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How do moving and other major life events impact mental health? A longitudinal analysis of UK children

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ABSTRACT

Research has suggested that children who move home report poorer mental health than those who remain residentially stable. However, many previous studies have been based on cross sectional data and have failed to consider major life events as confounders. This study uses longitudinal data from ALSPAC, a UK population based birth cohort study, and employs within-between random effects models to decompose the association between moving in childhood and poor mental health. Results suggest that while unobserved between-individual differences between mobile and non-mobile children account for a large portion of this association, within-individual differences remain and indicate that moving may have a detrimental impact upon subsequent mental health. There is heterogeneity in children’s response to moving, suggesting that a dichotomy of movers vs stayers is overly simplistic.

1. Introduction

Mental ill health is one of the largest contributors to the global burden of disease and a major global health priority, inflicting a number of health, social and personal burdens upon individuals (Whiteford et al., 2013). Between 10% and 20% of children and adolescents worldwide suffer from mental health problems (Kieling et al., 2011), a rise from 10% at the turn of the millennium (Meltzer et al., 2000). In the UK recent estimates put this figure between 12.5% (Beardsmore, 2015) and 20% (Fink et al., 2016), depending on the definition of poor mental health. Childhood and adolescence are critical developmental periods for identification and intervention of mental ill health because problems in early life are associated with both higher likelihood (Helliwell et al., 2015) and longer durations (Kovacs et al., 1984) of mental ill health in adulthood. Mental illness costs the UK between £70 and £100 billion a year, of which between £14 and £20 billion is spent on health and social care costs (Mental Health Foundation, 2015). The socioeconomic patterning of mental health has been studied in detail, with children growing up in households characterised by low family socioeconomic position (SEP) suffering from an elevated risk of problems compared to those in high SEP families (Reiss, 2013). However, beyond these broad patterns there is still a lack of understanding on the specific social pathways that may contribute to child mental illness.

1.1. Moving house and mental health

One potential social pathway is the role of place, and in particular how transitions between places may be linked to mental health. Evidence suggests that individuals who are exposed to moving house (commonly termed residential mobility), report poorer mental health than those who do not (Jelleyman and Spencer, 2008; Morris et al., 2016b). This is particularly true for children and adolescents; findings from the UK suggest that young families with children who move have poorer mental health than those who remain residentially stable (Tunstall et al., 2012, 2010), and that children and adolescents are particularly vulnerable to deleterious mental health effects of moving (Anderson et al., 2014; Flouri et al., 2013). Such effects are thought to operate through a number of pathways including weakened social ties (Pribesh and Downey, 1999), disturbance to social networks (Coleman, 1988), ‘social stress’ (Silver et al., 2002), household disruption (Haveman et al., 1991), social isolation (Stubblefield, 1955), and

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reductions in parent-child interactions (Anderson et al., 2014). Similar findings have been observed in the USA showing a dose response relationship (Susukida et al., 2015) and both short and long term negative associations between moving in childhood and subsequent mental health (Bures, 2003; Gilman et al., 2003; Oishi and Schimmack, 2010; Simpson and Fowler, 1994). However, it is important to note that there have been conflicting findings with a number of studies failing to replicate associations between moving and poor mental health (Gambaro and Joshi, 2016; Jelleymann and Spencer, 2008; Stoneman et al., 1999; Verropoulou et al., 2002). The conflicting results from these studies cannot be explained by a common research design or analytical method, though there is a suggestion that a detailed account of family circumstances may lessen the apparent association between moving and poor mental health. We return to this shortly.

1.2. Neighbourhoods and selective migration

Specific characteristics of places play an important role in the patterning of mental health in addition to individual and family level factors. Studies have long shown that mental health is poorer in socially, economically and environmentally deprived neighbourhoods compared to less deprived neighbourhoods (Kim, 2008; Leventhal and Brooks-Gunn, 2000; Mair et al., 2008). With respect to people moving through different types of neighbourhoods, studies have found that making moves to more depraved neighbourhoods is negatively associated with mental health over and above the effect of moving (Tunstall et al., 2014, 2012). This implies that varying exposure to such conditions is an important factor in the development of mental ill health under the assumption of instantaneous effects. The semi-randomised Moving To Opportunity (MTO) experiment in the USA has provided evidence that children who move from extreme high to low poverty neighbourhoods experience a reduction in mental health problems in adolescence and early adulthood (Bridge et al., 2012; Leventhal and Brooks-Gunn, 2003; Ludwig et al., 2012). The MTO studies have been substantially critiqued however (Clark, 2008; Manley and van Ham, 2012). There has long been debate as to whether the association between moving and poor mental health is due to a causal neighbourhood effect or a selection effect (Kawachi and Subramanian, 2007; Oakes, 2004) and very few studies have utilised data and methods suitable for answering such questions. However, recent studies using longitudinal data have suggested that the relationship between neighbourhood deprivation and poor health may be due to selective migration of unhealthy individuals into deprived areas, rather than a causal effect from neighbourhood deprivation to poor health (Jokela, 2015, 2014).

