
Volume 70
Issue 12
June 2010
ISSN 0277-9536
ELSEVIER
SOCIAL
SCIENCE
&
MEDICINE
an international journal
MOST CITED
Impact Factor
2.604
SOCIAL SCIENCE JOURNAL*

Social Science & Medicine 70 (2010) 2061–2069

Contents lists available at ScienceDirect

Social Science & Medicine

journal homepage: www.elsevier.com/locate/socscimed

Short report

Siblings and childhood mental health: Evidence for a later-born advantage

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A R T I C L E I N F O

Article history:
Available online 23 March 2010

Keywords:
Parental investment
Sibling competition
Birth order
Childhood mental health
Strengths and Difficulties Questionnaire
Avon Longitudinal Study of Parents and Children (ALSPAC)
UK

A B S T R A C T

The social and health sciences have often emphasised the negative impacts of large sibship size and late birth order on childhood. For example, it is now well established that, other things being equal, children in large families and/or with many older siblings, receive lower allocations of care time from both parents, are more likely to grow up in conditions of economic hardship, and, as a likely consequence, exhibit relatively poor educational and physical health outcomes. Few researchers have, however, quantitatively assessed how siblings may influence indicators of mental health, where it is conceivable that social interactions with siblings may have a positive influence. Here, using data from a large British cohort survey (the Avon Longitudinal Study of Parents and Children), we explored the effects of sibling configuration on the Strengths and Difficulties Questionnaire, as a multidimensional index for mental health problems. We demonstrate a significant socio-economic gradient in mental health between the ages of three and nine years, but little evidence for negative effects of large sibship size. Rerunning this analysis to examine birth order, a much clearer pattern emerged; the presence of older siblings was associated with relatively good mental health, while the presence of younger siblings was associated with relatively poor mental health. This suggests that being born into a large family, providing the child is not joined by subsequent siblings, may carry important benefits unconsidered by past research. We discuss possible interpretations of this pattern and the wider implications for understanding the family context of child development.

Introduction

Siblings, as far as most research in the social and health sciences is concerned, are bad news. Lawson and Mace (2009) for example, in a recent study of contemporary British families, demonstrate that sibship size has a strong negative influence on both maternal and paternal time allocation to child care activities over the first decade of life; number of siblings had a larger influence on this measure of parental investment than any other covariate considered, including socio-economic indicators and parental age (see also: Blake, 1989; Downey, 1995; Hill & Stafford, 1974, 1980). In the struggle, to feed, clothe and house more children, parents in large family households also report increased levels of economic hardship, even after adjustment for a range of factors including differences in income, education and ethnicity (Iacovou & Berthoud, 2006; Lawson & Mace, in press).

As a likely consequence of these deficits, children with many siblings perform worse on IQ tests and on formal educational assessments throughout life, a pattern recognised as one of the most stable relationships in the study of education (Blake, 1989; Downey, 1995, 2001; Lawson, 2009; Steelman, Powell, Werum, & Carter, 2002; Zajonc, 1976). Number of siblings also has an important negative effect on achieved socio-economic status in adulthood, particularly on wealth ownership (Kaplan, Lancaster, Bock, & Johnson, 1995; Keister, 2003, 2004). Keister (2003) for instance, demonstrates that number of siblings is a strong determinant of the likelihood of receiving a trust fund or an inheritance (see also: Cooney & Uhlenberg, 1992). Finally, siblings are also associated with deficits in childhood growth and achieved adult height, which may stem from reduced parental attention or early-life nutrition (Lawson & Mace, 2008; Li, Manor, & Power, 2004; Li & Power, 2004).

In most cases, later-born children are at the biggest disadvantage in terms of both the division of parental investment (Lawson & Mace, 2009; Price, 2008) and relatively poor educational and physical health outcomes (Kristensen & Bjerkedal, 2007; Lawson, 2009; Lawson & Mace, 2008; see also: Modin, 2002). This pattern may be explained by the simple fact that older siblings, being alive both before and after a child's birth, have an increased potential to dilute parental resources (Downey, 2001). It is also possible that parents systematically bias care by order of birth, reflecting cultural

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doi:10.1016/j.socscimed.2010.03.009

preferences for first and early-born children (for discussion on the ultimate origins of such preferences see: Hrdy & Judge, 1993; Jeon, 2008; Rosenblatt & Skoogberg, 1974). Few studies have investigated the role of siblings in mental health.

Siblings and childhood mental health

The existing mental health literature has rarely been directed by the resource dilution and life history perspectives on the family which emphasise investment competition between siblings and the resulting trade-offs between quantity and 'quality' of offspring (Becker & Lewis, 1973; Downey, 2001; Lawson, in press; Mace, 2007). Nevertheless, like physical and educational development, measures of mental health follow a socio-economic gradient; with a lower incidence of behavioural problems in children from high socio-economic status backgrounds (Dunn, Deater-Deckard, Pickering, O'Connor, & Golding, 1998; Ford, Goodman, & Meltzer, 2004; Green, McGinnity, Meltzer, Ford, & Goodman, 2005; McMunn, Nazroo, Marmot, Boreham, & Goodman, 2001). Parenting style and quality are also assumed to be important in limiting behavioural problems (e.g., Dunn et al., 1998). As such, the dilution of material and interpersonal investments associated with large sibship size and late birth order can be expected to lead to negative consequences for childhood mental health.

