California Center for Population Research
University of California - Los Angeles

The Changing Relationship Between Family Size and Educational Attainment Over the Course of Socioeconomic Development: Evidence From Indonesia

Vida Maralani

CCPR-017-04

November 2006
Last Revised: May 2007

California Center for Population Research
On-Line Working Paper Series

THE CHANGING RELATIONSHIP BETWEEN FAMILY SIZE AND EDUCATIONAL ATTAINMENT OVER THE COURSE OF SOCIOECONOMIC DEVELOPMENT: EVIDENCE FROM INDONESIA*

Vida Maralani

Robert Wood Johnson Foundation Health and Society Scholars

University of Pennsylvania

Last Revised: May 25, 2007

*Direct correspondence to: Vida Maralani, RWJF Health and Society Scholars, 3641 Locust Walk, Room 304, Philadelphia, PA 19104. Email: maralani@wharton.upenn.edu. This research was supported by The William and Flora Hewlett Foundation and the National Institute of Child Health and Human Development. I gratefully acknowledge the Robert Wood Johnson Foundation's Health & Society Scholars program for its financial support. I owe many thanks to Elizabeth Frankenberg, Jinyong Han, Robert Mare, William Mason, Douglas McKee, Judith Seltzer, Duncan Thomas, and Donald Treiman for their valuable comments and advice. I presented earlier versions of this paper at the International Sociological Association Research Committee 28 on Social Stratification and Mobility Meetings in Los Angeles, CA, August 2005 and the Population Association of America Annual Meeting in Boston, MA, April 2004.

THE CHANGING RELATIONSHIP BETWEEN FAMILY SIZE AND EDUCATIONAL ATTAINMENT OVER THE COURSE OF SOCIOECONOMIC DEVELOPMENT: EVIDENCE FROM INDONESIA

ABSTRACT

Studies from developed countries show a negative correlation between family size and children's schooling while results from developing countries show that this association can range from positive to neutral to negative depending on the context. The body of evidence suggests that this relationship changes as a society develops but this theory has been difficult to assess because the existing evidence requires comparison across countries with different social structures and at different levels of development. The world's fourth most populous nation, Indonesia has developed rapidly in recent decades. This context provides the opportunity to study these relationships within the same rapidly developing setting to see if and how these associations change. Results show that in urban areas the association between family size and children's schooling was positive for older cohorts while for more recent cohorts family size and schooling are negatively related. Models using instrumental variables to address the potential endogeneity of fertility confirm these results. In contrast, rural areas show no significant association between family size and children's schooling for any cohort. These findings highlight how the relationship between family size and children's schooling can differ within the same country and change over time as contextual factors evolve with socioeconomic development.

THE CHANGING RELATIONSHIP BETWEEN FAMILY SIZE AND EDUCATIONAL ATTAINMENT OVER THE COURSE OF SOCIOECONOMIC DEVELOPMENT: EVIDENCE FROM INDONESIA

INTRODUCTION

Numerous studies of educational attainment in the United States have shown that schooling is negatively correlated with sibship size. That is, children with fewer brothers and sisters obtain more schooling than those with more siblings. This negative association exists for many different measures of children's human capital including completed schooling, standardized test scores, and grades, and holds even after family socioeconomic characteristics are controlled (Blake 1989; Featherman and Hauser 1978; see Steelman, Powell, Werum and Carter 2002 for a comprehensive review of this literature). In the sociological literature, this finding is often explained using an argument of finite resources: parents have limited time, money, and patience to devote to the education of their children, and those with fewer children can invest more per child. This theory of resource dilution fits well with the classic notion of the quality-quantity trade-off in family economics (Becker and Tomes 1976, Becker 1991).[1]

In recent years, two lines of research have called into question this seemingly robust negative relationship between family size and children's schooling. First, some scholars have argued that this finding might be biased or spurious. If this negative association is explained by factors such as unobserved family characteristics that are not controlled in the analysis, then our understanding of the true relationship between these variables may be erroneous (Guo and VanWey 1999a). Moreover, few studies have addressed the concern that fertility and children's

[1] The competing "confluence theory" suggests that additional children lower a family's average intellectual environment, leaving children in larger sibships and those with higher birth order worse off (Zajonc and Markus 1975). Empirical analyses, however, have not found much support for this theory (Hauser and Sewell 1985; Retherford and Sewell 1991).

schooling are jointly determined, despite evidence suggesting this is the case in some settings or socio-historical periods (Wolfe and Behrman 1984; Caldwell, Reddy, Caldwell 1985; Axinn 1993). Indeed the very notion of a quality-quantity *trade-off* suggests that these processes are inter-related in complex ways that contradict the basic assumptions of regression analysis or fixed effects models. Using data from Norway and Israel, for example, recent research in economics shows that once fertility and child outcomes are modeled jointly, these variables display no meaningful association (Black, Devereux, and Salvanes 2005; Angrist, Lavy and Schlosser 2006). Qian (2006), in contrast, finds mixed results using data from China. She finds that an increase from zero to one sibling has a positive effect on children's schooling while an increase from one to two siblings has a negative effect. Further research that addresses these issues especially in a range of countries will advance our understanding of the extent of such potential bias.

Second, a growing literature on the nature of these relationships in the developing world shows that the negative correlation found so consistently in more developed countries is not necessarily generalizeable. Instead, the association between family size and children's outcomes varies greatly by time and place and ranges from negative to positive depending on the specific context (see Buchmann and Hannum 2001 for an overview of this literature). A few examples suffice to show the range. The evidence from Thailand and Brazil suggests a negative association between family size and educational attainment (Psacharopoulos and Arriagada 1989; Knodel, Havanon and Sittitrai 1990). Results from Vietnam show that the association is negative only for families with six or more children and effects are modest once other family characteristics are controlled (Anh et al. 1998). The evidence from Botswana and Kenya, on the other hand, suggests the reverse is true: educational attainment has a positive relationship with family size

(Gomes 1984; Chernichovsky 1985). Even within the same country, studies show that patterns differ by subgroup. Among Israeli Jews, for example, family size has a negative association with educational attainment. Among Israeli Muslims, who are less advantaged socioeconomically, live in less urban settings, have extended kinship networks, and have much higher fertility rates, family size and educational attainment are not associated (Shavit and Pierce 1991; but see Angrist, Lavy and Schlosser 2006).

Four ideas emerge from this array of evidence. First, the relationship between family size and educational attainment is related to a society's level of development, modes of production, and access to schooling, which in turn shape the relative influence of the family on the schooling of children (King 1987; Lloyd 1994; Desai 1995). In certain contexts or at certain stages of development, having more siblings to share household and labor market work may provide children with more resources for schooling. Thus, in some settings the quality-quantity trade-off may not hold and the desire to have better educated children may not necessarily lead parents to choose smaller families (Mueller 1984, Gomes 1984). These macro socioeconomic mechanisms relating family size and children's schooling might include the availability of schools, transportation and communication infrastructure, and participation in labor intensive production such as agriculture. These mechanisms might be most apparent when comparing less developed rural contexts with more developed urban ones.

Second, context-specific factors such as family organization and cultural roles determine wealth flows between parents and children, whether the burden of child rearing is limited to the nuclear family or extended across broader kin networks, whether and how much school-aged children work inside and outside the home, and whether these factors change as societies develop or as overall levels of educational attainment increase (Caldwell, Reddy, and Caldwell 1985). In societies where parents bear most of the cost of schooling and where the costs are high, we might expect a negative relationship between family size and educational attainment. In societies with extended kinship networks and lower costs of schooling, the relationship may be neutral or positive (Lloyd and Blanc 1996). The effects of family size on children's schooling may also differ within families by factors such age, sex, or birth order of children (Parish and Willis 1993; Lloyd and Gage-Brandon 1994). Context-specific mechanisms relating family size to children's schooling might include financial relationships within families, norms about how much schooling children should complete, parents' preferences by child sex or birth order, labor market returns to schooling, and the costs of schooling.

Third, the evidence highlights the interdependence of family size, family structure, and educational attainment. Family size and educational attainment are likely to be jointly determined, at least to some degree, with families choosing the level of fertility that is likely to produce children with the preferred level of education for a given family, context, or society. The relationship between family size and children's educational attainment can have demographic feedbacks as well. Small families may raise educational attainment, which in turn may lower fertility in the next generation. Moreover, if the effect of family size grows more negative or positive over time, then these aggregate demographic relationships may intensify or accelerate.[2]

Fourth, the relationship between family size and children's schooling might change as development brings changes in income, consumption, urbanization, migration, educational opportunities, gender roles, and family organization. Government policies to invest in women's schooling change women's relative position in families and society, and update norms about schooling girls. Of course, these updated norms also double the number of children families must

---

[2] See Preston (1976) for a precise demonstration of the relationship between the average number of children ever born for a cohort of women and the average sibship size of the offspring of those women.

educate relative to when only sons were schooled. Industrialization and modernization increase the returns and hence the demand for schooling. But higher levels of schooling may be cost prohibitive despite families' growing expectations for their children's education.

Socioeconomic development also brings increased differentiation between urban and rural areas, which might translate to differences in the association between family characteristics and children's schooling by geographical location. Urbanization improves access to schools by improving transportation and communication infrastructure, and allowing closer proximity to this infrastructure. As the formal job sector grows, urban residents have better access to these jobs, which require more investment in schooling. Although city dwellers might be richer on average, costs of living are also higher here making the marginal expense of an additional child different between urban and rural areas. The agrarian and informal economy of rural areas, in contrast, might allow families more flexibility in resource sharing and balancing home, work and school responsibilities. Educational opportunities, however, might be quite limited rural areas and the formal sector more difficult to enter.

Taken together, the evidence from developed and developing nations suggests that there is no axiomatic relationship between family size and children's schooling. Rather, this relationship varies by context and by levels of modernization and development. Much of the existing evidence, however, requires us to compare across different cultures and socioeconomic settings to determine the role of development in shaping the relationship between family size and children's schooling. Although it may be tempting to use an array of cross-sectional evidence from across the world to infer something about the role of socio-economic development, family organization or government policies, comparing across different cultures and socioeconomic contexts can be difficult and misleading (Thornton 2001). Comparisons across time within the same country help with this problem.

