California Center for Population Research
University of California - Los Angeles

Sibship Size, Family Organization and Children's Education in South Africa: Black-White Variations

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CCPR-045-05

October 2005

California Center for Population Research
On-Line Working Paper Series

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*An early version of this article was presented at the International Sociological Association Research Committee on Social Stratification and Mobility (RC28), Los Angeles, August 2005. I am grateful to Donald Treiman, William Mason, and Judith Seltzer for their helpful comments that greatly improve the manuscript. I thank Donald Treiman for providing access to his South Africa survey data.

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ABSTRACT

Recent studies suggest that the generally observed negative sibsize-education association is much less consistent in developing nations, partly because of different cultural customs reflected in family organization. Using data from a national survey in the early 1990s and from the 1996 census, the present study assesses the effect of number of siblings on education in South Africa. In a multi-level framework, I link family arrangements to the sibship size effect on education for two major population groups with distinctive family arrangements, Whites and Blacks. A negative effect exists for Whites, who have adopted a Western nuclear family system, whereas no effect is shown for Blacks, whose family life operates under extended family organization. The study goes beyond previous efforts by explicitly testing the hypothesis that it is extended family arrangements that protect children from negative sibship size effects: results show that the absence of a negative sibship size effect is restricted to extended households; in Black nuclear and fostering families, by contrast, the negative effect holds just as it does for White families. Sensitivity tests are performed to gauge the extent to which the observed sibship size effect is contaminated by endogeneity and a confounding birth order effect. Results suggest the observed differential sibship size effect is relatively robust for Blacks, whereas for Whites it tends to be exaggerated by endogenous factors.

SIBSHIP SIZE, FAMILY ORGANIZATION AND CHILDREN'S EDUCATION IN SOUTH AFRICA: BLACK-WHITE VARIATIONS[1]

INTRODUCTION

Sibship size, also referred to as the number of siblings or family size, is often studied as a determinant of children's schooling (Blake 1989; Blau and Duncan 1967; Featherman and Hauser 1978; Mare and Chen 1986). Considerable evidence has documented that sibship size has a negative effect on children's education in developed societies. The primary mechanism thought to account for this effect is the resource-dilution process: finite family resources are allocated thinly across a larger number of children, meaning that the available resources to each child are diluted (Downey 1995). In the developing world, however, the extent to which such a mechanism operates is conditioned by specific cultural, socioeconomic and political settings. For example, evidence in the developing settings suggests that the nuclear family organization as a main source of support tends to be a prerequisite for sibship size to have a negative effect (Lloyd 1994; Shavit and Pierce 1991).

Scholars and policymakers have long been determined to identify the familial conditions under which the detrimental effect of high fertility can be minimized. However, research incorporating both family organization and fertility structure is only an emerging field. Previous

[1] South Africa has four official racial groups: Blacks (indigenous Africans), Whites (mostly Europeans), Coloreds (mixed-race population) and Asians (Indian descendants). They constitute 76%, 13%, 8% and 3% of the population, respectively. The present study mostly focuses on Whites and Blacks in South Africa, who together account for almost 90% of the total population and provide the clearest contrast. Also, I restrict the sample because one of the datasets used here does not provide sufficient cases for the other two racial groups.

literature has largely neglected the role of extended families in providing support for children, and even fewer studies examine the role of family organization in mediating the sibship size effect on children's well-being. Rather, attention has been largely paid to the nuclear unit, by emphasizing the effect of parental presence and household headship (Anderson 2003; Townsend et al. 2002). Even the existing research that studies the mediating effect of family organization on sibship size is conducted in a less direct way by including a single variable indicating extended kinship (Shavit and Pierce 1991). The present study aims to fill this gap by explicitly testing the intermediate role of family arrangements in the sibsize-education association through estimating separate models by different types of family organization. The analysis is conducted in an African setting, where cultural customs and family arrangements are distinct from those of industrialized societies.

South Africa provides an instructive setting from the developing world for exploring the sibship size effect on children's education. From a racial stratification perspective, South Africa offers a distinctive social, economic and political context compared to the rest of sub-Saharan Africa. Research on South Africa, therefore, will enhance our understanding of the educational stratification process in a racialized developing society. In this diverse country, substantial racial disparities exist in almost every socioeconomic aspect, with the White-Black gap being the sharpest. These disparities were especially salient during the *apartheid* period, and they have more or less persisted even after the breakdown of the *apartheid* system (Treiman, Mckeever, and Fodor 1996; Treiman forthcoming). Black children have experienced disadvantageous socioeconomic conditions, primarily due to resource constraints. Meanwhile, the government consistently shifted educational costs to Black families (Maharaj, Kaufman, and Richter 2000). The availability of resources and their subsequent allocation within families, therefore, tends to

be crucial for Blacks' education. Considering the scarce resources available and the larger share of educational expenses for Blacks, one may well expect that children's education will be detrimentally affected by sibship size because the already limited resources are even more thinly diluted when there are more children. However, in Black families, the costs of raising children do not fall solely on parents, but are extended to related kin (Amoateng 2004). The effect of sibship size in South Africa, therefore, may exhibit a distinctive pattern compared to what has been generally observed in Western societies. Additionally, this effect is likely to vary across different racial groups marked by distinct family arrangements.

Using data from two sources—a national probability sample survey conducted in 1994 and a 10% sample from the 1996 census—I study the effect of sibship size on children's education among Whites and Blacks in South Africa, taking into account the mediating role of family organization. The present study also seeks to provide a comprehensive picture of racial differences in schooling based on two measures of education—ultimate educational attainment and current school enrollment. Additionally, in response to the claim that the sibship size effect is exaggerated by the confounding birth order effect, I incorporate the birth order aspect through decomposing sibship size by birth order and modeling the sibship size effect stratified by birth order. I further conduct sensitivity tests to assess the possibility that the sibship size effect and the role of family organization are plagued by endogenous parental quality-quantity calculation, and the selection into different family arrangements.

SIBSHIP SIZE AND CHILDREN'S WELLBEING IN DEVELOPING COUNTRIES

Studies conducted in Western industrialized societies show a clear negative effect of sibship size on children's educational attainment: each additional sibling reduces schooling by as much as one fifth of a year, even controlling for family socioeconomic background (Blake 1989; Blau and

Duncan 1967; Featherman and Hauser 1978; Mare and Chen 1986). The most widespread explanation for this inverse sibship size effect is the resource-dilution hypothesis (Blake 1989; Downey 1995). "Resources" include nonmaterial assets such as parental time, attention, and emotional support, as well as material assets such as financial investments in children's education and study environments. These resources are less effective as sibship size increases.

However, emerging evidence in the developing world suggests that this effect is not universal; rather, it varies in different socioeconomic, political and cultural contexts, and across different population subgroups (Buchmann and Hannum 2001; Lloyd 1993; Lloyd 1994; Lu and Treiman 2005a; Maralani 2004; Steelman et al. 2002). For example, Maralani (2004) and Gomes (1984) show that sibship size interacts with the level of socioeconomic development over time; and Lu and Treiman (2005a) suggest that political institution is a major intervening variable in the sibship size effect—when state policies emphasize educational equality, the negative effect on education disappears (for a detailed review, see Lu and Treiman 2005a and Steelman et al. 2002).

In a study by Shavit and Pierce (1991), the sibship size effect is shown to depend on family organization that determines resource flows in Israel. For Jews, who conform to the western mode of family organization, the sibship size effect on educational attainment remains detrimental because obligations are generally restricted to the nuclear family. In contrast, for Arabs, who inherit a culture of collective responsibility in which responsibilities for supporting children extend beyond parents to include extended kin, the negative effect disappears.

The Israeli example suggests that an extended kinship system can somewhat alleviate the resource competition generated by a large number of siblings. Such a support system enables large families to cope with the burden of high fertility through an increase in the overall available

resources, which subsequently leads to more extensive allocation of family resources. Similarly, Buchmann (2000), Desai (1992), Lloyd (1993) and Pong (1996) find children's wellbeing is negatively associated with the presence of siblings in Latin America, but is unrelated to sibship size in sub-Saharan Africa and south-east Asian, where extended family organization is extensive.

The present study aims to explicitly examine this mediating role of family arrangements in the sibship size effect by drawing on the unique case of South Africa. Previous studies have often examined this mediating effect implicitly by including a single variable indicating extended kinship. In contrast, I carried out more explicit analysis by examining interactions with family organization and sibship size, and by estimating separate models for each type of family arrangements.

THE CASE OF SOUTH AFRICA

South Africa is distinctive not only among African countries, but more generally in the contemporary world because its social and political institutions were organized exclusively on the basis of race until 1994. Particularly between 1948 and 1994, an *apartheid* system was legally constructed by the government to ensure the supremacy of the minority White population at the expense of other racial groups, particularly Blacks. As a result, the four racial groups constitute a clear socioeconomic hierarchy far larger than in any other multiracial countries, with Whites at the top, Blacks at the bottom, and Asians and Coloreds in between (Treiman et al. 1996; Treiman forthcoming). According to the 1999 South Africa October Household Survey, for example, Blacks make up 76% of the total population, whereas they account for 95% of the poor (Woolard 2002). Importantly, Blacks are distinctive from Whites with respect to fertility, family arrangements, and education.

Fertility

Africans, who have the lowest per capita income, have the highest fertility rate; the White population, which has the highest per capita income, has the lowest fertility rate (Chimere-Dan 1997; Swartz 2002). Following the demographic transition in the developing world, fertility began to decline among all population groups in South Africa since the 1960s. Meanwhile, measures were undertaken by government directly to increase the size of White population by officially encouraging White women to have more children while reducing the size of Black population (Chimere-Dan 1993). However, the Black fertility decline occurred at a much slower pace for Africans as compared to other racial groups (Swartz 2002). By 1990, Black fertility, measured by the total fertility rate (TFR), was around 4 to 4.5, dropping from about 6.5 some 30 years before, while the White fertility has undergone a sustained decline to a TFR of 1.9 (Caldwell and Caldwell 1993; Swartz 2002).

There are several explanations why the Black fertility did not respond more strongly to socioeconomic advancement and family planning programs. In African countries, Blacks are found to hold radically different notions about reproduction from those in the West: they desire children intensively and reject the desirability of reducing family size (Bledsoe, Banja, and Hill 1998; Preston-Whyte 1988). This is partly due to economic reasons such as dependence on children in old age and need of children's labor to provide for family subsistence (Kaufman 2000). Other explanations for the modest fertility decline posit profound cultural and social differences between Whites and Blacks in South Africa. The number of children also depends on what is rational to a woman. Research indicates that women's conjugal life depends heavily on their ability to produce children because having children elevates their status as women in the community, sometimes even bringing them material benefits (Kaufman 2000). Schoen (1978)

suggest another explanation: South Africa is a hierarchical society stratified by race. Thus, Blacks, even if they limit fertility, can hardly expect themselves and their offspring to achieve upward mobility. Since there are no gains from a small family, Blacks are not inclined to reduce fertility. The persistent high fertility may also be due to the oppression of Black women under *apartheid* (Chimere-Dan 1993): women are not yet able to achieve autonomy and obtain adequate education sufficient to use contraceptives and control fertility.

Consequently, government promotion of low fertility and the diffusion of birth-control information are unlikely to substantially reduce Black fertility, even though birth control is occasionally adopted to achieve birth spacing (Caldwell, Orubuloye, and Caldwell 1992). Importantly, these unique cultural and social contexts tend to alleviate the degree of endogeneity resulting from a quality-quantity tradeoff (Steelman et al. 2002), discussed below.

Family Organization[2]

South Africa is characterized by two prominent family systems: an extended family system, where family obligations are spread beyond nuclear unit (parents and their children) to include

[2] Traditional definition of extended families refers to extended families sharing the same household. Yet, along with modernization and urbanization, the definition has been modified to include extended families living apart but keeping contact by other means. While it is a possible trend, in South Africa, there is no empirical evidence that modified extended families have replaced traditional extended families. Also, most studies examine extended families in terms of the traditional type because information on non-coresident kin is usually not collected in surveys. The present study experienced similar difficulties, thus the traditional definition is used.

relatives, is clearly identified with Blacks[3]; and a nuclear family system is identified with Whites, with Coloreds and Asians lying between these two systems (Amoateng 2004; Thompson 1990). The nuclear families adopted by Whites are based on the cultural value of individualism, whereas African extended families are based on a value system that emphasizes interdependence and collectivism (Siqwana-Ndulo 1998). Given this cultural emphasis and the disadvantageous socioeconomic position held by Blacks, the African extended kinship operates as a way of coping with vulnerability by pooling resources and providing assistance when needed (Amoateng 1997; Makosana 2001; Thomas 1996).

