Birth Order and Educational Achievement : an Investigation of African American Families using the PSID

Abstract

Whether siblings of specific birth order perform differently has been an open empirical question for decades. In this work, we focus on black families and use the family tree structure of the PSID to examine two claims found in the literature: (1) whether being early in the birth order implies a distinct educational advantage; (2) whether there exists, within large families, a pattern of falling then rising attainment with respect to birth order, to the point that it becomes best to be last-born. Drawing from OLS and family fixed effects estimations, we find that being first-born confers a significant advantage.

1. Introduction

Whether siblings of specific birth order perform differently has been an open empirical question for decades. Surprisingly, economists have not paid much attention to this issue, although it is fundamental to our understanding of the intra-household allocation of resources.[1]

In this study, we prolong an important contribution to this literature by Hanushek (1992). Using a sample of low income, black families from Indiana in the early 1970s (the Gary Income Maintenance Experiment[2]), Hanushek analyzed the trade-offs between child quantity and quality by comparing the scholarly performances of children from smaller families with those from larger families. He noted that family size is expected to affect each child differently, depending on the child's birth order. Therefore, the finer test of the quantity/education trade-off involves comparing academic achievements across children of different birth order. Hanushek's paper advances that parents show no favoritism to first-born children, and that while being early in the birth order implies a distinct advantage, it is entirely due to the higher probability of coming from a small family. Following Lindert (1977), the paper also highlights, within large (more than five siblings) families, a distinct and sizeable pattern of falling then rising attainment with respect to birth order, to the point that it becomes best to be last-born.

---

[1] Birdsall (1979), Behrman (1986), Behrman and Taubman (1986), Kessler (1991). We elaborate on a few other studies in more details in the paper.

[2] See Kelly and Singer (1971), Kehrer (1979) for details on this experiment.

The empirical results presented here, drawn from the Panel Study of Income Dynamics (PSID), first present a close version of Hanushek's findings before challenging their robustness. While we agree that birth order may not have a major impact on education within small black families, we argue that, within larger families, it matters a great deal even after controlling for number of siblings. Specifically, we observe that being first-born or even second born does confer an advantage in those families, while being among the last-born actually confers none. We believe that the reason for the discrepancy between Hanushek's findings and ours is that we control for the age of the mother at childbirth.

We find that the age of the mother at childbirth is positively correlated with a child's education. At the same time, the age of the mother at childbirth is mechanically, positively correlated with a child's birth order. Therefore, the two effects of birth order and the age of the mother at childbirth may compete against each other. We show that the omitted variable bias results in a clear offset of the birth order effect. This provides a new key element in our understanding of siblings' competition for parental inputs

Arguably, a causal interpretation of the previous analysis may suffer from the potential correlation of some of our regressors with the error term in our OLS estimations. Total number of siblings, the age of the mother at childbirth, and other covariates (parental education, whether all siblings report, whether both parents report) could be correlated with unobservable socio-economic characteristics. In particular, the precise causal determination of early motherhood on children's academic outcomes has received considerable attention (*e.g.*, Geronimus, Korenman and Hillemeier, 1994;

Hofferth and Reid, 2003; Lopez-Turley, 2003), following an even larger debate on the consequences of early pregnancy on mothers themselves.

Obviously, the age of the mother at childbirth is linked to a number of variables that are expected to affect a child's educational attainment. Younger mothers are more likely to be single and have less human capital, which in turn may disproportionately impact the quality of the maternal inputs to their first-born etc. Also, there are adverse effects of unplanned motherhood that may dissipate over time (Bronars and Grogger, 1993). Yet, even if early motherhood does not *cause* lower educational attainment for a child, it is still possible that first-borns perform relatively better, conditional on early motherhood.