Although numerous studies have examined the changing association of family size and children's schooling across cohorts in the United States and Europe (Kuo and Hauser 1995; Hauser and Featherman 1976; de Graaf and Huinink 1992), few analysts have examined cohort trends in a developing country. The few studies that do consider changes within the same developing country generally use only two cohorts (e.g., Hermalin, Seltzer, and Lin 1982; Pong 1997; Sudha 1997; see Parish and Willis 1993 for an exception). These offer only a limited view of the development process. One way to improve our understanding of these relationships is to study these associations across many different birth cohorts within one rapidly developing country. Indonesia offers such a context.

Indonesia, the world's fourth most populous country, has experienced enormous socioeconomic growth in recent decades. Between 1970 and late 1990s, Indonesia had one of the fastest growing economies in the world. Over this period, the infant mortality and total fertility rate (TFR) declined dramatically, and life expectancy increased by almost 20 years (UNICEF 2000; Indonesia Central Bureau of Statistics 2003). There were also large gains in educational attainment, especially for women. Overall, participation in agriculture declined and industry grew. This setting provides a unique opportunity to examine how the relationship between family size and educational attainment changes as a society develops.

This paper examines the association between family size and composition and educational attainment for four cohorts of Indonesians born between 1948 and 1991. To preview the results, large families were associated with significantly better educational outcomes for the oldest urban birth cohort but are associated with significantly poorer educational outcomes in more recent urban cohorts. Instrumental variables analysis shows that this negative relationship in the recent urban cohorts remains even in models that address the possibility that fertility and children's schooling are jointly determined. In rural areas, in contrast, there is little evidence of a significant association between family size and children's schooling for any cohort.

THE INDONESIAN CONTEXT

Indonesia is an archipelago nation of thousands of islands, hundreds of ethnicities, and nearly 235 million inhabitants in 2007. Following 300 years of Dutch colonization and several years of Japanese occupation during World War II, Indonesia gained independence in 1949. During the 1950s and 1960s, most Indonesians lived in rural areas and poverty was endemic. The analysis starts with the 1948-1957 birth cohort, whose members were school age from the mid 1950s to the late 1960s. The Indonesian economy was largely agricultural in 1960, with about 75 percent of the labor force engaged in agriculture and 85 percent living in rural areas (Walton 1985; UNICEF and Government of Indonesia 1988). This was a period marked by severe economic strain, hyper-inflation, and infant mortality rates of about 150 per 1000 births (Hugo et. al 1987). Average life expectancy at birth was about 41 years, the per capita gross national product was $125 (USD), and only 27 percent of women ages 20 to 24 were literate (Hull 1987; Jones 1990).

A military coup in 1965 brought a shift in political power and a new focus on domestic development. Starting in 1967, the economy began to grow rapidly as the nation began to export oil and natural gas and attract foreign investment (Walton 1985). The 1958-1967 cohort was born during this time of transition and was school age just as the Indonesian economy began a period of massive growth. Between 1970 and the late 1990s, Indonesia went from being one of the poorest countries in the world to one at a middle level of socioeconomic development.

Indonesia used windfall profits from rising oil prices in the 1970s to finance an extensive educational expansion at the primary school level. From 1974 to 1979, the government built more than 61,000 primary schools and abolished tuition at public primary schools (Duflo 2001; Oey-Gardiner 1997). Primary school enrollments rose from 60 percent in 1974 to 94 percent by 1984 (Jones and Hull 1997). Meanwhile, the dramatic pace of socioeconomic development continued. Between 1970 and 1980, the infant mortality rate fell from 118 to 98 per 1000 births, life expectancy rose from 47 to 53 years, and female literacy grew from 47 to 66 percent (Iskandar 1997; Jones and Hull 1997; Firman 1997; Jones 1990). By the early 1980s, only 55 percent of the labor force participated in agriculture and per capita gross national product had risen to $447 (USD) (Hull 1987; Walton 1985). The 1968-1977 birth cohort was school age at the height of this period of growth and development.

After a period of turbulent oil prices and economic adjustment, the economy continued to grow in the 1980s. By the early 1990s, the infant mortality rate had declined to 66 per 1000, female literacy was 93 percent, primary school enrollment has risen to 98 percent, and 31 percent of Indonesia's population lived in urban settings (Jones and Hull 1997; Cobbe and Boediono 1993). The youngest birth cohort described below, born between 1978 and 1981, was school age in the late 1980s and 1990s after more than 20 years of sustained growth and development.

These four birth cohorts straddle another important part of Indonesia's development experience. Historically in Indonesia, fertility was quite high and exhibited an inverted U-shape pattern with women's educational attainment (women with some primary or middle school had higher fertility than women with no schooling or with 12 plus years). This pattern was stronger in urban than rural areas. Also, average fertility was higher in urban areas, and fertility exhibited a positive relationship with household income (Hull and Hull 1977; Cobbe and Boediono 1993).

The 1970s and 1980s marked the implementation of Indonesia's state-sponsored family planning program (Gertler and Molyneaux 1994). The program distributed contraceptives and educated women on how to use them, promoted two-child families, and encouraged women to postpone marriage. The program aimed to increase both contraceptive supply and spur contraceptive demand by changing fertility preferences and norms. Concomitant with sustained economic growth and improvements in women's socioeconomic position, the program was hugely successful. Between 1970 and the early 1990s, TFR declined from 5.6 to 2.9 children per woman (Jones and Hull 1997; Gertler and Molyneaux 1994). Women's reports of their ideal family sizes also declined from an average of 4.2 in 1972 to about 3.2 in 1987 (Jones 1990). By the 1980s, the inverted U-shape pattern of fertility by women's schooling began to flatten out and disappear, especially in urban areas. Moreover, the positive correlation between household wealth and fertility declined and a negative correlation emerged. The two oldest cohorts (1948-1957 and 1958-1967) were born before the implementation of the family planning program while the more recent cohorts (1968-1977 and 1978-1981) were born after.

The formal school system in Indonesia has four basic levels. Primary school covers grades one to six; junior secondary school covers grades seven to nine; senior secondary covers grades 10 to 12; and post-secondary covers all years from 13 upward. National final examinations separate each level of formal schooling. Urban-rural disparities in schooling were large in the 1960s and 1970s and have narrowed in more recent decades. Still, enrollment ratios after age 12 in rural areas in the 1990s are similar to those in urban areas in the 1970s (Oey-Gardiner 1997, Figure 8.8). Although primary schooling has become widespread throughout

Indonesia, higher educational levels remain quite expensive, and enrollment differences between urban and rural areas are substantial at these higher school levels (Oey-Gardiner 1997).

DATA AND METHODS

The analysis is based on the 1993 and 1997 waves of the Indonesian Family Life Survey (IFLS), a comprehensive longitudinal socioeconomic and health survey. The survey represents an area that includes 83% of Indonesia's population (See Frankenberg and Thomas (2000) for detailed documentation on the IFLS). The sample used in this study includes approximately 3,200 families with more than 13,000 living children. All variables are measured as of 1997, except for those respondents who died between 1993 and 1997 or households not found in 1997 (about 5% of the 1993 sample). For these cases I use the information provided in 1993, except for the schooling of respondents ages 19 or younger, which I leave missing because it is likely to be censored and systematically underreported when compared to the 1997 data. Overall, the data are quite complete and the variables used in the analyses include little missing data.

For each ever-married female respondent (Indonesia is a society with nearly universal marriage and essentially no nonmarital fertility) I assemble a full count of all children, alive or dead, coresident or living elsewhere. I also identify the schooling of each woman's husband, whether current or former. These women's children are the units of analysis. The woman and her husband are the parents and their children constitute the sibship.[3] For mothers, fathers and children, I measure educational attainment by years of completed schooling. I control for mother's age in all models, and report robust standard errors that correct for clustering of

[3] I use the terms sibship size and family size interchangeably. These terms mean all of a child's brothers and sisters plus the child himself/herself. When I count a child's siblings I use the sibship size minus one.

multiple children born to the same woman. I also control for province of residence, a key sampling stratum along with urban-rural status in all statistical models.

Because urban and rural areas are quite different in Indonesia, especially in terms of resources and infrastructure, I estimate separate models by rural status. To identify whether a child grew up in an urban or rural location, I use the mother's detailed migration history to determine each child's region of residence when he or she was age 12. In their retrospective reports, women report whether each of their places of residence was a village, small town, or big city. I classify villages as rural and small towns and big cities as urban areas.[4]

I assemble four birth cohorts spanning ages 16 to 49 in 1997. The first three cohorts are restricted to families with mothers who are at least age 41 and children ages 20 and older in order to measure completed fertility and completed schooling. This sample of children, born between 1948 and 1977, provides information on three broad cohorts: those born 1948-1957, 1958-1967, and 1968-1977. I examine the relationship between family size and completed education by cohort and urban-rural status using ordered probit models with completed education classified in categories as the dependent variable and family size and household characteristics as independent variables.[5] Children's education is categorized in the most basic levels of the Indonesian educational system: none, primary (1-6 years), junior secondary (7-9 years), senior secondary (10-12 years) and post secondary (13 or more years). To these three birth cohorts I add a fourth cohort of young children, those born 1978-1981 (ages 16 to 19 in 1997). Here, there are no restrictions on the mother's age and fertility is not completed in many cases. Because

[4] About two-thirds of Indonesia's total population and 70 percent of its urban population reside on Java. Java and the outer islands have a mix of urban and rural populations although the proportion urban differs across islands. In the 1990s, Java was about 40 percent urban compared to an average of 26 percent for the other areas (Firman 1997).

[5] About 7% of those in the 1968-1977 cohort were enrolled in school in 1997 (concentrated at ages 20 to 23). More than 90% of these respondents had completed grade 12 or higher, with a median and mode of 15 years of schooling. Therefore, nearly all are assigned to the highest education category despite their censored schooling.

many members of this most recent birth cohort have not completed their schooling, I use continuation beyond primary school as the measure of educational attainment for this sample.