Along with the industrialization and urbanization process, debate emerges on whether the Black family patterns are converging towards those of Whites. Some argue that family unity has been weakened as a result of rural-urban migration, separation of family members, and participation in urban modernization in the form of wage employment and formal education (Amoateng 1997). The African extended family arrangements thus tend to evolve into a nuclear type, which is perceived to meet the challenges of urban civilization better. However, abundant results demonstrate that the Blacks' cultural preference for extended living arrangements persists in spite of their participation in modernization and urbanization (Amoateng 2004; Russell 2004; Ziehl 2001). Smit (2001) illustrates situations under which Black families utilize traditional values and survival strategies actively to adapt to economic and social changes, and to ensure that modernization does not uproot their traditional way of life. As Burman and Fuchs (1986) demonstrated, in the early 1980s, just over half of African households were nuclear in structure,

[3] There are two types of extended family organization in South Africa: one is extended families where children live with both parents and extended kin; the other is fostering families where children live with only extended kin.

whereas most of the rest are extended in structure. Additionally, while there are claims that urban Black families are more likely to adopt the nuclear organization than are rural families, little rural-urban variation was found in an early 1990s study (Russell 1994).

Furthermore, evidence shows that a noticeable proportion of children live with neither parents but are fostered by extended family members, grandparents and uncles in particular. The elderly grandparents are able to provide resources by supplying the social pension they receive[4] (Case and Deaton 1998). This fosterage arrangement is another aspect of the African extended family arrangements, partly due to high rates of parents absence as a result of nonmarital childbearing, divorce, death of parents from HIV and labor migration (Gordon and Spiegel 1993; Niehaus 1994; Noumbissi and Zuberi 2001; Preston-Whyte 1993; van de Walle 1999). This arrangement serves a variety of functions, including socialization of the child, deepening the relationship among kin, and financial help during times of distress (McDaniel and Zulu 1996).

It should be acknowledged that choice of extended family arrangements may be endogenous to children's schooling: parents may choose to live with extended kin or foster their children to relatives in order to gain access to school or to obtain better education. Garg and

[4] Indeed, South Africa is one of the rare African countries where economic support for the elderly exists. All women aged 60 or more and men aged 65 or more are eligible for a significant social pension of about $100 per month. The availability of social pensions may affect living arrangements in that economically deprived families may move to live with pension recipients. Yet, this is less problematic in the current study because the pension system was operating fully in all areas only since 1993, about the time the data used here were collected; and African families were living in three generation households long before the pension became universal (Case and Deaton 1998).

Morduch (1998) address this issue using detailed fertility histories, but find little qualitative change. The present study will assess this endogeneity issue in the analysis.

Educational Institutions[5]

The implementation of *apartheid* in South Africa institutionalized separate and unequal development by racial groups with respect to education. Starting with the Bantu Education Act of 1953, educational policies, curriculum and pedagogical practices were designed specifically to ensure the political, economic and social domination of the White population over other racial groups (Constas 1997). Consequently, Black children have been provided with limited educational opportunities, and a separate education system with lower quality.

Specifically, Black schools were provided with drastically less funding than White schools—government spent approximately ten times as much on schooling for each White child as for each African child; colored and Asian children received intermediate amounts (Moll 1996; Maharaj et al. 2000). Until recently, Blacks were the only group in South Africa that had to pay in full for its education. The burden of funding African education, therefore, was largely placed on African families, the poorest part of the population, making Black family resources a key determinant of children's education. The unequal distribution of public educational resources, coupled with preexisting economic constraints in Black families, has resulted in limited educational opportunities and lower attainment for Blacks (Lam 1999; Thomas 1996; Treiman et al. 1996; Treiman forthcoming). For example, Zungu (1977) and Moll (1996) found that Blacks started school later, they had much lower enrollment rates, and most of those who attended did

[5] In South Africa, for each racial group, primary school took seven years (substandards A and B, standards 1-5) beginning at the age of six or seven, and secondary schooling five years (standards 6-10).

not progress to secondary school; by contrast, most Whites had completed twelve years of schooling. Costs of education are often reported by Blacks as the primary reason for not enrolling in school (Case and Deaton 1999). Also, Anderson (2000) suggests that the lower educational attainment for Blacks results partly from their slow progress in school: given the limited educational resources, grade repetition and school interruption are commonly observed for Black children.

In recent years, the government has attempted to alleviate the shortage of skilled Black labor through an expansion of Black education. This effort has led to a major increase in Black primary school enrollment, but the increase has not been extended to the secondary level (Maharaj et al. 2000). Only as recently as 1995 was education made compulsory for Blacks between age seven and sixteen, which is not yet achieved in South Africa (Nkabinde 1997).

Although enrollment rates have somewhat improved, the quality of Black schools remains a problem, reflected in uneven distribution of educational resources favoring Whites and low teacher-pupil ratios (Townsend et al. 2002). Case and Yogo (1999) show that school quality significantly affects educational attainment in South Africa.

HYPOTHESES

In the absence of adequate state funding in South Africa, Black families have played a crucial role in funding children's education. Considering the scarce resources available to Black families, and the larger burden of educational expenses placed on them, we may well expect that sibship size has a tremendous negative effect on children's education, by further diluting the already scarce resources.

However, among South African Blacks, the burden and financial costs of raising children are not borne exclusively by biological parents; rather, they may be shared by extended family

members. There is some evidence that extended family arrangements benefit children: they help provide a pool of adults who can supervise children, and offer positive reinforcement and help them with school work (enhance non-material resources); they help provide additional support and protection in case of the absence of the parents; and most importantly, they help offer financial assistance by resource pooling (enhance material resources). As a result, the extended arrangements raise the total resources available to the entire sibship, and subsequently allocate more broadly these familial educational resources. The resource-dilution process thus tends to be attenuated in these families, enabling a greater number of children to be educated than if only the resources of the children's biological parents are used. In other words, under such circumstances, sibship size tends to be less consequential in determining a child's education. Given that a significant number of Blacks live in extended living arrangements[6], I expect to see little or no association between sibship size and children's education for Blacks.

By contrast, in nuclear families where educational costs are restricted to parents only, no additional resources are available to mitigate the resource dilution associated with a large sibship size; that is, familial resources tend to be more thinly diluted as sibship size increases. This aggravated resource-dilution process implies a detrimental effect of sibship size in nuclear families. Given that Whites live predominantly in nuclear families, I speculate there is a negative

[6] Although both belong to the extended family arrangements, extended families (families with the nuclear unit and other relatives) and fostering families (families with only children and other relatives but no parents) are examined separately in this paper. The reasons for making this distinction are discussed later in this section. Throughout this paper, the term "extended family arrangements" is used to denote both types of families, whereas the term "extended families" refers only to families with both the nuclear unit and other relatives.

sibship size effect for Whites in general. These ideas can be summarized in the following hypotheses.

Hypothesis 1: In South Africa, there is no clear sibship size effect on education for Blacks in general due to the buffering effect of extended family arrangements.

Hypothesis 2: For Whites in general, sibship size is negatively associated with education because they rely nearly exclusively on the nuclear family organization.

To explicitly examine whether the lack of a sibship size effect for Blacks is due to their distinct extended family arrangements, I generated another set of hypotheses broken down by family organization[7]. Based on the logic presented above, I expect:

Hypothesis 3: For Blacks living in extended family arrangements, no sibship size effect exists.

Hypothesis 4: For Blacks living in nuclear families, there is a negative sibship size effect, as observed for Whites.

The above hypotheses deal with nuclear and extended family arrangements in general, whereas the following hypotheses make distinctions between two types of nuclear arrangements (single-parent nuclear vs. two-parent nuclear families) and two types of extended arrangements (extended families vs. fostering families).

First, the role of fostering families, one type of extended family arrangements, is mixed and needs to be investigated separately. In general, households are more likely to invest and usually invest more in more closely related children, because these children are considered to be more likely to provide transfers to the households later in life (Becker 1991). This suggests that

[7] These hypotheses are restricted to Blacks because nuclear and extended arrangements are both common for Blacks, while almost all Whites live in nuclear arrangements.

fostering families do not have many incentives to invest in foster children, and they tend to allocate fewer resources to foster children than to their own biological children. A recent study in South Africa shows that household relatedness to the focal child is closely and positively associated with household expenditures on that child (Anderson 2005).

Therefore, although fostering families are able to provide more resources than otherwise can be offered by children's nuclear families (particular when fosterage is chosen due to economic constraints), unequal allocation of these resources favoring biological children often occurs. This, in fact, means that foster families may actually invest fewer resources in fostering children than would comparable nuclear families, which is likely to aggravate the resource-dilution process and thus to lead to a negative sibship size effect. This unequal allocation implies the important role of parental presence: when parents are absent in fostering families, they lose control of the resource distribution process, resulting in smaller allocations to their own children. By contrast, in extended families, where both parents and extended relatives are present, the resources can be assumed to be evenly distributed among children.

Hypothesis 5: For Blacks living in fostering families, the sibship size effect tends to be different from that in extended families due to the absence of children's biological parents; that is, there is a negative sibship size effect in fostering families, whereas no such effect exists in extended families.

Second, the distinction between single-parent and two-parent nuclear families is critical. Single-parent families are found to be more common among Blacks than among Whites in South Africa (Zulu and Sibanda 2001). Also, as suggested by Biblarz and Raftery (1999: 323), "almost all existing theory about the consequences of family structure for children centers around the relationship between family type and resources": in both the developed and developing world

children from single parent families obtain fewer economic, social and cultural resources of the kind that facilitate success, presumably because income is reduced. The absence of a parent also implies fewer non-economic resources such as time and attention spent on children. As a result, children growing up with only one parent are disadvantaged across a broad range of outcomes (Case, Lin, and McLanahan 1999; McLanahan and Sandefur 1994; Townsend et al. 2002). Based on the resource-dilution hypothesis, the already limited resources in single-parent nuclear families are even more thinly diluted when sibship size is large, leading to a more detrimental sibship size effect on children's schooling.

Hypothesis 6: The sibship size effect is more pronounced in single-parent nuclear families than in two-parent nuclear families.

SIBSHIP SIZE, BIRTH ORDER AND THEIR INTERACTIVE EFFECT

While the sibship size effect assumes that parental resources are divided equally between all siblings within a family, some have argued that this is not necessarily the case. There has been increasing interest in the effects of sibling configurations other than sibship size, birth order in particular, which focus on within-family inequalities in resource allocation. However, the birth order effect has been much less reliably documented than the sibship size effect, and has often been shown to be negligible (Hauser and Sewell 1985; Kessler 1991; Kuo and Hauser 1997).

Additionally, some studies have claimed that the effect of sibship size has been exaggerated by the confounding effect of birth order (Black, Devereux, and Salvanes 2005). That is, being in a small sibship is found to be advantageous, but this effect may be due to the higher probability of being early in the birth order. Thus, according to this claim, the negative sibship size effect is actually an artifact of the negative birth order effect. Studies of this kind, however, are less satisfactory because they include both birth order and sibship size in a single model

without accounting for the multicolinearity between these two measures (correlations of 0.7 between such measures are not uncommon). This multicolinearity tends to moderate the sibship size effect.

A preferred way to study the effects of both factors is to decompose the sibship size effect into its meaningful components. This strategy can be implemented by splitting the number of siblings into the number of older siblings vs. younger siblings, as implemented by Chu, Yu, and Tsay (2004) and Post and Pong (1998). When entering the number of older siblings and that of younger siblings, the net effect of sibship size has been essentially controlled for. Yet, after decomposition, it is often hard to evaluate the effect of sibship size alone when the effects of number of older siblings and younger siblings are in the opposite direction. Hence, to examine the sibship size effect while controlling for the potential confounding effect of birth order, another possibility is to model the sibship size effect at each birth order, as implemented in Black et al. (2005). If the effect remains stable after controlling for birth order, then it is not likely to be a simple artifact.

To provide a more complete picture, the present study examines the extent to which findings including a single sibship size measure are contaminated due to the exclusion of birth order. Specifically, efforts are made to disentangle the possible interactive effects of sibship size and birth order through decomposition and through estimating the sibship size effect stratified by birth order. In addition, whenever possible, the effect of birth order on children's education is evaluated.