To date, the best available data come from two large national samples of UK families (Green et al., 2005; Meltzer, Gatward, Goodman, & Ford, 2000). The results of these analyses, not specifically designed to study the effects of sibling configuration, are difficult to interpret in the face of inconsistent conclusions and a generally poor regard for potential confounds. Meltzer et al. (2000) report that large sibship size is associated with increased prevalence of childhood mental health problems. This effect was largely driven by an increase in conduct disorders, with no significant relationship detected with emotional or hyperactivity problems in multivariate models. However, in a reanalysis of these data, adjusting for a wider range of covariates, Ford et al. (2004) reported no independent effects of sibship size (they also provided a wider discussion of the problem of highly interrelated risk factors for childhood psychological morbidity ignored in many early studies). Green et al. (2005) report that large sibship size was not associated with the overall prevalence of mental disorders, but was associated with increased conduct and emotional problems. However, effect estimates were not adjusted for related socio-economic and demographic factors. Green et al. (2005) also considered autistic spectrum disorders, with no effect of family size detected.

Using a distinct measure of peer-related mental health, Downey and Condron (2004), found that children in multiple child families were scored as having better social skills than only children in an American sample. This study, based on teacher ratings of child behaviour, adjusted the effects of family size for a range of socio-economic factors. Several studies specifically considering the development of 'theory of mind' have also reported that children in multiple child families tend to perform better for their age on theory of mind tasks (Peterson, 2000).

None of the main childhood mental health surveys have tested for the existence of birth order effects (Ford et al., 2004; Green et al., 2005; Meltzer et al., 2000), while Downey and Condron (2004) reported no difference in the effects of older and younger siblings on social skills. Gates, Lineberger, Crockett, and Hubbard (1986) reported a higher incidence of depression, anxiety and low self-concept in later-born children, but their findings can only be considered suggestive in the absence of multivariate analysis.

Here, we present new data on the influence of siblings on a multidimensional index for childhood mental health problems – the Strengths and Difficulties Questionnaire (Goodman, 1997, 2001). All data were sourced from the Avon Longitudinal Study of Parents and Children (ALSPAC), a large British cohort. Unlike most past studies we consider both the effects of sibship size and birth order simultaneously. Furthermore, replicating the methodology of our past research into related family structure effects on parental investment and child development in the ALSPAC sample (Lawson & Mace, 2008, 2009, in press), we used detailed longitudinal data to estimate relationships net of an unusually large range of important covariates.

Data and methods

The Avon Longitudinal Study of Parents and Children (ALSPAC)

ALSPAC is a uniquely detailed, ongoing cohort study designed to examine environmental and genetic influences on the health and development of British children. Study recruitment began in pregnancy, enrolling women who had an expected delivery date between April 1991 and December 1992 from the three main Bristol-based health districts of the former county of Avon. 14,472 pregnant women (14,062 live births) were recruited into the initial sample. Avon has a predominantly white population, and a mixture of rural and urban communities encompassing a broad socio-economic range (Golding, Pembrey, Jones, & the ALSPAC Study Team, 2001). A major advantage of ALSPAC is the exceptional frequency of data collection. Mothers complete up to three postal surveys a year, one relating to the characteristics of herself and the household in general and two relating to the child. The ALSPAC survey also includes data from other surveys including extraction from clinical records and school-based assessments and direct examination of children at specifically designed research clinics.

The analyses presented in this paper are based on available data from the first 10 years of data collection. All data considered were collected by self-completed questionnaires. Further information on the distribution of each independent variable over the study period and descriptive statistics at each wave can be found in Lawson and Mace (2008). Further information on data collection methodology can be found in (Golding et al., 2001). We refer readers to these publications for supplementary information on the cohort.

A number of exclusion criteria remove rare family configurations from our sample. Families where the study child is from a multiple birth (i.e., a twin or triplet), families recorded as experiencing the death of a child and families containing children unrelated to either the mother or her current partner (e.g., foster or adopted children) over the study period were all excluded. Cases where the study child's live-in 'mother figure' was ever recorded as other than the biological mother, as absent or in a same-sex relationship were also excluded. Cases of biological father absence after birth were included. We also included cases where the mother was recorded as in a new relationship with someone other than the biological father. However, we excluded rare cases where the mother reported unsure paternity of the study child or started a new relationship during this pregnancy. After implementing these criteria the key study sample contained 13,176 different families each containing a single study child.

Mental health: the Strengths and Difficulties Questionnaire

Assessments of mental health are based on the Strengths and Difficulties Questionnaire (SDQ), a widely used instrument for assessing psychological morbidity in children (Goodman 1997, 2001). The SDQ measures four domains of poor mental health status, on separate scales with five items each: emotional problems, hyperactivity, conduct problems and peer problems (Goodman, 1997, 2001). Responses to questions from the emotional problems, conduct problems, hyperactivity and peer problems subscales are added to give a total difficulties score (referred to as the 'TDS' or simply 'total score'), with a range of 0–40. This can be used as a dimensional outcome measure of mental health problems (Goodman & Goodman, 2009). The SDQ was assessed by parents at three points over the study period available at three years eleven months, six years nine months and nine years (these ages represent the target age of each questionnaire, with some unavoidable variation around the mean). Table 1 provides descriptive data on the SDQ scores along each subscale. For the total score there were 23,991 cases available for analysis on 9826 individual children.