Educational attainment is correlated with household wealth, especially in developing countries (Filmer and Pritchett 1999; Chernichovsky and Meesook 1985; Hull and Hull 1997). The analyses shown below do not control for household wealth because the data do not include an appropriate measure of family wealth for most individuals used in this study, namely a measure of family wealth when the individuals were school-aged. The IFLS does collect extensive information about assets, consumption, and expenditures at the time of the survey but most individuals used in this analysis were adults at that point (they were ages 16 to 49 in 1997). Thus, the measure of wealth comes after respondents' schooling is completed and is likely endogenous to their educational attainment. Because using current wealth to predict previous schooling is problematic, I report results from models that do not include this measure. Mother's and father's schooling, which serve as noisy proxies for family wealth (Hull and Hull 1977), are controlled in all models. Appendix A describes sensitivity tests that confirm that the results and patterns shown below are robust to omitted measures of household wealth.[6]

Another methodological issue involves the effect of mortality on the sample. Developing countries generally have high rates of infant and child mortality. This means that a full count of a woman's live births may not represent the actual number of living siblings a child had while growing up. Nor is it obvious which count represents the correct measure of sibship size. Infants who die soon after birth may not compete with older children for resources that relate to schooling. But sick infants may require much of their parents' attention, leaving older children with increased responsibilities within the household and less time for school. In the results

[6] I am grateful to Duncan Thomas for letting me borrow his coding of the IFLS household expenditure data for these sensitivity tests.

below, I use a count of living siblings as the measure of family size. This measure facilitates comparisons across these Indonesian cohorts, which have experienced steady declines in infant mortality. The full count of all live births would implicitly embed this changing pattern of infant mortality in the construct. In these Indonesian cohorts, including controls for the number of siblings who have died does not change the estimated effects of family size or any of the other independent variables in the analysis. In other settings, however, these different measures of family size may give different results.

Selective mortality may be a potential problem in at least two other ways. First, the oldest cohort (born 1948-57) may be positively selected because their mothers, who are on average in their 60s in 1997, survived long enough to appear in the IFLS sample. If the relationship between family size and children's schooling differs by mother's schooling, and mothers with less schooling are systematically censored from the sample, then the results may be biased. Fortunately, analyses confirm that there are no meaningful interactions between family size effects and women's schooling for the oldest cohort, which is the cohort most likely to suffer from selective survival of mothers. Second, the IFLS does not collect information on the schooling of children who have died more than 12 months before the survey, or detailed pregnancy histories from women over age 49. This means that children identified as the "oldest" may not be, strictly speaking, first born, if the women had any children born earlier who had died. Thus, measures of being the oldest boy or girl in a family represent being the oldest living child at the time of the survey.

Analyzing the relationship between family size and children's schooling also involves grappling with the difficult problems of unobserved heterogeneity and joint determination. If parents have characteristics not captured by the included covariates that are correlated with family size and children's schooling, or if parents choose their family size with an eye to how much schooling they would like to provide their children, then standard regressions of the type presented here provide biased estimates of the effect of family size on children's education. Although much of the research on the effects of family size on children's outcomes assumes that fertility is exogenous to children's outcomes (see DeGraff and Bilsborrow 2003 for a partial review of the literature in this regard), some recent research has tried to address these concerns through the use of fixed effects designs or by using twinning as a "natural experiment" that causes an exogenous increase in fertility.

Fixed effects models net out a constant individual or group level factor that is unobserved and is normally a component of the error term (see for example, Guo and VanWhey 1999a, which uses repeated observations on individuals, or Parish and Willis 1993, which uses differences between siblings).[7] The fixed effect approach assumes that the unobserved characteristics are not related to the dependent and independent variables in complex ways. In particular, the unobserved effect is assumed to be additive, meaning that the effect of the unobserved factor does not differ across different levels of the other regressors. To use a concrete example, this means assuming that the effect of parents' unobserved preferences for child's schooling does not, for example, differ by child's sex or birth order. The fixed effect approach also assumes that the effect of the unobserved factor on the dependent variable does not change over time.

The fixed effect approach, however, does nothing to resolve concerns about more complex relationships such as "trade-offs" or jointly determined processes. In this case, the analyst might model both processes together (as in a simultaneous equation model with

[7] See Phillips (1999); Downey, Powell; Steelman, and Pribesh (1999), and Guo and VanWhey (1999b) for a discussion of the relative merits of Guo and VanWhey's approach.

exclusion restrictions) or rely on an experiment that provides exogenous variation on one of the jointly determined processes without affecting the other one. Researchers in economics have utilized twinning as an instrument to measure the exogenous effects of fertility on outcomes such as labor force participation (Rosenzweig and Wolpin 1980a; Jacobsen, Pearce, and Rosenbloom 1999; Angrist and Evans 1998), women's economic outcomes (Bronars and Grogger 1994) and children's schooling (Rosenzweig and Wolpin 1980b; Black, Devereux, and Salvanes 2005; Angrist, Lavy and Schlosser 2006). The instrumental variables approach and the use of twins in particular have their own set of shortcomings.[8] But there is also a major practical problem with using twins as instruments—twinning is a rare event and one that is very difficult to capture in sample surveys. In this paper I use an alternative instrument that relies on women's report of miscarriages, which are more common events and widely available in survey data. I use this approach to test the robustness of the results to assumptions about joint determination. I discuss the approach in more detail below.

RESULTS

Changing Patterns Across Cohorts

Table 1 presents sample means and proportions for the full sample and each birth cohort. Families are relatively large in this Indonesian sample. In the first three birth cohorts, children have an average of four and a half living siblings. Average number of children ever born ranges

[8] Rosenzweig and Wolpin (2000) provide an excellent discussion of the limitations of using twins as instruments. Rosenzweig and Wolpin (1980a) point out that there are biological differences between twins themselves, and between twins and non-twin infants. Twins have substantially lower gestational age, lower birth weight, and more complicated deliveries. Twins might therefore be less healthy, on average, than singletons. And while twinning might represent the addition of an extra child in the short term, it is not equivalent to adding an additional child to a woman's completed fertility. The evidence shows that women adjust their subsequent fertility to compensate for the unexpected extra child. The completed fertility of women who have twins is substantially less than an average of one child higher and this difference has been declining over time and varies between groups (Jacobsen, Pearce, and Rosenbloom 1999; Bronars and Grogger 1994).

from about seven in the oldest cohort to about 6.4 in the 1968-1977 cohort. Family sizes are smaller in the most recent cohort but this is also the cohort in which some of the mothers are still in their peak childbearing years. In general, one and two child families are uncommon in Indonesia. In the first three birth cohorts, families with three, four, five, six, seven, and eight children are all relatively common. Despite a clear trend towards smaller families over time, the range across cohorts of the number of children born to Indonesian parents is quite wide and fairly high for the years covered by these data.

Average levels of schooling have increased steadily over the past forty years.[9] The proportion of those with no schooling has decreased steadily while the proportion of those with senior and post secondary schooling has increased in the cohorts with completed schooling (1948-1977). At the same time, differences in educational attainment by sex have gone from a 1.8 year advantage for males in the oldest cohort to parity in the most recent cohort.[10]

Table 2 shows the distribution of children's schooling by cohort and place of residence for the first three birth cohorts. For those living in urban areas at age 12, average levels of schooling have increased by about 1.6 years across the three birth cohorts (from about 9.3 years to 10.9 years). The median level of schooling has increased from completing junior secondary to completing senior secondary while the top of the distribution has also increased.[11] Most of the educational gains in the urban cohorts have come at the bottom of the education distribution. The bottom decile has gone from having only some primary schooling to completing that level and the 25th percentile has gone from completing primary to completing junior secondary.

[9] Means for the most recent cohort (1978-1981) are not directly comparable to the other three cohorts because schooling and family size are not completed for this younger group.

[10] Because schooling is censored for this cohort, it is impossible to know whether the gender gap will remain closed once this cohort completes its schooling. If boys have higher continuation probabilities into the higher levels of schooling, which the data suggest is still true for this recent IFLS cohort, then a small gender gap favoring boys may emerge by the time the cohort completes its schooling.

[11] About 7% of those in the 1968-1977 cohort are still enrolled in school and it is likely that the top decile will improve for this cohort once their schooling is completed.

In contrast, those who grew up in rural areas have substantially lower levels of average attainment and a more disadvantaged distribution of schooling across all cohorts. Average attainments are three to four years lower here than in the urban cohorts. The median level of schooling has remained stable at about six years across all rural cohorts although the interquartile range has shifted from straddling the primary grades to a range that includes junior and senior secondary grades. For the bottom rural decile, the floor has increased from no schooling to an average of three years. Those in the top decile in rural areas achieved relatively high attainments across all cohorts. Even in the top decile, however, few individuals who grew up in rural areas progressed beyond senior secondary.

Figure 1 shows completed schooling by number of siblings for those born between 1948 and 1977 (ages 20 to 49 in 1997). Respondents with zero or one sibling, a rare sibship size in this sample, have about one year less schooling than those with two to seven siblings. Those with eight or more siblings complete more schooling on average. Remarkably, children's educational attainment in the cross-section displays little relationship with family size for those with two to seven siblings, a substantial range by most demographic standards. Even more surprisingly, the cross-sectional data show a positive relationship between very large families and children's educational attainment.