DATA AND METHODS

Two Data Sets

Two dataset are used in the analysis, each with its own strengths and limitations, but which can provide a more complete picture when jointly studied. The first data come from a national probability sample survey of the adult population in South Africa conducted in the early 1990's, the Survey of Socioeconomic Opportunity and Achievement (SSOA) (Treiman, Moeno and Schlemmer 1998). Completed interviews were obtained from a stratified random sample of 9,086 persons age 20 and older, including the TVBC states[8]. After excluding a Black elite sample and appropriately weighting the data, a total sample of 8,714 can be regarded as representative of the South Africa population. This survey collected extensive life history information, such as family background and residential status when respondents were at age 14.

An important feature of the data is that it permits studying the effect of sibship size and other socioeconomic attributes measured when respondents were young on their ultimate educational attainment. This largely reduces the problem of temporal ambiguity: rather than resulting from education, the explanatory variables can be considered as the factors essentially influencing ultimate education. A limitation of the data, however, is that no retrospective

[8] During the 1960s and 1970s, 3.5 million of Blacks were officially assigned by the government to move away from the cities to their "homelands" according to their origin (which was frequently inaccurate). In the 1970s, four of the 10 homelands (the TVBC States: Transkei, Venda, Bophutatswana, and Ceskei) were set up to be independent from South Africa (Treiman forthcoming). Blacks living in the rural homelands and particularly the TVBC States were disproportionately impoverished even compared to the Black population as a whole. Under *apartheid*, the TVBC States were excluded from the South African census and often from other surveys.

information was collected on respondent's living arrangements. The data can only distinguish whether respondents lived with at least one of their parents when they were young.

Therefore, I use a second dataset: a 10% sample of the 1996 population census of South Africa, which was carried out two years after the end of *apartheid*. Including the former TVBC states, it was the first to cover all racial groups of the population equally and comprehensively. It also separated households of domestic servants from those of their employers. In addition, the census provides a much larger number of cases (N=3,621,201). Also, information on current living arrangements for school-aged children can be obtained from the household roster. Unlike the survey data that use educational attainment as the outcome variable, the census data can only use current enrollment status because most children will not have completed their schooling. Yet, it should be noted that examination of enrollment status may lead to a problem of censored observations. For example, we do not know if children who are currently attending school will drop out at some later stage. Also, this kind of analysis may obscure differences between permanent school leaving and short-term school interruptions. Ultimate educational attainment and current enrollment status, therefore, cannot be considered as identical measures of schooling. I thus expect that they will display somewhat dissimilar patterns. Yet, basing the analysis on both measures helps provide a more complete view of racial differences in schooling in South Africa.

Variables and Methods Using the Survey Data

Using the survey data, ordinary least square (OLS) techniques are used. First, I estimate the effects of sibship size for the overall South African population[9]. To check whether the sibship

[9] Censoring is a limitation for both the survey and census analyses, since some of the key variables (family arrangements and sibship size) are likely to vary over time. Ideally, event history analysis should be preferable. Yet, the data do not permit doing this.

size effect differs by race, I then include and test a set of race by sibship size interactions in the OLS regression. Finally, since the hypothesis of homogeneity by race is rejected, I estimate separate models for Whites and Africans, the two groups that have sufficient cases. The dependent variable is the total years of schooling completed, ranging from 0 to 17, and treated as a continuous variable.

Explanatory variables include basic background variables, such as gender, race and residential status. Because SSOA includes respondents of several cohorts who experienced different educational environments even during the *apartheid* period, I examine secular differences by including cohort as a categorical variable with 10-year intervals from age 20 to 79[10], rather than including age as a continuous variable. Gender and race are coded as discrete variables. Rural/urban residence is defined based on the place of residence at age 14. Town/city, squatter and peri-urban areas are coded as urban; villages, farms and mines are coded as rural; and other unknown or unreported information is coded as other.

I also control for retrospective family background. Parental education is measured by the total years of schooling completed by either father or mother, whichever is higher. Similarly, I

[10] Respondents 80 or older, the oldest cohort, are excluded because differential mortality by race and SES may bias the results (for example, Whites and high SES people tend to have fewer siblings and also survive to old age). Even the two oldest cohorts (50-69 and 60-79) may be influenced by differential mortality, leading to an undercount of sibship size in these cohorts. I include them to allow for sufficient number of cases, and I truncate the sibship size at 6 to adjust for this undercount, which is discussed later in this section. Handful of respondents younger than 20 are deleted because they may not have completed school, and according to the survey procedure people younger than 20 should not be included in the sample.

include parental ISEI when respondent was 14, measured by ISEI score of either father or mother, whichever is higher[11]. ISEI is a scale of occupational status, ranging in principle from 0 to 100 (Ganzeboom, De Graaf, and Treiman 1992). Since no retrospective information on family arrangements is available, the best I can do is to include a dichotomous variable indicating whether at least one parent was present in the household when the respondent was 14.

Household economic condition, a direct indicator of material resources, may be an important determinant of educational attainment, especially in developing countries (Filmer and Pritchett 1999). Because no such measure is directly provided, I construct a scale to represent household economic conditions when the respondent was 14. Ten items of household conditions are included in the scale[12]. These items constitute a single factor, with all of the items having high loadings (Cronbach's alpha is 0.85). To construct a final 0-1 economic scale, I standardize and average the items, and then transform the scale to a range from 0 to 1, with 1 indicating the highest level of economic status.

The key independent variable is sibship size. It was obtained from a question directly asked in the survey: "how many siblings did you have when you were young?". I further truncate it at sibship size 6 to reduce the leverage of the very small number of children from very large families[13]. It is treated both as a continuous and a discrete variable in the analysis.

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[11] The data originally collected occupational information based on ISCO88 (Ganzeboom and Treiman 1996), which was then transformed to ISEI in the data.

[12] The ten items are: water supply, toilet facilities, kitchen facilities, availability of dictionary, atlas, camera, telephone, gas/electric stove, refrigerator and servants when respondents were 14.

[13] Sensitivity analysis shows that results are very similar no matter where the sibship size is truncated, which is also the case for the census data and the rest of the analysis.

More than 10% of the data on parental education, occupation and economic condition are missing, and the deletion of these cases may lead to a loss of useful information. Hence, the mean level of these variables is substituted for the missing responses, and a separate dummy variable is created for each of the three variables indicating whether they are missing. Respondents older than 79 and younger than 20 (96 cases in total) as well as those missing in all the other variables (about 2%, 180 cases) are dropped, resulting in a sample of 8,438. Also, since the survey provides a stratified probability sample (containing multiple cases drawn from a single magisterial district within a subpopulation group, and these cases are potentially correlated), Stata survey estimation is utilized to adjust for the multi-stage design, by treating subpopulation groups as strata and magisterial district as PSU (StataCorp 2003). The data are also appropriately weighted to represent the general White and Black population in South Africa.

The model specification for racial differences in the sibship size effect on educational attainment is as follows:

$$EDU_{ij} = \beta_{0j} + \beta_{1j} SIB_{ij} + \sum_{p=1}^{P} \beta_{pj} X_{pij} + \varepsilon_{ij} \qquad (1)$$

where EDU_{ij} refers to the highest educational level achieved for the i th respondent in the j th racial group (j identifies the racial group membership of either Whites, $j = 1$, or Blacks, $j = 2$), SIB_{ij} refers to the number of siblings, X_{pij} refers to covariate p in the model such as age, gender, residential status and parental socioeconomic status with associated coefficient β_{pj}, and ε_{ij} to a random component. According to hypothesis 1 and 2, β_{11} would be expected to be significantly negative, whereas β_{12} would be insignificant.

Moreover, to take account of the confounding birth order effect, I first construct a continuous birth order variable, using the question: “How many siblings are older than you?”

together with the sibship size variable. I then decompose the sibship size into two measures, the number of older siblings and number of younger siblings. A similar set of models was estimated by including information on both sibship size and birth order, in order to explicitly model the effect of sibship size controlling for birth order.

Variables and Methods Using the Census Data

Analysis using the census requires some data preparation. First, because enrollment starts to significantly diverge at the secondary school level in South Africa (Maharaj et al. 2000), I restrict the sample to children aged 14 to 18 to study differentials in secondary school enrollment (children usually start secondary school at age 13 and complete it at age 18). Given that Black children may delay or repeat school and thus start secondary school a little later than age 13 (Anderson 2000), I established a lower limit at age 14 to avoid confounding enrollment of primary with secondary school. I limited my analysis to secondary rather than college education because children may not live at home during college years. Hence, the living arrangement information obtained from the household roster may not reflect the actual family arrangement that has been supporting the children. Additionally, the upper bound (age 18) was chosen to limit the extent to which older siblings have left the household and thus are not recorded[14].

Second, information on family arrangements and sibship size needs to be derived from the household roster. In the census and most other data sets, each household member reported only the relationship to the household head. This approach omits substantial information, and thus presents a limitation in examining the extent of relationships among household members.

[14] The 1996 census shows that on average more than 80% of people under age 18 live with at least one parent, while the fractions are much smaller for older children and fall to below 60% at age 25.

The present study deals with this problem by developing linking algorithms to obtain more accurate information on household composition for the focal children[15]. For example, if the child is recorded as the head, not very common in the data set (2.8% of the total subsample), I then look for and count individuals reported to be parents, grandparent, other relatives and siblings in the household; if the child is the child of the head (68.6%), I then look for and count individuals reported to be the head and spouse of the head as child's parent, parents of the head as child's grandparent, other children of the head as child's siblings, siblings, grandparent and other relatives of the head as child's extended family members; if the child is recorded as a sibling of the head (5%), I then look for and count individuals reported to be parents of the head as child's parent, grandparents of the head as child's grandparent, other relatives to the head as child's extended family members, and the head and other siblings to the head as child's siblings. If the child is the head's grandchild (15.1%), it is not clear whether the child of the head is the parent of the focal child or uncle/aunt of the child. Under such circumstances, I employ proxy information on whether parents of the child are still alive: when the focal child's parent is alive, children of the head considered to be focal child's parents[16]. The logic of deriving family composition is similar: head's other children, head's spouse, siblings and other relatives are considered to be the focal child's extended family members, and head's other grandchildren are

[15] Another limitation of the household roster needs to be acknowledged: relationship categories are very broad. For instance, the overly broad "children" and "parents" categories do not allow for distinctions between biological and step families.

[16] This may somewhat confound children living in fostering families with those in extended families, since a portion of fostering children whose parents are alive but absent is classified as living with parents.

considered to be the focal child's siblings[17]. Nuclear households are then defined as households with only children and their parents. Such households are further divided into single-parent and two-parent nuclear families. Extended households are defined as households with children, at least one parent and other extended relatives, and fosterage households are defined as those with children and children's relatives but neither parent. I do not consider situations where children are spouses, parents and grandparents of the head, because children age 14 to 18 are very unlikely to be in these categories and there are very few such cases (a total less than 4%, which probably are mainly errors). Also, I do not implement the linking algorithms in situations where children are other relatives or are not related to the household head (a total less than 5%) because such children's relationship to other adults is ambiguous. Instead, I treat all these ambiguous cases as a separate "other" unspecified category[18].

Using the census sample, children age 14 to 18 are the unit of analysis. Two-level random-intercept logit models and population-averaged (GEE) logit models are estimated to take account of the fact that multiple eligible children may be from the same household. I first estimate multilevel models for the total South Africa population. Given that I conjecture that the sibship size effect varies by race and family arrangements, I then include these variables in interaction terms and test the significance of these interactions. Since the hypothesis of

[17] In this situation, the number of siblings counted is the number of all grandchildren in the household, including not only child's real siblings but also cousins. No further distinctions can be made using the data.

[18] It should be acknowledged that some unspecified households, such as where children are labeled as other relatives, may be extended or fosterage households. This implies that the number of extended and fosterage households may be larger than my estimates.

homogeneity by race and family organization was rejected, separate models are estimated first separately for Whites and Blacks, and then respectively for Black children living in different types of households, to highlight the impact of race and family arrangements[19].

The dependent variable is a dichotomous variable indicating whether the focal child is currently enrolled in school, coded 1 if the child is currently enrolled or not currently enrolled but has completed a secondary education or more[20], and 0 otherwise. As in the survey analysis, basic demographic variables such as age, gender and rural/urban residence are controlled. Since parental education and occupation is not available for children not living with parents, the highest total years of schooling and occupational status (measured by ISEI score[21]) among adults

[19] Explanatory variables selected for inclusion in all models using the two datasets are based on previous studies of schooling. They are more or less similar in each data analysis in order to be comparable. A few more variables are added in the census analysis to improve the explanatory power.

