thus neglecting a wide-range of alternative (albeit not necessarily equal) social mobility indicators including changes in educational attainment, income, and housing characteristics (see Galobardes, Lynch & Smith, 2007, for comparable arguments).

In the current study, we therefore sought to extend Iveson & Deary's work and more comprehensively assess whether intergenerational social mobility predicts well-being in midlife. Specifically, we employed two large-scale prospective UK data cohorts with multiple features advantageous to answering the current question. Both cohorts include comparable prospectively-collected data on multiple aspects of SES (educational attainment, social class and dwelling size) in both child and adulthood, allowing assessment of the degree of change in social status, but no longer reliant on a single measure. Both cohorts also contain rich childhood characteristics data on intelligence, self-control and subjective well-being that enabled us to identify the potential role of these early life individual differences in explaining relationships between mobility and midlife well-being. Finally, both cohorts included measures of health as well as perceived financial difficulties at age 42, allowing examination of the role of these factors as distinct theoretical mediators of the relationship between mobility and midlife life satisfaction to be examined.

Method

Samples

The British Cohort Study (BCS; Elliot & Shepherd, 2006) and the National Child Development Study (NCDS; Power & Elliot, 2006) are nationally-representative British cohorts that each follow the lives of approximately 17,000 children (BCS: *N* = 16,569; NCDS: *N* = 17,416) born in Britain in a single week in 1970 (BCS) and 1958 (NCDS) respectively. Both surveys collect information on the socioeconomic circumstances, physical and educational development, and health and well-being of participants across life. To gauge childhood socioeconomic status (SES) and childhood characteristics we drew on information derived from parental interviews, teacher-ratings, and the individuals themselves from birth, age 5, and age 10 waves in the BCS and birth, age 7, and age 11 waves in the NCDS. We then linked this information to participant SES and well-being in adulthood. Our key outcome measure was life satisfaction measured at age 42. This time point was chosen because comparable measures were available for the same age in both the BCS and NCDS, because this age is similar to cohort members' parents when their SES was assessed and because this was an age at which we suspected that intergenerational changes in social indicators on life satisfaction would be most impactful given that participants were still in the labour market and would be established in their careers. We also examined midlife (age 42) self-rated health and perceptions of financial difficulties in both cohorts.

The analytical sample therefore included only participants who took part in the age 42 waves of the BCS (*N* = 9,841) and the NCDS (*N* = 11,419) and who provided a valid response to the life satisfaction, self-rated health, and financial difficulties questions at that point (*N* = 9,683 in BCS and *N* = 11,265 in NCDS). To avoid further reductions in sample size we used full-information maximum-likelihood estimation (FIML) to account for uncertainty due to missing data on covariates rather than excluding those missing data on these variables. Despite attrition, both cohorts have been shown to stay largely representative of the original samples over time with observable socioeconomic characteristics having weak predictive power in explaining patterns of retention (Hawkes & Plewis, 2006; Mostafa & Wiggins, 2015). In this study, we investigate sources of participant retention (vs. dropout) and provide supplementary analyses weighted to account for differential attrition as a function of sex, childhood socioeconomic status indicators (parental educational attainment, social class, and dwelling size), cognitive ability, self-control, and psychological distress (see Statistical Analysis section for details).

Measures

Intergenerational social mobility. Measures of SES were selected on the basis that comparable variables were available in both cohorts and for both participant's parents and the participant at age 42. Measures meeting these criteria were: social class, age when left formal education, and dwelling size (number of rooms in household). Questions about parents' and participants' income were also included, however, these were operationalised differently (i.e. gross vs. net income; banded vs. continuous measures) for each time-point and cohort. We therefore did not include income as a measure of SES because it was not possible to harmonize these variables. The final three indicators were coded so that high levels reflected higher SES. Parental social class was measured at the time of the participant's birth based on the father's occupation and at age 42 based on the participant's occupation. Occupations were classified into a six-point social class scale using the Registrar General's Social Classes scheme (where *1* = professional occupations (class I), *2* = managerial or technical occupations (class II), *3* = non-manual skilled workers (class IIInm), *4* = manual skilled workers (class IIIm), *5* = semiskilled workers (class IV), *6* = unskilled workers (class V; Office of Population Census & Surveys, 1980). To align with the direction of the education and dwelling size indicators this scale was reverse coded to run from *1* = unskilled workers to *6* = professional occupations. Education was measured using the age at which the participant/the participant's parents left school/education. Parental education was collected at birth in the BCS and when the child was 16 in the NCDS. In the NCDS parental education was measured using a 10-point scale (<13 years, 13-14, 14-15, 15-16, 16-17, 17-18, 18-19, 19-21, 21-23, >23 years) and all school leaving age measures in both cohorts were recoded to this 10-point ordinal scale. Parental education was derived from the average age at which the mother and father left school or if the school leaving age was only available for one parent then we used that measure. The number of rooms in the childhood home was measured by asking parents to indicate the number of rooms, excluding bathrooms, toilets, and kitchens, in the accommodation occupied by the household when the participant was aged five in the BCS and seven in the NCDS. Following the same format participants provided information on their dwelling size at age 42. Dwelling size was capped at a maximum of 12 or more rooms with less than .5% of the sample falling into this group at any point.