This gross relationship, however, hides substantial variation by cohort and place of residence. Figure 2 shows the same sample as Figure 1, stratified by birth cohort and urban-rural residence. The top panel provides a detailed specification of number of siblings. The bottom panel collapses adjacent categories to provide a more tractable grouping for analysis. Comparing the two graphs shows that the more parsimonious grouping stays true to the trends apparent in the more detailed representation. Figure 2 also highlights the dramatic increase in schooling in Indonesia over time and the large differences in schooling by urban-rural residence.[12]

In the first two urban birth cohorts (1948-1957 and 1958-1967) those with four or fewer siblings have lower educational attainment than those with five or more siblings. In the 1948-1957 cohort, for example, those with five or six siblings have an average of one and a half more years of schooling than those with fewer siblings while those with seven or more siblings have about two and a half more years of schooling, on average. This relationship is quite different in the 1968-1977 urban birth cohort. This cohort shows large gains in average levels of schooling overall and has a negative relationship between family size and children's schooling. Each larger category of siblings is associated with lower levels of average schooling, with a difference of about 1.5 years in average attainment between the smallest and largest sibship categories.

In the first two rural cohorts, there is a positive relationship between family size and completed schooling. In the 1948-1957 rural cohort those from the largest sibships have the highest levels of schooling, and this association is monotonically positive for those in the 1958-1967 rural cohort. In contrast, the most recent rural cohort exhibits no relationship between family size and educational attainment, a quite different pattern than has emerged for its urban counterpart. The patterns in Figure 2 suggest that the overall association between family size and children's schooling has been changing across Indonesian cohorts and differs by geographical area. The offsetting patterns in Figure 2 also explain the apparent lack of a pattern when family size and schooling are examined in the pooled cross-section (as in Figure 1).

Table 3 shows results from multivariate models that examine the relationship between family size and completed education controlling for other family characteristics.[13] For each

[12]None of the means for the 1968-1977 rural cohort are significantly different from each other. In all other cohorts, the mean for the 7+ category is significantly different from mean for 0-2 and 3-4 category and the mean from 0-2 is not significantly different from mean for 3-4.

cohort and place of residence I show both linear and categorical estimates of the effects of family size. All models control for mother's and father's education, child's sex, whether the child was the oldest child in the family, child's age, mother's age, and the regional province in which the household was located. Parents' education, child sex, and birth order are mediators that have been shown relevant in other studies of these relationships. The remaining variables control for sample composition. Models 3.1 and 3.2, shown in the first two columns, give estimates for the oldest urban birth cohort. For this cohort, the association between family size and children's schooling is positive. Holding all else constant at the sample mean, each additional sibling increases the probability of being in a highest education category by about 0.01. This positive association disappears in the 1958-67 urban cohort (Models 3.3 and 3.4). In this urban cohort, there is no meaningful relationship between family size and children's schooling net of family and individual characteristics. By the 1968-77 urban cohort, however, a negative association emerges (models 3.5 and 3.6). For this birth cohort, each additional sibling decreases the probability of being in a highest education category by about 0.013. The results are similar for both the linear and categorical specifications of family size. Over a span of three birth cohorts, the association between family size and children's schooling goes from positive to neutral to negative for those who lived in urban areas at age 12.

In these urban birth cohorts, daughters are disadvantaged in the oldest cohort but not in the more recent ones—a sign of the closing gender gap in schooling at least in urban areas. At

---

[13] The remainder of the analysis focuses on the additive effects of family size on children's schooling both to draw out the major contours of these relationships across cohorts and to provide tractable models for the IV analysis. This simplification ignores two interactions that are descriptively interesting but do not change the substantive patterns shown here. First, in all three urban cohorts, the association between family size and children's schooling differs by fathers' schooling. The effect of family size is diminished or made more negative when fathers are highly educated. This interaction is not significant in the first two rural cohorts, and is marginally significant in the 1968-77 rural cohort. Second, in some cohorts, fathers with more schooling provide an extra boost to their daughters' education. Other interactions such as family size by the sex of the index child or family size by mother's schooling show no meaningful patterns across cohorts (interaction models not shown).

the same time, being the oldest girl in a family is a disadvantage in more recent urban cohorts but not in the oldest cohort. This pattern might reflect emerging inequality within the family or capture a transitory period effect during a time of rapid school expansion. Girls who were born at the very beginning of this cohort may not have enjoyed the full benefit of the massive school expansion of this period. But the difference between the estimate for being the oldest boy (which is not significant) and being the oldest girl suggests an alternative possibility. Parents' attitudes towards schooling girls may have evolved over this period to the benefit of younger girls.[14]

Models 3.7 to 3.12 show the results for the three rural birth cohorts. In these rural cohorts, there is no significant association between family size and children's school once other family characteristics are controlled. The categorical specification for the oldest rural birth cohort (Model 3.8) suggests a marginally significant negative association of having three or four siblings relative to seven or more. The coefficients, however, are not jointly significant and the linear term is also not significant. This marginally positive association between larger families and children's schooling in the oldest rural cohort is consistent with the relationship found in the oldest urban cohort. Indeed, the overall pattern of results for the rural cohorts, although smaller in magnitude and not significant, is similar to that found in the urban cohorts.[15] I discuss possible explanations for these urban rural differences in the more detail below.

Daughters complete less schooling than sons across all rural cohorts, although this disadvantage diminishes over time. The marginal effect of being female on the probability of being in the highest education category shrinks from -0.027 to -0.01. Nonetheless, women have

[14] Overall, however, this association should be interpreted cautiously. Children in the most recent cohort are more likely to have siblings who survive and who are captured in the survey. Therefore children identified as the "oldest" in these cohorts are less likely to be positively selected than those identified as the oldest child in the oldest urban cohort (1948-57) when both mothers and children were older and mortality rates were higher.

[15] In a joint model, the two-way interactions of family size and cohort and rural status and cohort are significant. The two-way interaction of siblings and rural status is not significant. The three-way interaction of family size, cohort and rural status is also not significant (models not shown).

not made the same educational gains relative to men in rural areas as they have in urban settings, another example of differential associations by levels of socioeconomic development (in this case by urban-rural status within the same country). The association between being the oldest child and completed schooling shows two interesting patterns. The coefficient for being the oldest boy is negative in the middle rural cohort and for being the oldest girl negative in the younger rural cohort. These patterns suggest that as schooling became more widely available in Indonesia, boys first benefited more than girls and younger children benefited more than older children. In Indonesia, however, these patterns appear to be related to the availability of schools rather than strong preferences by birth order. This pattern does not exist for older boys in urban areas and dissipates for older girls by the 1968-1977 urban cohort.

Overall, these findings reproduce the trends shown in Figure 2 in a multivariate framework, particularly for the urban birth cohorts. Historically in Indonesia, children from large families living in urban areas obtained more schooling than those from smaller families, holding other family characteristics constant. In recent urban cohorts, this positive association between large families and children's schooling has disappeared and a negative association has emerged in the most recent cohort. In rural areas, there is no significant association between family size and educational attainment in any cohort once other factors are controlled.

The evidence so far suggests that by the 1970s and 1980s Indonesia had entered a stage of development, at least in urban areas, which favored smaller families. The results in Table 4 extend the analysis to the most recent cohort of children available in the 1997 IFLS, those ages 16 to 19 in 1997. Because many members of this cohort have not completed their schooling, these analyses rely on continuation beyond primary school rather than completed education. About three-fourths of IFLS children ages six to 19 were enrolled in school in 1997. In both rural and urban areas, at least 95 percent of children age eight to 12 were enrolled in school. After age 12, enrollment by age drops faster in rural areas, with about three-fourths of urban children still enrolled at age 16 compared to about half of rural children age 16.[16] While Indonesia has achieved nearly universal primary school enrollment, enrollment proportions drop rapidly at ages associated with secondary school and differ considerably between urban and rural areas.

Table 4 shows results from binary probit models that examine the likelihood of completing junior secondary schooling (completing grade nine) and entering senior secondary (completing at least grade 10) by urban-rural residence. The models also include a measure for having a child under age 6 present in the household. This measure is intended to capture whether having a very young sibling interferes with school investment for older siblings. Models are restricted to children ages 16 to 19 at the time of the survey so that only those who are old enough to have made these transitions are considered.[17] Because families are smaller in this cohort, the categorical version of family size is capped at five or more children (rather than at seven or more as in Table 3) for this birth cohort.

In the 1978-1981 urban birth cohort, larger families are associated with a lower likelihood of completing junior secondary schooling (model 4.1). Holding other covariates at the sample means, each additional living sibling is associated with a 0.02 decrease in the probability of entering junior secondary. Model 4.2 replicates this finding with a categorical version of number of siblings. Having two or fewer siblings (versus five or more) reduces the likelihood of transitioning from primary to junior secondary school. In contrast, family size and the likelihood of completing junior secondary are not significantly associated for children living in rural areas

[16] Only about 4% of children age seven to 19 worked while enrolled in school with the proportions slightly higher in rural than urban areas (4.7% versus 2.3%).
[17] These samples and transitions are not conditioned on completing at least 8th or 9th grade. The results do not change if the sample is restricted to only those who made the previous transitions.

at age 12 (models 4.3 and 4.4). The results are similar for the likelihood of entering senior secondary. [18] Holding other family and individual characteristics constant, the association of family size and entering senior secondary school is negative and significant in urban areas (models 4.5 and 4.6) but not significant in rural areas (models 4.7 and 4.8). Holding other covariates at the sample means, each additional living sibling reduces the predicted probability of entering senior secondary by about 0.03 in urban areas. Having a young child in the household is not associated with the likelihood of making either school transition. This result does not differ by the sex of the index child (interaction coefficients not shown).

Differences by sex and birth order are consistent with the patterns shown in Table 3. For urban cohorts, differences by birth order have disappeared and an advantage for girls has emerged. This advantage likely reflects girls' lower rates of school failure and grade retention rather than any emerging sex preference. Indeed, by senior secondary, this female advantage has become only marginally significant in urban areas. For rural areas, disadvantages by sex and birth order have largely disappeared, reflecting growing school opportunities for all children and particularly for girls in these areas (albeit only at these lower levels of schooling). The one exception to this pattern is that older rural girls are still marginally less likely to enter senior secondary school relative to their younger sisters.