Sibship size and birth order

Siblings were defined for the purpose of this study as all maternally related siblings resident with the study child (i.e., including siblings with different biological fathers, but excluding children from different mothers). Some mothers recorded additional children related only to their current partner (7% at the end of the study period), but in only in a very small percentage of families were such children coresident (1%).

Sibship size and birth order data (number of younger siblings and number of older siblings) were available in six questionnaires from the birth of the study child to 10 years of age (Lawson & Mace 2008). The final three points of data collection corresponded closely to the age at which mental health assessments were made (i.e., three years eleven months, seven years one month and 10 years). Thus, while number of older siblings is fixed at birth, number of younger siblings and total sibship size are entered into each analysis as time-varying determinants of mental health. Our analyses, therefore, considered sibling configuration as a dynamic rather than fixed state.

Half (51%) of the study children were first-borns, around a third (33%) were second-borns, and a significant number (16%) had two or more older siblings (i.e., third-born or later). A majority of children (53%) experienced the arrival of at least one younger sibling before the first assessment of mental health. By age 10, 43% of child had one younger sibling and 16% had two or more. At all points of data collection subsequent to the birth of the study child, modal sibship size was two (i.e., the child had one sibling). At 10 years, 54% were in sibships of two children, 27% in sibships of three children, 8% in sibships of four children and 2.4% in sibships of five or more children.

Covariates

We include mother's educational attainment (coded in pregnancy) as a time invariant measure of socio-economic position. In addition we use three time-varying measures – self-reported 'take home' household income, home ownership and neighbourhood quality. Two measures of social support were also incorporated, both based on questionnaires distributed to the mother in pregnancy. These measures were recorded only once and so could not be entered as time-varying variables. The social network score comprises ten items which ascertain the quality and frequency of social contact with friends and family and ranges from 0 to 30. The social support score measures perceived social support from family, friends and official agencies using a set of 10 items specifically designed for the cohort. The item presents statements relating to emotional, financial and instrumental support, with a summed overall score also ranging between 0 and 30. This measure showed a strong association with the mother's emotional well-being during pregnancy (Thorpe, Dragonas, & Golding, 1992). Both measures were banded into three groups of equal size, coded as 'low', 'medium' and 'high'. We also incorporated a banded measure of maternal emotional problems assessed in pregnancy (the Edinburgh Postnatal Depression Score: Cox, Holden, & Sagovsky, 1987). Mother's employment status and ethnicity were included as an additional dichotomous covariate terms in all models. Mother's employment status was measured at three intervals over the study period, with between 55 and 72% of mothers engaged in employment between a study child's age of three years eleven months and 10 years, respectively. A large majority (95%) of the ALSPAC population was recorded as white.

Fathers were coded as present provided the mother stated the study child had a biological father as the live-in 'father figure' at the time of the questionnaire. In cases where the father was coded as absent, the mothers were either coded as alone or with a new live-in partner. These data did not distinguish between different partners of the mother subsequent to the biological father of the study child. Father presence was assessed at the same intervals as sibship size data, with almost a quarter of mothers (24%) known to be separated from the father of the study child at 10 years (composed of 7% with new partners, 10% remaining single and 7% of unknown relationship status). A majority of parents were aged between 25 and 29 years at the birth of their study child, with a mean maternal age of 28.0 years (SD: 5.0) and mean paternal age of 30.7 (SD: 5.7).

Table 1
Strengths and difficulties score.

	Child age		
	3y11m	6y9m	9y0m
	$n = 8900$	$n = 7891$	$n = 7295$
Total difficulties score (TDS)	Mean (standard deviation)		
	8.85 (4.54)	7.45 (4.74)	6.79 (4.90)
Components			
Hyperactivity score	3.95 (2.30)	3.38 (2.36)	2.94 (2.25)
Emotional score	1.44 (1.50)	1.50 (1.67)	1.50 (1.76)
Conduct score	1.95 (1.40)	1.60 (1.46)	1.27 (1.42)
Peer score	1.51 (1.48)	1.05 (1.41)	1.11 (1.49)

Note that these values refer to the sample available at each study wave. They should not be directly interpreted as evidence of change over time due to selective attrition. Total *N*: TDS – 23,991 for 9826 individuals; hyperactivity score – 24,019 for 9826 individuals; emotional score – 24,020 for 9828 individuals; conduct score – 24,046 for 9829 individuals; peer score – 24,028 for 9829 individuals.

Analysis strategy

Relationships between sibship size, birth order and repeated assessments of childhood mental health were assessed using multi-level models for change over time (Singer & Willett, 2003). These models can be used to estimate multivariate relationships between time-varying categorical or continuous independent variables and a continuous dependent variable over time. Individual children are treated as level-two units and the timing of measures as level-one units. The major advantage of a multi-level modelling strategy is that it enables incorporation of all available outcome data, rather than restricting analysis to individuals with complete assessments at a specific subset of time points. In order to have unbiased estimates in the presence of missing data, it must be assumed that responses are missing at random; that is, the probability of any outcome measure being missed may depend on observed, but not unobserved, measures (Little & Rubin, 1987). Although we do not formally investigate this issue, given the large range of relevant independent variables considered in each model, it is likely that presented analyses conform to the missing at random assumption.