It would be very difficult to find compelling instrumental variables for all our potentially endogenous regressors. Therefore, to provide additional credibility to our results, we use a fixed effects model, which by construction removes variables that are constant *within* a family. As such, we take care of unobserved family-level heterogeneity, which may have previously led to bias. The results on birth order are broadly consistent with our initial ones. Hence, like Black, Devereux, and Salvanes (2004) in their analysis of family composition on children's education in Norway, we believe that our estimated birth order effects do not reflect omitted family characteristics.[3]

The rest of the paper is organized as follows. Section two presents our data. In section three, we closely replicate Hanushek's (1992) findings and then discuss how these results change with the inclusion of the age of the mother at childbirth; we provide consistency

[3] Black, Devereux and Salvanes (2004), p. 18. However, given our sample, we do not share their finding that "the effects appear to be of similar magnitude across families of different sizes." (*ibid*)

checks with different specifications of a family fixed effects model. Section four concludes the analysis.

2. Description of the data

Our data come from the Childbirth and Adoption History File (CAHF), a special supplemental file of the PSID. The CAHF contains details about childbirth and adoption events of eligible people living in a PSID family at the time of interview in any wave from 1985 through 2001. Eligible persons are defined as heads or wives of any age and other members of the family unit aged 12-44 at the time of the interview. These people are asked retrospective questions about their birth and adoption histories at the time of their first interview. In each succeeding wave these histories are updated.

The population examined here (henceforth the “index persons”) consists of all those for whom the CAHF sample contains records and links for the childbirth histories for at least one of their parents. The CAHF allows us to compile information on their birth order and the total number of children that their parent(s) report(s).

The index persons with missing information on their birth order or for whom the number of siblings is not ascertained are necessarily excluded from the sample. To make sure that all mothers have completed their fertility so that we correctly identify the total number of siblings, we further restrict the sample to those index persons whose mother was older than 44 in the last year she reported. Siblings are defined based on the childbirth histories of mothers.[4] In addition to the birth order and the number of siblings of the index persons, we have obtained additional demographic information on them and their parents from other PSID files using the unique individual identifiers that are present in the PSID main and supplemental files.

Notably, the PSID suffers from an important attrition bias. More educated people tend to stick with the questionnaire over longer periods of time while less educated ones do not; thus, it *appears* that education is decreasing over cohorts, which is of course untrue according to the U.S. Census. Since the first-born is older than other siblings by definition, this alone could, in theory, produce a spurious positive impact of being first-born on education. We have checked that this problem is of no consequence for our results.[5]

In order to match Hanushek's (1992) sample, in this paper, we focus on African American individuals. The resulting sample is very homogenous. Nearly 70% of the index persons are baby boomers (more than 80% in large families). The low education levels of both their mothers and fathers indicate that this sample belongs to the lower end

[4] The sample is twice as small if siblings are defined based on the childbirth histories of fathers, presumably because of the high proportion of single-mother households in the black community, as well as the high proportion of adult black males in the incarceration system. In case both parents report, we can identify between siblings and half siblings.

[5] First, it can be noted that if the problem was severe, we would observe a decreasing effect of birth order *per se* on education and not a specific first born effect. This is not the case. Second, the regressions presented in this article contain age controls that separate cohort from birth order effects. Further, our main results are robust to the exclusion of those age effects.

of the income distribution - more than 80% of the mothers do not have post secondary education (more than 90% in large families), which is quite similar to that used by Hanushek. In addition, since our purpose is to look at the effect of birth order, we have restricted attention to families with more than one child. Hanushek looked at the attainments of preschool and school children. However, in our data set, we can only look at years of education. Therefore, in this study, our focus here is on individuals who have presumably completed their education (*i.e.* index persons at least 22 years of age in 2001).[6] The summary statistics of our sample are presented in Table 1.

We found 2,887 index persons (929 distinct families) of black ethnicity, older than 22, with at least one other sibling, and whose mother has completed her fertility. About half of them have all their siblings reporting, and slightly less than half had both of their parents reporting their childbirth history.

[Insert Table 1]

3. Methods and Results

a. The first-born effect

We first use an OLS estimation with robust standard errors clustered by family unit (identified by the mother), which relaxes the independence assumption between the error terms and requires only that the observations be independent across clusters. Put

[6] Our results are robust to choosing other minimum ages in the mid 20s.

differently, the procedure allows for correlation of errors within a family.[7] We first test the hypothesis that being early in the birth order implies a distinct advantage, but that it is entirely due to the higher probability of coming from a small family.