Testing for the Potential Endogeneity of Fertility

The analyses above provide ample evidence for a correlation between family size and children's schooling but do not address the concern that these processes may be jointly determined. To assess the sensitivity of the results to assumptions about the exogeneity of fertility, I use

[18] The sibling categories in models 4.4 and 4.8 are not jointly significantly different from zero.

instrumental variables analysis (IV) as an alternative estimation strategy (Wooldridge 2002). One way to instrument completed fertility is to find a physiological (rather than behavioral) trait that is exogenous and heterogeneously distributed among individuals, which affects completed fertility but is independent of preferences for family size. As discussed above, twinning has been the most widely utilized instrument of this type in previous work. A few studies, however, have also used miscarriages to instrument fertility (Hotz, McElroy and Sanders 2005, Hotz, Mullin and Sanders 1997). Once maternal age and behaviors such as smoking and drinking are controlled, miscarriages represent an exogenous and random shock to a woman's fertility (Kline, Stein, and Susser 1989; Porter and Hook 1979; Leridon 1977). Miscarriages are involuntary, spontaneous fetal deaths that reduce the number of conceptions that result in live births, and therefore, represent lost fertility exposure time. (Casterline 1989; Bongaarts and Potter 1983; Leridon 1977). As described by Casterline (1989, pp. 81-82), "Accepted estimates of overall spontaneous loss rates of 20 percent of recognized pregnancies (Bongaarts and Potter, 1983) thus imply two and one-half months of time lost involuntarily for every live birth, on average. In most societies this represents a reduction of fertility of 5-10 percent from levels expected in the absence of pregnancy loss." Net of age and behavior, women who experience many miscarriages may fail to achieve their desired family size despite their preferences. Factors such as socioeconomic and nutrition status are not associated with miscarriage rates (Kline, Stein, and Susser 1989). [19]

Using miscarriages to instrument fertility is a different approach than using twins. Miscarriages represent an exogenous constraint on fertility while twins represent an exogenous

[19] About three-fourths of the women in the IFLS sample report no miscarriages. At the high end, about seven percent of the women report more than 20 percent of their total number of pregnancies end in miscarriage. A possible confounding factor is if women report abortions, which are legally prohibited in Indonesia except to save a woman's life, as miscarriages. Recall error is another important problem that I discuss in more detail below.

"extra" child. Women who experience many miscarriages may end up with lower fertility than they had hoped for while women with twins may end up with more children than desired. Ideally, one would test the robustness of the results to both these approaches. However, this is untenable in most sample surveys because twinning is a rare event. Miscarriages in contrast are a more common event, and recorded in many demographic surveys. When appropriate, miscarriages offer an alternative approach to measuring exogenous shocks to fertility.

Using the *number* of miscarriages a woman experiences as an instrument for completed fertility does not sufficiently control for underlying parental preferences about family size, which may be correlated with preferences about children's schooling (see Rosenzweig and Wolpin 1980a for a related discussion regarding twins). Those who desire very large families may have many pregnancies, and therefore experience more miscarriages because of their increased exposure to the risk of miscarriage (Kline, Stein, Susser 1989). To address this, I adjust number of miscarriages by regressing it on total number of pregnancies, pregnancies squared and pregnancies cubed. I then use the residual from this regression along with the residual squared and cubed as the instruments. This adjusts the count of miscarriages for underlying family size preferences in a flexible and nonlinear way.[20]

Two important limitations to using miscarriages to instrument fertility include women's knowledge of having had a miscarriage in the first place and recall error in the reporting of miscarriages. Women's knowledge and recall about miscarriages may vary by their level of schooling and urban-rural residence, and both variables are controlled in all models. Women

---

[20] There is no evidence that the *risk* of miscarriage increases with pregnancy number (Kline, Stein, and Susser 1989, Leridon 1976). Indeed in this sample, the average proportion of pregnancies ending in miscarriage is similar across different family sizes. Also, the number of living siblings is correlated but not collinear with either miscarriages or the proportion of pregnancies that end in miscarriage (ρ equals 0.13 and -0.02 respectively). In this sample, women's education is weakly positively correlated with miscarriages (ρ equals 0.14), and this variable is controlled in all models. Finally, although using whether a woman's first birth resulted in a miscarriage would be an even better IV strategy (Rosenzweig and Wolpin 2000) the IFLS sample is too small to support this approach, especially for the older cohorts.

may also forget to report miscarriages that occurred long ago as these events lose salience in their mental record of their reproductive history (Casterline 1989; Panis and Lillard 1994). In the current analysis, the issue of recall error seems particularly problematic for the older cohorts. I exclude these cohorts from the IV analyses and focus only on testing the robustness of the negative association between family size and children's schooling in the most recent cohorts. The endogeneity of fertility is also more likely to be relevant for more recent cohorts who experienced the widespread distribution of modern contraception after the 1970s.

The analyses in the previous section use ordered probit models because this specification offers a flexible and less parametric way to capture educational attainment. Measured in years of schooling, educational attainment is inherently discrete and ordered, and well-suited to this statistical modeling strategy. But combining IV with these types of nonlinear models is quite complex and generally not well identified. The best developed IV methods use two-stage least squares (2SLS) or binary probit models. Although in general ordered probit and OLS models do not necessarily produce the same results, for the case of educational attainment in this IFLS sample, the substantive results are the same for both methods. Appendix Table A.1 shows the OLS results corresponding to the ordered probit analyses described in Table 3. Therefore, I estimate the IV models using 2SLS and IV probit models. Instruments that are too weakly correlated with the endogenous regressor pose another methodological problem for IV (Hall, Rudebusch, and Wilcox 1996; Shea 1997; Staiger and Stock 1997; Bound, Jaeger and Baker 1995), but standard metrics show that the instruments used here are sufficiently strong. The IV models shown in Table 5 all have first stage F statistics larger than 12. The 2SLS models also have high values for the first stage partial R^2.

The full set of IV results is shown in Appendix Tables A.2 and A.3. Table 5 summarizes these results. For ease of comparison, I reproduce the estimates for number of siblings treated exogenously in the first row (from Table A.1 and Table 4). All models control for child's and mother's age and therefore mother's age at birth. The models do not control for smoking or drinking because these behaviors are rare among women in Indonesia. Fewer than 5% of IFLS women report having ever smoked, and controlling for this behavior does not change the results shown (models not shown). Alcohol consumption is even more uncommon (Indonesia is a predominantly Muslim country), and the IFLS does not include these measures.

Overall, the IV models show the same patterns as the models that treat fertility as exogenous but the coefficients are not estimated as precisely. For the 1968-1977 urban birth cohort, the instrumented effect of number of siblings is negative and marginally significant, suggesting that children with more siblings obtain less schooling in more recent urban cohorts. The effect of family size is also negative and marginally significant for the model predicting entry into senior secondary school. In the most recent urban birth cohort, children with more siblings are less likely to make this transition. For those cohorts in which the concern of the endogeneity of fertility and a spurious negative effect of family size on children's schooling is most likely, the IV results support the results of the analyses that treat fertility as exogenously determined. The IV results lack power for the two recent rural cohorts. Here, none of the models yield a significant relationship between family size and children's schooling although the patterns are suggestive and similar to those in the earlier models.

DISCUSSION AND SUMMARY

In Indonesia, the relationship between family size and children's schooling is not uniformly positive or negative. Rather, there are important differences by cohort and urban-rural residence. For rural cohorts, there is little evidence of a statistically significant association between family size and children's schooling, a finding that is consistent in all birth cohorts. In contrast, in urban areas, this association evolved from positive to neutral to negative over a span of thirty years. For the 1948-1957 urban cohort, larger families were associated with more schooling. This positive relationship diminished over time and a negative relationship developed in the 1968-1977 cohort and remains for the most recent birth cohort (1978-1981). Although the estimated effects are smaller and not statistically significant, this evolving pattern is also apparent across the rural cohorts, suggesting that as rural areas continue to develop these associations may change.

In both rural and urban areas, differences in schooling by sex and birth order also changed over time. In the oldest urban cohort and in all rural cohorts, girls obtained less schooling than boys. This disadvantage, however, has disappeared over time in urban areas and diminished over time in rural ones. Similarly, disadvantages associated with being the oldest child also diminished and disappeared across cohorts. These results highlight the Indonesian government's success in expanding schooling opportunities, especially for girls. The speed with which sex and birth order differences in schooling have narrowed reflects Indonesia's relatively more egalitarian family organization, at least relative to other East Asian countries.

Standard analyses of the relationship between family size and children's schooling make strong assumptions about the relationship between family size and parents' preferences for children's schooling. Although IV analysis requires its own set of strong assumptions, the results presented here suggest that, at least in the most recent urban cohorts, the negative relationship between family size and children's schooling is not sensitive to assumptions about the exogeneity of fertility. Ideally, one would confirm this result using other instruments such as twins, but household surveys such as the IFLS do not have large enough samples to support this type of analysis, especially for a cohort design like the one used here. Recent studies using twin data from developed countries find no relationship between family size and children's schooling once these variables are modeled jointly. Rosenzweig and Wolpin (1980a), in contrast, found a negative relationship using data from India and Qian (2006) finds mixed results. Few other studies have explored this question in a developing setting.

This pattern of associations changing from positive to neutral to negative in urban areas suggests that changes in socioeconomic and demographic conditions brought by development can alter the ways in which families benefit or impinge on children. Differences between more socioeconomically developed urban areas and less developed rural ones provide additional support for this explanation. The contrast between urban and rural areas, for example, serves as another measure of differences in transportation and communication infrastructure, school opportunities, and labor market opportunities.

Although testing specific mechanisms explaining this changing relationship is beyond the scope of the current study (and the data are limited in this regard), rapid development in Indonesia has changed numerous aspects of the economy, family organization, and educational opportunities, offering some candidate explanations for the patterns shown above. For the oldest urban cohort, it seems likely that widely shared preferences for larger families, combined with better family resources such as education and occupation and much better accessibility of schools offered those in urban areas the ability both to have more children and to provide those children with some schooling. Moreover, in this older cohort, average attainments were still relatively low such that the cost of schooling itself was low. It also seems likely that schooling offered better payoffs for urban residents than the largely agrarian based residents of rural areas.