Modelling data in this also way requires contemporaneous data on independent and dependent variables, a feature not strictly met by the temporal distribution of variables considered in the presented analyses (e.g., mental health assessments made at 3 years 11 months, 6 years 9 months and 9 years, while household income was assessed twice at 3 years 11 months and 7 years 1 month). To overcome this issue we used the closest available measure of time-varying independent variables to the months when outcome data were recorded. This serves as a reasonable approximation given the small gaps in convergence between measures (particularly for sibling configuration and mental health data), and the relatively short total analysis period (just over five years).

Our analysis strategy followed three steps. First, for the total SDQ score and each of its subscale scores, we determined unconditional growth models which established the overall relationship of the outcome with time (age of the study child in years). This gave an impression of how assessed mental health varied as children grew older and allowed us to adjust for this pattern in later multivariate models. Secondly, we specified univariate associations between each independent variable and outcome variables to get a general sense of the relationships in the data. These univariate models only included adjustment for the relations between time and the outcome. For each independent variable, effects were estimated by both a main effect term (effect on 'initial status', i.e., point of first measurement) and an interaction term with time (effect on rate of change per year). Statistical significance of each predictor term was assessed (as in standard linear regression) by dividing the regression coefficient by its standard error and 95% confidence intervals were calculated. For the sake of brevity, these univariate associations are not presented here, but can be consulted in Lawson (2009).

Table 2
Final multivariate model: total difficulties score (TDS).

	Initial status (at 3y11m)		Rate of change (per year)	
	Coefficient (B)	95% CI	Coefficient (B)	95% CI
Intercept[a]	10.64***	10.12–11.16	−0.33***	−0.42 to −0.29
Family structure				
Sibship size (ref: 1)				
2	0.38**	0.11–0.65	−0.14**	−0.23 to −0.05
3	0.15 ns	−0.17 to 0.47	−0.11*	−0.21 to −0.01
4	0.07 ns	−0.39 to 0.53	−0.12 ns	−0.25 to 0.01
5+	0.15 ns	−0.55 to 0.85	−0.18 ns	−0.38 to 0.02
Sex of Child (ref: male)				
Female	−0.84***	−1.02 to −0.66	–	–
Mother's age (ref: <25)				
25–29	−0.40**	−0.68 to −0.12	–	–
30–34	−0.55***	−0.84 to −0.26	–	–
35+	−0.79***	−1.15 to −0.43	–	–
Father's age (ref: <25)				
25–29	–	–	–	–
30–34	–	–	–	–
35+	–	–	–	–
Father figure status (ref: biological father)				
Mother alone	0.29*	0.02–0.56	–	–
Unrelated male	−0.50**	0.15–0.85	–	–
Socio-economic measures				
Mother's education (ref: <O-level)				
O-level	−0.40**	−0.68 to −0.12	0.02 ns	−0.05 to 0.09
A-level	−0.84***	−1.14 to −0.54	0.05 ns	−0.03 to 0.13
Degree	−1.19***	−1.54 to −0.84	0.15***	0.07–0.23
Household income (ref: <£200 per week)				
£200–299	−0.07 ns	−0.31 to 0.17	–	–
£300–399	−0.38**	−0.64 to −0.12	–	–
£400+	−0.46***	−0.73 to −0.19	–	–
Neighbourhood (ref: <V. Good)				
V. Good	−0.37***	−0.50 to −0.24	–	–
Home ownership (ref: rented)				
Mortgaged/Buying	−0.60***	−0.89 to −0.31	–	–
Owned	−0.46*	−0.87 to −0.05	–	–
Social support				
Social network score (ref: low)				
Med	−0.34 ns	−0.69 to 0.01	–	–
High	−0.79***	−1.12 to −0.46	–	–
Social support score (ref: low)				
Med	−0.59***	−0.82 to −0.36	–	–
High	−1.17***	−1.46 to −0.88	–	–
Other				
Ethnicity of child (ref: white)				
Non-white	–	–	–	–
Maternal employment (ref: no)				
Yes	–	–	–	–
Maternal emotional problems (ref: low)				
Med	1.05***	0.84–1.26	–	–
High	2.35***	2.12–2.58	–	–

ns – Non-significant; $^{*}p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$.
Predictor terms which failed to reach statistical significance at any comparison are excluded from the final model (–).
Model fit (pseudo-R^2): within-Person (over time) – 0.29; initial status – 0.19; rate of change – 0.02.
Final N – 16,526.
[a] The estimated mean value for initial status and rate of change for the group with the baseline values for every factor included in the model.

Finally, multivariate models were constructed to assess the effects of family structure, net of influential covariates. These models were constructed in a stepwise fashion for the total score and each of its subscale scores (to allow for the possibility that different aspects of mental health are influenced independently by family structure). All variables relating to family structure (except relative sibling age) were entered in the initial block. This model was then reduced by a backwards procedure removing predictor terms that did not reach significance at the $p < 0.05$ level. All family structure variables maintained in the model at this stage were carried forward to a final presented model. The second block entered all remaining variables (social support, maternal emotional problems, etc). Predictor terms were maintained if $p < 0.05$ or their presence affected notable change on any of the family structure coefficients. In multi-level models for change, total outcome variation was partitioned into several within- and between-person variance components. For each of these components, a pseudo-R^2 statistic can be calculated based on the reduction of this term from 'unconditional models' containing only a constant and age terms (Singer & Willett, 2003). These pseudo-R^2 statistics are presented to estimate the fit of final models to the data. Final multivariate models were then also rerun replacing the independent variable of sibship size with number of younger and number of older siblings to assess the effects of birth order on mental health. All analyses were carried out using MLwiN 2.02 (Rasbash, Browne, Healy, Cameron, & Charlton, 2005).