1.3. The impact of life events on mental health

One aspect that has been commonly overlooked in studies of mental health and migration is that of major life events (Morris et al., 2016b). It has long been accepted that adverse life experiences play a role in the onset of psychological conditions (Rutter, 1981), with a body of research suggesting that adverse life events and childhood adversity are strongly associated with poor subsequent mental health (Dong et al., 2005; Felitti et al., 1998). This finding is consistent across a range of events including union dissolution (Strohschein et al., 2005), parental death (Trotta et al., 2015), childhood abuse (Varese et al., 2012), unemployment and job loss (Paul and Moser, 2009), and ‘total’ adversity (Trotta et al., 2015). Experience of each of these events has been associated with poorer mental health and the experience of such social adversity in adolescence has also been linked to poor mental health development trajectories into adulthood (Rajaleid et al., 2016). In a detailed analysis of mental health survey data from 21 countries, Kessler et al. (2010) found that all childhood adversities examined were associated with psychiatric disorders, and that this association increased with multiple adversities.

1.4. Limitations of the literature

Many studies of moving and mental health do not examine the impact of life events, raising questions over the validity of findings. Moving is not an exogenous process: rather, it is a highly complex set of processes that are influenced by a wide range of factors, including major life events (Morris, 2017), which also directly influence mental health. However, few studies examining the impact of moving on mental health have considered the occurrence of major life events. Because of the robust relationship between life events and both moving and mental health, excluding such events is likely to introduce bias due to unobserved confounding and raises the possibility of spurious associations between moving and mental health (Morris et al., 2016b). Those studies that have accounted for life events find that their occurrence has a strong attenuating effect on associations between moving and poor mental health, either eliminating this association (Dong et al., 2005; Gambaro and Joshi, 2016) or heavily attenuating it (Flouri et al., 2013; Tunstall et al., 2015). These findings suggest that while moving house may be a constituent part of a child’s story, it is likely the events that lead to moves rather than the moves themselves which are the main reason for differences between movers and stayers (Gambaro and Joshi, 2016).

1.5. Study aim

In this study we test the longitudinal association between moving in childhood and subsequent mental health, and determine whether this association is robust in the presence of major life events as confounders or if these induce spurious associations. Further, we decompose the association between moving and mental health to determine the extent to which any changes in mental health are due to the effects of moving or due to factors that are more common in mobile than non-mobile children. The use of longitudinal data allows us to model directional associations and provides greater freedom to draw causal inferences than the cross-sectional approaches that remain widespread throughout the literature. Given the ongoing health selection debate over whether moving causes poor mental health or poor mental health causes people to move, such a longitudinal approach allows us to correctly structure and test for temporal effects of moving to mental health.

2. Data and methods

2.1. Data source

We use data from the Avon Longitudinal Study of Parents and Children (ALSPAC). All pregnant women resident in the (former) Avon Health Authority area in South West England with an expected date of delivery between April 1991 and December 1992 were eligible to enrol. The full sample consists of 14775 live births. After birth, data were primarily collected from the study mothers and then from children via regular self-completion questionnaires and hands on assessments from the age of seven. The ALSPAC cohort is largely representative of the UK population when compared with 1991 Census data; however, there is under representation of ethnic minorities, single parent families, and those living in rented accommodation. For full details of the cohort profile and study design see Boyd et al. (2013) and Fraser et al. (2013).2

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1 The study website contains details of all the data that is available through a fully searchable data dictionary (http://www.bristol.ac.uk/alspac/researchers/access/).