Between 1970 and 1980, primary schooling became nearly universal in urban areas and grew by nearly 30 percentage points in rural areas (Oey-Gardiner 1997). Given this infusion of schooling at the primary level it seems likely that families' educational aspirations grew. But the expansion in educational infrastructure was substantially slower at the secondary level. The per-student cost of secondary schooling remains about three times that of primary schooling while the per-student cost of tertiary schooling is about 13 times as high (UNESCO 1999). Parents, rather than other family members or broader kin, pay the family's portion of these costs, which can be sizeable especially for poorer families. Even before the economic crisis of 1997-1998 (an experience that affected subsequent waves of the IFLS but not the samples used here) Indonesia could not finance the expansion in post-primary schooling the nation hoped to achieve (World Bank 1998). As the education distribution continues to shift upwards, parents with more children may be unable to afford the cost of post-primary schooling, at least for some of their children. At the same time, changing norms about schooling girls increases the number of children families have to educate relative to older cohorts.

Differences between urban and rural areas might be explained by several factors. Before the 1970s, parents in rural areas had limited options for schooling children. As Lloyd (1994, p. 9) argues, some threshold level of development is necessary before family size can even influence children's schooling. In an area with no schools, having many or few siblings is irrelevant for children's schooling. Also, the local agrarian economy meant that the returns to schooling were likely low relative to those living in urban areas. In more recent years primary schooling has become widely available, free, and compulsory. Although most children now complete primary school, the education distribution in rural areas has yet to shift up to the more expensive secondary level. It remains an open question whether higher schooling costs, changing labor market conditions, change or stability in family organization, and growing migration rates will produce similar or different patterns in rural areas in the future.

These results and candidate explanations are not meant to argue that socioeconomic development inevitably shapes family and school patterns that are negatively correlated, whether causally or because those who value education highly will have fewer children. Indeed, one of the most important lessons of this literature has been that patterns differ greatly by context. These results instead speak to the diversity of possible relationships and to the interplay of macro socioeconomic factors, government policies, norms and preferences, and familial and local mechanisms in producing patterns of stratification by place and family structure.

Appendix A. Robustness of Results to Including Household Wealth

The IFLS has limited information on household wealth during childhood for adult respondents. All candidate measures (household floor and water supply at age 12, hospital birth, mother's wedding dowry, father's occupation when respondent was a child) are collected for only parts of the sample, are missing for a large portion of the sample, or are poorly measured. The survey does collect extensive information on assets, consumption and expenditures at each wave, but these measures come after schooling is completed, at least for the vast majority of the respondents that make up the cohorts used in the analysis. Nonetheless, the possibility that omitting wealth might bias the results is compelling enough to warrant checking the robustness of results in whatever ways possible.

The expenditure data are the most complete and reliable wealth data available in the survey and offer the best way to measure long-term household wealth. Expenditure data are often less volatile than income measures, more reliable than asset measures, and a fairly good representation of wealth across the life course. I address the problem of temporal ordering of household wealth to education as follows. For respondents in the youngest birth cohort, who were ages 16 to 19 in 1997, I include 1993 log per capita expenditures of their mother's household in all models. These children were age 12 to 15 in 1993 so this is a fairly good measure of family wealth when these respondents were school age. For the remaining cohorts, I exclude individuals who coreside with their mothers and are not enrolled in school and include 1993 log per capita expenditures for their mothers' household in all models. For most cohorts, this reduces the sample size by no more than 18 percent. For the 1968 to 1977 cohort, however, this reduces the rural sample by about 27 percent and the urban sample by 35 percent. This subsection of the sample is far from random: the coresiding individuals have fewer siblings on average and come from families with lower average expenditures in 1993. Nonetheless this approach offers a way to test the robustness of the results to controls for family wealth.

I present a full set of OLS, binary probit and IV results in Appendix Tables A.4, A.5, and A.6. The results are remarkably robust to the inclusion of household wealth, despite the reduced sample size. As expected, in all cohorts, 1993 log per capita expenditures of the mother's household has a strong, statistically significant positive association with children's schooling. But the association between family size and children's schooling is largely robust to controlling for this measure of wealth. For the youngest cohort, where the data are the most complete and appropriate in terms of time ordering, the results are the same. Including wealth does not change the associations between family size and children's schooling in either the exogenous or instrumented models. The results for the other cohorts are also quite consistent. The only meaningful difference is in the estimate for the 1968-1977 urban cohort where the exogenous model that omits wealth show a significant negative coefficient but the model that includes wealth shows a smaller negative coefficient that is not significant. The instrumented model that includes wealth for this cohort produces a result that is very close to the instrumented model without wealth included (p-value was marginally significant before and is now significant at $p<.109$). Given that this young urban cohort has the largest portion of young adults coresiding with their parents (and these individuals are excluded in the models that include wealth) the difference between the two sets of results is not surprising. In fact, the consistency of the instrumented result for this cohort is remarkable. Moreover, the overall pattern of results—the association between family size and children's schooling going from positive to neutral to negative across cohorts—remains unchanged with or without a measure of family wealth.

REFERENCES

Angrist, Joshua, Victor Lavy, and Analia Schlosser 2006. "Multiple Experiments for the Causal Link between the Quantity and Quality of Children." Massachusetts Institute of Technology, Department of Economics Working Paper 06-26: http://ssrn.com/abstract=931948.

Angrist, Joshua D. and William N. Evans. 1998. "Children and Their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size." *The American Economic Review* 88:450-477.

Anh, Truong Si, John Knodel, David Lam and Jed Friedman. 1998. "Family Size and Children's Education in Vietnam." *Demography* 35:57-50.

Axinn, William G. 1993. "The Effects of Children's Schooling on Fertility Limitation." *Population Studies* 47:481-493.

Becker, Gary S. and Nigel Tomes. 1976. "Child Endowments and the Quantity and Quality of Children." *The Journal of Political Economy*, 84(4, Part II): S143-S162.

Becker, Gary S. 1991. *A Treatise on the Family*. (2nd Ed.) Cambridge: Harvard University Press.

Bongaarts, John and Robert G. Potter. 1983. *Fertility, Biology, and Behavior: An Analysis of the Proximate Determinants*. New York: Academy Press.

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

DeGraff, Deborah S. and Richard E. Bilsborrow. 2003. "Children's School Enrollment and Time at Work in the Philippines." *The Journal of Developing Areas* 37:127-158.

Blake, Judith. 1989. *Family Size and Achievement*. Berkeley: University of California Press.

Black, Sandra E., Paul Devereux, and Kjell Salvanes. 2005. "The More the Merrier? The Effect of Family Size and Birth Order on Children's Education." *Quarterly Journal of Economics* 120:669-700.

Bound, John, David A. Jaeger and Regina Baker. 1995. "Problems with Instrumental Variables Estimation When the Correlation Between the Instruments and the Endogenous Explanatory Variables Is Weak." *Journal of the American Statistical Association* 90:443-450.

Buchmann, Claudia and Emily Hannum. 2001. "Education and Stratification in Developing Countries: A Review of Theories and Research." *Annual Review of Sociology* 27:77-102.

Caldwell, John C., P. H. Reddy and Pat Caldwell. 1985. "Educational Transition in Rural South India." *Population and Development Review*, 11:29-51.

Casterline, John B. 1989. “Collecting Data on Pregnancy Loss: A Review of Evidence from the World Fertility Survey.” *Studies in Family Planning* 20:81-95.

Chernichovsky, Dov. 1985. “Socioeconomic and Demographic Aspects of School Enrollment and Attendance in Rural Botswana.” *Economic Development and Cultural Change* 33:319-332.

Chernichovsky, Dov and Oey Astra Meesook. 1985. “School Enrollment in Indonesia.” World Bank Staff Working Paper Number 746, The World Bank, Washington DC.

Cobbe, James and Boediono. 1993. “Education, Demographics, the Labor Market, and Development: Indonesia in the Process of Transition?” *Journal of Asian and African Studies* 28:1-29.

de Graff, Paul M. and Johannes J. Huinink. 1992. “Trends in Measured and Unmeasured Effects of Family Background on Educational Attainment and Occupational Status in the Federal Republic of Germany. *Social Science Research* 21:84-112.

Desai, Sonalde. 1995. “When Are Children from Large Families Disadvantaged? Evidence from Cross-National Studies.” *Population Studies* 49:195-210.

Douglas B. Downey, Brian Powell, Lala Carr Steelman, and Shana Pribesh. 1999. “Much Ado about Siblings: Change Models, Sibship Size, and Intellectual Development: Comment on Guo and VanWey.” *American Sociological Review* 64:193-198.

Duflo, Esther. 2001. “Schooling and the Labor Market Consequences of School Construction in Indonesia: Evidence from an Unusual Policy Experiment.” *The American Economic Review* 90:795-813.

Featherman, David L., and Robert M. Hauser 1978. *Opportunity and Change*. New York: Academic.

Filmer, Deon, and Lant Pritchett. 1999. “The Effect of Household Wealth on Educational Attainment: Evidence from 35 Countries.” *Population and Development Review* 25:85-120.

Firman, Tommy. 1997. “Patterns and Trends of Urbanization: A Reflection of Regional Disparity.” Pp. 101-117 in *Indonesia Assessment: Population and Human Resources*, edited by Gavin W. Jones and Terence H. Hull. Singapore: Institute of Southeast Asian Studies.

Frankenberg, Elizabeth, and Duncan Thomas. 2000. “The Indonesian Family Life Survey (IFLS): Study Design and Results from Waves 1 and 2. RAND DRU-2238/1-NIA/NICHD.

Gertler, Paul J. and John W. Molyneaux. 1994. “How Economic Development and Family Planning Programs Combined to Reduce Indonesian Fertility.” *Demography* 31:33-63.

Gomes, Melba. 1984. "Family Size and Educational Attainment in Kenya." *Population and Development Review* 10:647-660.

Guo, Guang and Leah K VanWhey. 1999a. "Sibship Size and Intellectual Development: Is the Relationship Causal?" *American Sociological Review* 64:169-187.