[20] Some children, especially Whites, may complete secondary school a little early, thus those not currently enrolled may have already finished secondary school but drop out at tertiary level (about 3% in the sample). To capture secondary school enrollment status, both children currently enrolled and those with a secondary and higher education are considered to be enrolled.

[21] ISEI scores were created from the detailed occupational categories used by the census, which closely approximate the 1988 ISCO categories (Ganzeboom and Treiman 1996). However, unlike most of the other variables, which have relatively few missing cases, many cases have missing data on occupation, and hence missing ISEI codes, due to unemployment. This is particularly true of Blacks, about half of whom have missing ISEI scores. Because unemployed Blacks would be likely to work at low level jobs if they were employed, when ISEI codes are

older than 25 in the household are used as a proxy for socioeconomic background. Adult older than 25 is chosen because they have completed their schooling. The choice also avoids the complication of autocorrelation when the focal children's education is the highest in the household and thus included in both sides of the equation. As in the survey analysis, the highest adult socioeconomic status is used because I assume people of the highest status would tend to maximize the positive effect of their status in the household. Total annual income is also included as a control, given it is a strongly associated with total financial resources available to the focal child. This variable is measured by summing annual income earned by all people older than 18 in the household[22]. To treat it as a single continuous variable, I assign the amount of the midpoint of each of the annual income categories. The natural log of the per capita income variable is finally included as a covariate in the model. Additionally, given that migrants may contribute to household resources through remittances and may draw on household support during economic crises, whether the household has sent out migrants is included as a dummy

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missing for Blacks, I substitute the mean ISEI score for unskilled laborers (ISCO88 Group 9), and also include a dummy variable indicating that the ISEI score was assigned. This strategy is not adopted for Whites since few cases are missing.

[22] Household income may also depend on earnings of children, which in turn may be affected by whether they are in school. However, in the census data, only 1.4% of all children younger than 18 reported income, which suggests that endogeneity between children's schooling and income is not a serious concern. Thus, income earned by children younger than 18 in the household is not included in constructing the total income variable.

variable[23]. The two crucial variables, namely family arrangements and sibship size, are created using the household roster, as described earlier. Family organization is a five-category discrete variable (two types of nuclear and extended arrangements plus the other category). Sibship size is measured by the number of siblings in the household[24], ranging from 0 to 16, which is also truncated at 6. It is treated both as a continuous and a discrete variable.

[23] This is especially important for single-parent nuclear families because by including migration status, it is possible to make distinctions between single-parent families resulting from death, divorce of parent, and labor migration of a parent. The familial resources available in these two types of families tend to differ significantly.

[24] Because the relationship code does not distinguish biological children from step and adopted children, the number of siblings captures both biological and non-biological siblings. Also, I use the number of siblings rather than the number of all children in the household, because it is less direct to locate children other than siblings who may share the resources. Using all children essentially assumes equal allocation of resources among children, which may not be the case especially in fostering families. In addition, the number of siblings is an individual-level variable while number of all children is at household-level. As discussed below, if the number of all children is used, it will be dropped in fixed-effect estimation, which makes the sensitivity tests impossible. Moreover, an analysis comparing the number of siblings and of all school-aged children (age 6 to 22) shows that the two are very similar, a majority (about 90%) within a difference less than 2. Finally, I need to acknowledge a potential problem: only siblings currently living in the household are counted, because information on non-coresident siblings is not available.

Individuals younger than age 14 and older than age 18 are discarded, and so are households that are missing on the above variables. This results in an overall population of 366,014[25]. The data are weighted to adjust for undercounting, especially for Blacks (weights are included in the census file). The model specifications for the random-intercept and GEE models are similar and the substantive interpretations do not change. I then present the results from two-level logit models as implemented in Stata[26]. Specifically, the model specification for race and family arrangements differentials in the sibship size effect on enrollment is as follows[27]. At the child-level:

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[25] The 1996 census individual-data file originally contains 3,621,201 cases. I restrict the analysis to family households, excluding institutions and hostels (112,117 cases deleted). Also, 394 cases are dropped where there are discrepancies between the individual data file and the corresponding household data file. I truncate household size to 20 which comprise 99.8% of the sample (8,045 cases are dropped). Restricting the analysis to children age 14 to 18, 3,126,683 cases are deleted. Finally, 7,948 cases missing on variables included in the analysis are dropped (excluding highest household ISEI; see discussion above). This results in a total population of 366,014.

[26] Random-effect models are presented because they provide context-specific estimates and are advantageous in modeling individual-level covariates (sibship size in the present case). They are also useful when the cluster size (eligible children from each household) is small, which is the case in the current study (cluster sizes of 1 and 2 account for more than 80% of the cases). Results in this paper show the coefficients from both models are very close, with coefficients from GEE models usually shrinking toward zero. GEE tables will be provided upon request.

[27] The separate analyses by family arrangements are restricted to Blacks because Whites live disproportionately in nuclear arrangements. Also, I do not present and interpret results from

$$Logit(Enroll_{ijkl}) = \beta_{0jkl} + \beta_{1jkl} SIB_{ijkl} + \sum_{p=1}^{P} \beta_{pjkl} X_{pijkl} \quad (2)$$

where the left-hand side is the logit of enrollment in secondary school for the i th child in the j th type of household, the k th racial group and l th household (the index j ranges from 1 through 5 representing single-parent nuclear, two-parent nuclear, extended, fostering, and other types of household, respectively; k is either 1 or 2, representing Whites and Blacks, respectively; l ranges from 1 to the total number of household in the sample). SIB_{ijkl} denotes the number of siblings, and X_{pijkl} refers to child-level covariates p such as age and gender. These are assumed to have fixed slopes across households. At the household-level, β_{0jkl} is a random intercept determined by p household-level covariates from Z_{1jkl} to Z_{pjkl} (in the analysis, the covariates are place of residence, three measures of household socioeconomic status, and whether the household sent out migrants) for household l with associated coefficient γ_{pjk}:

$$\beta_{0jkl} = \gamma_{0jk} + \sum_{p=1}^{P} \gamma_{pjk} Z_{pjkl} + \delta_{0jkl} \quad (3)$$

According to hypothesis 3 through 6, β_{112j} and β_{122j} is expected to be significantly negative, as observed elsewhere, whereas β_{132j} is expected to be insignificant due to the buffering effect of extended families. I also expect β_{142j} to be significantly negative due to the uneven allocation in fostering families. Finally, β_{112j} is expected to be larger in absolute value than β_{122j} due to the intensified resource-dilution process in single-parent families.

"other" types of arrangements because several possibilities are included in this category. Generally, results show this group is the least advantaged.

Furthermore, to take into account the confounding birth order effect, I construct two continuous variables: number of older siblings and number of younger siblings from the household roster. A similar set of multilevel models is then estimated by replacing the single sibship size variable with the two decomposed measures. These two measures are also truncated at 6. In order to explicitly examine the sibship size effect while controlling for birth order, I estimate the effect of sibship size stratified by birth order.

ENDOGENEITY: A METHODOLOGICAL ISSUE

The endogeneity problem arises from two sources. First, the causal relationship between sibship size and schooling may result from the fact that parents make fertility decisions based on their expectations regarding the cost of children, often referred to as the quality-quantity tradeoff (Steelman et al. 2002). In other words, parents may choose to have fewer children in order to provide each of them with better education. In this case, the true effect of sibship size would be greatly attenuated. A few studies address such issue using suitable data and sophisticated statistical methods: some conclude that the sibsize effect is artifactual (Black et al. 2005; Guo and van Way 1999), whereas others find the effect to be real (Conley and Glauber 2004; Downey et al. 1999; Kuo and Hauser 1997; Philips 1999). As shown in the South Africa background section, the quality-quantity tradeoff seems to be relatively uncommon in Black families, where fertility level remains quite high and fertility decisions tend to be based on factors exogenous to children's well-being. Despite of this, the present study seeks to address this aspect of endogeneity.

Another aspect of the endogeneity problem, which is not very well documented, arises from the selection of family arrangements based on parents' expectations regarding children's welfare. In other words, parents may choose to live with extended family members only when

they perceive that it can help children gain better education, either because they can draw on resources from extended kin, or because extended kin live close to schools or close to better schools. In this case, the buffering effect of extended family arrangements tends to be biased.

As discussed above, the effect of sibship size is likely to be contaminated by unmeasured family characteristics, particularly in terms of endogenous parental quality-quantity calculations and endogenous parent's expectations that lead to the selection of specific family arrangements. Hence, I carried out some sensitivity tests to take account of the endogeneity problem. Specifically, I use the census data to estimate household-specific fixed-effect (FE) models, which absorb the influence of endogenous family characteristics that are fixed within households. By comparing them with corresponding random-effect (RE) models, which assume no bias due to endogeneity, I am able to examine potential bias in RE estimates due to unmeasured selection bias: if results for corresponding models are consistent, the unobserved endogenous factors are quite unlikely to account for the sibship size effect, and vice versa. Yet, a disadvantage of FE modeling is that all covariates identical for each child within a household are dropped. For this reason, I focus my interpretation on RE estimates, and estimate FE models only for comparison purposes. Importantly, by estimating conditional FE models, households with only one eligible child are deleted from the analysis. To compare results from two sets of models, they should be based on the exact same cases. I thus estimate a new set of RE models restricted to the same observations as the FE models and then compare corresponding coefficients. This approach is useful, because when doing this, it is also possible to compare the same method with different subsets of the data. Given that half of the focal children come from households with only one eligible child, it is informative to learn how the RE models from the full-sample and restricted-sample differ.

RESULTS

Descriptive Statistics

Table 1 and Table 2 present sample means and percentages for the two data sets. The general patterns are as expected. The racial gap in socioeconomic background is large, with the most significant difference lying in total household annual income in the census sample. Total income for Whites is almost ten times as for Blacks, which reflects resource constraints for Blacks. The economic condition scale in the survey sample tells a similar story: while Whites achieve an average score of 0.8 out of 1, Blacks score only a little more than 0.2.

With respect to family arrangements, a clear racial difference emerges. More than 80% of Whites live in nuclear families, what is true of only about 50% of Blacks. Also, more than 30% of the Blacks reside in extended family arrangements, with 24% and 7.5% in extended and fostering families, respectively. Since several unspecified households in the "other" category tend to be extended in structure, the proportion of Black children in extended family arrangements may in fact be even higher. This is consistent with Burman and Fuchs's (1986) finding that just over half of African households adopt the nuclear structure, while most of the rest are extended in nature. In terms of sibship size, the racial difference in fertility is confirmed by both data sets. The survey sample shows that Black respondents had about four siblings, in comparison to about three for Whites. The census sample demonstrates an even larger fertility gap, with Blacks and Whites having an average of 3 and 1.3 siblings respectively. The larger gap in the census sample is probably due to the different pace of fertility decline in more recent cohorts. The fertility pattern, in general, is in line with Caldwell and Caldwell's (1993) finding that Black fertility was around 4 to 4.5, whereas White fertility was about 1.9 in the 1990s.

Differential educational opportunities and attainment are evident in both data sets. In the survey sample, Whites' total years of schooling averages almost twice that for Blacks. For school-aged children in the census sample, Whites have a secondary school enrollment of about 96%, whereas Blacks average about 8% lower. Although the enrollment gap appears to be not as large as the attainment gap, given the large share of Black school-aged population in South Africa, the number of Black children who are not currently enrolled in secondary school is tremendous.

Sibship size and Educational Attainment Using the Survey Data

Table 3 presents OLS models predicting respondents' educational attainment using the survey sample. Examination of correlation matrices confirms that multicolinearity is not substantial. Column 1 shows the results for the whole sample, which are generally in the expected direction. Males average four-tenth of a grade more than females. People in younger cohorts attain more schooling, as do people raised in urban areas and in socioeconomically advantaged families. Parental presence at age 14 seems to have no effect, presumably due to the little variation of this variable. Racial differences in educational attainment are sharp: Blacks are predicted to have one and half years less schooling than Whites, net of other factors, with Coloreds and Asians receiving intermediate amounts of education. The crucial sibship size effect, however, is neither significant nor quantitatively large. Since this gross relationship may hide substantial variations by race, I then estimate a model including interactions with race and sibship size. An adjusted-Wald test suggests that the interaction terms are jointly significant at 0.05 level ($F(3,349) = 2.76$). Hence, I conclude that the effect of sibship size varies by race, and all results are reported separately for Whites and Blacks in the next two columns.