Results

Mental health over the study period (unconditional growth models)

Unconditional growth models estimate overall relationships of each behavioural score with child age (linear functions are estimated only to keep models easy to compute and compare directly). For the total difficulties score, initial status (i.e., at 3 years 11 months) was estimated at 8.83 (CI: 8.74–8.92, $p < 0.001$) decreasing at −0.40 units per year (CI: −0.42 to −0.38, $p < 0.001$) indicating the prevalence of behavioural problems decreased as children aged (at least between the ages of 3 years and 9 years). This pattern was confirmed for all subscale scores: Hyperactivity Score – initial status 3.96 (CI: 3.91–4.01, $p < 0.001$), rate of change −0.19 (CI: −0.20 to −0.18, $p < 0.001$); Emotional Score – initial status 1.45 (CI: 1.42–1.48, $p < 0.001$), rate of change −0.01 (CI: −0.02 to 0.00, $p < 0.005$); Conduct Score – initial status 1.97 (CI: 1.94–2.00, $p < 0.001$), rate of change −0.13 (CI: −0.14 to −0.12, $p < 0.001$); Peer Score – initial status 1.46 (CI: 1.43–1.49, $p < 0.001$), rate of change −0.09 (CI: −0.10 to −0.08, $p < 0.001$).

A consideration of univariate associations revealed that independent variables were significantly associated with initial status more often than rate of change effects; in other words, relationships between each covariate and mental health rarely varied significantly by age of child. Socio-economic measures, measures of social support and maternal emotional problems were significant in every univariate model at high levels of significance. Family structure variables demonstrated a mixed pattern of association across the total and subscale scores (for details see Lawson, 2009).

Sibship size

Sibship size failed to show a consistent relationship with childhood mental health. Table 2 presents the final multivariate model for the total score, including effect estimates for sibship size and all covariates retained in the final model (excluded predictor terms failing to reach statistical significance at any comparison are indicated by a dash). Table 3 summarises the sibship size effects for each component score of the total score, adjusted for all retained covariates (full model outputs can be consulted in Lawson, 2009).

For the total score, significant main effects and rate of change effects were retained in the final model for some comparisons but these effects run in the opposite direction (Table 2). For example, initial status effects suggested a higher total score for children in

Table 3
Final mental health score models for sibship size (a) total difficulties score (b) hyperactivity score (c) emotional problems score (d) conduct problems score (e) peer problems score.

	Sibship size (ref: 1)	Initial status (at 3y11m)		Rate of change (per year)	
		Coefficient (B)	95% CI	Coefficient (B)	95% CI
(a) Total difficulties score	2	0.38**	0.11–0.65	−0.14 **	−0.23 to −0.05
	3	0.15 ns	−0.17 to 0.47	−0.11 *	−0.21 to −0.01
	4	0.07 ns	−0.39 to 0.53	−0.12 ns	−0.25 to 0.01
	5 +	0.15 ns	−0.55 to 0.85	−0.18 ns	−0.38 to 0.02
(b) Hyperactivity score	2	0.27***	0.14–0.40	−0.09 ***	−0.13 to −0.05
	3	0.15 ns	−0.01 to 0.31	−0.08 ***	−0.12 to −0.04
	4	0.06 ns	−0.16 to 0.28	−0.08 **	−0.14 to −0.02
	5 +	−0.05 ns	−0.39 to 0.29	−0.05 ns	−0.15 to −0.05
(c) Emotional problems score	2	0.17 ***	0.09–0.25	–	–
	3	0.10 *	0.01–0.19	–	–
	4	0.07 ns	−0.05 to 0.19	–	–
	5 +	0.13 ns	−0.06 to 0.32	–	–
(d) Conduct problems score	2	0.15 ***	0.08–0.22	–	–
	3	0.19 ***	0.11–0.27	–	–
	4	0.19 ***	0.08–0.30	–	–
	5 +	0.24 ***	0.08–0.40	–	–
(e) Peer problems score	2	−0.21 ***	−0.31 to −0.11	−0.04 **	−0.07 to −0.01
	3	−0.23 ***	−0.35 to −0.11	−0.04 **	−0.08 to 0.00
	4	−0.19 ns	−0.36 to −0.02	−0.04 ns	−0.09 to 0.01
	5 +	0.07 ns	−0.19 to 0.33	−0.16 ***	−0.24 to −0.08

Models contain control variables for additional aspects of family structure and parental resources (see Table 2 and Lawson, 2009).
ns – Non-significant; $^*p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$.
Predictor terms which failed to reach statistical significance at any comparison are excluded from the final model (–).
Final N: TDS – 16,526; Hyperactivity – 18,512; Emotional – 19.307; Conduct – 17,757; Peer – 15,066.

sibships of two compared to children with no siblings, but this effect was reversed by a negative rate of change effect within a few years. Children in sibships of four or more did not significantly differ from children with no siblings.

