J Popul Econ (2015) 28:937–964
DOI 10.1007/s00148-015-0556-x

ORIGINAL PAPER

Intra-household allocation of family resources and birth order: evidence from France using siblings data

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Received: 8 December 2013 / Accepted: 9 April 2015 /
Published online: 13 May 2015

Abstract We examine the effect of birth order on education, occupation, and parental transfers using four cross sections of the French Wealth surveys conducted between 1992 and 2010. Estimates from ordered models confirm the presence of a first born advantage in education and occupation, the latter persisting to a lesser extent after controlling for education. Strikingly, parents are on average more likely to make transfers to first-born children, although the vast majority provides cash or property gifts to all of their children. This first-born advantage in transfers is uncorrelated with the likelihood of having attained a higher education or better occupation. Overall, our findings suggest that in France, the mechanism supporting the first born advantage may not stem from confluence effects or family resource dilution.

Keywords Birth order · Education · Occupation · Siblings · Intergenerational transfers

JEL Classification D1 · I2

1 Introduction

Recent research shows a consistent pattern across countries and cultures revealing a first-born advantage in education and, to a lesser extent, earnings in developed

Responsible Editor: Responsible Editor: Junsen Zhang

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countries (Black et al. 2005; Kantarevic and Mechoulan 2006; Booth and Kee 2009; de Haan 2010). In the developing world, the evidence is more mixed with some studies suggesting first-born children are actually less favored (Ejrnaes and Pörtner 2004; Edmonds 2006; Emerson and Souza 2008). Interestingly, the mechanism underlying the phenomenon of birth order-dependent outcomes has not been identified with certainty and sparks an ongoing debate among both economists and psychologists.

For example, Zajonc's early confluence theory states that first-born children benefit from having mainly adult influences around them in their early years (Zajonc 1976). Alternatively, resource dilution (or depletion) theory posits that later-born children come in an environment where family resources, including attention and quality time spent with parents, must be shared with multiple children so that they receive systematically fewer inputs relative to their elder siblings (Downey 2001; Price 2008; de Haan et al. 2014). More recent hypotheses include the idea that birth order relates to differences in IQ (Black et al. 2011) or that parents exert a birth order-dependent discipline in which those who are born later face more lenient environments (Hotz and Pantano 2015). In the USA, Lehmann et al. (2014) show that birth order differences may be explained by changes in maternal behavior and parental inclinations. These various explanations do not have to be exclusive.[1]

Our contribution to this literature is twofold. First, we propose an empirical analysis of the role of birth order on both educational attainment and occupation in the French context. We use repeated cross-sectional surveys conducted in 1992, 1998, 2004, and 2010 on samples of about 10,000 households with information on children living either with their parents or in an independent dwelling. We are thus able to account for both observed and unobserved heterogeneity at the sibship level in our regressions. Second, we complement the traditional birth order analysis on siblings' achievements with an investigation of the role of birth order in parental transfers. While our investigation is limited to the provision of inter vivos gifts made till the time of the survey (which therefore excludes bequests), the combination of data about education, occupation, and transfers allows us to uncover new insights about parents' responsibility in conferring a birth order-dependent advantage among siblings over the life cycle.

We confirm the pattern found in other Western cultures that being first born confers an advantage in education and occupation. Part of the occupational edge actually persists when controlling for education. Incidentally, this advantage is not specifically transmitted from first-born parent to first-born child. Another insight is that, conditional on unequal sharing, parents are more likely to make inter vivos transfers to first-born children, although the vast majority provides cash gifts to all of their children. Therefore, it appears that parents, on average, have a persistent, conscious positive bias toward first-born children. Our results are particularly striking since they show that parents do not offset the educational and occupational advantage of first-born children through cash and property gifts. To the extent that the mechanisms accounting for the first born advantage early in life are similar to those accounting for parents' preference toward first-born children later in life, our results call into question traditional theories

[1] There are also reasons why the last born could, in theory, be favored, e.g., parental earnings increase over the life cycle (hence, there would be fewer resources for the first-born children if households face liquidity constraints) and older mothers are more experienced.

that explain the first born advantage through elements that parents have little control over or no awareness of.

The remainder of this contribution is organized as follows. In the next section, we describe the French data sets used for the empirical analysis. Section 3 presents the ordered Probit specifications used to study education and occupation and the results are discussed in section 4. We focus on the relationship between birth order and financial or property transfers made to children in section 5. Finally, section 6 concludes.

2 Data and descriptive statistics

2.1 The French Wealth surveys

To study the effect of birth order on the intra-household allocation of family resources, the empirical strategy considered in this paper is based on the construction of samples matching respondents (parents) with their various children. Specifically, we consider four repeated cross-sectional data sets on household wealth conducted in France by the National Institute of Statistics and Economic Studies (INSEE) in 1992, 1998, 2004, and 2010.[2] The primary objective of the Wealth surveys is to describe the situation of households with respect to financial, real-estate, and professional assets as well as debts. These surveys are used to observe the distribution of assets as well as the different types of asset holding patterns across households living in France. They also shed light on factors accounting for wealth accumulation over the life cycle like gifts and inheritances. These surveys have relatively large sample sizes, with more than 10,000 interviewed households per survey.[3]

An interesting feature of the Wealth surveys is that they include a specific module on children of the respondents who live in an independent dwelling. The questionnaires also provide some characteristics of those children living with their parents (the respondents) at the date of the survey. Thus, we proceed in the following way to construct parents-children samples. First, for each survey, we construct a sample of parents by selecting household heads (including their spouses, if any) having at least one child, either at home or living elsewhere. Secondly, we construct a sample of children that combines all children living with their parents and children living outside. For each child, the survey provides a detailed description of the sibship composition.[4]

The last step consists in matching the parent and child samples for each survey and then in combining the 4-year-specific matched samples. We restrict the final sample in the following way. First, we drop all children under age 24 (34,932 observations deleted). Our assumption is that at age 24, we know the final level of education of each child. For the sake of robustness, we also estimate regressions with children aged

[2] Databases are available free of charge for use in research from the French Data Archives for social sciences (Réseau Quetelet, http://www.reseau-quetelet.cnrs.fr/spip/?lang=en).

[3] The number of respondents is, respectively, 9530 in 1992, 10,168 in 1998, 9692 in 2004, and 15,006 in 2010. In the 2010 Wealth survey, the sample includes 2218 households living in overseas departments. These households were excluded for comparability purposes over the period.

[4] Although we can always identify recomposed families, we treat co-resident half siblings as pure siblings since we do not know which siblings are not those biological children of the reference person. All our results hold when removing recomposed families from the sample.

over 16.[5] As most of these younger children have not yet completed their schooling at the time of the survey, we account for this form of censoring in our estimation procedure. Second, we exclude mothers aged under 45 since they may have not completed their fertility (667 observations deleted).[6] Third, to avoid possible measurement errors, we choose to exclude children whose father or mother age at birth was under 14 (316 observations deleted).

After excluding children with missing values for education (N=997), the matched sample comprises 41,688 parent-child pairs corresponding to 18,219 families. There are 8213 children (3468 families) in the 1992 survey, 9684 children (4099 families) in the 1998 survey, 9300 children (4127 families) in the 2004 survey, and 14,491 children (6525 families) in the 2010 survey.[7] By using repeated cross sections, our sample includes birth cohorts born from 1920 to 1986. The mean age of the selected children is 39 years: 11.1 % were born before 1950, 25 % between 1950 and 1959, 34.1 % between 1960 and 1969, 22.4 % between 1970 and 1979, and 7.5 % since 1970. We consider the three following outcomes to study the effect of birth order on the intra-household allocation of family resources: education, occupation, and financial transfers from parents.

Concerning education, there is no information about years of schooling or age at the end of schooling for children living on their own in the Wealth surveys. Instead, in the 1998, 2004, and 2010 questionnaires, parents indicate the highest level of education of each child according to the following five ordered categories: no diploma, less than high school, high school, undergraduate, and graduate-postgraduate studies. In the 1992 survey, the two upper categories were merged into one, so that we only know whether children have completed more than high school.