[Insert Table 2a]

Columns 1 and 2 of Table 2a provide the justification for that claim. In Table 2a, we regress completed education on our regressors. In column 1, we omit the number of siblings; therefore, the significant coefficient on firstborn reflects not only the birth order effect but also the probability of coming from a small family. In column 2, the inclusion of the number of siblings makes the coefficient on firstborn insignificant. These two findings therefore provide a replication of Hanushek's results.

In column 3, however, we include the age of the mother at childbirth and find a strongly positive and significant effect of being first-born on years of education. This effect is confirmed when including the father's characteristics in column (4), where both a father and a mother report. The insignificance of the variable indicating whether all siblings report gives us confidence that we are not biasing our results by including families where not all siblings report; obviously, this is especially the case for large families.[8] The results presented here are for the impact of being first-born in families of

[7] We also used random effects procedures, but since they yield almost identical results as those with the family clustered standard errors, those are not reported.

[8] Hanushek reported a similar observation with his famtest variable.

more than one child. We obtain similar results on the impact of being first or second born in families of three and more. Further checks into the data show that mostly the large families drive our results on birth order (with results being more homogenous among whites).

The reason why the inclusion of the age of the mother at childbirth makes the coefficient on first-born significant is clear: the age of the mother at childbirth is mechanically, positively correlated with the birth order of a child and even more strongly across large families. On the other hand, we see that it is positively correlated with child's education. Then, if having a high birth order carries a negative impact on education, the two effects of birth order and the age of the mother at childbirth compete against one another. Therefore the coefficient on firstborn in column 2 reflects an omitted variable bias.[9] The results hold for both males and females (though among large families, a first-born girl gets a greater advantage relative to her sisters than a first-born boy relative to his brothers) and for mothers with or without more than a high school education. We found similar patterns for whites as well. Lastly, being last-born confers no advantage for any family size once the age of the mother at childbirth is factored in.

Conley offers the following argument: "in terms of parental investment, the cup starts to run dry as we go down the line (…) Parental resources, it appears, are allotted on a first come, first-served basis."[10] Alternatively, in our specific sample, since many black

---

[9] The correlation coefficient between firstborn and age of mother at childbirth is −0.44 in the sample used in Table 2a columns (1)-(3).

[10] Conley (2004), p. 69.

families are headed by single mothers, it might be that first born children are better off because on average, they live with both parents for longer than later born children. This is consistent with the coefficient showing presence of both parents is positive and significant in all specifications of Table 2a. There are not enough never married mothers in our sample to investigate this claim further, but interestingly, our results hold when only considering children whose mothers were unmarried at the time of their birth.

As noted earlier, a causal interpretation of the age of the mother at childbirth would hinge on the assumption of its exogeneity. Without instrumental variables or a treatment vs. control quasi experiment, it is hard to say more. The age of the mother at childbirth, itself positively correlated with birth order, could easily proxy for other unobserved variables such as level of human capital and parental resources.[11]

To address this problem, our fixed-effect estimation (Table 2b) removes family characteristics and unobserved family-level heterogeneity.

[Insert Table 2b]

The estimation over the entire sample offers only modest support to our earlier results since the coefficients on first-born are again positive, but not always significant at the 5% level. However, when restricted to larger families, the coefficients on first-born become larger, more significant and robust to all specifications. We illustrate this pattern with

[11] This is further evidenced by the fact that including a dummy variable indicating whether the mother was married at the time of childbirth instead of (and, obviously, also along with) age of mother at childbirth also results in a highly significant first born coefficient. Results are available upon request.

families of more than five siblings. A detailed interpretation of the first-born variable in that context is given in Table 2b. We also found the same greater advantage of a first-born girl relative to her sisters than a first-born boy relative to his brothers as mentioned earlier. Most of the effect comes from those large families with substantial spacing (age difference between first and last born).