When separate models are estimated, the effects of family socioeconomic background more or less remain, whereas effects of other explanatory variables diverge from the overall patterns. The gender gap in educational attainment does not exist for Blacks, which is consistent with previous findings (Case and Deaton 1999; Lam 1999; Thomas 1996). Also, cohort does not have much impact for Whites, but for Blacks younger cohorts receive significantly more schooling. This may be due to the fact that Whites have been constantly better-educated, whereas education for Blacks has undergone improvements over time. Crucially, sibship size has a negative effect for Whites but not for Blacks, as expected from Hypothesis 1 and 2[28].

Sibship size, Family Arrangements and Enrollment Using the Census

To examine whether the racial difference in the sibship size effect can be attributed to different family arrangements, the census sample is then used. Likelihood-ratio test suggests that the hypothesis of homogeneity across racial groups is rejected at 0.01 level ($LR\ \chi^2(3) = 12.89$). Hence, I first estimate a similar set of models by race as in Table 3. Results are reported in Table 4. The effects of demographic and socioeconomic background for the total South Africa population are generally in the expected directions, with a few exceptions.

The African disadvantages in secondary school enrollment seem to disappear, after controlling for covariates such as demographic and household characteristics. Blacks are more

[28] I also estimate similar models treating sibship size as aggregated discrete variables. Results are consistent with those from the continuous specification. For Whites, the coefficients of the sibship size dummy variables are mostly significant and become increasingly negative. In contrast, for Blacks, the coefficients of dummy variables are mostly insignificant and do not show a clear pattern in magnitude. These results are not shown here. Tables will be provided upon request.

likely to be enrolled, even compared to Whites. This result is consistent with Sibanda's (2004) finding of significantly lower odds of dropping out in secondary school among Africans using the same data. Sibanda considers this to be an unexpected finding given the racially segregated educational policy, and leaves it unexplained. To provide an interpretation, I estimate a series of models, as shown in Appendix A. Model 1 includes only race as the explanatory variable. It clearly shows a racial gradient of enrollment with Whites on top, and Blacks and Coloreds at the bottom. After controlling for individual demographic characteristics in Model 2, the White advantages more or less remain. However, when household socioeconomic status is taken into account in Model 3, the White advantages over Blacks disappear. This finding clearly suggests the White-Black gap in enrollment is predominately due to racial differences in household socioeconomic status: when Blacks and Whites experience similar socioeconomic status, Blacks are as likely to be enrolled as are Whites. This reverse pattern may partially reflect government's promotion of universal basic education, in particular focusing on increasing the number of Black schools, since the end of the *apartheid* period. This enables an increasing number of economically sufficient families to provide schooling for children. In contrast, during most of the *apartheid* period, there were not adequate schools for Blacks, which denied enrollment of a large number of Black children, even though their families could afford the expenses. However, as in Table 3, we see a persistent racial gap in ultimate educational attainment. These two findings are not contradictory, given that the two are different measures of education. Educational attainment relies not only on financial resources, but on non-material resources such as family cultural capital, parental time and attention. These factors are difficult to control for, but they may affect children's cognitive development, school performance and subsequently their school progression. Enrollment status, by contrast, disproportionately depends on financial resources, especially at

the primary and secondary school level. Children are often able to stay in school as long as their families can afford it. Low school quality for Blacks may be another explanation: poor quality leads to poor performance and thus lower attainment. Yet another explanation is the different time periods covered in two samples: while the census was collected just after *apartheid* and covers school-aged children benefiting from the education expansion, the survey sample covers respondents with completed education in a much wider age range.

When separate models are estimated, shown in the last two columns in Table 4, the effects of explanatory variable more or less remain, with a few exceptions. For Whites, males and females are equally likely to enroll in secondary school, whereas for Blacks, males are more likely to be enrolled. This seems to be contradictory to results from the survey sample that no gender disparities exist in ultimate attainment. Several studies, however, support my finding that in South Africa girls are less likely to be enrolled; but as long as they enroll in school, they are not disadvantaged anymore because they perform better than boys (Fuller and Liang 1996). Also, the rural-urban enrollment pattern is different from that of educational attainment: a majority of Blacks reside in rural areas where Black schools are more available, they thus may be as likely to enroll as urban children. Yet, probably due to the low quality of rural Black schools, being in rural areas is not as beneficial to educational attainment as being in urban areas. This pattern is reversed for Whites, which is probably due to a similar reason—their predominant urban residence. Having migrants in the household appears to be beneficial for Black children's schooling but not for White children (Lu and Treiman 2005b). This might be because labor migration in South Africa is particularly prevalent among Africans and is done in order to enhance the standard of living of families left behind via remittances. By contrast, migration is quite uncommon among White families, and may be due to reasons other than economic

considerations. In addition, the direct effect of family arrangements on Black children's enrollment is less clear: children in two-parent nuclear families tend to be better off than those from single-parent nuclear families, whereas the two types of extended family arrangements are not distinguishable from nuclear families with respect to socioeconomic status. This might result because the main effect of extended family arrangements effect, the availability of more resources, is absorbed by the household socioeconomic characteristics in the model, specifically the income variable. Yet, it is important to include this variable because of its strong predictive power on educational outcomes. Therefore, I estimate models predicting enrollment from family arrangements in order to purge the effect of family socioeconomic status. Results clearly show that being in two-parent nuclear families and extended families predicts greater chances of enrollment, followed by single-parent and fostering families. The impact of being in extended families is not distinguishable from that of two-parent families (a Wald test shows a difference of -0.041 with a standard error of 0.040), whereas the effect of being in fostering families is detrimental. These results tend to support the beneficial role of being in extended families versus other families, in particular fostering families. The role of family arrangements will be more clearly demonstrated in term of their mediating effect in sibship size later in the analysis.

Sibship size has a negative impact for Whites, but no such effect exists for Blacks (Hypothesis 1 and 2). The results are consistent with those from the survey sample. The coefficient for Whites is only marginally significant, perhaps because of the small variability in their enrollment status. The overall enrollment rates for Whites are as high as 95% due to their higher economic status and relatively lower schooling expenses. This result is in line with findings from other studies which suggest that in families with better economic conditions the sibship size effect may not exist because these families are able to afford all children no matter

how many they have (Blake 1989). This is also consistent with the insignificant coefficient of total household income for Whites, similar to findings of Case and Deaton (1999): economic resources do not have much impact on White children's enrollment because their relative expenses are so low that funding is not a constraint for Whites; but it can be a serious constraint for the much poorer Blacks. In contrast, sibship size is negatively associated with Whites' educational attainment. This may be because educational attainment depends also on familial intellectual resources, which are more thinly diluted as sibship size increases regardless of financial resources in the household.

To examine the mediating role of family arrangements on the sibship size effect, I test the hypothesis of homogeneity in sibship size effect across family arrangements for Blacks. This hypothesis is rejected by a Likelihood-ratio test (LR $\chi^2(4) = 375.1$, $p<0.001$), I thus estimate separate models for Blacks by family organization, reported in Table 5[29]. The effects of gender, age, presence of migrants and family socioeconomic status are mostly as expected. The effect of sibship size appears to be sensitive to family arrangements: the detrimental effect of sibship size holds for children in both types of nuclear families and fostering families, while it disappears for children in extended families. These findings are consistent with Hypotheses 3 through 5. Extended families are able to buffer the negative effect of sibship size by making more resources available; nuclear families, in contrast, suffer from resource constraints associated with large sibship size because educational costs are restricted to parents only. In addition, although extended and fostering families both belong to the extended arrangements, the beneficial role is

[29] I also estimate similar models treating sibship size as aggregated discrete variables. Results are consistent with those from the continuous specification. These tables will be provided upon request.

restricted to extended families, whereas the coefficient of sibship size for foster children is significantly negative and large[30]. Fostering families have less incentive to allocate and tend to allocate fewer resources to children other than their biological ones. To test Hypothesis 6, I combine the parameter estimates and associated standard errors from two-parent and single-parent nuclear families, and find support for the more detrimental sibship size effect in single-parent families due to the aggravated resource-dilution process (a Wald test shows a difference of 0.015 with a standard error of 0.006, $p<0.05$)[31, 32].

[30] I further estimate models separately for foster families with own biological children versus those without; the sibship size effect is significantly negative in both families. The effect is more detrimental in fostering families with own children, but the difference is only marginally significant. The result suggest that fostering families have weak incentives to invest in foster children, which does not depend much on whether they have own children.

[31] This result suggests that selection into fostering families tends to be based largely on factors exogenous to education outcomes, such as death of parents. If fostering arrangement is endogenously selected based on parental expectation, the negative sibship size effect in such families should be greatly reduced.

[32] To examine whether these results are obtained only for the specific cohort of children studied (who may be distinctive in that they were teenagers at the breakdown of *apartheid*), I replicate the analysis to other cohorts of children: one during the *apartheid* period, using the 1985 census, and one during the post-*apartheid* period, using the 2001 census. Results are relatively consistent across the three cohorts (coefficients remain qualitatively the same with a few quantitative changes), except that the sibship size effect for White children in the 2001 census sample is only marginally significant, probably due to their high level of enrollment that approaches being

Interactive Effect of Sibship Size and Birth Order

To evaluate the confounding effect of birth order, I first estimate OLS regression models based on Table 3 using the survey sample, with only the sibship size variable changed. Results are shown in Table 6. The correlation between birth order and sibship size is about 0.6, not high enough to create multicolinearity problem. Therefore, I am able to estimate Models 1, 3 and 5 by including both sibship size and birth order in a single model. Models 2, 4 and 6 further explore the interactive effect of these two variables by decomposing sibship size into its birth order components, namely the number of younger siblings and older siblings. To maintain consistency, all these variables are truncated at 6. Compared to Table 3, the effects of all the other covariates on educational attainment more or less hold.

The effect of sibship size across Models 1, 3 and 5 is very similar as in the previous models that do not control for birth order: the sibship size effect is negative for Whites, but the effect disappears for Blacks and the overall population. Also, coefficients for the continuous birth order variable are not significant, suggesting the absence of a clear birth order effect in South Africa. When further decomposing the sibship size effect in Models 2, 4 and 6, both the number of older and the number of younger siblings have negligible effects for the overall population and for Blacks, but they have a detrimental effect in White families. Considered together, the negative effects of number of younger and older siblings for Whites essentially reflect an overall negative sibship size effect. In sum, results from the survey sample suggest a relatively robust sibship size effect, independent of the birth order effect.

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universal. Overall, the consistent patterns suggest that the process works similarly independent of the cohorts being studied.

A similar analytic strategy is adopted using the 1996 census. The correlation between birth order and sibship size approaches 0.9. In this case, including both variables will likely contaminate the results due to multicolinearity. Hence, I estimate similar models as in Table 4 and 5 by substituting sibship size with the number of younger and older siblings (also truncated at 6). Results are reported in Table 7. The effects of all the other covariates on enrollment status more or less hold. In general, the number of younger siblings exerts a slightly positive effect, whereas the number of older siblings appears to have a negative effect. These results suggest a negative birth order effect; that is, parents are more likely to invest in early-borns, presumably in hopes of obtaining support early from these older children[33]. However, for Blacks in extended families, the negative effect of having older siblings is essentially reduced to zero, which tends to support the buffering effect of extended family arrangements.

Since the number of younger and older siblings exert effects in the opposite direction, no clear inference regarding the effect of sibship size can be drawn. To examine the sibship size effect while controlling for birth order, I estimate separate models at each birth order, wherever there is a sibship size effect observed in Table 4 and 5. Table 8 presents the results. In general, the coefficients of sibship size in each subsample remain significantly negative, with an exception of Whites in the fourth, fifth and sixth birth order, and Blacks in fostering families at the sixth birth order. This lack of association for Whites may be caused by the extremely small

[33] Birth order is irrelevant to final educational attainment but negatively associated with enrollment. This may be because parents ensure older children's schooling but only to a certain level they consider necessary (probably secondary school). To obtain an earlier return on their investment and finance younger children's education, parents do not encourage early-borns to proceed further, so that ultimate attainment does not differ among children in the family.

number of cases in these cells (only a few hundred), which leads to unstable results. Also, children in these three cells come from large families, which are quite uncommon for Whites. These large families may be different from other White families in a way that enhances educational opportunities. The lack of effect for children at birth order 6 in fostering families may have a similar explanation: families fostering more than 6 children other than their own may be quite distinctive in a way that promotes schooling. Overall, results from the census sample also show that the sibship size effect remains strong after controlling for birth order; thus, it is not a simple artifact of the birth order effect.