Following the approach employed by Iveson & Deary (2017), to produce our indicators of intergenerational social mobility we calculated the difference in age leaving education, social class, and dwelling size between the participant at age 42 and his/her childhood level of the corresponding measure. For example, a participant classified as professional class (6) at age 42 and whose father was a non-manual skilled worker (4) would receive a score of +2 on intergenerational mobility in social class.

Life satisfaction. At age 42, individuals in both cohorts were presented with the following text: *Here is a scale from 0-10 where 0 means that you are completely dissatisfied and 10 means that you are completely satisfied. Please enter the number which corresponds with how satisfied or dissatisfied you are about the way life has turned out so far*. The same scale was used to assess life satisfaction at ages 30 (in the BCS) and 50 (NCDS). Single-item measures of life satisfaction are often used in previous work looking at social mobility and subjective well-being (e.g. Hadjar & Samuel, 2015; Iveson & Deary, 2017), exceed the threshold of .70 reliability (Lucas & Donellan, 2012), and have been shown to have high validity akin to multiple item satisfaction with life scales (Cheung & Lucas, 2014). Average life satisfaction at 42 was 7.36 (SD = 2.00) in the BCS and 7.28 (SD = 1.92) in the NCDS.

Self-rated health. At age 42, participants rated their current health on a scale from *poor* (1) to *excellent* (4 in the NCDS and 5 in the BCS). Single-item indicators of self-rated health provide a global summary of health which have been shown both to exceed objective measures such as physical measures derived from blood assays in predicting mortality (Ganna & Ingelsson, 2015) as well as to provide comparable predictions for healthcare usage and hospitalizations as multi-item subjective health measures (DeSalvo, Fan, McDonnell, & Fihn, 2005). Average self-rated health was 3.61 (SD = 1.07) in the BCS and 3.08 (SD = .76) in the NCDS. The majority of the difference in self-rated health levels between the two cohorts is likely attributable to the different rating scales used in both studies as noted above.

Perceived financial difficulties. At age 42, participants in both cohorts were asked a version of the following question: “How well would you say you personally are managing financially these days?” to which the following responses were available: *1* = living comfortably, *2* = doing alright, *3* = just about getting by, *4* = finding it quite difficult and *5* = finding it very difficult. On average participants indicated that they were 'doing alright' financially in both the BCS (*M* = 2.16, *SD* = 0.98) and the NCDS (*M* = 2.03, *SD* = 1.01).

Covariates

Childhood cognitive ability. At age 10, BCS cohort members completed the 120-item British Ability Scales; which are comprised of four subscales (Elliott, Murray & Pearson, 1978). In the Word Definitions subscale, children were asked to give a meaning for each of 37 words that increased in complexity. In the Word Similarities task, children were presented with 21 three-word lists and were required to provide another category exemplar as well as a name to unite the items. In the two non-verbal subscales, children were asked to recall 34 series of digits that increased in difficulty (Digit recall) and to complete the missing section of 28 incomplete patterns (matrices). The BAS has high internal reliability (Cronbach's α = .93) as well as convergent validity with more recently established measures of cognitive ability such as the Wechsler Intelligence Scale for Children and the Stanford-Binet Intelligence test (Elliott et al., 1978; McCallum & Karnes, 1987). The mean BAS score for the current sample with available data (*n* = 7,133) was 61.33 (*SD* = 13.02). At age 11, NCDS cohort members completed an individual general ability test, conducted by their teacher. The teacher presented children with 80 general ability items (40 verbal and 40 non-verbal). For both sets of items, children were presented with four stimuli (words in the verbal tests, shapes and symbols for the non-verbal items) linked either logically, semantically or phonologically. Children demonstrated their understanding of this association by selecting a fifth item from a second list of five alternatives. Performance on the task was quantified using a total tally of correct responses, giving a final score between 0 and 80 with average childhood cognitive ability of 44.72 (*SD* = 15.59) in the NCDS sample with available data (*n* = 9,705). This test has high levels of test-retest reliability (Cronbach's α = .94).

Childhood self-control. In the BCS self-control was measured from teacher ratings taken at age 10, using a subset of items from the Child Developmental Behaviors questionnaire. Using a visual analogue scale, teachers marked the extent to which the child demonstrated a particular behavior from "not at all" (1) to "a great deal" (47). Based on prior work using this scale (Daly, Delaney, Egan & Baumeister, 2015), we employed 9-items related to attentional control (e.g. "child becomes bored during class", "pays attention in class") and perseverance (e.g. "shows perseverance", "completes tasks"). Total self-control scores could vary from 1 to 47 and mean self-control was 31.37 (*SD* = 9.98; *n* = 7,722). At ages 7 and 11, teachers of children in the NCDS were asked to make a number of yes/no evaluations of children's behaviour using the Bristol Social Adjustment Guide (BSAG; Stott, 1969) and 13 phrases relating to 'inconsequential behaviour' (e.g., "cannot attend or concentrate for long", "does not know what to do with himself, can never stick at anything long"), were used to index childhood self-control in the current sample. The average of scores at ages 7 and 11 was calculated by summing the number of items to which teachers responded yes to and the product re-coded so that higher scores reflect more self-control. Mean self-control was 11.70 (*SD* = 1.65; range 1-13; *n* = 10,814). Although self-control was high in these samples, mean self-control is comparable with previous work wherein these measures have been shown to correlate strongly (*r* > .7) with contemporary self-control scales and to predict economic life outcomes (Daly et al., 2015).