The second outcome investigated in this paper is the type of occupation held by children. The classification in the four Wealth surveys includes the following categories: farmer, self-employed, manager, intermediate occupation, white-collar worker, unskilled/skilled worker, and other occupation. In our empirical analysis, we will only consider the subsample of children having one of the following occupations: manager, intermediate, white-collar worker, or unskilled/skilled worker. With this selection, the ranking of occupations is obvious. Conversely, it seems much more difficult to compare the farmer and self-employed categories (which are very heterogeneous) with other occupations.

There are a few questions about financial transfers from parents to children in the Wealth surveys. First, parents indicate whether they have provided any help to their non-coresident children through the following form: financial gift for a specific event, regular gifts, payment of housing rent, or financial loan. The main difficulty here is that we do not know who benefits from the transfer within the sibship. The timing of these transfers is also poorly documented since we know if these transfers have been made during the period of schooling, after that period, or both. Second, in the 1998, 2004, and 2010 surveys, parents indicate whether they have made large gifts in the form of cash or property to their children for those living in an independent dwelling. If any, we know exactly which children within the sibship have received those transfers.[8] We will therefore focus on those transfers to shed light on the intra-household reallocation of parental resources.

[5] In the Wealth surveys, there is no information on education for children under 17.

[6] Also, we choose to exclude the few observations (N=413) in which the household head was under 45.

[7] In 2004, the questionnaire did not include any information on education and occupation of children living with their parents. This explains the smaller number of observations for the 2004 survey.

[8] However, there is no information on the amounts received by children.

2.2 Descriptive statistics on education and occupation

In Fig. 1, we describe the pattern of education by birth cohort and gender. Education in France has substantially increased during the second part of the twentieth century: the percentage of children having more than high school has increased from around 20 % for the older cohorts to about 60 % for the youngest ones. The last few decades have also been marked by greater growth in the participation of girls in higher education relative to boys. The higher proportion of girls with more than high school is apparent for cohorts born after 1950.

As shown in Fig. 2, occupational patterns for those having a job have also considerably changed for the selected cohorts. The proportion of unskilled workers has been divided by more than two over the period, from 23.8 % for cohorts born before 1945 to around 10 % for cohorts born since 1975. At the same time, the share of men in this category has increased substantially.[9] Since the proportion of women having a white-collar occupation has remained rather stable (around 50 %), this implies that the shift in occupations for women has essentially occurred from factory worker status to the intermediate/clerical and managerial categories. Obviously, one reason explaining this shift toward more qualified occupations is the higher educational attainment of women.

We consider the following set of family characteristics to explain both education and occupation outcomes. Concerning children, we include a gender dummy, a set of birth cohort dummies, number of siblings, and birth order. Concerning parents (head of the household), we introduce age at birth, type of family dummies (parents living in a couple, lone-parent family, blended family), and education coded with five categories.[10] We also construct a dummy variable which is equal to one when parents have given financial transfers to any of their children through either regular or irregular cash gifts or other transfers, exclusive of loans. This covariate is expected to pick up the effect of parental wealth on education and may be seen as a proxy for the distinction made by Becker and Tomes (1986) between poor and rich families. In the presence of liquidity constraints, only rich families should be able to make the wealth-maximizing investments in their children's education.

Table 1 highlights the composition of the sample by level of education of the children. On average, educational attainment is higher for girls and for children from the youngest cohorts, while it is negatively correlated with the number of siblings and birth order. Children from lone-parent families are less likely to succeed in school. As expected, there is a large and positive correlation between educational attainment and parental education: 59.7 % of children whose parents have no diploma have themselves no diploma, while this proportion is less than 10 % among children having completed more than high school. Children are also much more educated when they have rich parents. The proportion of families having made financial transfers to their children is 62.8 % among highly educated children compared to 30.1 % for children without any diploma.

[9] Among workers, the male-female ratio was 2.5 for cohorts born before 1945, 4 for cohorts born before 1955 and 1959, 6 for cohorts born between 1975 and 1979, and even 8 for cohorts born since 1980.

[10] Regional dummies and dummies for size of urban unit, which are expected to explain differences in educational supply, will also be taken into account.

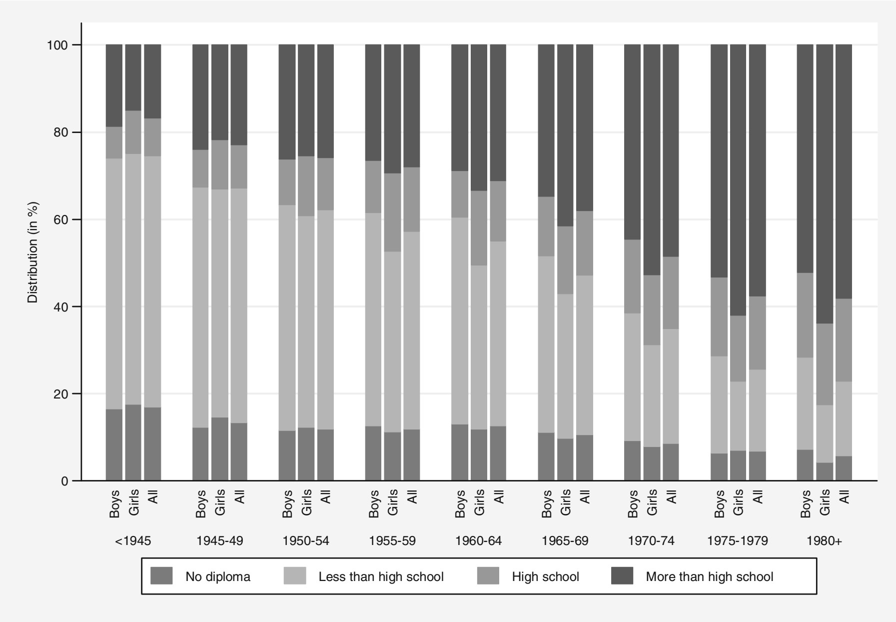

Fig. 1 Education of children, by birth cohort and gender. Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

In Table 2, we focus on the role played by the composition of the sibship on education. Having numerous brothers and sisters is negatively correlated with the

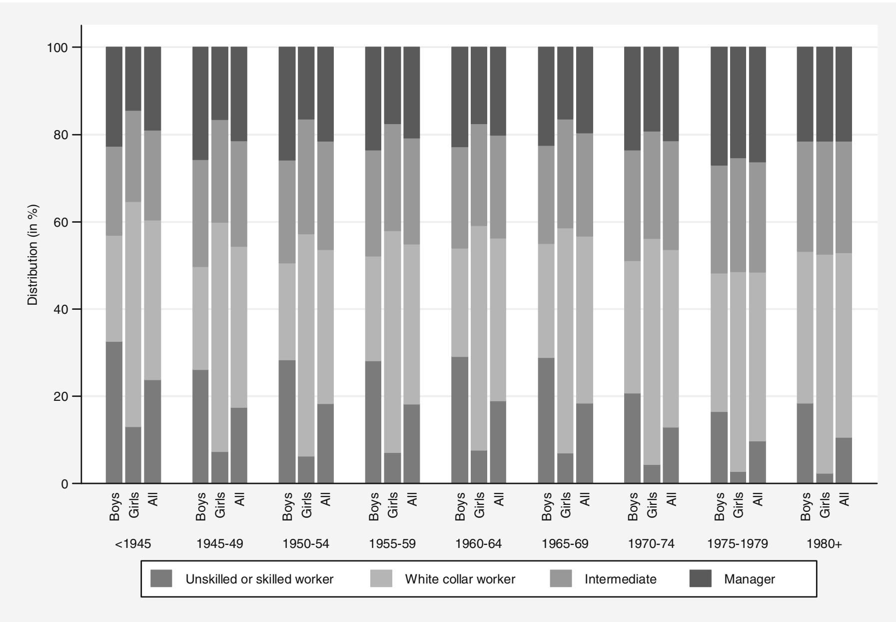

Fig. 2 Occupation of children, by birth cohort and gender. Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