____. 1999b. "The Effects of Closely Spaced and Widely Spaced Sibship Size on Intellectual Development: Reply to Phillips and to Downey et al." *American Sociological Review* 64:199-206.

Hall, Alastair R., Glenn D. Rudebusch, and David W. Wilcox. 1996. "Judging Instrument Relevance in Instrumental Variables Estimation." *International Economic Review* 37:283-98.

Hauser, Robert M., and William H. Sewell. 1985. "Birth Order and Educational Attainment in Full Sibships." *American Educational Research Journal* 22:1-23.

Hauser, Robert M. and David L. Featherman. 1976. "Equality of Schooling: Trends and Prospects" *Sociology of Education* 49:99-120.

Hermalin, Albert, Judith A. Seltzer, and Ching-Hsiang Lin. 1982. "Transitions in the Effect of Family Size on Female Educational Attainment: The Case of Taiwan. *Comparative Education Review* 26:254-270.

Hotz, V. Joseph, S. McElroy, and Seth Sanders. 2005. "Teenage Childbearing and Its Life Cycle Consequences: Exploiting a Very Natural Experiment." *Journal of Human Resources*, 40:683-715.

Hotz, V. Joseph, Charles H. Mullin and Seth G. Sanders. 1997. "Bounding Causal Effects Using Data From a Contaminated Natural Experiment: Analyzing the Effects of Teenage Childbearing." *The Review of Economic Studies* 64 (No. 4, Special Issue):575-603.

Hugo, Graeme J., Terence H. Hull, Valerie J. Hull, and Gavin W. Jones. 1987. *The Demographic Dimension in Indonesian Development.* Singapore: Oxford University Press.

Hull, Terence H. 1987. "Fertility Decline in Indonesia: An Institutionalist Interpretation." *International Family Planning Perspectives* 12:90-95.

Hull, Terence H. and Valerie J. Hull. 1977. "The Relation of Economic Class and Fertility: An Analysis of Some Indonesian Data." *Population Studies* 31:43-57.

Indonesia Central Bureau of Statistics (BPS), Population Statistics. Available at www.bps.go.id/sector/population.

Iskandar, Meiwita B., 1997. "Health and Mortality." Pp. 205-231 in *Indonesia Assessment: Population and Human Resources*, edited by Gavin W. Jones and Terence H. Hull. Singapore: Institute of Southeast Asian Studies.

Jacobsen, Joyce P., James Wishart Pearce III, and Joshua L. Rosenbloom. 1999. “The Effects of Childbearing on Married Women’s Labor Supply and Earnings: Using Twin Births as a Natural Experiment.” *The Journal of Human Resources* 34:449-474.

Jones, Gavin W. 1990. “Fertility Transitions Among Malay Populations of Southeast Asia: Puzzles of Interpretation.” *Population and Development Review* 16:507-537.

Jones, Gavin W. and Terrence H. Hull, eds. 1997. *Indonesia Assessment: Population and Human Resources*, Singapore: Institute of Southeast Asian Studies.

King, Elizabeth. 1987. “The Effect of Family Size on Family Welfare: What Do We Know?” Pp. 373-411 in *Population Growth and Economic Development: Issues and Evidence*, edited by D. Gale Johnson and Ronald D. Lee. Madison, WI: The University of Wisconsin Press.

Klein, Jennie, Zena Stein and Mervyn Susser. 1989. *Conception to Birth: Epidemiology of Prenatal Development*. New York: Oxford University Press.

Knodel, John, Napaporn Havanon and Werasit Sittitrai. 1990. “Family Size and the Education of Children in the Context of Rapid Fertility Decline.” *Population and Development Review* 16:31-62.

Kuo, Hsiang-Hui Daphne and Robert M. Hauser. 1995. “Trends in Family Effects on the Education of Black and White Brothers.” Sociology of Education 68:136-160.

Leridon, Henri. 1977. *Human Fertility: The Basic Concepts*. Translated by Judith F. Helzner. Chicago: The University of Chicago Press.

Lloyd, Cynthia. 1994. “Investing in the Next Generation: The Implications of High Fertility at the Level of the Family.” Pp. 181-202 in *Population and Development: Old Debates and New Conclusions*, edited by Robert Cassen. New Brunswick, NJ: Transaction Publishers.

Lloyd, Cynthia B. and Ann K. Blanc. 1996. “Children’s Schooling in sub-Saharan Africa: The Role of Fathers, Mothers, and Others.” Population and Development Review 22:265-298.

Lloyd, Cynthia B. and Anastasia J. Gage-Brandon. 1994. “High Fertility and Children’s Schooling in Ghana: Sex Differences in Parental Contributions and Educational Outcomes.” *Population Studies* 48:293-306.

Mueller, Eva. 1984. “Income, Aspirations, and Fertility in Rural Areas of Less Developed Countries.” Pp. 121-150 in *Rural Development and Human Fertility*, edited by Wayne A. Schutjer and C. Shannon Stokes. New York: Macmillan Publishing Company.

Oey-Gardiner, Mayling. 1997. "Educational Developments, Achievements and Challenges." Pp. 135-166 in *Indonesia Assessment: Population and Human Resources*, edited by Gavin W. Jones and Terence H. Hull. Singapore: Institute of Southeast Asian Studies.

Panis, Constatijn W. A. and Lee A. Lillard. 1994. "Health Inputs and Child Mortality: Malaysia." *Journal of Health Economics* 13:455-489.

Parish, William L. and Robert J. Willis. 1993. "Daughters, Education, and Family Budgets Taiwan Experiences." *The Journal of Human Resources* 28:863-898.

Phillips, Meredith. 1999. "Sibship Size and Academic Achievement: What We Now Know and What We Still Need to Know: Comment on Guo and VanWey." *American Sociological Review* 64:188-192.

Pong, Suet-Ling. 1997. "Sibship Size and Educational Attainment in Peninsular Malaysia: Do Policies Matter?" *Sociological Perspectives* 40:227-242.

Porter, Ian H. and Ernest B. Hook. 1980. *Human Embryonic and Fetal Death*. New York: Academic Press, Inc.

Preston, Samuel H. 1976. "Family Sizes of Children and Family Sizes of Women." *Demography* 13:105-114.

Psacharopoulos, George and Ana Maria Arriagada. 1989. "The Determinants of Early Age Human Capital Formation: Evidence from Brazil." *Economic Development and Cultural Change* 37: 683-708.

Qian, Nancy. 2006. "Quantity-Quality: The Positive Effect of Family Size on School Enrollment in China." Working paper available at: www.econ.brown.edu/fac/Nancy_Qian/Papers/qq_pap&tab-2006_01.pdf

Retherford, Robert D. and William H. Sewell. 1991. "Birth Order and Intelligence: Further Tests of the Confluence Model." *American Sociological Revie*w 56:141-158.

Rosenzweig, Mark R. and Kenneth I. Wolpin. 1980a "Testing the Quantity-Quality Fertility Model: The Use of Twins as a Natural Experiment." *Econometrica* 48:227-240.

____. 1980b. "Life-Cycle Labor Supply and Fertility: Causal Inferences from Household Models" *The Journal of Political Economy* 88:328-348.

____. 2000. "Natural 'Natural Experiments' in Economics." Journal of Economic Literature 38:827-874.

Shavit, Yossi and Jennifer L. Pierce. 1991. "Sibship Size and Educational Attainment in Nuclear and Extended Families: Arabs and Jews in Israel." *American Sociological Review* 56:321-330.

Shea, John. 1997. “Instrument Relevance in Multivariate Linear Models: A Simple Measure.” *The Review of Economics and Statistics* 79:348-352.

Staiger, Douglas and James H. Stock. 1997. “Instrumental Variables Regression with Weak Instruments” *Econometrica* 65:557-586.

Steelman Lala Carr, Brian Powell, Regina Werum, and Scott Carter. 2002. “Reconsidering the Effects of Sibling Configuration: Recent Advances and Challenges.” *Annual Review of Sociology* 28:243-269.

Sudha, S. 1997. “Family Size, Sex Composition and Children’s Education: Ethnic Differences Over Development in Peninsular Malaysia.” *Population Studies* 51:139-151.

Thornton, Arland. 2001. “The Developmental Paradigm, Reading History Sideways, and Family Change.” *Demography* 38:449-46.

UNESCO Institute for Statistics, Organization for Economic Co-operation and Development. 1999. “Financing Education – Investments and Returns: Analysis of the World Education Indicators 2002.” Paris: UNESCO Publishing.

UNICEF. 2000. “The State of the World’s Children.” Available at www.unicef.org/sowc00.

UNICEF and the Government of Indonesia, 1988. “Situation Analysis of Children and Women in Indonesia.” Jakarta: UNICEF.

Wooldridge, Jeffrey M. 2002. “Econometric Analysis of Cross Section and Panel Data.” Cambridge: MIT Press.

Walton, John. “Industrial Development and Trade.” Pp.43-65 in *Understanding Indonesia*, edited by Leslie Palmier, Aldershot, Hampshire, England and Brookfield, Vermont: Gower Publishing Company.

Wolfe Barbara L. and Jere R. Behrman. 1984. “Who is Schooled in Developing Countries? The Roles of Income, Parental Schooling, Sex, Residence, and Family Size.” *Economics of Education Review* 3:231-245.

World Bank. 1998. “Education in Indonesia: From Crisis to Recovery.” Report No. 18651-IND, Washington D.C.

Zajonc, Robert B. and Gregory B. Marcus. 1975. “Birth Order and Intellectual Development.” *Psychological Review* 82:74-88.