Childhood distress. Five-items from the teacher-rated Child Developmental Behaviors questionnaire were operationalised to index childhood distress in the BCS. These included statements such as "afraid of new things/situations", is "worried and anxious about many things" and "behaves nervously". Responses to each of these items (from 1 to 47) were averaged to create a composite measure of distress (*M* = 17.02, *SD* = 9.52, *n* = 7,548). These items typically measure neuroticism, which is a robust correlate of depressive symptoms as well as longevity and quality of life (Lahey, 2009). In the NCDS, childhood distress was operationalised using scores from the depression 'syndrome' of the BSAG. Phrases on this dimension related to depression (e.g. "Expression is miserable, depressed"), apathy (e.g. "Apathetic (just sits))" and lethargy (e.g. "Lacks physical energy"). Scores across ages 7 and 11 were averaged and the mean score on this scale was 1.95, where high scores reflect greater distress (*SD* = 1.25; range 0 - 14; *n* = 10,814).

Statistical analysis

We employed Mplus (Version 8.3; Muthén & Muthén, 2019) and covariance Structural Equation Modelling (SEM) to examine the association between intergenerational mobility and life satisfaction, with both variables modelled as continuous. SEM was selected as the most appropriate analytic approach primarily because it allows mobility to be more accurately captured as a formative construct derived from changes in multiple aspects of SES, but also because it enables multiple potential pathways between intergenerational mobility and life satisfaction to be modelled simultaneously. Mobility was treated as a formative rather than a reflective construct because the direction of causality is from the indicators (intergenerational changes in socioeconomic status indicators) to the latent variable (social mobility) rather than vice versa. For example, those who become more educated than their parents tend to show intergenerational increases in social status in adulthood. In contrast, moving up the social hierarchy in adulthood does not necessarily generate improvements in educational attainment. As such, we produced a formative construct labelled 'intergenerational social mobility' with three formative indicators: intergenerational change in educational attainment, change in social class, and change in dwelling size. This construct therefore allowed us to model the relationship between each key variable at age 42 and overall intergenerational mobility, which was modelled on a continuum ranging from high levels of downward mobility to high levels of upward mobility. Covariances between the formative indicators were included in the model.

In each cohort our intergenerational social mobility formative composite was first used to predict life satisfaction at age 42 in a model that adjusted for participant gender only. To ascertain goodness of model fit we examined the root mean-square error of approximation (RMSEA) and the comparative fit index (CFI). The RMSEA is sensitive to the number of parameters included in the model and assesses how well a model "with unknown but optimally chosen parameter values, [would] fit the population covariance matrix if it were available" (Browne & Cudeck, 1993). The Tucker-Lewis Index (TLI) also takes the number of parameters into account by drawing on the chi square to degrees of freedom ratio to assess the change in fit per degree of freedom. The CFI takes sample size into account and compares the fit of the sample covariance matrix with the null model assuming all variables are uncorrelated. RMSEA values lower than .06 and TLI and CFI values of .95 or greater indicate good fit. Next, to understand the potential role of childhood factors we included further adjustment for the role of childhood intelligence, self-control, and distress in forecasting adult life satisfaction. We apply the same approach to understanding the relation between social mobility and midlife self-rated health and perceived financial difficulties.

Next, we examined the direct effect of social mobility on life satisfaction and indirect effects via self-rated health and financial difficulties in a path model. This model is depicted in Figure 1. Indirect paths were examined simultaneously and tested by estimating the product of path coefficients. The statistical significance of the indirect effects was estimated using 10,000 bootstrap samples created by drawing random samples with replacement from the original data to produce bias-corrected estimates of the 95% confidence intervals. Confidence intervals not including zero are statistically significant. In supplementary analyses weighted to account for differential attrition we use maximum likelihood estimation with robust standard errors as this approach facilitates the use of survey weights in Mplus. In supplementary analyses we also incorporate composite constructs assessing parental and participant SES with each composite assessed using three indicators gauging educational attainment, social class, and dwelling size. These analyses aimed to test the sensitivity of our mediation models to the inclusion of controls for the effects of parental and participant SES.

Finally, we examined the association between social mobility and longitudinal change in life satisfaction. Specifically, in the BCS we tested whether there was an association between social mobility and changes in life satisfaction between ages 30 and 42 adjusting for life satisfaction at age 30. This analysis is a conservative test as it captures only mobility-related changes in life satisfaction occurring in the second half of the overall period of adulthood examined (i.e. age 30 to 42) and discounts potential links between mobility and life satisfaction from ages 18 to 30. In the NCDS it was possible to test whether intergenerational mobility up to age 42 forecasts future life satisfaction at age 50 controlling for satisfaction at age 42. This is also a conservative test because it controls for those aspects of the association between social mobility and life satisfaction that persist over time (i.e. from age 42 to 50 years). Accordingly, only contributions of mobility to future life satisfaction that are additional to this persistent link are estimated.