Table 1 Descriptive statistics of the sample

Variables	No education	Less than high school	High school	More than high school	All
	Means	Means	Means	Means	Means
Characteristics of children					
Female	0.484	0.456	0.542	0.530	0.499
Age	40.573	40.876	38.023	37.092	39.037
Number of siblings	3.445	2.586	2.013	1.842	2.321
Number of sisters	1.717	1.281	1.013	0.921	1.156
Birth order	2.608	2.198	1.953	1.830	2.071
Characteristics of parents					
Head's age at birth	27.819	27.551	27.993	28.555	28.013
Parents living in couple	0.443	0.549	0.607	0.678	0.594
Lone-parent family	0.487	0.399	0.333	0.270	0.352
Blended family	0.069	0.052	0.060	0.051	0.055
Head's education					
No diploma	0.597	0.347	0.193	0.094	0.258
Primary	0.234	0.371	0.284	0.169	0.270
Secondary	0.120	0.213	0.312	0.266	0.237
High school	0.022	0.042	0.107	0.133	0.083
High school	0.027	0.027	0.104	0.338	0.152
Family rich (transfers to children)	0.301	0.380	0.459	0.628	0.474
Number of children	4437	15,869	6039	15,343	41,688

Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

probability of being highly educated. The proportion of children having completed more than high school is for instance equal to 45.6 % for children with only one brother or sister, 40.1 % with two siblings, 32.3 % with three siblings, and 25.8 % with four siblings.[11] Simple correlations suggest a negative relationship between educational attainment and birth order, net of the influence of sibship size. If we consider for instance the case of two-children families, the proportion of children with more than high school is 46.8 % for the first-born child against 44.1 % for the second-born child. For families with three children, the same proportions are equal to 41.8 % for the eldest, 39.6 % for the second born, and 38.6 % for the last born.[12]

[11] However, compared to two-child families, the proportion of having high education is lower for one-child families (42.3 % compared to 45.6 %).

[12] For families with four or five children, the proportion of children with more than high school education appears to be U-shaped with respect to birth order in the USA (Hanushek 1992). Kantarevic and Mechoulan (2006) analyze this puzzling stylized fact and find that it stems from the confounding factor of the mother's age at birth. That is, when age of the mother at birth is controlled for, the inverse U-shape relationship disappears.

Table 2 Distribution of education of children, by size of sibship and birth order

Size of sibship	Birth order	Education				Number of observations
		No diploma Percent	<High school Percent	High school Percent	>High school Percent	
1	1	7.1	34.0	16.6	42.3	3775
2	1	5.7	31.6	15.9	46.8	6764
2	2	6.3	33.4	16.2	44.1	5756
2	All	6.0	32.4	16.0	45.6	12,520
3	1	7.4	35.2	15.6	41.8	4173
3	2	7.8	37.3	15.4	39.6	3840
3	3	8.7	36.6	16.2	38.6	3119
3	All	7.9	36.3	15.7	40.1	11,132
4	1	10.2	41.5	13.9	34.4	1740
4	2	12.0	44.2	12.5	31.3	1664
4	3	10.1	45.4	14.1	30.4	1519
4	4	12.1	39.5	15.4	33.0	1305
4	All	11.1	42.8	13.9	32.3	6228
5	1	16.6	46.5	9.9	27.1	721
5	2	17.3	44.3	12.5	26.0	707
5	3	20.4	45.8	9.8	23.9	681
5	4	18.4	46.4	9.2	26.0	630
5	5	15.6	46.3	12.3	25.7	544
5	All	17.7	45.8	10.7	25.8	3283
All		10.6	38.1	14.5	36.8	41,688

Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

We find similar results when investigating the role of birth order on occupation. For two-children families, the proportion of children in managerial occupations is 3.6 points higher for first-born compared to later-born children (26.8 against 23.2 %). The gap between the first-born and last-born children amounts to 4 points for three-child families (24.5 against 20.5 %) and around 3.5 points for families with four and five children. Conversely, it seems more difficult to conclude that a high birth order carries a direct negative impact on occupation. Indeed, later-born siblings could have fewer chances to obtain better jobs because they tend to be less educated on average.

These descriptive statistics must be interpreted cautiously. For instance, sibship size is expected to be negatively correlated with parental education, which should have a direct effect on a child's educational achievement. In the same way, the influence of birth order on either education or occupation does not take into account the fact that later-born children belong to younger cohorts, meaning that they are likely to achieve higher levels of education according to the increasing trend presented in Fig. 1. As a consequence, we turn to an econometric framework to study the role of birth order on education, occupation, and receipt of parental transfer.

3 Econometric strategy

In what follows, we present our estimation strategy for the educational outcome. The methodology used for studying occupational attainment is similar as both outcomes are ordered, and we turn to Probit and linear probability regressions when studying transfers. Let E_{ji} be the level of education a child i living in a family j, with i=1, …, N_j and j=1, …, J. The latent variable measuring the propensity of attaining a certain educational level, which is denoted by E^*_{ji}, is expected to depend on a set of explanatory variables X_{ji} that includes family characteristics of both parents and children. We rely on the following linear specification:

$$E^*_{ji} = X_{ji}\beta + \delta_j + \varepsilon_{ij} \tag{1}$$

with β a vector of parameters to estimate. The random perturbation is decomposed into one family-specific unobserved heterogeneity term δ_j and a pure random error ε_{ji}. The first term picks up measurement errors associated with parental characteristics along with unobserved parental factors.[13] Unobserved factors specific to children, like their abilities, are picked up by the error term ε_{ji}. This term may also reflect specific parental preferences for some of their children. Both δ_j and ε_{ji} are assumed to be normally distributed such that $\delta_j \sim N(0;\sigma^2_\delta)$ and $\varepsilon_{ji} \sim N(0;1)$. By definition, the latent variable E^*_{ji} is unobserved. However, the Wealth surveys indicate the level of education E_{ji} completed by each child. For a given level k, with k=1, …, K, the relationship between E_{ji} and E^*_{ji} is:

$$E_{ji} = k \ \text{ if } \ \mu_{k-1} < E^*_{ji} \leq \mu_k \tag{2}$$

with μ_k a set of threshold values.[14] Under the assumption that both δ_j and ε_{ji} are uncorrelated with X_{ji} (meaning that there is no correlation between the family-specific component and the selected covariates), the corresponding model is a random effect ordered Probit model (see Greene and Hensher 2010). Denoting by Φ the cumulative distribution function of the normal distribution, the probability for a child to achieve the level of education k is:

$$\Pr\left(E_{ji} = k\right) = \Phi\left(\mu_k - X_{ji}\beta - \delta_j\right) - \Phi\left(\mu_{k-1} - X_{ji}\beta - \delta_j\right) \tag{3}$$

The likelihood of the model $L = \prod_j \Pr\left(E_{j1} = e_{j1}, \ldots, E_{jN_j} = e_{jN_j}\right)$ depends on the distribution of the unobserved heterogeneity term δ_j and is thus estimated using quadrature techniques (Butler and Moffitt 1982; Frechette 2001).

There are five ordered categories for education in the Wealth surveys performed in 1998, 2004, and 2010: no diploma (E_{ji}=1), less than high school (E_{ji}=2), high school (E_{ji}=3), undergraduate (E_{ji}=4), and graduate or postgraduate (E_{ji}=5). In the 1992 survey, there is only one category above high school for children living on their own.[15] So, these children are characterized by either E_{ji}=4 or E_{ji}=5, which can be summarized as $E_{ji} \geq 4$. The corresponding probability $\Pr(E_{ji} \geq k|\delta_j)$=

[13] This includes for instance the degree of parental altruism which should be positively correlated with investment in the human capital of children (Becker and Tomes 1986).