Similar mixed effects were found on the hyperactivity score (Table 3). Sibship sizes of two, and to a lesser extent three, were associated with more emotional problems compared to single child families. However, sibship sizes of four or five plus were not significantly different from only child families. Conduct problems followed a clearer trade-off pattern with incremental increases in sibship size associated with more problems. Peer problems followed the reverse pattern with increased sibship size associated with reduced problems, particularly in later childhood, as indicated by negative rate of change effects. Fig. 1 displays the mixed effects of sibship size graphically when, for simple comparison, only main effects are fit for the total score and all subscale scores. These estimates illustrate the mean association of sibship size and mental health scores across the study period from age three to nine (i.e., variation by age of child is not considered).

Birth order

Analysing the effects of siblings on mental health by sibling age rather than total sibship size provided a much clearer pattern of results across measures (Table 4). The presence of older siblings was associated with a reduction in problem scores; indicating clear evidence of a general later-born *advantage* in childhood mental health. In contrast, the presence of younger siblings was associated higher problem scores (driven principally by an increase in emotional and conduct difficulties). Thus, the benefits of having older siblings could be offset by the arrival of younger siblings. Fig. 2 provides a graphical illustration of the opposing effects of having older and younger siblings (main effects only). All reported effect estimates for number of older/younger siblings are mutually adjusted.

Significant rate of change effects indicated that some associations varied with the age of the child (Table 4). For example, increased emotional problems associated with the presence of younger siblings were most evident in young children (i.e., reduced over time by a negative rate of change effect). While for the conduct score, increased problems associated with the presence of younger siblings were more evident in later childhood.

Covariates

Consistent with prior research (Dunn et al., 1998; Ford et al., 2004; Green et al., 2005; McMunn et al., 2001), the total score showed a clear socio-economic gradient across the study period. Improved maternal education, household income, home ownership status and neighbourhood quality all demonstrated independent negative effects on the prevalence of mental health difficulties in childhood (Table 2). This pattern was shared by all subscale scores, albeit to varying degrees with, for example, emotional problems showing relatively weak socio-economic effects (see Lawson, 2009).

Higher levels of maternal social support and better social networks were also strongly associated with reduced mental health problems on all measures. Maternal employment failed to be retained in most final models. However, employed mothers reported that their children had slightly lower peer problems. Children of mothers recording high depressive symptomology in pregnancy had higher levels of difficulty on all measures (see also Dunn et al., 1998). These effects remained significant even after socio-economic and demographic variables had been taken into account and were consistently the strongest predictors of childhood mental health across all measures considered.

Children of older mothers had improved mental health on the total and subscale scores. Boys tended to have more mental health problems than girls across the study period, represented both in the total score and subscale scores for hyperactivity, conduct problems and peer problems. However, girls were scored as having higher levels of emotional problems than boys. In general, the absence of father figures and particularly the presence of unrelated father figures was associated with increased mental health problems relative to children with biological fathers recorded as present. However, for several of the SDQ subscale scores these effects failed to reach significance. The negative effects of unrelated father presence compared to biological father presence (see also McMunn et al., 2001) appeared to be driven largely by an increase in hyperactivity problems.

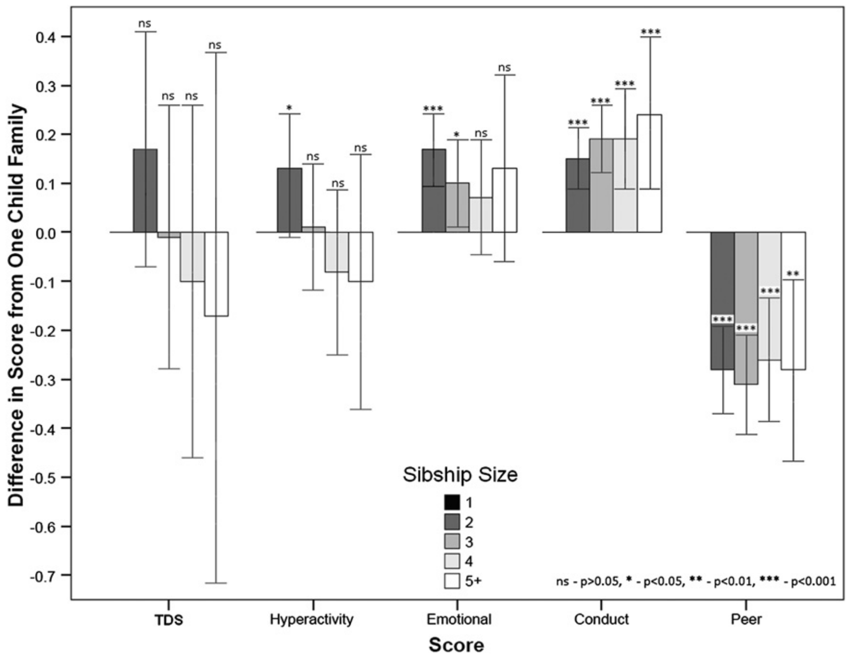

Fig. 1. Difference in mental health scores over the study period as a function of sibship size (main effects only – confidence intervals are set at 95%).

Table 4
Final mental health score models for sibling age configuration (a) total difficulties score (b) hyperactivity score (c) emotional problems score (d) conduct problems score (e) peer problems score.