Missing data. In the analytical sample, 17.2% of covariate variable values were missing in the BCS and 5.5% in the NCDS. In addition, 9.1% of values on the social mobility measures were missing in the BCS and 11.2% in the NCDS (see Table 1 for sample size for each variable). To appropriately address missing data and avoid biased parameter estimates we fitted all SEM models using full-information maximum-likelihood estimation (FIML).

FIML estimation accounts for uncertainty due to missing data by utilizing all available cases in the likelihood function. FIML has been shown to produce consistent parameter estimates and standard errors under the missing at random assumption and has a superior statistical profile to traditional approaches such as listwise and pairwise deletion (Enders & Bandalos, 2001).

We conducted attrition analyses because a notable portion (42% of the BCS sample and 35% of the NCDS sample) of the initial sample assessed at birth were missing data on our outcome measures which were assessed at age 42. Specifically, we examined linkages between participant retention (vs. dropout/outcome data unavailable) and childhood covariates: parental social status, educational attainment, dwelling size and child sex, cognitive ability, self-control, and distress levels. We estimated the standardized difference (d) between those retained in the sample and those without outcome data. This analysis showed that, compared to those without outcome data, those included in our analyses had higher levels of cognitive ability (BCS: d = .33; NCDS: d = .34) and self-control (BCS: d = .23; NCDS: d = .22) and lower levels of psychological distress (BCS: d = -.10; NCDS: d = -.20). Those included in the study also came from families with higher levels of socioeconomic status than those without available outcome data (social class – BCS: d = .19; NCDS: d = .12; educational attainment – BCS: d = .13; NCDS: d = .11; dwelling size – BCS: d = .09; NCDS: d = .06) (all p values < .01). We also found that females made up a larger portion of those included in the study (BCS: 52.2%; NCDS = 50.9%) compared to those without available outcome data (BCS: 43.9%; NCDS = 44.1%) (p values for differences < .01).

Given this evidence of selective attrition we generated inverse probability weights to test the sensitivity of the study results to selection bias. To do this we used available covariate data to predict the probability of retention in the study. Specifically, we ran a logistic regression analysis where retention (non-missingness) was estimated as a function of baseline/childhood characteristics. This analysis showed that retention levels were significantly higher amongst female participants, those with higher childhood cognitive ability, and self-control in both cohorts (see Table S1 in the Supplementary Materials for further details). In contrast, childhood distress and SES indicators did not consistently predict retention. The relationship between childhood factors and retention did not differ markedly depending on whether missing covariate data was handled using mean imputation or Montecarlo integration. We therefore used the calculated probability of retention from our analyses using mean imputation to construct an inverse probability weighting variable for each cohort. By upweighting observations with a high probability of attrition the application of these weights ensures a more valid reflection of the original sample characteristics. We applied this weighting variable in supplementary analyses (reported below) to gauge the sensitivity of our results to correction for bias due to selective attrition.

< INSERT TABLE 1 ABOUT HERE >

Results

Across both cohorts, all three measures of SES increased from childhood (T0) to adulthood (T1) indicating the sample as a whole experienced upward intergenerational mobility, as shown in Table 1. Participants completed on average two additional years of education (BCS = 1.98, NCDS = 2.14) than their parents and lived in houses with .39 (BCS = .39, NCDS = .38) more rooms than their childhood homes. The most common occupational social class of participants' fathers was 'manual skilled worker' (BCS: 44.5%; NCDS: 50.4%) whereas the most common class of participants was 'managerial and technical' (BCS: 44.6%; NCDS: 37.7%). Correlations between aggregate SES at T0 and T1 were moderate

(BCS, average r = .42; NCDS, average r = .41), consistent with prior work showing stability of SES between generations in the UK (Blanden, Goodman, Gregg & Machin, 2004). Perceived financial difficulties were negatively correlated with life satisfaction scores (BCS: r = -.38, p < .001; NCDS: r = -.33, p < .001) and life satisfaction was positively correlated with self-rated health in both cohorts (BCS: r = .33, p < .001; NCDS: r = .24, p < .001) (see Tables S2 and S3 for correlation matrix examining individual variables).

Intergenerational social mobility and life satisfaction at age 42

Intergenerational social mobility positively predicted midlife life satisfaction in both the BCS (β = .19, 95% *CI* = [.17, .21], p<.001) and the NCDS (β = .15, 95% *CI* = [.13, .16], p<.001) in a model that adjusted for participant gender only. In both the BCS and NCDS our life satisfaction SEM models fit the data well (BCS: RMSEA = .02, CFI = .95, TLI = .93; NCDS: RMSEA = .02, CFI = .95, TLI = .93). Further adjustment for individual differences in childhood traits (cognitive ability, self-control, psychological distress) had little impact on these associations, reducing the magnitude of coefficients by less than 10% on average (see Table 2 & Table S4).