[14] We set $\mu_0 = -\infty$ and $\mu_K = +\infty$ and assume that the thresholds are strictly ordered ($\mu_{k-1} < \mu_k$).

[15] For co-resident children, the 1992 Wealth survey indicates whether they have completed the undergraduate level or the graduate/postgraduate level.

$\Pr(E^*_{ji}>\mu_{k-1}|\delta_j)$ is $\Pr(E_{ji}\geq k|\delta_j)=1-\Phi(\mu_{k-1}-X_{ji}\beta-\delta_j)$. We account for these censored observations in our regressions:

$$\Pr(E_{j1}=k)=\frac{(1-\mathbb{I}_{ji})*[\Phi(\mu_k-X_{ji}\beta-\delta_j)-\Phi(\mu_{k-1}-X_{ji}\beta-\delta_j)]}{+\mathbb{I}_{ji}*[1-\Phi(\mu_{k-1}-X_{ji}\beta-\delta_j)]} \quad (4)$$

with $\mathbb{I}_{ji}$ being a dummy variable equal to one for a censored observation and zero otherwise. This random effect ordered Probit model with censoring is, as before, estimated by a maximum likelihood method. As usual, when considering a random effect specification, the crucial assumption is that the family characteristics X_{ji} do not depend on the family specific component δ_j.

For the sake of robustness, we have also estimated fixed effect ordered regressions. Since standard demeaning techniques cannot be applied to non-linear fixed effect models like the ordered Probit specification, we turn to the minimum distance estimator proposed in Das and van Soest (1999). Their strategy consists in estimating first a set of fixed effect Logit models à la Chamberlain (1980). The ordered-dependent variable E_{ji} is converted in a set of dummy variables E^k_{ji} such that $E^k_{ji}=1$ if $E_{ji}\geq k$ and $E^k_{ji}=0$ if $E_{ji}<k$ for each $k=2, \ldots, K$. Estimation of $K-1$ conditional fixed effect models provides efficient estimates of the corresponding vectors of parameters β^k. In a second stage, a classical minimum distance estimator is implemented to form a unique vector β from the different estimators $(\beta^2, \ldots, \beta^K)$.

4 Results for education and occupation

4.1 Birth order and education

Estimates from random effect ordered Probit models for children aged at least 24 are presented in column 1 of Table 3. The selected covariates include individual characteristics of both generations, and we introduce both family size and birth order dummies as additional control variables to be as flexible as possible.[16]

Children's education increases with parental age at birth.[17] Children from blended families and, to a lower extent, from lone parent families exhibit lower education achievement than children from intact families. As expected, there is a high positive correlation in education between parents and their children. Less educated parents may be more likely to face liquidity constraints, which would prevent them from investing optimally in the human capital of their progeny (Becker and Tomes 1986). However, in France, education is by and large almost free and children from poor families typically receive bursaries. Other possible explanations for this intergenerational correlation are that education of children is affected by some unobservable parental characteristics or that there is a causal relationship from parents' to children's education as evidenced in France by Maurin and McNally (2008).

[16] We have also estimated ordered regressions using the birth order index proposed by Booth and Kee (2009). These additional results, which are available upon request, lead to similar conclusions.

[17] In our regressions, we control for the head's age at birth since we do not always have information on both spouses for each child. Recall that the parent sample is constructed by assembling information on the head's characteristics and spousal characteristics, if any.

Table 3 Random effect ordered Probit estimates of education

Variables	(1) Boys and girls aged at least 24	(2) Boys aged at least 24	(3) Girls aged at least 24	(4) Boys and girls aged at least 17
Characteristics of children				
Female	0.185*** (13.82)			0.204*** (16.08)
Birth cohort				
<1945	Ref	Ref	Ref	Ref
1945–1949	0.119*** (2.67)	0.132** (1.99)	0.155** (2.41)	0.109** (2.48)
1950–1954	0.225*** (5.15)	0.203*** (3.18)	0.323*** (5.14)	0.209*** (4.84)
1955–1959	0.250*** (5.70)	0.114* (1.81)	0.461*** (7.35)	0.231*** (5.33)
1960–1964	0.200*** (4.51)	0.052 (0.82)	0.402***(6.36)	0.184*** (4.21)
1965–1969	0.339*** (7.48)	0.172*** (2.67)	0.562*** (8.74)	0.331*** (7.43)
1970–1974	0.546*** (11.69)	0.388*** (5.89)	0.819*** (12.32)	0.582*** (12.75)
1975–1979	0.712*** (14.10)	0.544*** (7.63)	0.998*** (13.81)	0.733*** (15.20)
1980+	0.607*** (11.52)	0.342*** (4.65)	1.002*** (13.35)	0.555*** (11.24)
Number of siblings				
0	Ref		Ref	
1	0.060** (1.97)	0.045 (1.05)	0.083* (1.84)	0.057** (1.96)
2	0.001 (0.02)	0.021 (0.47)	−0.017 (−0.35)	−0.002 (−0.05)
3	−0.131*** (−3.35)	−0.085 (−1.63)	−0.173*** (−3.13)	−0.116*** (−3.08)
≥4	−0.406*** (−9.75)	−0.337*** (−6.08)	−0.488*** (−8.38)	−0.398*** (−9.91)
Birth order				
First-born	Ref	Ref	Ref	Ref
Second-born	−0.198*** (−12.39)	−0.189*** (−7.73)	−0.220*** (−8.76)	−0.191*** (−12.53)
Third-born	−0.317*** (−13.60)	−0.295*** (−8.38)	−0.366*** (−10.29)	−0.315*** (−14.29)

Table 3 (continued)

Variables	(1) Boys and girls aged at least 24	(2) Boys aged at least 24	(3) Girls aged at least 24	(4) Boys and girls aged at least 17
Fourth-born	–0.363*** (–10.92)	–0.405*** (–8.14)	–0.369*** (–7.34)	–0.368*** (–11.68)
Fifth-born and more	–0.493*** (–11.88)	–0.534*** (–8.74)	–0.555*** (–9.11)	–0.481*** (–12.27)
Characteristics of parents				
Head's age at birth	0.029*** (15.00)	0.028*** (10.59)	0.032*** (11.76)	0.029*** (16.04)
Lone-parent family	–0.166*** (–7.73)	–0.159*** (–5.69)	–0.175*** (–6.01)	–0.209*** (–10.18)
Blended family	–0.480*** (–12.09)	–0.460*** (–8.76)	–0.515*** (–9.33)	–0.516*** (–13.87)
Head's education				
No diploma	Ref	Ref	Ref	Ref
Primary	0.574*** (21.14)	0.578*** (16.53)	0.605*** (16.62)	0.561*** (21.45)
Secondary	0.992*** (34.23)	0.969*** (25.58)	1.090*** (27.33)	0.972*** (35.36)
High school	1.625*** (40.98)	1.681*** (31.93)	1.711*** (30.76)	1.609*** (42.74)
>High school	2.275*** (61.46)	2.334*** (46.15)	2.421*** (45.31)	2.270*** (64.93)
Family rich (transfers to children)	0.386*** (19.36)	0.379*** (14.67)	0.429*** (15.73)	0.364*** (19.27)
Number of children	41,688	20,894	20,794	47,665
Number of families	18,219	13,282	13,208	20,224
Log likelihood	–49,614.1	–25,142.8	–25,188.6	–53,959.7

Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

Each regression also includes a set of regional and size of urban unit dummies. Significance levels are, respectively, 1 % (***), 5 % (**), and 10 % (*)

On average, girls are more educated than boys. Education increases with successive birth cohorts, reflecting the overall improvement in schooling over the second part of the last century.[18] As expected, we find an inverse relationship between the number of siblings (this covariate being potentially endogenous) and children's educational performance when there are three siblings and more. A commonly suggested explanation is that parents have finite levels of time and financial resources, so that these resources have to be diluted among children as sibship size increases. This corresponds to the well-known trade-off between child quantity and child quality emphasized in Becker and Lewis (1973).