2 Ethical approval for the study was obtained from the ALSPAC Ethics and Law Committee and the Local Research Ethics Committees.
2.2. Outcomes

We utilise two measures of mental health, the Strengths and Difficulties Questionnaire\(^3\) (SDQ) and the Development and Well-Being Assessment\(^4\) (DAWBA).

2.2.1. Strengths and difficulties questionnaire

The SDQ is used to assess child emotional and behavioural difficulties (Goodman et al., 2000b). Study mothers reported on the SDQ for children on four occasions; child ages 4, 7, 10, and 12 years. The SDQ is one of the most widely used questionnaires for evaluating psychological well-being amongst children and consists of 25 items that covers common areas of emotional and behavioural difficulties (emotional symptoms; conduct problems; hyperactivity/inattention; peer relationship problems; and prosocial behaviour). Responses to each question are on a three-point scale: not true, somewhat true and certainly true, coded to scores of 0, 1 and 2 respectively. Given our sample size and potential issues with using individual subscales of the SDQ (Goodman et al., 2010) we use total SDQ score defined as the count of problems on the first four scales (median: 7; inter-quartile range: 6). The Cronbach’s alpha for the total SDQ scales at each measurement occasion was 0.639; 0.668; 0.670; 0.681.

2.2.2. Development and well-being assessment

The DAWBA is used to assess the presence of a range of child psychiatric disorders (Goodman et al., 2000a) and therefore provides a different measure of mental health to the SDQ. DAWBA information was collected from responses to a series of open and closed questions reported by the child’s mother at three occasions; child ages 7, 10, and 13 years, which are then coded by a computer programme to predict the probability of clinical diagnoses of 15 disorders. Six probability bands were calculated and these were reversed due to software model parameterisation requirements with the lowest two categories combined due to low numbers, resulting in an ordered categorical measure giving the likelihood of diagnosis of a psychiatric disorder in decreasing order: > 70%; ~ 50%; ~ 15%; ~ 3%; ~ 0.5%.

2.3. Exposure

Our exposure variable is the experience of moving house. Study mothers were asked to report if they had moved since the previous questionnaire, providing a binary response at each occasion coded 1 if a family had moved since the last occasion and 0 if not. We also included a variable indicating the total number of moves made in the analytical period prior to the current occasion to determine whether there were any threshold effects. Given the low numbers of study families moving more than once between measurement occasions we were unable to use an absolute measure of residential move frequency.

Table 1 and 2 show the changes in SDQ and DAWBA scores throughout the study period. Mean SDQ scores reduce with age from 8.73 at 4 years to 6.16 at 12 years. Similarly, overall distributions within the DAWBA bands reduced with age.

2.4. Covariates

A range of covariates relating to moving and mental health were used as controls in our models (Table 3). Time invariant covariates include sex and ethnicity of child, maternal and paternal age at birth, highest parental education, highest parental social class based on occupation, family structure prior to the analytical period, subjective financial difficulty, maternal anxiety, maternal depression during childhood, housing tenure, and number of residential moves made prior to the analytical period. Time variant covariates include child age in years and age squared to account for non-linear effects of age, neighbourhood deprivation as measured by quintiles of the Index of Multiple Deprivation (IMD), and a range of adverse life events as reported by the study mothers. These include separation from partner, divorce from partner, marriage, child birth, death within the family, and job loss of either the mother or her partner. Modelling separation and divorce separately allows for differentiation between married and non-married union dissolutions. Given the short time intervals between measurement occasions it was not plausible to utilise an absolute measure of frequency of events where these occurred more than once due to low numbers.

Participants included in the analytical samples were less likely to have experienced parental separation and divorce but more likely to have experienced sibling birth and parental job loss than those excluded. Children included in analyses also came from more educated, higher social class and stable families, were less likely to be non-white or live in deprived neighbourhoods, and more likely to have younger parents, have experienced more financial difficulty, and have mothers who experience depression and anxiety than those excluded (Table 3).

### Table 1

Occasion level information for SDQ sample.

<table>
<thead>
<tr>
<th>Occasion</th>
<th>Sample size</th>
<th>Age in years mean (SD)</th>
<th>Moved n (%)</th>
<th>Mean SDQ score (SD)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>6033</td>
<td>3.94 (0.12)</td>
<td>1701</td>
<td>8.73 (4.52)</td>
</tr>
<tr>
<td>2</td>
<td>5497</td>
<td>6.78 (0.10)</td>
<td>1040</td>
<td>7.39 (4.75)</td>
</tr>
<tr>
<td>3</td>
<td>4399</td>
<td>9.63 (0.11)</td>
<td>1012</td>
<td>6.58 (4.76)</td>
</tr>
<tr>
<td>4</td>
<td>2770</td>
<td>11.7 (0.10)</td>
<td>460 (16.61)</td>
<td>6.16 (4.64)</td>
</tr>
</tbody>
</table>

SD, Standard deviation.