< INSERT TABLE 2 ABOUT HERE >

Mediators of the relationship between social mobility and life satisfaction

Social mobility was negatively predictive of perceived financial difficulties at age 42 in both cohorts (average β = -.22), as shown in Table 2. Intergenerational change in social hierarchy was also predictive of midlife self-rated health in both cohorts (average β = .13). Adjustment for childhood individual differences produced a notable reduction in the association between social mobility and self-rated health, reducing this link by 21% in the BCS and 45% in the NCDS.

Our mediation analyses, shown in Table 3, revealed a significant indirect effect between intergenerational social mobility and life satisfaction through perceived financial difficulties in both cohorts (BCS: β = .05, 95% *CI* = [.05 , .06]; NCDS: β = .05, 95% *CI* = [.05, .06]). Similarly, self-rated health explained part of the association between mobility and life satisfaction (indirect effect in BCS: β = .03 95% *CI* = [.02, .03]; NCDS: β = .01, 95% *CI* = [.01, .02]). In total, 44% / 43% of the association between social mobility and life satisfaction was mediated by participant ratings of financial difficulties and health in the BCS / NCDS. In supplementary analyses we also examined life satisfaction at age 50 in the NCDS and showed that the mediating role of self-rated health and financial difficulties differed little when this more extensive time horizon was examined (see Table S5).

< INSERT TABLE 3 ABOUT HERE >

<INSERT FIGURE 1 ABOUT HERE>

Sensitivity tests

Applying inverse probability weights to account for differential attrition as a function of childhood factors (i.e. childhood cognitive ability, self-control, distress, and parental SES indicators) did not affect the point estimates of the relationship between intergenerational social mobility and life satisfaction, self-rated health, and financial difficulties (see Table S6). Similarly, our weighted mediation analysis differed minimally from our main estimates (see Table S7).

Including further adjustment for a formative construct comprised of three childhood SES indicators (i.e. parental education, social class, and dwelling size) did not diminish any of the key associations outlined in Figure 1, as shown in Table S8. In this model childhood SES was positively associated with life satisfaction at age 42 in both cohorts. The inclusion of a formative construct assessing participant SES at age 42 (educational attainment, social class, dwelling size) produced a decline of 37% in the magnitude of the total effect of intergenerational social mobility on life satisfaction (β = .16 reduced to β = .10). This reduction could be attributed to a reduction in the direct effect of mobility on life satisfaction whilst indirect effects remained unchanged. Participant SES at age 42 was robustly related to life satisfaction in both cohorts, as can be seen in Table S8.

Intergenerational social mobility and longitudinal changes in life satisfaction

Social mobility positively predicted changes in life satisfaction from age 30 to 42 years in the BCS (β = .11, 95% *CI* = [.09, .13], $p<.001$) in a model that adjusted for participant gender, childhood individual differences, and satisfaction levels at age 30 (see Table S9). We also found that intergenerational mobility forecasted increases in life satisfaction from age 42 to 50 in the NCDS (β = .07, 95% *CI* = [.05, .08], $p<.001$) in a model that adjusted for gender, childhood individual differences and satisfaction at age 42.

Discussion

The social gradient in well-being has been firmly established (Diener & Oishi, 2000; Kahneman & Deaton, 2010), yet whether shifts in social circumstances between generations are related to subjective well-being has remained uncertain. To test this, we drew on two large-scale birth cohorts, for which changes in social position between generations were modelled as a formative construct capturing prospectively assessed changes in educational attainment, social class and dwelling size from childhood to adulthood. Taken together our findings showed that one standard deviation of movement in social circumstances between generations was associated with a 0.17 SD predicted increase in life satisfaction at age 42. This pattern extended to longitudinal changes in life satisfaction – in both cohorts mobility continued to predict satisfaction at ages 42 (BCS) and 50 (NCDS) when life satisfaction in earlier adulthood was accounted for.

These findings differ from those reported in a similar approach by Iveson & Deary (2017) where social mobility from parental occupational social class was not found to predict either life satisfaction or health in later life. We posit that the divergences between the two reports arise because key adult outcomes were assessed here much earlier in life when mobility is likely to be more influential and because we modelled mobility as a formative construct formed from multiple indicators (i.e. changes in education, social class, and dwelling size), rather than focusing only on a specific indicator of socioeconomic status. It is reasonable to consider that the well-being impact of changes in individual indicators of socioeconomic status may vary across the lifespan (i.e. education benefits may be most tangible in early adulthood, whereas social class may be most pertinent in midlife when careers are usually most established). The advantage of the current approach of modelling mobility as a formative construct gauged via multiple indicators therefore was to capture the mobility-well-being association across a set of indicators to ascertain their combined link to well-being at age 42. Using this approach, associations between intergenerational mobility and midlife well-being were robust across two UK cohorts born in 1958 and 1970.