Our random effect random effect estimates show that the coefficients associated with the birth order dummies are all negative and significant at the 1 % level. The average latent level of education is equal to 2.07 at the means of the sample. Net of the influence of family characteristics, being second born rather than first born reduces this predicted outcome by 9.6 %. The marginal effects are, respectively, −15.3 % for the third born, −17.5 % for the fourth born and −23.8 % for the fifth born. In columns 2 and 3 of Table 3, we estimate the same regressions by gender. Both for boys and girls, we find very similar results concerning the negative impact of birth order on education. So, our first finding is that later born children achieve lower education levels compared to first-born children in France. In what follows, we perform several robustness checks to assess the validity of this result.

First, we have excluded so far children aged between 17 and 23 as most of them had not completed their schooling at the time of the survey.[19] This may be problematic since younger children are more likely to have lower birth order. In our framework, it is straightforward to account for enrolled children as they correspond to censored observations: they will end their education with at least the level which has been recorded during the interview. Drawing on (4), we account for censoring of enrolled children and re-estimate the random effect ordered Probit model on the sample of children aged at least 17 (47,655 observations). As shown in column (4), inclusion of those younger children has very little influence on our estimates. In particular, educational attainment is again declining with birth order at the 1 % level.

Second, since the effect of birth order on education may be non-linear (Ejrnaes and Pörtner 2004; Booth and Kee 2009), we estimate our regressions on children aged at least 24 for various family sizes (from two to five children) with dummy variables for birth order. Our results appear in Appendix Table 7 (panel A1). For families with two children (column 2), the coefficient for the second-born child is negative (−0.246) and significant. With three children (column 3), the coefficient for the third-born child is around twice higher than that for the second-born child (−0.335 against −0.177). The situation is slightly different for larger families (columns 4 and 5). With four children for instance, second-born, third-born, and fourth-born children are less educated than first-born children, but there is no significant difference for the last three-born children.

Third, we relax the assumption of exogeneity between the covariates and the specific family component and turn to fixed effect regressions by family size using the

[18] We have also considered a piecewise linear function for the child's age, by adding both age and birth cohort dummies interacted by age in the regression.

[19] We have also estimated the random effect ordered Probit regression on the subsample of families whose youngest child is at least 24. The corresponding estimates are very close to those reported in column 1 of Table 3, with a negative and significant correlation between birth order and education.

minimum distance estimator previously described (panel A2, Appendix Table 7).[20] We expect similar results since, conditional on family size, birth order is orthogonal to the family-specific error component.[21] When considering all family sizes (column 1), we obtain negative coefficients for the various birth order dummies although the estimates are not really different for birth order exceeding two. Also, the fixed effect estimates by family size (columns 2 to 5) are very similar to those obtained with the random effect specification. Altogether, our estimates confirm that in France first-born children achieve higher education compared to other siblings net of the influence of parental characteristics.

Finally, we investigated whether there exists an intergenerational transmission of a first-born effect by using the information, uniquely available in the Wealth surveys, on whether parents are themselves first born or not. Because the information on parents' birth order is not available in the 1998 wave of the survey, we did not include that covariate in the estimations presented in our paper. The corresponding sample comprises 32,004 children (14,120 families). We introduce in the random effect ordered regression a set of interactions terms crossing the child's birth order dummies with the parental first-born dummies. Our results, not reported, show that there is no statistically significant intergenerational birth order effect.

4.2 Birth order and occupation

We now investigate the possibility of an effect of birth order on occupation. First, birth order may have a direct influence on occupation. This occurs if birth order is associated with personality traits. For instance, one could imagine that first borns are more motivated to leave the parental home when reaching adulthood and devote more efforts to finding a good job. Also, it may be that parents are more likely to help their first-born children in accessing good jobs, e.g., by going through their own professional networks.

Second, birth order should have an indirect impact through education since occupational status is strongly correlated with educational attainment. In that scenario, controlling for the child's education in a regression explaining occupation would make the birth order coefficient insignificant. A difficulty here is that our measure of education is not as precise as desired. Ideally, we would like to have not only the last grade completed by the child but also their field study. It is well acknowledged that there are substantial differences by gender in tertiary education: female students are more often involved in humanities and arts while male students are more numerous in engineering, manufacturing, construction, science, mathematics, and computing fields. However, this information is unavailable in the Wealth surveys. Thus, the correlation between birth order and occupation (net of education level) may be affected by the child's gender because of unobserved differences in study field.

We focus on the subsample of children aged at least 24 being either manager, in intermediate profession, white-collar worker, or unskilled/skilled worker to investigate the role of birth order on occupation (31,701 children). Given this selection and the fact

[20] By definition, parental characteristics which remain constant at the sibship level are picked up by the family fixed effect and are thus excluded from the regression.

[21] We thank an anonymous referee for this insight.

Table 4 Random effect ordered Probit estimates of occupation

Variables	(1) Boys and girls aged at least 24	(2) Boys and girls aged at least 24	(3) Boys aged at least 24	(4) Girls aged at least 24
Characteristics of children				
Female	0.007 (0.49)	0.024* (1.65)		
Birth cohort				
<1945	Ref	Ref	Ref	Ref
1945–1949	0.118** (2.41)	0.117** (2.42)	0.146** (2.12)	0.140* (1.79)
1950–1954	0.038 (0.82)	0.030 (0.66)	0.028 (0.43)	0.085 (1.14)
1955–1959	−0.087* (−1.91)	−0.099** (−2.18)	−0.077 (−1.20)	−0.080 (−1.10)
1960–1964	−0.193*** (−4.25)	−0.231*** (−5.05)	−0.222*** (−3.47)	−0.247*** (−3.38)
1965–1969	−0.383*** (−8.38)	−0.431*** (−9.33)	−0.437*** (−6.75)	−0.449*** (−6.10)
1970–1974	−0.478*** (−10.21)	−0.536*** (−11.30)	−0.562*** (−8.47)	−0.551*** (−7.30)
1975–1979	−0.535*** (−10.70)	−0.627*** (−12.33)	−0.720*** (−10.04)	−0.556*** (−6.92)
1980+	−0.688*** (−13.34)	−0.793*** (−15.01)	−0.880*** (−11.80)	−0.747*** (−9.00)
Number of siblings				
0	Ref	Ref	Ref	Ref
1	0.032 (1.05)	0.027 (0.90)	0.009 (0.20)	0.062 (1.32)
2	−0.045 (−1.41)	−0.022 (−0.70)	−0.043 (−0.94)	−0.003 (−0.06)
3	−0.063* (−1.70)	−0.024 (−0.65)	−0.038 (−0.73)	−0.009 (−0.15)
≥4	−0.292*** (−7.45)	−0.194*** (−4.91)	−0.262*** (−4.74)	−0.105* (−1.70)
Birth order				
First-born	Ref	Ref	Ref	Ref
Second-born	−0.038** (−2.26)	−0.045** (−2.55)	−0.041 (−1.55)	−0.040 (−1.40)
Third-born	−0.050** (−2.13)	−0.055** (−2.14)	−0.038 (−1.00)	−0.083** (−1.97)

Table 4 (continued)

Variables	(1) Boys and girls aged at least 24	(2) Boys and girls aged at least 24	(3) Boys aged at least 24	(4) Girls aged at least 24
Fourth-born	−0.087*** (−2.61)	−0.088** (−2.39)	−0.096* (−1.77)	−0.057 (−0.94)
Fifth-born and more	−0.072* (−1.81)	−0.071 (−1.56)	−0.060 (−0.91)	−0.132* (−1.79)
Education				
No education	Ref	Ref	Ref	Ref
Less than high school	0.502*** (16.30)	0.447*** (14.58)	0.402*** (9.73)	0.711*** (12.83)
High school	1.414*** (39.91)	1.278*** (36.07)	1.338*** (26.96)	1.608*** (24.89)
More than high school	2.728*** (77.05)	2.431*** (67.55)	2.572*** (49.24)	2.880*** (41.21)
Parental controls	No	Yes	Yes	Yes
Number of children	31,701	31,701	16,761	14,940
Number of families	16,176	16,176	11,378	10,607
Log likelihood	−32,887.0	−32,242.2	−17,441.9	−13,078.6

Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

Each random effect regression also includes a set of regional and size of urban unit dummies. Significance levels are, respectively, 1 % (***), 5 % (**), and 10 % (*)

that women have a lower probability of working compared to men, the proportion of women in this subsample is equal to 47.1 % (compared to 49.9 % in the whole sample, see Table 1). We turn to random and fixed effects ordered models (without censoring) to explain children's occupational choices.[22] In column 1 of Table 4, we control for the following characteristics: gender, birth cohort, number of siblings, birth order, and education. Since education itself depends on birth order, this means that our estimates will indicate the direct effect of rank within the sibship (net of education).