\(^5\) For further information on the DAWBA see (http://www.dawba.info).

\(^6\) By capturing all omitted individual-specific factors, the random intercept controls for time-invariant unobserved heterogeneity.

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\(^3\) For further information on the SDQ see (http://www.sdqinfo.org).

\(^4\) For further information on the DAWBA see (http://www.dawba.info).
The random intercept and slope residuals are normally distributed with $\sigma_{u0}^2$ the variance between intercepts, $\sigma_{u1}^2$ the variance between slopes, and $\sigma_{u01}$ the covariance between intercepts and slopes. Because the random slope on the house move variable allows the change in SDQ to differ for individuals in years in which they move it permits an examination of whether children exhibit greater heterogeneity in SDQ in years when they move home than in years when they are residentially stable.

DAWBA responses are fitted using multilevel ordered logistic regression, specified as:

$$
\logit\{Pr(y_{ij} \leq s|\beta, \sigma^2)\} = \log\left\{ \frac{Pr(y_{ij} > s)}{1 - Pr(y_{ij} > s)} \right\} = \beta_1 x_{ij} + \beta_2 y_{ij} + \beta_3 x_{ij} + \epsilon_{ij},
$$

$$
s = 1, 2, 3, 4, 5, \mu_y - N(0, \sigma_{\epsilon}^2)
$$

where $y_{ij} = s$ denotes the ordinal response for occasion $i$ of individual $j$, $s = 1, 2, 3, 4, 5$ denotes the five DAWBA bands of decreasing likelihood of psychiatric disorder diagnosis, $\beta_1$ denotes the coefficient of a time variant variable $x_{ij}$ (i.e. moving), $\beta_2$ denotes the coefficient of a time

<table>
<thead>
<tr>
<th>Occasion</th>
<th>Sample size</th>
<th>Age in years mean (SD)</th>
<th>Moved n (%)</th>
<th>DAWBA band</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>5255</td>
<td>7.46 (0.14)</td>
<td>1255 (23.88)</td>
<td>&gt; 70%</td>
</tr>
<tr>
<td>2</td>
<td>4024</td>
<td>10.59 (0.21)</td>
<td>1144 (28.43)</td>
<td>~ 50%</td>
</tr>
<tr>
<td>3</td>
<td>2545</td>
<td>13.80 (0.17)</td>
<td>367 (14.42)</td>
<td>~ 15%</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>39 (1.53)</td>
<td>~ 3%</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>61 (2.4)</td>
<td>~ 0.5%</td>
</tr>
</tbody>
</table>

Table 2
Occasion level information for DAWBA sample.

<table>
<thead>
<tr>
<th>Occasion</th>
<th>Sample size</th>
<th>Age in years mean (SD)</th>
<th>Moved n (%)</th>
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</tr>
</tbody>
</table>

Table 3
Sample characteristics for SDQ and DAWBA samples.
invariant variable $X_2$ (i.e. sex) that is held constant across the log-odds contrasts (that is, the ordered categories of DAWBA diagnosis) where variable $X_2$ satisfies the proportional odds assumption, $\beta_1$, denotes the coefficient of a time-invariant variable $X_2$ that is allowed to vary across the log-odds contrasts where variable $X_3$ fails to satisfy the proportional odds assumption, and $\kappa_i$ denotes the cut points $\kappa_{i-1}, \kappa_i,$ that is the constant associated with each DAWBA band. $u_i$ is a normally distributed person-level random effect.