			Initial status (at 3y11m)		Rate of change (per year)	
			Coefficient (B)	95% CI	Coefficient (B)	95% CI
(a) Total difficulties score	Number of older siblings (ref: 0)	1	−0.12 ns	−0.34 to 0.10	–	–
		2	−0.67***	−0.99 to −0.35	–	–
		3+	−1.00 ***	−1.57 to −0.43	–	–
	Number of younger siblings (ref: 0)	1	0.32***	0.13–0.51	–	–
		2	0.27 ns	−0.01 to 0.55	–	–
(b) Hyperactivity score	Number of older siblings (ref: 0)	1	0.25***	0.14–0.36	−0.09***	−0.12 to −0.06
		2	−0.19*	−0.36 to −0.02	−0.02 ns	−0.06 to 0.02
		3+	−0.22 ns	−0.51 to 0.07	−0.03 ns	−0.10 to 0.04
	Number of younger siblings (ref: 0)	1	–	–	–	–
		2+	–	–	–	–
(c) Emotional problems score	Number of older siblings (ref: 0)	1	−0.10**	−0.17 to −0.03	–	–
		2	−0.21***	−0.31 to −0.11	–	–
		3+	−0.25 **	−0.43 to −0.07	–	–
	Number of younger siblings (ref: 0)	1	0.27 ***	0.20–0.34	−0.03**	−0.05 to −0.01
		2+	0.32***	0.19–0.45	−0.05 **	−0.09 to −0.01
(d) Conduct problems score	Number of older siblings (ref: 0)	1	0.11***	0.04–0.18	–	–
		2	0.10 ns	0.00–0.20	–	–
		3+	−0.11 ns	−0.29 to 0.07	–	–
	Number of younger siblings (ref: 0)	1	0.10**	0.03–0.17	0.03 ***	0.01–0.05
		2+	0.16*	0.03–0.29	0.03 ns	0.00–0.06
(e) Peer problems score	Number of older siblings (ref: 0)	1	−0.22***	−0.29 to −0.15	–	–
		2+	−0.20***	−0.30 to −0.10	–	–
		3+	−0.13 ns	−0.32 to −0.06	–	–
	Number of younger siblings (ref: 0)	1	−0.12***	−0.19 to −0.05	–	–
		2+	−0.12 *	−0.22 to −0.02	–	–

Models contain control variables for additional aspects of family structure and parental resources (see Table 2 and Lawson, 2009).
ns – non-significant; $^{*}p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$.
Predictor terms which failed to reach statistical significance at any comparison are excluded from the final model (–).
Final *N*: TDS – 16,158; Hyperactivity – 18,702; Emotional – 15,536; Conduct – 15,536; Peer – 14,741.

Discussion

Siblings and childhood mental health

We found no consistent pattern between sibship size and childhood mental health (Fig. 1). The only evidence for predicted quantity–quality trade-off effects was on conduct problems (see also Meltzer et al., 2000). Peer problems were actually reduced in the presence of siblings. Other measures displayed mixed and largely non-significant associations with sibship size. This mix of patterns occurred despite the existence of strong socio-economic gradients in mental health for all measures, even though large families in the same cohort faced higher levels of economic hardship (Lawson & Mace in press). It is also generally considered that good parenting practice leads to positive child mental health outcomes (e.g., Dunn et al., 1998). So it is surprising that the presence of siblings did not reduce mental health through decreasing parental time investment (Lawson & Mace, 2009). These findings

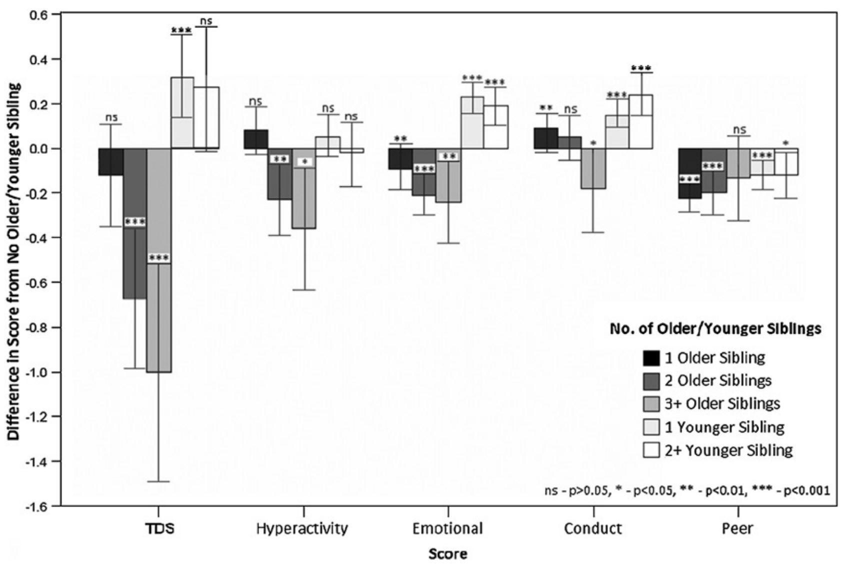

Fig. 2. Difference in mental health scores over the study period as a function of number of older/younger siblings (main effects only – confidence intervals are set at 95%).

suggest that siblings offset their negative effects on parental resource dilution by other means.