Tests for Endogeneity: Results from Corresponding Fixed-Effect and Random-Effect Models

I conduct sensitivity tests to examine the degree of endogeneity due to parental quality-quantity calculations and selection into different family arrangements. Based on models in Table 4 and 5, I estimate corresponding household FE models using the census to control for pre-existing stable family characteristics, given that more than one eligible child may be from the same household. Results are shown in Appendix B1 and B2. As we see in these models, covariates that are constant within households are dropped from these models. Also, the FE modeling excludes all children from households with only one eligible child, thereby reducing the number of cases in the analysis (about 50% of the whole sample). To be able to compare results across models, I thus estimate a new set of RE models restricted to the same cases as the FE models. The results are shown in Appendix C1 and C2.

First, comparisons between the restricted and full-sample RE models reveal highly consistent patterns in the effects of covariates, sibship size in particular: there is a significant sibship size effect for Whites, and for children in Black nuclear and fostering families, but no

such effect exists in Black extended families. These effects also remain similar in magnitude. The analysis, therefore, is shown to be robust across sub-samples.

Next, comparisons of coefficients from the FE models and corresponding RE models show more or less similar patterns with a few discrepancies[34]. In particular, Hausman specification tests (Hausman 1978), which assume that unmeasured family characteristics are orthogonal to observed regressors, suggest that two sets of models are consistent for Blacks in general and for Blacks in all four types of family arrangements (the respective Hausman χ^2 s are: 9.74 with a d.f of 7; 1.39 with a d.f. of 2; 1.64 with a d.f. of 2; 0.76 with a d.f. of 3; and 0.93 with a d.f. of 3.). However, the orthogonal assumption is rejected by Hausman tests for Whites in general (Hausman χ^2 is 79.74 with a d.f of 6.), which suggests the endogeneity tends to exist so that the FE model taking account of endogenous factors fits better than the RE model. I finally turn to the individual coefficients to determine where selectivity renders specific effects spurious. Importantly, the sibship size effect persists in the FE models for Blacks in general, for children in

[34] In both the RE and FE models, the relative advantages of Blacks in school enrollment after controlling for socioeconomic status disappear, compared to results in Table 4. This may be explained by the sample differences: the restricted sample only contains children from households with multiple eligible children. This means that the previously advantages may be essentially driven by households with only one eligible child (about 50% of the sample). These households likely benefit children because they are relatively small in size with only one child aged 14-18 competing for resources pertinent to secondary education. Also, given the high fertility of Blacks, these relatively small households may be positively selected in some way that enhances schooling, even compared to their White counterparts.

extended and fostering families[35]. The consistency of sibship size coefficients is further supported by Wald test (for Blacks, the difference in coefficients is 0.005 with a standard error of 0.019. The differences (standard errors) for two types of extended family arrangements are 0.003 (0.074) and 0.005 (0.096), respectively). These results reinforce earlier findings in this paper on the role of extended family arrangements in mediating the sibship size effect. The effect of sibship size for Whites, however, fails to survive after controlling for unobserved heterogeneity: it is significantly negative in the RE model, but not in the FE model. This discrepancy implies the presence of endogenous quality-quantity tradeoffs in White families. After using the FE models to absorb the exaggerating effect of the endogeneity, the sibship size effect is essentially reduced to zero. In contrast, the sibship size effect observed for Blacks turns out to be quite robust to parental calculation and selection of family arrangements, at least in two types of extended family arrangements. This may be because South African Whites have adopted a Western cultural system of childbearing with an emphasis on the quality of children, whereas Blacks' traditional cultural system desires children for exogenous reasons such as child labor and old age security.

SUMMARY AND DISCUSSION

[35] The coefficients of sibship size in two types of nuclear families are dropped because this variable is constant within household (it is identical across eligible children in the same household), as shown in Table B2. It is thus hard to examine the endogeneity problem in Black nuclear families. Yet, given that the results are relatively robust for the overall Black population and Blacks in two types of extended family arrangements, it is not very likely that the sibship size effect in the nuclear portion of Black population is seriously plagued by selection bias: if it is, we can barely see a robust effect for the general Black population.

This article has examined the effect of sibship size on children's schooling, focusing in particularly on the intermediate role of family arrangements in this sibsize-education association. South Africa was drawn on as the case study, given its racially stratified patterns in socioeconomic status and cultural customs reflected in family arrangements. Importantly, the present study explicitly tests this mediating effect by distinguishing different types of family arrangements, and meanwhile takes account of methodological issues to evaluate the robustness of the findings. In addition, two measures of schooling and two data sets are used to offer a more comprehensive picture of educational patterns in South Africa.

In general, the results support my hypotheses. Sibship size has different impacts on the schooling of Blacks and Whites in South Africa: it is negatively related to both ultimate educational attainment and school enrollment for Whites, but no such effect exists for Blacks as a whole. The differential sibship size effect is largely mediated by the distinctive family arrangements representing Whites and Blacks. Specifically, the lack of effect for Blacks is attributable to their extended family arrangements, in which costs of childrearing are shared by a wide range of relatives beyond parents by resource-pooling. As a result, the availability of family resources increases and the subsequent resource-dilution process is attenuated. In contrast, Whites are not protected from this effect because they rely almost exclusively on resources from the nuclear family. These conclusions are drawn because the absence of a negative sibship size effect is restricted to Blacks in extended families, whereas the negative effect of many siblings holds for Blacks in nuclear and fostering families as it does for Whites. Interestingly, although both belong to extended family arrangements, the effect is negative in fostering families because these families operate differently from extended families in resource allocation. It is likely that fostering families have fewer incentives to invest in children other than their own because they

do not expect much return from these children; meanwhile, the absence of parents means they are unable to control this unequal allocation in fostering families. Additionally, children reared in single-parent nuclear families are even more vulnerable to the sibship size effect than those in two-parent nuclear families, due to aggravated resource constraints associated with single parenthood. Overall, the analysis highlights the importance of focusing critically on the socioeconomic and cultural contexts in which family effects operate when studying the effect of sibship size. I view the above findings as evidence of the beneficial role that extended family arrangements play in coping with socioeconomic constraints in developing countries and especially in the contexts of high fertility.

Beyond these observations, sensitivity tests are carried out to examine how endogeneity and the confounding birth order effect may alter the above conclusions. Contrary to some claims, I find the sibship size effect is not a simple artifact of the birth order effect, at least in the two samples I have examined. The inclusion of the birth order component either as a continuous variable or as a decomposition hardly alters the sibsize-education association. When modeling the sibship size effect stratified by birth order, this negative effect remains strong in most cases. Birth order itself, in contrast, has no clear influence on educational attainment, although it is negatively related to enrollment status, reflecting the specific strategies employed by parents in South Africa. Moreover, the sibship size effect for Blacks is relatively robust to the endogenous quality-quantity tradeoff and selection into family arrangements, especially for Blacks in extended family arrangements. For Whites, however, the observed negative sibship size effect tends to be inflated by endogenous family characteristics, most likely the quality-quantity calculations made by parents. These disparities may be due to the distinctive cultural systems of childrearing adopted by Whites and Blacks: while the former value the quality of children over

the quantity, the latter desire children mostly for reasons exogenous to children's well-being. The findings of a spurious sibship size effect for Whites suggests the importance of addressing the endogeneity issues when studying sibship size effects, particularly for populations influenced by Western values of childbearing[36].

Besides the above findings, the analysis reveals some general educational patterns in South Africa, in terms of racial, gender and background differences in two measures of schooling. Furthermore, this study explores a way of using household rosters to construct typologies of family arrangements among household members other than the household head. Distribution statistics show that this method is relatively accurate in identifying family arrangements. Yet, it should be acknowledged that this linking method is somewhat limited by the ambiguities of less detailed relationship categories and the ability of the data to capture non-coresident family members.

The findings of this paper have several implications for the well-being of children in South Africa. With the observed strong effect of household resources on enrollment, revising the options for waiving or reducing school costs among poor Blacks may lead to an unprecedented increase in Black's education, and thus reduce the racial gap in schooling. Also, the quality of Black education deserves attention from policymakers. As demonstrated in the analysis, while the racial enrollment gap can be largely explained by resource differences, the attainment gap persists even after controlling for family socioeconomic status. This may result from factors such as school quality and familial nonmaterial resources. While household nonmaterial resources such as cultural capital and social capital can barely be efficiently influenced by state policies, improving the quality of Black schools seems to be a viable approach for narrowing the racial

[36] For example, Guo and VanWay (1999) find no sibship size effect for the U.S. population.

gap in educational attainment. Additionally, the diversity of family arrangements for Blacks implies that public policies may affect a different group of people than they have originally targeted. While most policies consider the nuclear family to be the fundamental unit, the South African evidence clearly illustrates the prevalence of extended family arrangements. Hence, to appropriately identify the vulnerable children of targeting subsidies, a policy that targets children instead of parents would be more efficient, particularly in places where fostering families are common. Furthermore, given the beneficial role of extended families in coping with family constraints, this type of arrangements should be protected for Blacks during the post-*apartheid* recovery period in South Africa. To the extent one believes in the ability of extended families to minimize the negative impact of high fertility and improve the well-being for Black children, the adoption of this arrangement would have considerable positive implications in reducing racial inequalities among this generation of South African children. For comparativists, as a final note, there is a corollary to these findings in developing settings where extended family arrangements are prevalent.

APPENDICES

Appendix A. Random-intercept Logit Models of Current School Enrollment on Race and Other Control Variables, 1996 Census (N=366,014). (Standard Errors in Parentheses)

Independent variables	Model 1	Model 2	Model 3
Race (ref. Whites)			
Asian	-0.522***	-0.559***	-0.321***
	(0.054)	(0.056)	(0.056)
Colored	-1.420***	-1.447***	-0.671***
	(0.037)	(0.037)	(0.039)
Blacks	-0.947***	-0.734***	0.054
	(0.032)	(0.034)	(0.036)
Male (ref. female)		0.052***	0.119***
		(0.012)	(0.013)
Age		-0.420***	-0.525***
		(0.005)	(0.005)
Urban residence (ref. rural)		0.459***	0.040*
		(0.014)	(0.016)
Highest adult education in HH			0.379***
			(0.003)
Highest adult ISEI in HH[a]			0.040***
			(0.007)
Missing on highest adult ISEI			0.351***
			(0.019)
Total HH annual income (ln)			0.001
			(0.002)
Intercept	3.506***	9.910***	6.976***
	(0.032)	(0.084)	(0.094)
Percent of variance explained between households	46.1***	49.3***	35.8***
	(0.003)	(0.003)	(0.003)
Log-likelihood	-117482.1	-112530.4	-100734.9

[a] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.