Importantly, we also observed that these associations remain when intelligence, self-control, and childhood distress are adjusted for, demonstrating that the association between intergenerational social mobility and well-being could not be attributed to the impact of these key childhood traits, which are typically not measured in studies examining the health and well-being impact of intergenerational mobility. Including these early life traits in our models most markedly impacted upon the association between social mobility and self-reported health. Our findings suggest that failing to consider childhood self-control, intelligence, and distress could result in estimates of the relationship between intergenerational mobility and midlife health ratings that are exaggerated by over 50% on average. However, adjusting for early life traits only had a minimal impact on the association between mobility and life satisfaction.

Our findings also shed light on some of the reasons why social mobility tends to forecast subjective well-being. Approximately 30% of the indirect effect of mobility on well-being was explained by improved reported general health levels in midlife. As such, the beneficial implications of social mobility for adult health (Cundiff et al., 2017) appear to spill over to positively impact life satisfaction. Those who improved their social standing relative to their parents also perceived fewer financial pressures in midlife and this accounted for the remainder of the indirect effect. We note with interest that our supplementary analyses showed that this mediatory pathway remained when socioeconomic circumstances at age 42 were adjusted for, indicating that perceptions of one's financial position operates as a route between mobility and well-being regardless of a person's objective material circumstances. This does not mean that material changes in one's finances are not a determinant of well-being, but rather that a significant aspect of the pathway between changes in one's social circumstances and a person's satisfaction with their life depends upon how these changes are perceived. This may be analogous to arguments that a key step in the causal pathway between social disadvantage and downstream health is an individual's perceptions of their social environment and whether they perceive them to be threatening (Matthews & Gallo, 2011).

This may not be the only class of psychological mechanism which can explain why mobility co-varies with life satisfaction. There are a number of social comparison mechanisms – perceptions of how we compare to others, past selves or future expectations (Kraus, 2018) - which may contribute but which remain unaccounted for in the current analyses. Expectancy-based accounts, for example, posit that well-being is improved once living standards increase beyond a certain expected standard. The social position of those who experience upward mobility may compare favorably to those that the person cares most about such as their family and long-standing friends. Experiencing the respect and admiration of this important social group may enhance feelings of social acceptance and well-being (Anderson, Kraus, Galinsky, & Keltner, 2012). As such, the psychological experience of social mobility may be informed by one's childhood social position that operates as a baseline comparator or is used to define social comparisons with important others. In line with this, it is now well-established that subjective status and comparisons with others are important predictors of physical health (Cundiff & Matthews, 2017; Gugushvili, Jarozs, & McKee, 2019; Rivenbark et al., 2019; Zell, Strickhouser & Krizan, 2018). It remains to be determined how much of the association between intergenerational mobility and subjective well-being arises because of changes in subjective social status relative to others or past selves.

Demonstrating a consistent beneficial effect of intergenerational social mobility on well-being is particularly important because elsewhere it has been argued that upward social mobility may actually undermine well-being. This is key, for example, to the original "dissociative hypothesis" (Sorokin, 1959) which posits that mobility is a stressful and alienating social experience (Friedman, 2014). More recently, it has been proposed that because social class is an entrenched aspect of social identity, movements between social classes will elicit 'status-based identity uncertainty' which should decrease subjective well-being (Simandan, 2017). Mechanisms of this kind would predict a negative association between mobility and subjective well-being which is not supported by the current finding that intergenerational mobility is positively associated with midlife subjective well-being. However, it is important to note here that whilst a strength of the current analytical approach is that it allows mobility to be captured as a function of multiple indicators of SES it does not permit differences between upward and downward mobility to be examined. Dissociating the direction of mobility may be of interest to assess in future work given that it has been shown that there may be different consequences for upward vs. downward mobility (e.g. Zhao et al., 2017) and it is well recognised in the psychological literature that losses and negative outcomes are weighted more heavily than wins and positive outcomes (Baumeister et al., 2001; Kahneman & Tversky, 1979).

Limitations

There are some other limitations of the current study to consider. Life satisfaction, self-rated health and perceived financial difficulties were all measured concurrently at age 42, reducing our ability to distinguish the directionality of the relationships between these variables. This timing was necessary to ensure that our assessment of participant health and financial circumstances was made after an extensive period of social mobility had occurred (i.e. between childhood and age 42). However, when we examined satisfaction with life as an outcome eight years after the assessment of our mediating variables, the indirect effect of general health ratings and perceived financial difficulties was unchanged. It is also the case that both cohorts are considerably more ethnically homogeneous (majority white) than the actual UK population and it is therefore not appropriate to generalise these findings to more ethnically diverse populations. Future work is needed to remedy this.

We should also consider issues relating to how mobility is modelled. It was not possible to adjust for both parental and participant SES simultaneously because this would mean our intergenerational mobility measure would be perfectly multi-collinear with these variables. Nonetheless, when adjusting separately we find that although both parental and participant SES are both predictive of life satisfaction, only participant SES reduced the mobility-well-being link. Moreover, this was not able to explain the associations between social mobility and life satisfaction observed in the current study. We also did not apply a diagonal reference model which is sometimes viewed as the "gold standard" for mobility research. Principally this is because this paper was designed to directly extend on the approach employed by Iveson & Deary (2017) and because simulation research has shown that diagonal reference models may generate results that force mobility effects to zero (Fosse & Pfeffer, 2019).