As expected, occupation is highly correlated with education. Children having completed more than high school have access to much better occupations compared to low-educated children. We find a negative correlation between occupation and sibship size, at least for children having at least three siblings. Another result is that the coefficients associated to the birth order dummies are all significant and negative. There is clearly a disadvantage for late-born children compared to first borns although the various estimates obtained for the second born, third born, fourth born, and fifth born are not statistically different from each other.

At the same time, compared to education, we note that the *t* values associated with the birth order dummies are much lower for occupation. In fact, both the coefficients and levels of significance strongly increase when estimating the same regression without the child's level of education: they are around four times higher in absolute values. Yet, these findings show that the first borns' higher average education is not the exclusive reason explaining why first borns have better occupations. As shown in column 2, our results are robust to the inclusion of parental characteristics which do not really affect the influence of birth order on occupation.

We also estimated gender-specific regressions. As shown in columns 3 and 4, the birth order coefficients remain negative but they are hardly significant, especially for boys (only one coefficient is significant at the 10 % level). It may be that the birth order effect on occupation observed for boys is essentially accounted for by education. At the same time, as previously emphasized, we do not have enough information to pick up possible differences in study fields between men and women in our regression. Another interpretation is that the birth order effect holds in families comprising both boys and girls. Additional results, not reported, show that this is indeed the case. When selecting the subsample of families with at least one boy and one girl, the various birth order coefficients are all significant at the 5 % level.

We also estimated ordered regressions for occupation by family size (panel B1, Appendix Table 7). The coefficient for the second-born child is negative and significant for two-children families. With three children, the coefficients associated to birth order are negative, but only the third-born child estimate is significant. Conversely, there is no significant effect for families with either four or five children.[23] In panel B2 of Appendix Table 7, we report estimates from fixed effect ordered models to relax the

[22] We also estimated our regressions on the subsample of families where all siblings are employed and reach similar conclusions concerning the role of birth order. We also investigated the relationship between unemployment and family characteristics, with a focus on non-co-resident children interviewed either in 1998, 2004, or 2010 given data constraints. When estimate a random effect Probit model to explain the probability for a child to be unemployed, we find no significant relationship between the probability of being unemployed and birth order.

[23] For three children and higher families, results from Wald tests show that the birth order coefficients are not jointly significant.

assumption of independence between the family level unobserved heterogeneity term and the explanatory variables. Net of the positive impact of education on occupation, the fixed effect estimates confirm the negative correlation between birth order and occupation, this time for any family size.[24] This persistence of a birth rank effect in occupational attainment goes against the findings by Bertoni and Brunello (2013), which show evidence of a catch-up occupational effect among later-born children in a sample drawn from several European countries.[25]

To summarize, our random and fixed effect estimates show that first-born children are more educated and reach better occupations on average compared to later-born siblings, the latter effect persisting to a lesser extent after controlling for education. Since these characteristics translate into higher earnings, an inferred finding of our empirical analysis is that birth order entails substantial financial inequalities between siblings. Our next inquiry is therefore to assess whether parents attempt to reduce these differences in the economic well-being of their children through those financial transfers they may give over the life cycle.

5 Financial transfers and inequalities among siblings

Over the last three decades, economists have emphasized the importance of private family transfers, either financial or in the form of services. As discussed in the survey by Laferrère and Wolff (2006), two main models have been suggested to explain the motivation for these transfers. They are either based on altruism or on reciprocity.

According to the altruistic model proposed by Becker (1991), parents care about the situation of their children. They take into account the well-being of their children when maximizing their own utility function and private transfers are a means to redistributing money between generations. As a consequence, parents should give more money when they have high incomes and when their children experience economic hardships. Further, reducing the income of the donor parent by one euro and increasing the income of a recipient child by one euro should reduce the amount transferred by one euro (Altonji et al. 1997).[26] Also, they should make larger transfers to those siblings with fewer resources. It follows that unequal sharing is expected within the family under altruism unless parents attach more importance to the utility function of some of their children.[27]

[24] At first sight, it could be a little surprising that the random and fixed effect estimates differ since there should be no correlation between birth order and family-level unobservables conditional on family size. However, both regressions are not estimated on the same sample of children. The fixed effect ordered models rely on the estimate of conditional Logit models, so that sibships in which children all have the same occupation are excluded.

[25] Specifically, these authors find that first borns get more education but also argue that by age 30, no birth rank effect persists across siblings. They explain their findings by claiming that later borns are more risk taking than first borns. It remains to be explained why later borns would not outpace first borns after age 30, however.

[26] Hence, the difference in transfer income derivatives has to be equal to minus one, a property called redistributive neutrality (see Laferrère and Wolff 2006).

[27] Alternative explanations have been suggested to explain equal sharing within the family. For instance, parents may suffer from a psychic cost when deviating from an equal allocation of resources (Wilhelm 1996).

The other model involves some reciprocity between generations with several mechanisms of exchange. A first scenario is when parents make financial transfers in exchange (explicit or not) of services and visits provided by their children (Cox 1987; Cox and Rank 1992). When there is no market substitute for the child's attention, the transfer amount can increase with the child's income contrary to what is expected under altruism. Children may be induced to enter into family exchange when they face liquidity constraints (Cox 1990). Parents will lend money to their children and this family loan is paid back later at an interest rate potentially above that of the financial market. Rather than being a substitute for private consumption, parental transfers are a form of investment in the mutuality model of Cigno (1991, 1993). There is a family contract such that adults transfer resources to their children and children pay back to their parents the loan they have previously received when being young.

This part of our empirical analysis studies the determinants of financial transfers made to children. When interpreting our results, it should be kept in mind that we only consider a subset of the various parental transfers flowing to children since we neglect other forms of irregular financial transfers and bequests for which information on recipients is missing in the Wealth surveys. Yet, property and cash transfers are the most significant form of inter vivos transfers as they correspond to substantial transmissions of parental wealth. In 2006, according to the Direction Générale des Finances Publiques, the total amount of taxable transfers was equal to 39.4 billion euros, which is around 50 % lower than the total amount of taxable bequests for the same year (58.9 billion euros). However, there is no obvious way to compare major transfers and other more informal forms of financial transfers as the latter are not taxable and thus not subject to official records.[28]

In what follows, we rely on the 1998, 2004, and 2010 surveys due to the lack of information on recipients in the 1992 survey. Our empirical analysis is based on a sample comprising 32,856 children living by themselves and aged at least 24 (14,322 families). As shown in Table 5, the proportion of children having benefitted from a parental transfer is around 20 % over the period. There are substantial differences in the gift rate over the period, which we attribute to both increased parental wealth across the successive cohorts and changes in French tax incentives for inter vivos transfers. The proportion of recipients was 9.9 % in 1998, 16.8 % in 2004, and 28 % in 2010. This increase in the last survey is likely explained by the adoption of the French TEPA law (bill on labor, employment, and purchasing power) in August 2007, with an increase of tax exemptions for financial transfers made to children from 50,000 to 150,000 euros per child.