Unfortunately, the coefficients on age or on age of the mother at childbirth are uninformative in those fixed effects regressions. Deviations from family means for age convey the same information as deviations from family means for age of mother at childbirth. We thus do not provide separate estimations with age and age of mother at childbirth. The age of the mother at first childbirth is certainly relevant but here, it is differenced out.

b.The "last borns effect" in large families

We now test the hypothesis that within large (more than 5 siblings) families, the middle born do the worst and that the last-born do better than everyone else.[12]

[Insert Table 3a]

---

[12] This pattern is nicely illustrated on pages 101-102 (figures 1 and 2) of Hanushek's article. We were also able to reproduce the pattern in pages 103-104 (figures 3 and 4) through a regression of education on the probability of being first or second born (with and without our other variables), yet without including the number of siblings. The coefficient was significant at the 1% level, confirming the importance of coming from a small family. Interestingly, when including age of mother at childbirth, the coefficient doubles in magnitude while retaining 1% statistical significance.

Table 3a replicates those findings. The variables of interest are a dummy indicating whether the child is among the first three, and two dummy variables (whether among the first three, and whether fourth and above) interacted with birth order. Those last two interaction terms capture the trends of falling scholastic performances from first to fourth child, and the rising performance of children fourth and up as presented by Hanushek and in the context of time allocation, by Lindert.

Column 1 provides estimates without total number of children. Like in Table 2a, the coefficients are a mix of birth order and sibling size and are therefore misleading. In column 2, when controlling for the number of siblings, we do find that exact same significant upward trend from the fourth child on as in Hanushek. We also find a significant dummy variable indicating worse outcomes on average for children born fourth and up, which in this context should be interpreted as an intercept, and a weakly significant negative slope for the first to third children. In column 3, when controlling for the age of the mother at childbirth, we find a much more significant negative trend from first to third, a significant negative dummy for children born fourth and up, but a non significant coefficient for the upward trend of later born children. We also found that being the first-born in those families increases years of education by a little more than half a year. The interpretation of those results is similar to the ones presented earlier and the same qualifications apply.[13]

[13] Note that a regression restricting the sample to both parents present at childbirth (with or without father's characteristics) does not yield meaningful results: the sample becomes too small. Only 43% of families with more than 5 siblings report with both mother and father present, which leaves only 592 observations from 114 families. Even the number of siblings is insignificant in such a small sample.

To summarize, the apparent advantage of being last-born in large families is entirely attributable to the rising age of the mother at childbirth, again with the caveat that the age of the mother at childbirth actually picks other elements that are also correlated with it and may affect the outcome of interest. We note that the presence of a father at childbirth seems irrelevant in those families, which may explain Hanushek's result on the apparent lack of impact of divorce, even if, in the overall sample used in Table 2a, the presence of a father is positively correlated with the children's education.[14]

Finally, in Table 3b, the fixed effects estimation again provides strong support for the negative impact of birth order on education and confirms the absence of any upward trend from middle born to last-born. The same specifications also revealed a greater advantage of first-born girls relative to their sisters than first-born boys relative to their brothers.

[Insert Table 3b]

4. Conclusion

We have shown how the omission of the age of the mother at childbirth effect results in an underestimation of the impact of being first-born and an overestimation of the impact of being last-born. At this point however, the age of the mother at childbirth

[14] We also directly checked that formal divorce counts little compared to parental separation, confirming earlier findings on the psychological well-being of children when parents split (Amato and Booth, 1997).

should be interpreted broadly as a proxy for a set of maternal inputs. Most importantly, a fixed-effects estimation confirmed the presence of a positive first-born effect (especially that of a first-born female relative to her sisters[15]) and the absence of either specific middle born or last-born effects among large families.

From a policy perspective, any effort to treat individuals differently on the basis of birth parity seems unfeasible. Because of the endogeneity of the mother's age at first birth, and without any source of exogenous variations, we cannot ascertain with enough confidence at this point the benefits of a policy that would offset the determinants of early motherhood in order to promote black children's education.