We extend each of these models to utilise a within-between random effect multilevel modelling approach (Bell and Jones, 2015), a method gaining popularity in the analysis of residential and neighbourhood transitions (Morris et al., 2016a; Nieuwenhuis et al., 2016). In longitudinal model formulations such as those above the coefficient of interest $\beta_i$ represents the combined effect of two separate components or ‘effects’ of the variable $X_i$. The within-between random effect model allows these components to be estimated separately. In our case the first component compares the outcomes of one person who experienced a move to a different person who experienced no move. This is termed the ‘between person’ component as it estimates the difference in outcomes between different people. The second component compares the outcomes of a person who experienced a move at one time to themselves at a different time at which they experienced no move. This is termed the ‘within person’ component as it estimates the difference in outcomes within the same person at different time points. The difference between the between and within components is important. The between person component compares the outcomes of movers and stayers, who are likely to differ across a range of individual-specific factors (confounders), and so it can only be used to infer associational differences. The within-person component compares the outcomes of the same person at different occasions and so by using each person as their own control unit provides a less biased measure of how a change in exposure relates to a change in outcome. The within component therefore provides greater freedom to infer causal interpretations to associations (i.e. that moving causes subsequent changes in mental health).

All models are fitted in the MLwiN software using Markov Chain Monte Carlo (MCMC) methods (Browne, 2015) as implemented in the MLwiN software and called from within Stata v14 (StataCorp, College Station, TX, USA) using the runmlwin function (Leckie and Charlton, 2013). All results are presented as mean draws from parameter estimates with 95% Credible Intervals. We present results from three models: (1) results adjusted for covariates but with life events excluded to assess the impact of moving on mental health; (2) inclusion of life events to test for bias due to time-varying confounding; and (3) inclusion of life events to test for bias due to unobserved confounding caused by the exclusion of time varying life events.

3. Results – do movers and stayers differ?

3.1. SDQ

Results from the SDQ analysis are presented in Table 4. In the covariate adjusted model (Model 1) there is strong evidence that moving is associated with higher SDQ scores, but no evidence for a cumulative effect of moves. A random slope is included in model 1 on the move coefficient, and the negative intercept slope covariance provides evidence that there is heterogeneity in SDQ between movers and stayers; residentially stable children have more variable mental health than movers. Neighbourhood deprivation was unrelated to child SDQ scores while the coefficient for sex shows that boys had higher SDQ scores than girls, as did children with older mothers and those who had experienced greater financial difficulty. Children with highly educated parents had lower scores than those whose parents had lower levels of education scores, and a similar trend was found for parental occupational social class. Maternal self-reported depression and anxiety were strongly associated with increased child SDQ scores, although it is possible that a study mother’s responses on child mental health may in part reflect her own mental health. Children who had experienced parental separation prior to the analytical period had higher SDQ scores than those from stable two parent families, but interestingly there was no clear difference between children from stable single parent families and those from stable two parent families. There was no evidence for associations between SDQ and child age, ethnicity, housing tenure or paternal age at birth.

Model 2 decomposes the association between moving and SDQ scores into between and within parts to determine the extent to which the association observed in Model 1 is being driven by differences between individuals and/or differences within individuals caused by moving. The large between component suggests that the difference in SDQ scores between mobile and non-mobile children is being heavily driven by unobserved between-child differences. That is, the higher SDQ scores of mobile children are due in part to unmeasured factors; mobile children have poorer mental health regardless of the fact that they have moved. Turning to the within component of Model 2, there is a positive association between moving and SDQ score, suggesting that a given child moving leads to an increase in subsequent SDQ score. The coefficient is smaller than that for the between component, implying that unobserved between-child factors provide a greater contribution to the higher SDQ scores of mobile children than the experience of making a residential move. However, the fact that there is a positive within-child component indicates that moving may lead to poorer child mental health. The random slope is now included for the within-individual move coefficient; the negative intercept slope covariance suggests that children are more similar in terms of SDQ in periods in which they move than in periods in which they are stable. That is, there is greater heterogeneity amongst non-movers than movers. Because the slope is based upon the within-child effect of moving which is itself associated with higher SDQ, the results suggest that children with lower SDQ scores in stable periods experience disproportionately worse mental ill health because of moving than children with higher SDQ scores. That is, moving is more damaging for healthier children with fewer behavioural difficulties than those whom already have greater behavioural difficulties. Failing to include a random slope on the within-individual move variable would force the incorrect assumption that the effect of moving on SDQ is uniform and that all children respond to moving in the same way.