A final limitation was the self-report nature of our mediating variables. We make sense of our findings above by inferring that the physical health impact of mobility has a corollary effect on well-being, however an alternative explanation is that self-reported health captures a subtype of well-being related to concerns over one's health and that it is these concerns which are captured in the indirect pathway. Without simultaneous objective and self-report measures of health it is not possible to fully differentiate these alternatives, although it is important to note that the differential impact of adjusting for childhood characteristics on the mobility-health and mobility-wellbeing association makes it unlikely that our measure of health reduces entirely to a subdomain of life satisfaction. We also finish by noting that we anticipated that it would be the interpretation of one's physical health and financial situation, rather than objective changes in one's circumstances that would be crucial in translating the impact of social mobility on well-being. This expectation is supported by the observation that participants' perceptions of their financial situation continued to mediate this pathway when other markers of SES in adulthood were controlled for.

Conclusions

Collectively, these data provide evidence that intergenerational mobility is associated with well-being in adulthood. This relationship was comparable across cohorts and remained when key psychological traits were adjusted for. Our mediation analyses provided further support for arguments for life-course approaches to health (Marmot et al., 2001) because they show that improving one’s social circumstances positively predicted adult health, an important corollary of which is higher satisfaction with life. Upward mobility was also associated with fewer perceived financial pressures which in turn predicted enhanced well-being. Importantly, our findings support the possibility that efforts to promote upward social mobility may yield well-being dividends in mid-life. Finally, the present study provides considerable support for the oft-invoked assumption that improvements in living standards and health associated with upward mobility may have a positive effect on subjective well-being. The potential salutary consequences of social mobility also further highlight the importance of concerns that rates of upward intergenerational mobility are both slowing and not equally accessible to all areas of society.

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Table 1. *Descriptive Statistics for Participants in the 1970 British Cohort Study (BCS) and the 1958 National Child Development Study (NCDS).*

	BCS		NCDS	
Variable	*M*(*SD*) / %	*n*	*M*(*SD*) / %	*n*
Female (%)	52.00	9,683	50.90	11,265
Age parents left education[a]	16.37 (1.67)	8,868	15.50 (1.52)	8,251
Age participant left education[a]	18.35 (2.41)	9,651	17.64 (1.98)	11,027
Social class at birth[b]	3.36 (1.22)	8,885	3.23 (1.22)	10,181
Social class at age 42[b]	4.11 (1.23)	8,193	3.94 (1.25)	9,504
Dwelling size (childhood home)[c]	4.66 (1.27)	7,800	4.81 (1.32)	9,820
Dwelling size (home at age 42)[c]	5.05 (1.73)	9,640	5.19 (1.71)	11,216
Childhood cognitive ability[d]	61.32 (13.03)	7,132	44.72 (15.59)	9,705
Childhood self-control	31.37 (9.98)	7,722	11.70 (1.65)	10,814
Childhood distress	17.03 (9.53)	7,548	1.95 (1.25)	10,814
Financial difficulties at age 42[e]	2.16 (0.98)	9,683	2.03 (1.01)	11,265
Subjective health at age 42[f]	3.61 (1.07)	9,683	3.08 (0.76)	11,265
Life satisfaction at age 42[g]	7.36 (2.00)	9,683	7.28 (1.92)	11,265

[a] Calculated using mid-points from the 10-point school/education leaving age scale (<13 years, 13-14, 14-15, 15-16, 16-17, 17-18, 18-19, 19-21, 21-23, >23 years)
[b] Six-point social class scale coded from 1 (*unskilled workers*) to 6 (*professional occupations*).
[c] Measured by number of rooms used in the home excluding bathrooms/toilets and kitchen.
[d] Assessed using the British Ability Scales (BAS; range = 0 - 120) in the BCS and general ability test (range = 0 - 80) in the NCDS.
[e] Rated from 1 (*living comfortably*) to 5 (*finding it very difficult*).
[f] Rated from 1 (*poor*) to 5 (*excellent*) in the BCS and from 1 (*poor*) to 4 (*excellent*) in the NCDS.
[g] Rated from 0 (*completely dissatisfied*) to 10 (*completely satisfied*).

Table 2. *Standardized Path Coefficients of the Association between Intergenerational Social Mobility and Life Satisfaction, Self-rated health, and Perceived Financial Difficulties at age 42 in the 1970 British Cohort Study (BCS; N = 9,683) and the 1958 National Child Development Study (NCDS; N = 11,265).*

	Life satisfaction		Self-rated health		Financial difficulties	
Study	BCS	NCDS	BCS	NCDS	BCS	NCDS
	β 95% *CI*	β 95% *CI*	β 95% *CI*	β 95% *CI*	β 95% *CI*	β 95% *CI*
Intergenerational social mobility[a]	.19 [.17, .21]	.15 [.13, .16]	.14 [.11, .16]	.11 [.09, .13]	-.20 [-.22, -.18]	-.22 [-.24, -.20]
+ Childhood traits[b]	.18 [.15, .20]	.13 [.12, .15]	.11 [.09, .14]	.06 [.04, .08]	-.18 [-.20, -.16]	-.19 [-.21, -.17]

Note. All estimates are statistically significant at the p<0.001 level. 95% confidence intervals presented in brackets.