Using twins as an instrument, Black, Devereux and Salvanes (2004) advance that "the family size itself has little effect on the quality of each child but more likely impacts the marginal children through the effect of birth order"[16] in their Norwegian sample. Therefore, while we tentatively agree with the idea that smaller family sizes are responsible for a rise in scholastic performances over cohorts among African Americans in the U.S., we would like to emphasize that this effect is compounded by a corresponding increase in the proportion of first-born children.

---

[15] It would be interesting to link this finding with the well-known result that first born daughters are more likely to care for parents than any other sibling. The causal mechanism is of course ambiguous at this point.

[16] Black, Devereux and Salvanes (2004), p. 2.

References

Amato, P. R. and A. Booth. 1997. *A Generation at Risk: Growing Up in an era of Family Upheaval.* Cambridge, Mass: Harvard University Press.

Birdsall, N. 1979. Siblings and Schooling in Urban Columbia. PhD Dissertation, Yale University, Department of Economics.

Behrman, J. R. 1986. Birth Order, Nutrition and Health; IntraHousehold Allocation in Rural India. Mimeo. University of Pennsylvania.

Behrman, J. R. and P. Taubman. 1986. Birth Order, Schooling, and Earnings. *Journal of Labor Economics* 4(3): S121-45.

Black, S. E., P. G. Devereux, K. G. Salvanes. 2004. The More the Merrier? The Effect of Family Composition on Children's Education. NBER Working Paper No. w10720, Forthcoming in the *Quarterly Journal of Economics*.

Bronars, S. G., and J. Grogger. 1994. The Economic Consequences of Unwed Motherhood: Using Twin Births as a Natural Experiment. *American Economic Review* 84: 1141-1156.

Conley, D. 2004. *The Pecking Order*. Pantheon Books, New York.

Geronimus A. T., S. Korenman, M. M. Hillemeier. 1994. Does young maternal age affect child development? Evidence from cousin comparisons in the United States. *Population and Development Review* 20 : 585-609.

Hanushek, E. A. 1992. The Trade-off between Child Quantity and Quality. *Journal of Political Economy* 100 : 84-117.

Hofferth S. L., and L. Reid. 2002. Early Childbearing and Children's Achievement and Behavior over Time. *Perspectives on Sexual and Reproductive Health* 34 : 41-49.

Kehrer K. C. 1979. The Gary Income Maintenance Experiment. *Journal of Human Resources* 14 : 431-33.

Kelly, T. F. and L. Singer. 1971. The Gary Income Maintenance Experiment: Plans and Progress. *American Economic Review* 61 : 30-38.

Kessler, D. 1991. Birth Order, Family Size, and Achievement: Family Structure and Wage Determination. *Journal of Labor Economics* 9(4): 413-26.

Lindert, P. H. 1977. Sibling Position and Achievement. *Journal of Human Resources* 12 : 198-219.

Lopez-Turley, R. N. 2003. Are Children of Young Mothers Disadvantaged Because of Their Mother's Age or Family Background? *Child Development*, 74: 465-474.

Table 1: Descriptive Statistics

	Two or more siblings				Five or more siblings			
	Mean	S.D.	Min.	Max.	Mean	S.D.	Min.	Max.
Index Person:								
Education (years)	**12.19**	**1.84**	**1**	**17**	**11.95**	**1.90**	**1**	**17**
Male	0.47	0.50	0	1	0.46	0.50	0	1
Age (in 2001)	39.34	8.61	23	78	42.29	7.15	23	70
Number of siblings	6.01	3.12	2	15	8.68	2.37	6	15
First-born	0.21	0.41	0	1	0.08	0.27	0	1
Birth order > 3	0.43	0.50	0	1	0.72	0.45	0	1
All siblings present	0.48	0.50	0	1	0.24	0.43	0	1
Both parents report	0.43	0.50	0	1	0.43	0.50	0	1
Mother:								
Education (years)	10.52	2.63	3	17	9.55	2.64	3	17
Age at childbirth	25.60	6.18	15	48	27.60	6.47	15	47
Mother married when child was born	0.63	0.48	0	1	0.69	0.46	0	1
Father:								
Education (years)	9.35	3.42	1	17				
Age at childbirth	29.81	7.35	17	65				
# observations	2,887[a]				1,376			
# families	929				407			

The sample includes individuals who are black, with at least one sibling, over 22 years of age, and with mothers who are at least 44 years of age.