Comparison of movers and stayers (Supplementary Table S3) shows that in the SDQ sample children who experience major life events are more likely to move than children who experience no event. Clearly, excluding life events from Model 2 may have led to a biased within-child component due to time varying confounding by life events. Model 3 includes life event data to assess this confounding. Remarkably, the inclusion of major life events did not attenuate the
within-child effect of moving on SDQ scores, providing evidence that the within-child moving association in Model 2 was not being driven (confounded) by the occurrence of life events. The fact that there is a positive within-child component even after consideration of major life events permits greater freedom for drawing causal inferences from the data and suggests that moving may have a subsequent negative effect on child mental health. The negative intercept slope covariance remains, indicating that the longitudinal association between moving and mental health depends in part on mental health ‘pre-move’. Considering the effects of major life events, the only robust association is reported between sibling birth and increased SDQ scores.

3.2. DAWBA

Table 5 displays the results from the DAWBA analysis where the coefficients represent the log-odds of being in DAWBA band or lower (i.e. having better mental health). The "/cut1 constant" coefficients represent the log-odds of being in each decreasing DAWBA band when all else is zero, for example "/cut1 constant" refers to the log-odds of being in the second DAWBA band or lower compared to the highest band. Proportional odds assumptions were satisfied for all variables except for sex and financial difficulties, the effects of which were allowed to vary in strength across the across the log-odds contrasts (the DAWBA bands). The "/cut male" and "/cut financial difficulty" coefficients therefore represent the (non-proportional) log-odds of being male and increased financial difficulty respectively across each of the DAWBA categories. For example, the coefficient "/cut1 male" refers to the log-odds of being in the second DAWBA band or lower compared to the highest band for males compared to females. Like the SDQ results there is an association between moving and poorer mental health diagnosis was consistent across the log-odds contrasts. Again, no associations were observed between mental health and either cumulative or current neighbourhood deprivation. Age was associated with poorer mental health, while children from separated families were more likely to be in higher DAWBA bands than those from stable two parent families. There was evidence that children with
lower social class parents had poorer mental health, although those with degree educated parents had poorer levels of mental health. No associations were observed between DAWBA categorisation and ethnicity, maternal age, paternal age or housing tenure. Boys were less likely to be classified in a lower DAWBA band than girls, as were children who had experienced a greater number of financial difficulties in early childhood. These associations showed a non-proportional trend by which coefficients were larger in each successive band.

The results from Model 2 show that most the association between moving and poorer mental health is again due to unobserved between-child differences. The between-child component shows that there is strong evidence that mobile children have poorer underlying levels of mental health than non-mobile children. The point estimate for the within-individual moving coefficient is negative but Credible Intervals are wide, providing only suggestive evidence that moving may lead to subsequently poorer mental health as measured by DAWBA. There was a positive coefficient for the between-child IMD component suggesting that the weak within-child association observed in Model 2 may be attenuated by exposure to moving and other life events. The inclusion of life events attenuates the within-child moving association and reduces statistical support, suggesting that the weak within-child association observed in Model 2 may be confounded by major life events. Regarding life events themselves, robust associations with mental health are observed only for sibling family death and maternal job loss.

<table>
<thead>
<tr>
<th>Table 5</th>
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</thead>
<tbody>
<tr>
<td>Regression results from DAWBA sample.</td>
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<tr>
<td></td>
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<tr>
<td></td>
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<tr>
<td>Residential move</td>
</tr>
<tr>
<td>RM between component</td>
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<tr>
<td>IMD between component</td>
</tr>
<tr>
<td>IMD within component</td>
</tr>
<tr>
<td>Cumulative moves</td>
</tr>
<tr>
<td>Separation</td>
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<tr>
<td>Divorce</td>
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<tr>
<td>Marriage</td>
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<tr>
<td>Sibling birth</td>
</tr>
<tr>
<td>Family death</td>
</tr>
<tr>
<td>Maternal job loss</td>
</tr>
</tbody>
</table>

RM, Residential move; IMD, Index of multiple deprivation; Q1, Quintile 1; VPC, Variance partition coefficient; DIC, Bayesian deviance information criterion.
birth and maternal job loss, both of which are associated with poorer mental health. The coefficient for family death is larger than for any other life event but Credible Intervals are wide, reflecting the small number of cases.