* p < .05, ** p < .01, *** p < .001

Appendix B1. Fixed-effect Logit Models of Current School Enrollment on Sibship Size and Control Variables, 1996 Census. (Standard Errors in Parentheses)

Independent variables	Overall SA population	Whites only	Blacks only
Race (ref. Whites)			
Asian	-0.577		
	(1.233)		
Colored	-0.102***		
	(0.002)		
Blacks	0.226		
	(0.855)		
No. of siblings	-0.025***	-0.187	-0.016
	(0.003)	(0.375)	(0.016)
Male (ref. female)	0.097***	0.006	0.080**
	(0.029)	(0.231)	(0.031)
Age	-0.473***	-1.113***	-0.504***
	(0.010)	(0.092)	(0.010)
Family arrangements (ref. single-parent nuclear)			
Two-parent nuclear	0.334*	0.194***	0.357*
	(0.145)	(0.015)	(0.153)
Extended	-0.082	-0.367*	-0.014
	(0.104)	(0.147)	(0.088)
Fostering	-0.400**	—[a]	-0.518***
	(0.154)		(0.157)
Other	-0.809***	-1.113***	-0.842***
	(0.122)	(0.046)	(0.125)
Log likelihood	-45416.5	-1939.0	-43308.0
N	179,870	10,654	152,081

[a] This category is dropped due to no within-group variance.
* p < .05, ** p < .01, *** p < .001

Appendix B2. Fixed-effect Logit Models of Current School Enrollment on Sibship Size and Control Variables by Types of Family Arrangements, 1996 Census. (Standard Errors in Parentheses)

Independent variables	BSN [a]	BTN [a]	BE [a]	BF [a]
No. of siblings	—[b]	—[b]	-0.025	-0.061**
			(0.076)	(0.024)
Male (ref. female)	0.101***	0.008	0.099***	0.242*
	(0.013)	(0.068)	(0.008)	(0.123)
Age	-0.498***	-0.554***	-0.513***	-0.435***
	(0.022)	(0.022)	(0.024)	(0.041)
Log likelihood	-10395.1	-10401.9	-10381.0	-3418.5
N	36,806	42,219	40,477	11,169

[a] BSN, BTN, BE and BF columns from the left to right are short for "Black single-parent nuclear", "Black two-parent nuclear", "Black extended" and "Black fostering" families, respectively.
[b] Sibship size is dropped due to no within-group variance.
* p < .05, ** p < .01, *** p < .001

Appendix C1. Random-intercept Logit Models of Current School Enrollment on Sibship Size and Control Variables, 1996 Census, Restricted to the Same Subsample as Fixed-effect Models. (Standard Errors in Parentheses)

Independent variables	Overall SA population	Whites only	Blacks only
Child-level:			
Race (ref. Whites)			
Asian	-0.612***		
	(0.096)		
Colored	-0.838***		
	(0.072)		
Blacks	0.055		
	(0.067)		
No. of siblings	-0.022***	-0.233***	-0.011
	(0.007)	(0.058)	(0.007)
Male (ref. female)	0.042*	-0.021	0.217***
	(0.020)	(0.105)	(0.022)
Age	-0.466***	-0.995***	-0.486***
	(0.007)	(0.048)	(0.008)
Family arrangements (ref. single-parent nuclear)			
Two-parent nuclear	0.243***	0.387*	0.261***
	(0.032)	(0.183)	(0.036)
Extended	-0.043	-0.210	-0.009
	(0.032)	(0.223)	(0.035)
Fostering	0.024	-0.416	0.015
	(0.049)	(0.671)	(0.052)
Other	-0.761***	-1.608***	-0.806***
	(0.038)	(0.237)	(0.041)
Household-level:			
Urban residence (ref. rural)	0.156***	-0.060	0.117***
	(0.027)	(0.197)	(0.029)
Highest adult education in HH	0.407***	0.276***	0.391***
	(0.005)	(0.034)	(0.005)
Highest adult ISEI in HH[a]	0.019	0.124**	0.023***
	(0.011)	(0.041)	(0.001)
Missing on highest adult ISEI	0.198***	0.617**	0.192***
	(0.031)	(0.204)	(0.035)
Total HH annual income (ln)	0.013**	0.009	0.015***
	(0.003)	(0.017)	(0.004)
HH sent out migrants (ref. no migrants)	0.283***	-0.337	0.300***
	(0.027)	(0.301)	(0.028)
Intercept	6.237***	16.434***	6.640***
	(0.151)	(0.925)	(0.143)
Percent of variance explained between households	46.8***	67.7***	52.3***
	(0.005)	(0.016)	(0.004)
Log-likelihood	-47204.2	-1944.6	-43620.6
N	179,870	10,654	152,081

[a] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.

* p < .05, ** p < .01, *** p < .001

Appendix C2. Random-intercept Logit Models of Current School Enrollment on Sibship Size and Control Variables by Types of Family Arrangements, 1996 Census, Restricted to the Same Subsample as Fixed-effect Models. (Standard Errors in Parentheses)

Independent variables	BSN [a]	BTN [a]	BE [a]	BF [a]
Child-level:				
No. of siblings	-0.079***	-0.061***	-0.023	-0.066***
	(0.016)	(0.014)	(0.022)	(0.013)
Male (ref. female)	0.192***	0.187***	0.091*	0.267***
	(0.043)	(0.044)	(0.041)	(0.077)
Age	-0.460***	-0.506***	-0.453***	-0.463***
	(0.015)	(0.016)	(0.015)	(0.028)
Household-level:				
Urban residence (ref. rural)	-0.009	0.120*	0.077	0.108
	(0.059)	(0.056)	(0.052)	(0.103)
Highest adult education in HH	0.389***	0.410***	0.345***	0.362***
	(0.010)	(0.010)	(0.010)	(0.017)
Highest adult ISEI in HH[b]	0.041	0.052*	0.012	0.043***
	(0.030)	(0.026)	(0.023)	(0.009)
Missing on highest adult ISEI	0.094	0.063	0.053	0.623***
	(0.070)	(0.066)	(0.059)	(0.131)
Total HH annual income (ln)	0.016*	0.020***	0.027***	0.027*
	(0.007)	(0.002)	(0.004)	(0.012)
HH sent out migrants (ref. no migrants)	0.429***	0.071	0.189***	0.255**
	(0.050)	(0.064)	(0.050)	(0.088)
Intercept	6.289***	7.145***	6.459***	5.928***
	(0.278)	(0.284)	(0.268)	(0.512)
Percent of variance explained between households	52.4***	57.1***	53.3***	49.3***
	(0.008)	(0.008)	(0.008)	(0.015)
Log-likelihood	-10467.8	-10779.2	-11065.3	-3215.9
N	36,806	42,219	40,477	11,169

[a] BSN, BTN, BE and BF columns from the left to right are short for "Black single-parent nuclear", "Black two-parent nuclear", "Black extended" and "Black fostering" families, respectively.
[b] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.
* $p < .05$, ** $p < .01$, *** $p < .001$

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TABLES

Table 1. Sample Means and Percentages Separately for Whites and Blacks, SSOA 1994.

	Whites	Blacks
Discrete variables	Percent (%)	Percent (%)
Gender		
Male	57.8	50.6
Cohort		
20-29	28.1	41.1
30-39	23.5	24.0
40-49	21.2	14.9
50-59	13.1	10.0
60-69	9.6	6.8
70-79	4.7	3.3
Residential status at age 14		
Urban	68.6	49.5
Rural	9.5	37.3
Other	21.9	13.3
Whether at least one parent present at age 14		
Yes	94.7	92.9
Missing on parental education		
Yes	11.1	7.8
Missing on parental ISEI when r was 14		
Yes	7.8	21.5
Missing on household economic condition when r was 14		
Yes	7.8	14.3

	Whites		Blacks	
Continuous variables	Mean	(S.D.)	Mean	(S.D.)
No. of siblings (excluding focal children)	3.2	(2.48)	3.9	(2.61)
Birth order	2.6	(1.82)	3.6	(2.10)
No. of younger siblings	2.3	(1.76)	3.0	(1.96)
No. of older siblings	1.1	(1.92)	1.1	(1.87)
R's total yrs of schooling	12.3	(2.26)	6.4	(4.25)
Parental education	10.8	(3.10)	4.3	(3.93)
Parental ISEI when r was 14	44.2	(16.23)	28.3	(11.38)
HH economic condition when r was 14	.80	(0.22)	.24	(0.20)
N	2,086		3,759	

Table 2. Sample Means and Percentages Separately for Whites and Blacks, 1996 Census.

	Whites	Blacks
Discrete variables	Percent (%)	Percent (%)
Gender		
Male	50.0	48.8
Current residential status		
Urban	92.1	36.2
Whether household sent out any migrants		
Yes	3.2	29.9
Family arrangements		
Single-parent nuclear	11.4	25.3
Two-parent nuclear	70.0	27.9
Extended	10.8	24.0
Fostering	1.0	7.5
Other	7.0	15.2
Missing on highest adult ISEI in HH		
Yes	14.8	55.9
Current school enrollment		
Enrolled	95.5	88.2

	Whites		Blacks	
Continuous variables	Mean	(S.D.)	Mean	(S.D.)
No. of siblings (excluding focal children)	1.3	(1.00)	2.9	(2.13)
Birth order	1.8	(0.90)	2.8	(1.73)
No. of younger siblings	0.8	(0.90)	1.8	(1.73)
No. of older siblings	0.5	(0.68)	1.0	(1.26)
Age	15.9	(1.42)	15.9	(1.41)
Highest adult education in HH	12.9	(1.81)	10.3	(2.50)
Highest adult ISEI in HH	48.2	(18.71)	26.1	(12.85)
Total annual income in HH	105,023	(127,688)	13,614	(31,444)
N	27,245		299,213	

Table 3. OLS Regression of Years of Schooling on Sibship Size and Control Variables, SSOA 1994. (Standard Errors in Parentheses)

Independent variables	Overall SA population	Whites only	Blacks only
Race (ref. whites)			
Asian	-0.574*		
	(0.264)		
Coloreds	-1.262***		
	(0.223)		
Blacks	-1.558***		
	(0.216)		
No. of siblings	0.035	-0.137***	0.002
	(0.027)	(0.036)	(0.038)
Male (ref. female)	0.406***	0.334**	0.232
	(0.118)	(0.117)	(0.175)
Cohort (ref. 20-29)			
30-39	-0.759***	0.246	-0.844***
	(0.153)	(0.184)	(0.210)
40-49	-1.376***	0.369*	-2.004***
	(0.191)	(0.150)	(0.264)
50-59	-1.518***	0.118	-2.070***
	(0.220)	(0.171)	(0.213)
60-69	-2.307***	0.547**	-3.507***
	(0.263)	(0.203)	(0.311)
70-79	-2.304***	-0.027	-3.480***
	(0.319)	(0.291)	(0.368)
Residential status (ref. urban)			
Rural	-0.625***	0.221	-0.500*
	(0.189)	(0.167)	(0.240)
Other	-0.237	-0.097	-0.311
	(0.154)	(0.163)	(0.264)
Parental education	0.341***	0.286***	0.331***
	(0.020)	(0.026)	(0.026)
Missing on parental education	-0.271	0.239	-1.290*
	(0.345)	(0.146)	(0.528)
Parental ISEI when r was 14 [a]	0.083	0.040	0.229**
	(0.050)	(0.052)	(0.072)
Missing on parental ISEI when r was 14	-0.438*	-0.054	-0.462*
	(0.214)	(0.228)	(0.217)
Economic condition when r was 14	3.327***	2.027***	5.254***
	(0.400)	(0.268)	(0.571)
Missing on economic condition at 14	-0.377	0.632***	-1.173***
	(0.208)	(0.193)	(0.258)
Parent present at age 14 (ref. nonpresent)	-0.209	-0.078	-0.077
	(0.224)	(0.221)	(0.304)
Intercept	6.696***	7.474***	4.772***
	(0.469)	(0.393)	(0.495)
R^2	0.564	0.289	0.438
N	7,336	2,086	3,759

[a] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.
* p < .05, ** p < .01, *** p < .001

Table 4. Random-intercept Logit Models of Current School Enrollment on Sibship Size and Control Variables, 1996 Census. (Standard Errors in Parentheses)

Independent variables	Overall SA population	Whites only	Blacks only
Child-level:			
Race (ref. Whites)			
Asian	-0.314***		
	(0.057)		
Colored	-0.591***		
	(0.040)		
Blacks	0.162***		
	(0.037)		
No. of siblings	-0.017***	-0.061*	-0.001
	(0.004)	(0.030)	(0.004)
Male (ref. female)	0.101***	0.016	0.264***
	(0.013)	(0.056)	(0.014)
Age	-0.506***	-0.884***	-0.500***
	(0.005)	(0.025)	(0.005)
Family arrangements (ref. single-parent nuclear)			
Two-parent nuclear	0.231***	0.298***	0.242***
	(0.020)	(0.090)	(0.021)
Extended	-0.060**	-0.232*	-0.021
	(0.021)	(0.114)	(0.022)
Fostering	0.012	-0.389	-0.002
	(0.031)	(0.250)	(0.030)
Other	-0.863***	-1.587***	-0.892***
	(0.022)	(0.108)	(0.022)
Household-level:			
Urban residence (ref. rural)	0.101***	0.184	0.010
	(0.016)	(0.100)	(0.017)
Highest adult education in HH	0.380***	0.218***	0.343***
	(0.003)	(0.017)	(0.003)
Highest adult ISEI in HH[a]	0.038***	0.009***	0.035***
	(0.007)	(0.002)	(0.008)
Missing on highest adult ISEI	0.331***	0.445***	0.312***
	(0.019)	(0.102)	(0.020)
Total HH annual income (ln)	0.009***	0.005	0.009***
	(0.002)	(0.008)	(0.002)
HH sent out migrants (ref. no migrants)	0.280***	-0.261	0.297***
	(0.017)	(0.149)	(0.017)
Intercept	6.700***	14.375***	6.853***
	(0.096)	(0.483)	(0.090)
Percent of variance explained between households	46.2***	68.2***	52.6***
	(0.004)	(0.011)	(0.006)
Log-likelihood	-99179.3	-5655.8	-91075.5
N	366,014	27,245	299,213