[a] Except for ‘father’s education’ (1,245 observations) and ‘father’s age at childbirth’ (1,243 observations).

Table 2A

OLS REGRESSION WITH DEPENDENT VARIABLE: COMPLETED EDUCATION

	(1)	(2)	(3)	(4) (d[both]=1)
d[first-born]	0.209	0.1	0.312	0.295
	(0.08)***	(0.082)	(0.087)***	(0.133)**
total number of siblings		−0.084	−0.1	−0.118
		(0.021)***	(0.022)***	(0.031)***
age of the mother at childbirth			0.043	0.056
			(0.009)***	(0.021)***
d[Male]	−0.294	−0.291	−0.303	−0.401
	(0.074)***	(0.074)***	(0.073)***	(0.107)***
Age	0.136	0.172	0.166	0.14
	(0.031)***	(0.032)***	(0.031)***	(0.05)***
Age^2	−0.001	−0.002	−0.001	-9×10^{-4}
	(4×10^{-4})***	(4×10^{-4})***	(4×10^{-4})***	(7×10^{-4})
Mother's education	0.154	0.135	0.143	0.157
	(0.02)***	(0.02)***	(0.02)***	(0.032)***
d[all siblings report]	0.264	0.072	0.158	−0.157
	(0.104)**	(0.11)	(0.111)	(0.155)
d[both parents report]	0.343	0.382	0.341	
	(0.095)***	(0.093)***	(0.094)***	
Father's education				0.052
				(0.023)**
d[half siblings]				−0.252
				(0.178)
Age of father at childbirth				10^{-4}
				(0.016)
Constant	7.033	7.065	5.933	5.938
	(0.66)***	(0.656)***	(0.697)***	(1.073)***
R^2	0.078	0.09	0.105	0.16
# observations	2887	2887	2887	1239
# family clusters	929	929	929	407

(1)-(4): all mothers have completed their fertility (age>44), all respondents are black, assumed to have completed their education (age>22). *: 10% significance; **: 5% significance; ***: 1% significance. Robust standard errors clustered by family.

Table 2B

FIXED EFFECTS LINEAR REGRESSION WITH DEPENDENT VARIABLE:

COMPLETED EDUCATION (ALL COEFFICIENTS: DEVIATIONS FROM FAMILY MEANS)

	Families with more than one sibling		Families with more than five siblings	
	(1)	(2)	(3)	(4)
d[first-born] †	0.184	0.172	0.343	0.425
	(0.083)**	(0.096)*	(0.172)**	(0.19)**
d[Male]	-0.319	-0.32	-0.246	-0.246
	(0.069)***	(0.069)***	(0.096)***	(0.096)***
Age of the mother at childbirth		-0.002		0.01
		(0.008)		(0.01)
Constant	12.30	12.364	12.034	11.742
	(0.047)***	(0.228)***	(0.063)***	(0.294)***
R^2	0.01	0.01	0.01	0.01
# observations	2887	2887	1376	1376
# family clusters	929	929	268	268

(1)-(4): all mothers have completed their fertility (age>44), all respondents are black, assumed to have completed their education (age>22); (3)-(4): all respondent from families > 5 siblings.

No estimation run with age because including both age of the mother at childbirth and age is redundant in a family fixed effects regression.

*: 10% significance; **: 5% significance; ***: 1% significance.