3.3. Is the effect of moving uniform?

One assumption of our analyses thus far is that movers and stayers are homogenous groups with one to one associations with mental health, that is, staying is a positive experience and moving is a negative experience for all children. Previously, the literature has treated this assumption as intuitive, however there is evidence that the effects of moving can vary depending on whether a move (or lack of move) is desired or undesired (Woodhead et al., 2015). We were unable to formally test for effect moderation of moving or staying by preference due to a lack of available data on family moving preferences. However, to examine whether the effects of moving are homogenous we created a quasi-move preference variable based upon neighbourhood deprivation transitions whereby moves were classified as preferred if they were to less deprived neighbourhoods, neutral if they were to similarly deprived neighbourhoods, and unwanted if they were to more deprived neighbourhoods. The results of analyses on the SDQ and DAWBA measures using the quasi-preference variable are displayed in Table 6 (full model results in Supplementary Table S4 and S6). The results suggest that the effect of moving on subsequent mental health is non-uniform. On both the SDQ and DAWBA measures, there was only suggestive evidence that children who made ‘preferred’ moves experienced poorer subsequent mental health than children who did not move, with Credible Intervals covering zero. However, on both measures children who made ‘neutral’ moves had poorer subsequent mental health compared to children who did not move, although the statistical support for this association on the DAWBA measure was reduced when life events were accounted for. The point estimate in the SDQ sample for ‘unwanted’ moves was higher than for the other preference groups suggesting that unwanted moves may have the greatest impact upon subsequent mental health compared to other types of moves, but the wide credible intervals suggest that this could be statistical noise due to low numbers (Supplementary Table S5). DAWBA numbers for this group were lower still (Supplementary Table S7) resulting in a noisy estimate with wide credible intervals. These results suggest that moving is not a uniform experience for children, but may depend in part on the nature of moves. If we make the assumption that moves to less deprived neighbourhoods are preferred moves (a fair assumption given the significant economic cost that is associated with moving to more affluent areas) and that those to more deprived areas are unwanted moves (a reasonable assumption based upon the desire of people to make moves up the neighbourhood ladder (Bolt et al., 2009)), then our quasi-move preference variable can be considered to capture some underlying preference towards moving. It is however unlikely that these results point to neighbourhoods or ‘place’ as an effect modifier given the lack of deprivation effects observed in the main analyses over and above the effect of moving, but they do highlight the presence of heterogeneity within movers and imply that a simple dichotomy of movers vs stayers is overly simplistic.

<table>
<thead>
<tr>
<th>SDQ</th>
<th>Model 1: Covariate adjusted</th>
<th>95% CI</th>
<th>Model 2: Life event adjusted</th>
<th>95% CI</th>
</tr>
</thead>
<tbody>
<tr>
<td>No move</td>
<td>Ref. Cat</td>
<td>0.015</td>
<td>−0.009 to 0.040</td>
<td>0.014</td>
</tr>
<tr>
<td>Quasi-preferred move</td>
<td>0.034</td>
<td>0.008 to 0.059</td>
<td>0.032</td>
<td>0.007 to 0.058</td>
</tr>
<tr>
<td>Quasi-neutral move</td>
<td>0.037</td>
<td>−0.007 to 0.078</td>
<td>0.036</td>
<td>−0.008 to 0.080</td>
</tr>
<tr>
<td>Quasi-unwanted move</td>
<td>−0.188</td>
<td>−0.376 to 0.006</td>
<td>−0.161</td>
<td>−0.354 to 0.041</td>
</tr>
<tr>
<td>DAWBA</td>
<td>No move</td>
<td>Ref. Cat</td>
<td>−0.247</td>
<td>−0.465 to −0.016</td>
</tr>
<tr>
<td>Quasi-preferred move</td>
<td>0.091</td>
<td>−0.351 to 0.546</td>
<td>0.133</td>
<td>−0.314 to 0.586</td>
</tr>
</tbody>
</table>

Ref. Cat, Reference category.