[a] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.
* p < .05, ** p < .01, *** p < .001

Table 5. Random-intercept Logit Models of Current School Enrollment on Sibship Size and Control Variables by Types of Family Arrangements, 1996 Census. (Standard Errors in Parentheses)

Independent variables	Black single-parent nuclear	Black two-parent nuclear	Black extended	Black fostering
Child-level:				
No. of siblings	-0.050***	-0.036***	-0.016	-0.060***
	(0.008)	(0.008)	(0.011)	(0.013)
Male (ref. female)	0.230***	0.170***	0.158***	0.299***
	(0.028)	(0.028)	(0.030)	(0.051)
Age	-0.503***	-0.481***	-0.451***	-0.503***
	(0.011)	(0.011)	(0.011)	(0.019)
Household-level:				
Urban residence (ref. rural)	-0.081*	0.074*	0.054	0.032
	(0.035)	(0.033)	(0.035)	(0.063)
Highest adult education in HH	0.356***	0.350***	0.312***	0.321***
	(0.006)	(0.006)	(0.006)	(0.010)
Highest adult ISEI in HH[a]	-0.001	0.075***	0.022	-0.070
	(0.002)	(0.015)	(0.016)	(0.036)
Missing on highest adult ISEI	0.150***	0.117**	0.092*	0.673***
	(0.043)	(0.040)	(0.040)	(0.081)
Total HH annual income (ln)	0.008	0.012**	0.020***	0.042***
	(0.004)	(0.005)	(0.005)	(0.007)
HH sent out migrants (ref. no migrants)	0.381***	0.062	0.165***	0.220***
	(0.032)	(0.040)	(0.035)	(0.056)
Intercept	7.064***	6.696***	6.381***	6.828***
	(0.181)	(0.179)	(0.189)	(0.333)
Percent of variance explained between households	51.0***	56.6***	52.7***	46.5***
	(0.007)	(0.007)	(0.006)	(0.011)
Log-likelihood	-22342.5	-22004.3	-20102.9	-6806.7
N	75,808	82,553	71,780	22,471

[a] The ISEI variable was divided by 10 to allow for more digits to the coefficient, which shows the effect of a 10 score change in ISEI scale.

* p < .05, ** p < .01, *** p < .001

Table 6. OLS Regression of Years of Schooling on Sibship Size, Birth Order and Control Variables, SSOA 1994. (Standard Errors in Parentheses)

	Overall population		White only		Black only	
Independent variables	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Race (ref. whites)						
Asian	-0.548*	-0.533*				
	(0.266)	(0.266)				
Coloreds	-1.252***	-1.237***				
	(0.217)	(0.217)				
Blacks	-1.543***	-1.538***				
	(0.230)	(0.230)				
No. of siblings	0.010		-0.147***		-0.033	
	(0.052)		(0.044)		(0.082)	
Birth order	0.025		-0.031		0.055	
	(0.058)		(0.031)		(0.080)	
No. of younger siblings		0.016		-0.145***		0.004
		(0.032)		(0.036)		(0.045)
No. of older siblings		0.001		-0.130***		-0.039
		(0.043)		(0.036)		(0.067)
Male (ref. female)	0.380**	0.380**	0.290*	0.293*	0.239	0.238
	(0.135)	(0.135)	(0.120)	(0.120)	(0.200)	(0.200)
Cohort (ref. 20-29)						
30-39	-0.841***	-0.838***	0.178	0.175	-0.962***	-0.957***
	(0.173)	(0.173)	(0.213)	(0.212)	(0.233)	(0.233)
40-49	-1.399***	-1.399***	0.280	0.274	-2.008***	-2.013***
	(0.196)	(0.196)	(0.156)	(0.155)	(0.280)	(0.280)
50-59	-1.628***	-1.627***	-0.041	-0.040	-2.173***	-2.173***
	(0.233)	(0.234)	(0.187)	(0.188)	(0.343)	(0.343)
60-69	-2.286***	-2.285***	0.490*	0.489*	-3.573***	-3.577***
	(0.281)	(0.280)	(0.223)	(0.220)	(0.335)	(0.335)
70-79	-2.236***	-2.232***	-0.075	-0.079	-3.652***	-3.653***
	(0.345)	(0.345)	(0.333)	(0.332)	(0.407)	(0.407)
Residential status (ref. urban)						
Rural	-0.519**	-0.517**	0.197	0.192	-0.425	-0.424
	(0.200)	(0.200)	(0.168)	(0.167)	(0.259)	(0.257)
Other	-0.229	-0.230	-0.096	-0.106	-0.301	-0.300
	(0.162)	(0.163)	(0.173)	(0.174)	(0.296)	(0.298)
Parental education	0.342***	0.341***	0.294***	0.294***	0.328***	0.327***
	(0.021)	(0.021)	(0.031)	(0.031)	(0.028)	(0.028)
Missing on parental education	-0.134	-0.133	0.224	0.238	-1.095*	-1.096*
	(0.270)	(0.271)	(0.163)	(0.164)	(0.465)	(0.465)
Parental ISEI when r was 14 [a]	0.118*	0.118*	0.049	0.048	0.243**	0.243**
	(0.050)	(0.050)	(0.049)	(0.050)	(0.076)	(0.077)
Missing on parental ISEI when r was 14	-0.196	-0.198	-0.195	-0.192	-0.203	-0.205
	(0.168)	(0.169)	(0.223)	(0.224)	(0.199)	(0.199)
Economic condition when r was 14	3.255***	3.246***	1.862***	1.841***	5.029***	5.028***
	(0.408)	(0.407)	(0.299)	(0.299)	(0.616)	(0.616)
Missing on economic condition at 14	-0.370	-0.369	0.565*	0.551*	-1.118***	-1.119***
	(0.222)	(0.222)	(0.234)	(0.233)	(0.296)	(0.296)
Parent present at age 14 (ref. nonpresent)	-0.232	-0.227	-0.165	-0.160	-0.211	-0.203
	(0.223)	(0.223)	(0.237)	(0.241)	(0.311)	(0.312)
Intercept	6.613***	6.692***	7.795***	7.730***	4.847***	4.955***
	(0.529)	(0.514)	(0.541)	(0.532)	(0.577)	(0.552)

R^2	0.558	0.558	0.302	0.303	0.431	0.431
N	7,336		2,086		3,759	

[a] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.
* p < .05, ** p < .01, *** p < .001

Table 7. Random-intercept Logit Models of Current School Enrollment on Number of Younger and Older Siblings and Control Variables, 1996 Census. (Standard Errors in Parentheses)

Independent variables	Whites	Blacks	BSN[a]	BTN[a]	BE[a]	BF[a]
Child-level:						
No. of younger siblings	0.051	0.067***	0.048***	0.085***	0.017***	0.006***
	(0.053)	(0.006)	(0.012)	(0.012)	(0.002)	(0.001)
No. of older siblings	-0.350***	-0.154	-0.258***	-0.217***	-0.147	-0.241***
	(0.068)	(0.087)	(0.015)	(0.016)	(0.090)	(0.030)
Male (ref. female)	0.026	0.318***	0.279***	0.210***	0.175***	0.355***
	(0.084)	(0.017)	(0.034)	(0.036)	(0.036)	(0.061)
Age	-1.285***	-0.631***	-0.656***	-0.648***	-0.574***	-0.621***
	(0.067)	(0.008)	(0.016)	(0.017)	(0.016)	(0.028)
Family arrangements (ref. single-parent nuclear)						
Two-parent nuclear	0.404**	0.263***				
	(0.141)	(0.027)				
Extended	-0.394*	-0.121***				
	(0.180)	(0.027)				
Fostering	-0.622	-0.087*				
	(0.397)	(0.039)				
Other	-2.559***	-1.154***				
	(0.215)	(0.030)				
Household-level:						
Urban residence (ref. rural)	0.314*	0.033	-0.060	0.104*	0.077	0.076
	(0.160)	(0.021)	(0.044)	(0.044)	(0.046)	(0.078)
Highest adult education in HH	0.320***	0.443***	0.472***	0.481***	0.405***	0.404***
	(0.031)	(0.005)	(0.010)	(0.011)	(0.010)	(0.017)
Highest adult ISEI in HH[b]	0.138***	0.039***	-0.001	0.079***	0.016	-0.083
	(0.034)	(0.010)	(0.022)	(0.020)	(0.020)	(0.045)
Missing on highest adult ISEI	0.635***	0.382***	0.222***	0.193***	0.105*	0.745***
	(0.163)	(0.026)	(0.055)	(0.054)	(0.052)	(0.103)
Total HH annual income (ln)	0.001	0.010**	0.012	0.033***	0.039***	0.065***
	(0.015)	(0.003)	(0.007)	(0.008)	(0.008)	(0.011)
HH sent out migrants (ref. no migrants)	-0.463	0.354***	0.399**	0.109*	0.192***	0.251***
	(0.237)	(0.022)	(0.041)	(0.054)	(0.046)	(0.069)
Intercept	21.100***	8.645***	9.007***	8.895***	8.155***	8.400***
	(1.182)	(0.125)	(0.260)	(0.266)	(0.267)	(0.463)
Percent of variance explained between households	69.0***	52.5***	51.8***	57.3***	53.8***	48.5***
	(0.032)	(0.007)	(0.016)	(0.015)	(0.016)	(0.032)
Log-likelihood	-5574.5	-90032.8	-21992.6	-21653.1	-19888.8	-6744.1
N	27,245	299,213	75,808	82,553	71,780	22,471

[a] BSN, BTN, BE and BF columns from the left to right are short for "Black single-parent nuclear", "Black two-parent nuclear", "Black extended" and "Black fostering" families, respectively.
[b] The ISEI variable was divided by 10 to allow for more significant digits in the coefficient, which shows the effect of a 10 score change in ISEI scale.
* p < .05, ** p < .01, *** p < .001

Table 8. Random-intercept Logit Models of Current School Enrollment on Sibship Size Stratified by Birth Order, 1996 Census [a]. (Standard Errors in Parentheses)

Sibship size effect	Overall	Whites	BSN[b]	BTN[b]	BF[b]
First child	-0.071***	-0.137**	-0.145***	-0.094***	-0.167***
	(0.009)	(0.053)	(0.018)	(0.020)	(0.045)
N=162,744	112,421	12,291	17,754	14,660	5,618
Second child	-0.192***	-0.420***	-0.279***	-0.225***	-0.228***
	(0.011)	(0.075)	(0.040)	(0.035)	(0.037)
N=132,305	81,165	9,511	17,751	19,169	4,709
Third child	-0.205***	-0.496***	-0.240***	-0.164***	-0.136*
	(0.013)	(0.086)	(0.025)	(0.023)	(0.055)
N=112,168	69,932	4,219	15,575	19,287	4,155
Fourth child	-0.224***	-0.009	-0.282***	-0.220***	-0.309***
	(0.018)	(0.279)	(0.035)	(0.029)	(0.064)
N=79,633	48,945	999	11,685	14,801	3,203
Fifth child	-0.254***	-0.816	-0.316***	-0.277***	-0.330***
	(0.026)	(0.503)	(0.054)	(0.047)	(0.093)
N=48,670	29,678	175	7,372	9,194	2,251
Sixth child and higher	-0.244***	3.229	-0.279*	-0.169*	-0.287
	(0.053)	(4.217)	(0.125)	(0.074)	(0.191)
N=38,571	23,873	50	5,671	6,442	2,535
Total no. of cases	366,014	27,245	75,808	82,553	22,471

[a] Multi-level models are run separately by birth order controlling for a similar set of covariates as in Table 4 and 5; each coefficient represents the effect of sibship size on the outcomes of children in that particular birth order. Only coefficients of sibship size are presented, and other covariates are omitted from the table.
[b] BSN, BTN and BF columns are short for "Black single-parent nuclear families", "Black two-parent nuclear families" and "Black fostering" families, respectively.
* $p < .05$, ** $p < .01$, *** $p < .001$