[a] Intergenerational social mobility is modelled as a linear composite of formative indicators (i.e. intergenerational changes in educational attainment, social class, and dwelling size).

[b] Models include further adjustment for childhood traits: cognitive ability, self-control, and child distress.

Table 3. *Standardized Path Coefficients of the Direct, Indirect, and Total Effects of Intergenerational Social Mobility on Life Satisfaction at age 42 in the 1970 British Cohort Study (BCS; N = 9,683) and the 1958 National Child Development Study (NCDS; N = 11,265).*

	Life satisfaction	
Study	BCS	NCDS
	β [95% *CI*]	β [95% *CI*]
Total effect	.18 [.15, .20]	.13 [.11, .15]
Total direct effect	.10 [.07, .12]	.07 [.05, .09]
Total indirect effect	.08 [.07, .09]	.06 [.06, .07]
via self-rated health	.03 [.02, .03]	.01 [.01, .02]
via financial difficulties	.05 [.05, .06]	.05 [.05, .06]

Note. All estimates are statistically significant at the $p<0.05$ level. 95% confidence intervals presented in brackets. Models are adjusted for gender, childhood cognitive ability, self-control, and distress.

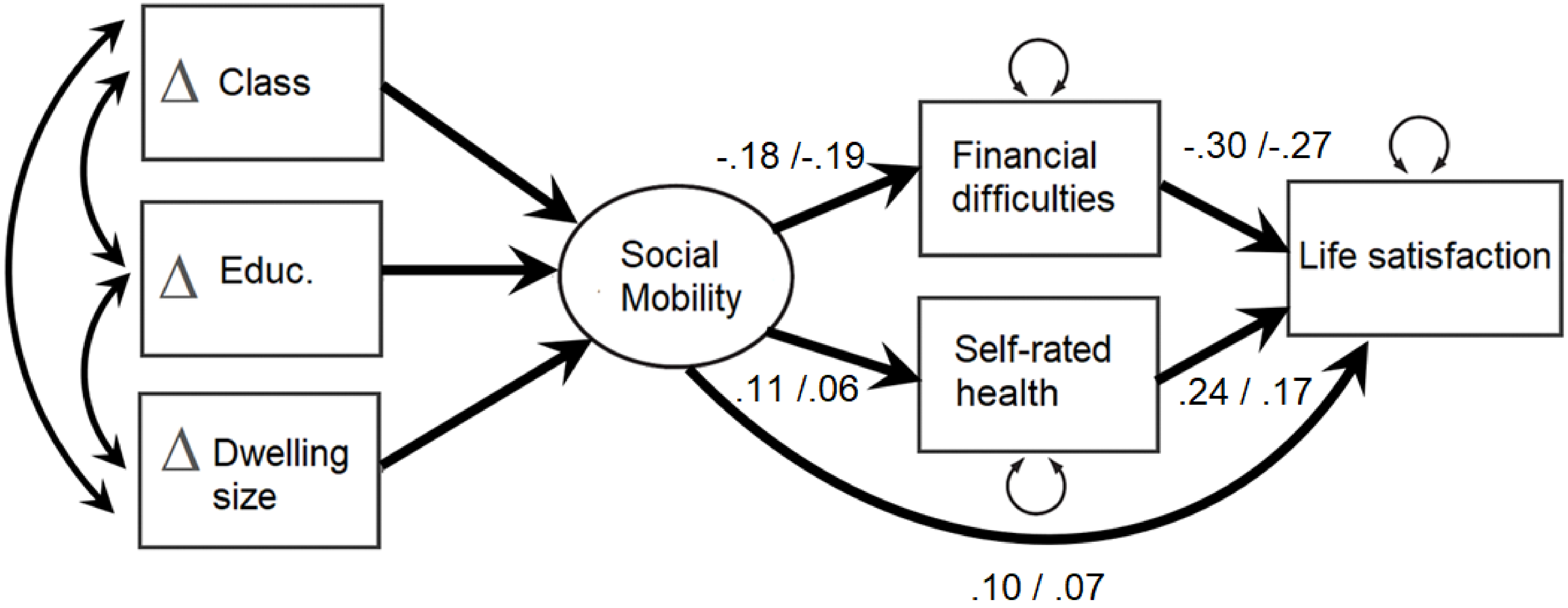

Figure 1. Conceptual Diagram of the Mediation Channel from the Formative Intergenerational Social Mobility Construct to Life Satisfaction at age 42 through Perceived Financial Difficulties and Self-Rated Health. Standardized coefficients for key paths in the BCS / NCDS are presented on the left and right respectively. All coefficients are statistically significant at the .001 level.

Note: Intergenerational Social Mobility was assessed by changes in social class (Δ Class), educational attainment (Δ Educ.), and dwelling size from childhood to age 42. Covariates included in the final model (i.e. gender, intelligence, self-control, and distress) are excluded for illustration purposes.