In France, the probability of receiving money from parents is negatively correlated with the size of the sibship. The proportion of recipients is 27.1 % with one child, 23.6 % with two, around 19 % with three or four, and 11.6 % with five or more (see Table 5). An interesting question is whether parents help all of their

[28] In the French Wealth surveys, there is no information on the amount per transfer for regular or irregular cash gifts.

Table 5 Receipt of financial transfers

Variables	Year of survey			All
	1998	2004	2010	
Proportion of children receiving money from parents				
All children	0.099	0.168	0.280	0.199
Number of siblings				
0	0.213	0.217	0.340	0.271
1	0.118	0.205	0.314	0.236
2	0.086	0.150	0.277	0.194
3	0.080	0.188	0.256	0.185
≥4	0.059	0.096	0.190	0.116
Proportion of families with unequal sharing				
All families	0.191	0.090	0.086	0.105
Families with at least two children	0.282	0.119	0.112	0.140
Number of siblings				
1	0.290	0.090	0.101	0.128
2	0.291	0.127	0.124	0.149
3	0.250	0.211	0.104	0.160
≥4	0.250	0.118	0.171	0.176
Number of children	8893	9418	14,545	32,856
Number of families	3874	4127	6321	14,322

Source: authors' calculations, INSEE Wealth surveys 1998, 2004, 2010

children when they provide financial transfers. Among the 2848 parents who have made a transfer, only 299 of them have chosen to share their resources unequally by making transfers to only some of their children (10.5 %). Since by definition equal sharing is always observed for one-child families, we calculate the same proportion on the subsample of families with at least two children and find a proportion of 14 %. It follows that equal sharing (not in terms of amounts, but in the sense of all siblings receiving at least some financial transfer) is a behavior which is very frequently observed in France: more than 85 % of donors give money or property to all of their children.

Another result is that unequal sharing is more frequently observed as the sibship size increases. The corresponding frequencies are 12.8 % with two children, 14.9 % with three, 16 % with four, and 17.6 % with more than four. Several explanations may account for this positive correlation. First, it may be that parents cannot afford to give money to all of their children in large families, especially if they are liquidity constrained. Second, if we assume that parents only give money to children old enough, then unequal sharing will be more likely in large families because our cross-section data provides an incomplete picture of intergenerational transfers (younger children being expected to receive money later in life). Third, the probability of an uneven distribution

Table 6 Random effect Probit and fixed effect Logit estimates of transfer receipt

Variables	(1)	(2)	(3)
Characteristics of children			
Female	−0.016 (−0.22)	−0.029 (−0.30)	−0.023 (−0.14)
Age			
Less than 30	Ref	Ref	Ref
30–39	2.220*** (11.79)	0.459*** (2.81)	0.837** (2.19)
40–49	3.450*** (15.94)	0.267 (1.57)	0.484 (0.93)
50+	4.541*** (18.25)	0.170 (0.87)	0.152 (0.23)
Number of siblings			
0	Ref		
1	−1.146*** (−6.76)	Ref	
2	−2.229*** (−11.49)	0.028 (0.23)	
3	−2.435*** (−9.96)	−0.086 (−0.53)	
≥4	−3.358*** (−11.50)	−0.176 (−0.88)	
Birth order			
First-born	Ref	Ref	Ref
Second-born	−0.639*** (−7.54)	−0.919*** (−7.87)	−0.862*** (−5.33)
Third-born	−0.991*** (−7.06)	−1.119*** (−6.43)	−1.318*** (−4.70)
Fourth-born	−1.046*** (−5.08)	−0.920*** (−3.66)	−1.051** (−2.53)
Fifth-born and more	−1.290*** (−4.64)	−0.804** (−2.57)	−1.078* (−1.82)
Education			
No education	Ref	Ref	Ref
Less than high school	0.317 (1.62)	0.271 (1.33)	0.543 (1.21)
High school	0.331 (1.58)	0.290 (1.29)	0.472 (1.00)
More than high school	0.492** (2.46)	0.160 (0.77)	0.271 (0.61)
Characteristics of parents			
Head's age at birth	0.181*** (15.41)	0.003 (0.28)	
Lone-parent family	−0.330*** (−2.64)	0.092 (0.82)	
Blended family	−2.145*** (−6.39)	0.123 (0.46)	
Head's education			
No diploma	Ref	Ref	
Primary	0.367* (1.87)	0.003 (0.02)	
Secondary	0.682*** (3.36)	−0.061 (−0.36)	
High school	1.896*** (7.70)	−0.084 (−0.39)	
>High school	3.741*** (17.26)	0.160 (0.90)	
Random/fixed effects	Random	Random	Fixed
Number of children	32,856	825	825
Number of families	14,322	299	299
Log likelihood	−7414.5	−496.5	−252.4

Source: authors' calculations, INSEE Wealth surveys 1998, 2004, 2010

Each random effect regression also includes a set of regional and size of urban unit dummies and year-specific survey dummies. Significance levels are, respectively, 1 % (***), 5 % (**), and 10 % (*)

may increase with family size if parents give money to their adult children when the latter experience financial problems.[29]

Next, we investigate the role of birth order on the probability of receipt of a parental transfer. For that purpose, we consider the child sample (32,856 observations) and explain the probability for a child to receive a transfer using Probit and Logit models that control for unobserved heterogeneity at the sibship level. Results are shown in Table 6. In column (1), we estimate a random effect Probit model and account for the possible indirect effect of birth order through the educational attainment of children. We find that the probability of receiving a transfer does not depend on gender. Older children are more likely to have received a transfer, which is also the case for children with many siblings and children having completed more than high school. The estimate for birth order is negative, meaning that later-born children are less likely to receive some financial assistance. Concerning the role of parental characteristics, the probability of a transfer increases with age and education of the head, while it is significantly lower among non-intact families.

Our results suggest that parents do not use financial transfers to offset the effect of birth order. At the same time, most parents do not share their wealth unequally between their various children. As a consequence, we decide to focus on the subsample of families in which parents select some of their children as recipients. By definition, this leads to an exclusion of the following types of families: those with only one child, those without any transfer and those where all children receive money. This reduces the sample size to 825 children, corresponding to 299 families. We first estimate a random effect regression for those families characterized by an unequal allocation of family resources (column 2). As expected, the birth order effect is still found for those families where parents share their financial resources unequally.[30]

In column (3), we allow for some correlation between the family-specific component and the covariates and estimate a fixed effect model. The corresponding specification is the conditional Logit model described in Chamberlain (1980). By definition, the conditional likelihood approach excludes those families where parents make transfers to all of their children and those where parents do not provide any cash gift. The fixed effect estimates confirm that within the sibship, first-born children have a significantly higher probability of receiving transfers compared to their younger siblings. As we control for the child's age in our regression, this first-born effect for private transfers does not come from a censoring issue which would occur if children receive at a certain age and if there is a larger proportion of first born who reach this age in our sample.

[29] Assuming that the probability of a child to be in financial problems is p, then the probability of transferring to all children is p^n for a family with n children. The probability of an uneven distribution is $1-p^n-(1-p)^n$, which is an increasing function of n.

[30] Children aged between 30 and 39 are more likely to receive a gift from their parents, while the other characteristics have no significant influence (presumably due to small sample size).

In order to further assess whether parents compound (or offset) the effect of birth order through their transfer decisions, we decided to re-estimate both the education and occupation regressions on the subsample of families characterized by unequal sharing. Our results, not reported here, have nonetheless to be interpreted with caution given the small sample size. Both for education and occupation, the random effect ordered regressions show that the birth order dummies play no significant role. This, however, does not evidence a substitution effect between an advantage conferred earlier in life and one conferred later on. Including interaction effects between the presence of unequal transfers among siblings and the birth order dummies in the whole child sample revealed no significant offset. This suggests that the advantage in transfers is uncorrelated with the likelihood of receiving more education or having better occupations than the other siblings.