† Strictly speaking, the coefficient on d[first-born] measures a variable that takes value [# siblings who report – 1]/[# siblings who report] if the individual is a first-born, – 1/[# siblings who report] if the individual is not a first-born and a first-born reports, and 0 if no first-born reports. For families where a large enough number of siblings reports, this is roughly equivalent to a 0/1 dummy. We do not have enough observations to run separate regressions for all # siblings who report. Yet, regressions restricted to those families where only 6 to 8 siblings report yield similar results as those in columns (3)-(4). Additional regressions on all races families where either 4, or 5 to 6, or 6 to 7 (and obviously 4 to 6, 4 to 7 and 5 to 7) siblings report yield comparable results as those above. There are not enough observations where more than 7 siblings report for further checks. Black, Salvanes and Devereux (2004) have enough observations to run siblings size by siblings size regressions in their Norwegian sample and report similar results.

Table 3A

OLS REGRESSION WITH DEPENDENT VARIABLE: COMPLETED EDUCATION

	(1)	(2)	(3)
d[birth order>3]	−0.521	−0.888	−0.919
	(0.305)*	(0.288)***	(0.289)***
Birth order*d[birth order>3]	0.03	0.118	0.035
	(0.027)	(0.037)***	(0.047)
Birth order*d[birth order<4]	−0.241	−0.224	−0.308
	(0.115)**	(0.115)*	(0.12)**
total number of siblings		−0.128	−0.1
		(0.048)***	(0.048)**
age of the mother at childbirth			0.042
			(0.017)**
d[Male]	−0.236	−0.244	−0.252
	(0.12)**	(0.119)**	(0.117)**
Age	0.193	0.204	0.21
	(0.065)***	(0.062)***	(0.063)***
Age^2	−0.002	−0.002	−0.002
	(8×10^{-4})**	(8×10^{-4})**	(8×10^{-4})**
Mother's education	0.106	0.108	0.108
	(0.032)***	(0.03)***	(0.03)***
d[all siblings report]	0.145	0.054	0.087
	(0.193)	(0.195)	(0.198)
d[both parents report]	0.228	0.283	0.26
	(0.16)	(0.157)*	(0.158)
Constant	6.49	6.918	5.9
	(1.286)***	(1.265)***	(1.38)***
R^2	0.046	0.06	0.068
# observations	1376	1376	1376
# family clusters	268	268	268

(1)-(3): all mothers have completed their fertility (age>44), all respondents are black, assumed to have completed their education (age>22), all respondent from families > 5 siblings.
*: 10% significance; **: 5% significance; ***: 1% significance. Robust standard errors clustered by family.

Table 3B

FIXED EFFECTS LINEAR REGRESSION WITH DEPENDENT VARIABLE: COMPLETED EDUCATION (ALL COEFFICIENTS: DEVIATIONS FROM FAMILY MEANS)

	(1)	(2)†
d[birth order>3]‡	-0.666	-0.651
	(0.3)**	(0.3)**
Birth order*d[birth order>3] ‡	0.031	0.069
	(0.027)	(0.053)
Birth order*d[birth order<4] ‡	-0.225	-0.184
	(0.106)**	(0.117)
d[Male]	-0.245	-0.244
	(0.096)**	(0.096)**
age of the mother at childbirth		-0.021
		(0.025)
Constant	12.52	12.873
	(0.244)***	(0.493)***
R^2	0.011	0.012
# observations	1376	1376
# family clusters	268	268

(1)-(2): all mothers have completed their fertility (age>44), all respondents are black, assumed to have completed their education (age>22), all respondent from families > 5 siblings.

No estimation run with age because including both age of the mother at childbirth and age is redundant in a family fixed effects regression.

*: 10% significance; **: 5% significance; ***: 1% significance. Robust standard errors clustered by family.

†: (2)-(3): joint test of d[birth order>3] and Birth order*d[birth order<4] significant at the 10% level.
‡: The same caveats on the interpretation of these coefficients as in Table 2b apply. Additional regressions restricted to those families where only 6 to 8 siblings report yield comparable results as above. So do regressions on all races families where 5 to 6, 5 to 7 or even 6 to 7 siblings report. There are not enough observations where more than 7 siblings report for further checks.