Interactions with siblings may provide some supportive or socialising benefit to children who otherwise may spend more time in isolation from their peers. Downey and Condron (2004) found that children in multiple child families had better social skills than only children, but found little difference between children within multiple child families by number of siblings. This pattern of results is very similar to the presented findings for peer problems in this study (Fig. 1). Thus, there is some evidence that exposure to siblings has a positive influence on social maturation. Perhaps, by becoming accustomed to sharing resources with at least one other child, children with siblings learn how to navigate social relationships more easily (Downey & Condron, 2004; see also Peterson, 2000).

Considering mental health problems by the relative age of siblings presents a much clearer pattern. Across all measures, with the exception of the conduct score, older siblings were associated with reduced mental health problems (Fig. 2). These results reverse the pattern of later-born disadvantage found consistently across previous analyses of the ALSPAC cohort on parental time investment (Lawson & Mace, 2009) and physical health and educational domains of child development (Lawson & Mace, 2008; Lawson, 2009). As far as we are aware, this study is the first to show a broad trend of *later-born advantage* in childhood mental health. Previous research has either not tested for birth order effects, or has done so only with very poor consideration of potential confounding factors.

With the lack of previous research on this topic, and in direct contraction to the expectations of the resource dilution and life history theoretical frameworks which have guided our previous studies of sibling effects, it is difficult to provide a definitive explanation for this result. Nevertheless, the current findings suggest that social interaction with older siblings may hold important mental health benefits over and above their negative effects on parental resource dilution. In a recent study using a small subset of ALSPAC cohort, Grass, Jenkins, and Dunn (2007) found that self-reported affectionate relationships between siblings had a protective effect on adjustment to stressful life events. Thus, older siblings may be more effective in providing a buffering role to social stress.

Alternatively, it is possible that the existence of older children may ensure that children are born into a household environment that is already socially and emotionally prepared for family life and so more conducive to positive mental health outcomes, even though time and money are in shorter supply. Consistent with this interpretation, transitions to parenthood are often associated with decreased marital satisfaction as both the new mother and father adjust and negotiate the responsibilities of raising children (reviewed in: Johns & Belsky, 2008). Later-born children are born at a time when such transitions have already taken place and stability may be improved.

In contrast, the presence of younger siblings was associated with increased difficulty scores, particularly for emotional and conduct problems (Fig. 2). This result is consistent with a number of studies noting the difficulty of adjusting to a new sibling (Baydar, Greek, & Brooks-Gunn, 1997; Dunn & Kendrick, 1980; but see Strohschein, Gauthier, Campbell, & Kleparchuk, 2008). It also gives some support for the Alder's theory of birth order in which early-born children are seen to suffer feelings of 'dethronement' with the arrival of younger siblings (see: Gates et al., 1986). The behavioural problems of these children are thus seen as a stress response to the sudden arrival of a competitor for parental investment.

Limitations

One limitation of the analyses presented in this manuscript is that mental health scores were based on parent ratings. Parent-rated measures of childhood mental health may be open to perception biases in the presence of other children in the household. Future research should, therefore, consider whether the reported birth order pattern holds up for alternative ratings of mental health, such as teacher-rated scores. It is not obvious, however, how parent ratings could create a perception bias of low incidence of mental health problems in later-borns. For example, it is conceivable that because children's mental health problems tend to decline with age, having older children in the house may bias the mother towards feeling her children in general have fewer problems. But the reduction of assessed mental health with age is quite modest for most measures, so such a perception bias would have to be very strong to account for the results reported here. Future research also needs to directly estimate the role of parenting quality and economic hardship in mediating family size and birth order effects on child mental health outcomes. The consequences of birth spacing, middle-born status and interactions between sibling configuration and mother-father relationships also remain unexplored by the current study. Many of these advancements could be achieved through future analysis of the ALSPAC cohort.

Conclusions

Research from across the social and health sciences has emphasised the negative impacts of large sibship size and late birth order on child development. Siblings are envisaged as competitors for parental investment which reduce individual allocations of parental attention and financial investment (Becker & Lewis, 1973; Downey, 2001; Lawson, in press; Mace, 2007). Here, using the same methodology and study sample used to confirm the existence of anticipated deficits in other dimensions of parenting and child development, we show that childhood mental health bucks the trend. There was no generalised relationship between sibship size and childhood mental health, with only conduct problems following the predicted trade-off pattern; increasing in incidence incrementally as sibship size increased. Moreover, in contrast to physical and educational development we found no evidence of a predicted later-born disadvantage; instead children born into large families were assessed as having better mental health than first-borns or children born into small families. Thus, it appears that alternative mechanisms to family resource division mediate associations between sibling configuration and childhood mental health. This result highlights the importance of considering multiple measures of child development and supports the largely folk hypothesis that siblings can play an important role in social maturation and emotional support.

Acknowledgements

We are extremely grateful to all the families who took part in this study, the midwives for their help in recruiting them, and the whole ALSPAC team, which includes interviewers, computer and laboratory technicians, clerical workers, research scientists, volunteers, managers, receptionists and nurses. We thank Jon Heron for his essential assistance with data management. The UK Medical Research Council, the Wellcome Trust and the University of Bristol provide core support for ALSPAC. This research was funded by the Economic and Social Research Council. We thank Anna Goodman and three anonymous reviewers for constructive comments on an early version of this manuscript.

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