We performed several additional robustness checks. First, we estimated linear probability models to explain the receipt of financial transfer since the conditional likelihood estimation throws away most observations. Using the sample of 32,856 children, results from either random or fixed effect linear regressions show the same negative correlation between birth order and the receipt of a financial transfer.[31] Second, we selected a subsample of older parents, age 70 and above. As they grow older, it is less likely that they will leave a transfer differential if they ever want to equalize transfers among siblings. Again, we find negative coefficients for the birth order dummies, which rules out the possibility that parents give more money to their older children first and compensate the later born children afterwards.

It is worth making the connection between our results and the theoretical literature on intergenerational transfers. Clearly, our estimates go against the predictions of the altruistic framework according to which parents should help their less well-off children more. Instead, results from the random effect specification show a positive correlation between the probability of receiving money and education, meaning that parents are more likely to transfer resources to children with a higher permanent income. Also, in most cases, parents give money to all of their children rather than favoring those with higher needs. This is consistent with previous evidence on the pattern of inter vivos transfers in France (Wolff 2000). An explanation of the birth order effect could be that parents invest more in the first born so as to receive support and care during old age.

By devoting more resources to the first born, it may be that parents expect some compensation later. Their behavior could be interpreted as a way to induce and secure the provision of transfers from children in case of parental needs. In France, empirical studies on caregiving decisions have shown that time-related transfers to elders were influenced by the situation of parents and the children's availability (Jellal and Wolff 2002). Looking at families with two children, Fontaine et al. (2009) suggest the existence of different expectations in terms of filial duty according to the birth rank and the gender of each child.[32] Unfortunately, the lack of information about upstream transfers in our data does not allow us to further investigate whether the higher

[31] These additional results are available upon request. In the fixed effect linear regression, both gender and education of the child have no significant influence on receiving money from parents.

[32] A few papers have investigated the existence of interactions between siblings in long-term care decisions (Heidemann and Stern 1999; Engers and Stern 2002; Byrne et al. 2009; Fontaine et al. 2009).

investment made by parents in first-born children is effectively related to some intertemporal exchange.

6 Conclusion

Using French surveys covering the period from 1992 to 2010, we have shown that first-born children are more likely to achieve higher levels of education, better occupations, and higher likelihood of receiving financial transfers, although there are only few families where parents do not make transfers to all of their children. Our results challenge those theories that explain the well-documented birth order advantage through mechanical effects. While first-born children spend more time with exclusive adult surroundings and benefit longer from undiluted parental resources, parents do have the opportunity to equalize transfers later in life and even to offset the first-born effect through asymmetric transfers. Far from it, our data indicates that parents are more likely to make transfers to the first born but the advantage in transfers is uncorrelated with the likelihood of having higher education or better occupation.[33]

While it is possible that parents do not realize that first born children get a natural advantage early on, it is less likely that parents would be so myopic later in life. Assuming the mechanism at play is similar at both stages, our results shed new light on the possible source of the birth order effect and we must be prepared to consider a persistent bias over the life cycle toward first born children. While our findings on financial transfers clearly rule out a parental motivation based on altruism, it may be that parents invest more in their first born as part of an exchange strategy. In a highly forward-looking perspective, parents could devote more resources to first born in order to receive caregiving during old age. In Japan for instance, Kureishi and Wakabayashi (2010) have shown that first-born children were more likely to live with their elderly parents and that childcare assistance was one of the factors explaining the residential location choice of siblings.[34]

Our results offer new perspectives for research on intra-family transfers. However, a deeper analysis of the allocation of parental resources between siblings will require detailed information on all private transfers that flow within the family, whether descending or ascending, in-kind (through contact and services), or financial. With such detailed data over the life cycle of each family member, it would be possible to better understand why parents favor some of their children. At that stage, a large number of hypotheses await further testing.

Acknowledgment We are indebted to two anonymous reviewers for their very helpful comments and suggestions on previous versions of this paper. We also thank Hippolyte d'Albis, Carole Bonnet, Laurent Gobillon, Robert Pollak, Anne Solaz, and seminar participants at the Paris Seminar in Economic Demography, McMaster University, the University of Guelph, Ryerson University, and L'Université du Québec in Montréal. Any remaining errors are ours.

[33] Data limitations do not allow us to go further. In particular, we would like to know the amounts at stake to quantify the transfer differentials and compare those to other siblings' outcome differentials. While we can only observe the presence of financial transfers, we are not able to compare amounts as well as other forms of cash gifts across siblings.

[34] At the same time, there may be strategic considerations within the sibship to avoid the burden of providing care to parents. See in particular Konrad et al. (2002) and Stern (2014).

Appendix

Table 7 Ordered Probit estimates for education and occupation, by family size

Variables	(1) All families	(2) Two-children families	(3) Three-children families	(4) Four-children families	(5) Five-children families
Panel A. Education					
A1. Random effect ordered estimates					
Birth order					
First-born	Ref	Ref	Ref	Ref	Ref
Second-born	−0.198*** (−12.39)	−0.246*** (−9.50)	−0.176*** (−6.08)	−0.226*** (−5.27)	−0.045 (−0.68)
Third-born	−0.317*** (−13.60)		−0.335*** (−8.35)	−0.292*** (−5.74)	−0.244*** (−3.27)
Fourth-born	−0.363*** (−10.92)			−0.314*** (−4.67)	−0.173* (−1.88)
Fifth-born	−0.493*** (−11.88)				−0.096 (−0.81)
Family controls	Yes	Yes	Yes	Yes	Yes
A2. Fixed effect ordered estimates					
Birth order					
First-born	Ref	Ref	Ref	Ref	Ref
Second-born	−0.277*** (−11.44)	−0.262*** (−5.78)	−0.260*** (−6.18)	−0.428*** (−6.90)	−0.022 (−0.23)
Third-born	−0.406*** (−10.75)		−0.450*** (−6.86)	−0.552*** (−6.96)	−0.298*** (−2.68)
Fourth-born	−0.345*** (−6.34)			−0.500*** (−4.39)	−0.034 (−0.25)
Fifth-born	−0.412*** (−5.73)				0.191 (1.03)
Family controls	Yes	Yes	Yes	Yes	Yes

Table 7 (continued)

Variables	(1) All families	(2) Two-children families	(3) Three-children families	(4) Four-children families	(5) Five-children families
Panel B. Occupation					
B1. Random effect ordered estimates					
Birth order					
First-born	Ref	Ref	Ref	Ref	Ref
Second-born	−0.038** (−2.26)	−0.088*** (−3.16)	−0.033 (−1.00)	0.046 (0.93)	−0.131* (−1.69)
Third-born	−0.050** (−2.13)		−0.084** (−1.97)	0.003 (0.05)	−0.042 (−0.50)
Fourth-born	−0.087*** (−2.61)			−0.008 (−0.12)	−0.063 (−0.63)
Fifth-born	−0.072* (−1.81)				−0.032 (−0.26)
Family controls	Yes	Yes	Yes	Yes	Yes
B2. Fixed effect ordered estimates					
Birth order					
First-born	Ref	Ref	Ref	Ref	Ref
Second-born	−0.193*** (−4.85)	−0.245*** (−3.09)	−0.123* (−1.71)	−0.050 (−0.51)	−0.303** (−1.99)
Third-born	−0.290*** (−4.79)		−0.308*** (−2.77)	−0.172 (−1.39)	−0.397** (−2.24)
Fourth-born	−0.360*** (−4.23)			−0.325* (−1.87)	−0.520** (−2.36)
Fifth-born	−0.328*** (−2.95)				−0.727*** (−2.58)
Family controls	Yes	Yes	Yes	Yes	Yes

Source: authors' calculations, INSEE Wealth surveys 1992, 1998, 2004, 2010

Significance levels are, respectively, 1 % (***), 5 % (**), and 10 % (*). The full set of covariates is described in Table 3 for education and in Table 4 for occupation

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