4. Discussion

Our results shed new light on the relationship between moving and mental health and make several contributions to the literature. First, decomposing the association between moving and poor mental health allows us to demonstrate for the first time the relative similarities and differences of movers and stayers. Our results show that most the association between moving and mental ill health is driven by unobserved between-individual factors; that is, that movers and stayers are fundamentally different types of people with regards to mental health. These unobserved between-individual factors account for most of the difference in mental health between mobile and non-mobile children. The implication is that mobile children as a group have a greater underlying propensity towards poor mental health than non-mobile children. These unobserved between-individual differences remained even in the presence of a wider range of variables than has been considered in many previous studies. Given that we control for many family level socio-economic and demographic characteristics, it is possible that these unobserved differences may relate to factors such as child personality, genetics or family level wellbeing. Second, the decomposition provides to our knowledge the first time the relative similarities and differences of movers and stayers given the lack of deprivation effects of moving, or staying, for this association on the DAWBA measure was reduced when life events were accounted for. The point estimate in the SDQ sample for ‘unwanted’ moves was higher than for the other preference groups suggesting that unwanted moves may have the greatest impact upon subsequent mental health compared to other types of moves, but the wide credible intervals suggest that this could be statistical noise due to low numbers (Supplementary Table S5). DAWBA numbers for this group were lower still (Supplementary Table S7) resulting in a noisy estimate with wide credible intervals. These results suggest that moving is not a uniform experience for children, but may depend in part on the nature of moves. If we make the assumption that moves to less deprived neighbourhoods are preferred moves (a fair assumption given the significant economic cost that is associated with moving to more affluent areas) and that those to more deprived areas are unwanted moves (a reasonable assumption based upon the desire of people to make moves up the neighbourhood ladder (Bolt et al., 2009)), then our quasi-move preference variable can be considered to capture some underlying preference towards moving. It is however unlikely that these results point to neighbourhoods or ‘place’ as an effect modifier given the lack of deprivation effects observed in the main analyses over and above the effect of moving, but they do highlight the presence of heterogeneity within movers and imply that a simple dichotomy of movers vs stayers is overly simplistic.
amongst movers.

Interestingly, we did not observe any within-individual associations between neighbourhood deprivation and mental health, consistent with other studies that have utilised similar designs (Jokela, 2015, 2014) or examined the effect of deprivation over and above the impact of making a residential move (Gambino and Joshi, 2016). This suggests that it may be the move itself rather than changes to neighbourhood context that matter for child mental health when moving through neighbourhoods. This finding is in contrast to some studies which have suggested moving to more deprived neighbourhoods is associated with poorer mental health (Leventhal and Brooks-Gunn, 2003; Ludwig et al., 2012). There are various possibilities for these disparate findings; it may be that the associations observed in these studies have been driven more by the act of moving than the specific deprivation trajectories; they may have suffered from unobserved confounding that we were able to avoid by using rich cohort data and an appropriate analytical approach; or it could be due to systematic differences in the samples. Regardless of the precise mechanism that is driving these differences, our findings highlight the importance for future studies to determine the extent to which any observed associations between deprivation transitions and health capture the effects of moving rather than effects of changing neighbourhoods. The results of our quasi-preference analysis suggest that moving may not be a uniform experience for all children, conforming to findings from other UK studies (Turnstall et al., 2014, 2012). However, statistical support for differences between our quasi-preference groups was weak and future studies with larger sample sizes may be able to explore this area further.

The major limitation of this work relates to the accuracy of the SDQ and DAWBA measures, which do not provide perfect measures of child psychiatric illness compared to clinician diagnoses. However, such diagnoses are unavailable to our study, and there is a large body of work validating the SDQ and DAWBA measures as diagnostic tools in population data (Goodman et al., 2000a, 2006b). Another limitation is that we were unable to decompose movers and stayers into groups defined by moving/staying preference. However, this assumption is more likely to be satisfied in studies such as ours that involve children instead of adults because young children are not involved in the moving decision making process and do not have the vested beneficial interests that adults do when considering a residential move. Our analyses separating moves into positive and negative experiences based upon neighbourhood deprivation transitions suggest that failure to decompose movers may ignore heterogeneity within this group, although this analysis did not directly assess moving preference. A further limitation is the possibility that the within-child associations were driven by school changes (Winsper et al., 2016); residential moves may be accompanied by non-compulsory school changes which could further increase the environmental change that children must adapt to and the erosion of child social networks.

In conclusion, our findings suggest that moving in childhood is associated with subsequent behavioural difficulties. Decomposing this association into its constituent between and within individual components shows that while most of the association is due to unobserved differences between ‘movers’ and ‘stayers’, there is evidence that moving house may have an independent, albeit smaller detrimental impact upon subsequent child mental health. This finding holds for behavioural but not psychiatric measures of mental health, suggesting that the relationship between moving and poor mental health may be dependent on the measure of mental health under examination.

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Appendix A. Supporting information

Supplementary data associated with this article can be found in the online version at doi:10.1016/j.healthplace.2017.06.004.

References