Table 1. Sample Means and Proportions (standard errors in parentheses), IFLS 1997

	Born 1948-1957	Born 1958-1967	Born 1968-1977	Born 1978-1981	All Sample Children
No. living siblings	4.6	4.6	4.4	3.8	4.3
	(0.07)	(0.04)	(0.03)	(0.04)	(0.02)
No. sibs who have died	1.4	1.2	1.0	0.68	1.0
	(0.05)	(0.03)	(0.02)	(0.03)	(0.01)
Avg. No. Children Ever Born (row 1 + row 2 + 1)	7.0	6.8	6.4	5.48	6.3
Family Size (living children)					
1	0.03	0.03	0.02	0.02	0.02
2	0.06	0.06	0.05	0.10	0.06
3	0.11	0.11	0.13	0.18	0.13
4	0.14	0.14	0.18	0.20	0.17
5	0.14	0.16	0.18	0.17	0.17
6	0.17	0.17	0.16	0.13	0.16
7	0.11	0.13	0.12	0.09	0.11
8	0.11	0.10	0.09	0.06	0.09
9	0.07	0.07	0.05	0.04	0.06
10 plus	0.05	0.04	0.03	0.02	0.03
Child's education (years)	6.5	7.1	8.5	8.3	7.8
	(0.13)	(0.09)	(0.07)	(0.06)	(0.04)
Child's education					
None	0.14	0.11	0.05	0.02	0.07
Elementary	0.52	0.49	0.41	0.33	0.43
Jr. Secondary	0.11	0.11	0.17	0.28	0.17
Sr. Secondary	0.15	0.19	0.27	0.34	0.25
College	0.07	0.09	0.10	0.02	0.08
Gender gap in educ. (M - F)	1.8	1.1	0.8	0.2	.87
Mother's education (years)	1.7	2.3	3.3	4.0	3.0
	(0.08)	(0.06)	(0.05)	(0.08)	(0.03)
Father's education (years)	3.6	4.0	4.9	5.5	4.6
	(0.11)	(0.07)	(0.06)	(0.09)	(0.04)
Age of mother	67.1	60.6	52.5	43.8	54.7
	(0.22)	(0.15)	(0.12)	(0.17)	(0.11)
Age of child	43.4	33.9	24.5	17.5	28.0
	(0.08)	(0.05)	(0.04)	(0.02)	(0.09)
Rural residence at age 12	0.73	0.73	0.72	0.70	0.72
Oldest boy in family	0.29	0.25	0.20	0.23	0.23
Oldest girl in family	0.31	0.26	0.22	0.24	0.25
Child < age 6 in sibship	0.00	0.00	0.05	0.21	0.06
Number of obs.	1,612	3,632	5,329	2,770	13,343

Notes: Descriptive data are weighted using sample probability weights. Because schooling and family size are not completed for those in the 1978-81 cohort these means are not directly comparable to the other three cohorts.

Table 2. Children's Education Distribution (Years of School Completed) by Cohort, IFLS 1997 (N=10,573)

	Percentiles						
	10	25	50	75	90	Mean	N
Urban 1948-57	3	6	9	12	15	9.3	559
Urban 1958-67	3	6	12	12	17	10.0	1203
Urban 1968-77	6	9	12	12	15	10.9	1911
Rural 1948-57	0	2	6	6	12	5.4	1053
Rural 1958-67	0	3	6	9	12	6.0	2429
Rural 1968-77	3	6	6	12	12	7.6	3418

Notes: Table includes only cohorts with completed education. Descriptive data are weighted using sample probability weights.

Table 3. Ordered Probit Models Predicting Completed Education by Cohort and Rural Residence, IFLS 1997, (Total N=10,573, robust standard errors in parentheses)

Cohort	Urban 1948-57		Urban 1958-67		Urban 1968-77		Rural 1948-57		Rural 1958-67		Rural 1968-77	
Model	3.1	3.2	3.3	3.4	3.5	3.6	3.7	3.8	3.9	3.10	3.11	3.12
No. of sibs (linear, 0-14)	0.064* (0.025)		0.010 (0.020)		-0.052* (0.017)		0.025 (0.020)		0.005 (0.015)		-0.012 (0.013)	
No. of sibs (categories):												
0-2		-0.363† (0.188)		0.016 (0.137)		0.323* (0.121)		-0.150 (0.136)		-0.037 (0.103)		0.105 (0.089)
3-4		-0.441* (0.193)		-0.184 (0.116)		0.198† (0.113)		-0.223† (0.127)		-0.024 (0.096)		0.010 (0.082)
5-6		-0.408* (0.156)		-0.068 (0.119)		0.188† (0.110)		-0.184 (0.136)		-0.097 (0.095)		-0.031 (0.085)
7 plus (reference)												
Mother's education	0.071* (0.024)	0.073* (0.023)	0.106* (0.012)	0.107* (0.012)	0.129* (0.012)	0.129* (0.012)	0.136* (0.023)	0.139* (0.024)	0.116* (0.014)	0.117* (0.014)	0.116* (0.010)	0.116* (0.010)
Father's education	0.077* (0.018)	0.080* (0.018)	0.109* (0.012)	0.109* (0.012)	0.090* (0.010)	0.091* (0.010)	0.126* (0.017)	0.124* (0.017)	0.100* (0.010)	0.100* (0.010)	0.106* (0.009)	0.107* (0.009)
Female	-0.478* (0.160)	-0.489* (0.161)	-0.113 (0.092)	-0.125 (0.092)	-0.083 (0.071)	-0.087 (0.071)	-0.533* (0.126)	-0.538* (0.125)	-0.385* (0.071)	-0.383* (0.072)	-0.168* (0.053)	-0.169* (0.053)
Oldest boy	0.041 (0.148)	0.012 (0.150)	-0.004 (0.098)	-0.011 (0.099)	-0.113 (0.077)	-0.098 (0.076)	0.128 (0.101)	0.112 (0.101)	-0.139* (0.066)	-0.145* (0.067)	-0.029 (0.057)	-0.042 (0.057)
Oldest girl	0.058 (0.126)	0.022 (0.131)	-0.261* (0.094)	-0.261* (0.091)	-0.140† (0.078)	-0.120 (0.078)	0.042 (0.105)	0.034 (0.105)	-0.079 (0.066)	-0.087 (0.066)	-0.178* (0.051)	-0.192* (0.051)
Log likelihood	-701	-699	-1415	-1412	-2145	-2146	-1132	-1131	-2825	-2823	-3970	-3968
N	559		1203		1911		1053		2429		3418	

Notes: Cut-point parameters not shown. All models control for child's age, mother's age, and province in which household is located (coefficients not shown).
*p < .05; † p<.10

Table 4. Binary Probit Models Predicting School Enrollment and Continuation Beyond Primary School for the 1978-81 Cohort (age 16-19 in 1997), IFLS

	Completed Jr. Sec.				Entered Sr. Sec.			
	Urban		Rural		Urban		Rural	
Model	4.1	4.2	4.3	4.4	4.5	4.6	4.7	4.8
No. of siblings (linear)	-0.087*		-0.017		-0.089*		-0.026	
	(0.032)		(0.023)		(0.030)		(0.026)	
No. of siblings:								
0-2		0.457*		0.195†		0.315*		0.212†
		(0.166)		(0.114)		(0.143)		(0.122)
3-4		0.196		0.132		0.198		0.132
		(0.127)		(0.093)		(0.122)		(0.098)
5 plus (reference)								
Mother's education	0.135*	0.135*	0.120*	0.121*	0.085*	0.086*	0.115*	0.116*
	(0.020)	(0.020)	(0.015)	(0.015)	(0.017)	(0.017)	(0.015)	(0.015)
Father's education	0.081*	0.081*	0.123*	0.122*	0.087*	0.086*	0.117*	0.116*
	(0.016)	(0.016)	(0.012)	(0.012)	(0.015)	(0.015)	(0.013)	(0.013)
Female	0.344*	0.332*	-0.064	-0.062	0.213†	0.204†	-0.002	-0.001
	(0.131)	(0.131)	(0.092)	(0.092)	(0.119)	(0.119)	(0.101)	(0.101)
Oldest boy in family	0.081	0.063	0.126	0.100	0.215	0.239†	-0.022	-0.043
	(0.155)	(0.160)	(0.105)	(0.106)	(0.135)	(0.135)	(0.113)	(0.112)
Oldest girl in family	-0.066	-0.062	0.012	-0.017	0.040	0.077	-0.213†	-0.235*
	(0.156)	(0.157)	(0.103)	(0.103)	(0.136)	(0.135)	(0.115)	(0.115)
Child under 6 in household	0.173	0.162	-0.131	-0.099	0.177	0.127	-0.142	-0.118
	(0.157)	(0.162)	(0.105)	(0.102)	(0.147)	(0.145)	(0.111)	(0.109)
Constant	-0.937*	-1.332*	-1.149*	-1.377*	-1.561*	-1.869*	-1.589*	-1.825*
	(0.480)	(0.511)	(0.311)	(0.348)	(0.444)	(0.478)	(0.340)	(0.379)
Log likelihood	-441	-441	-932	-930	-548	-550	-765	-764
N	1059	1059	1711	1711	1059	1059	1711	1711

Notes: All models control for child's age (in single years), mother's age and province of residence (coefficients not shown). Sibling categories in models 4.4 and 4.8 are not jointly significantly different that zero. *p < .05; † p<.10

Table 5. Instrumental Variables Results for Models Shown in Tables 3, 4, and A.1, IFLS 1997

Cohort	Urban 1968-77	Urban 1978-81	Urban 1978-81	Rural 1968-77	Rural 1978-81	Rural 1978-81
Dependent Variable	Yrs of school completed (linear)	Complete Jr. Sec. (0/1)	Enter Sr. Sec. (0/1)	Yrs of school completed (linear)	Complete Jr. Sec. (0/1)	Enter Sr. Sec. (0/1)
No. of Siblings treated as exogenous	-0.138* (0.050)	-0.087* (0.032)	-0.089* (0.030)	-0.017 (0.040)	-0.017 (0.023)	-0.026 (0.026)
No. of Siblings instrumented	-0.260† (0.144)	-0.110 (0.085)	-0.143† (0.079)	-0.167 (0.117)	-0.070 (0.049)	-0.054 (0.055)
F (IV first stage)	18.18	--	--	24.78	--	--
Partial R^2 (IV first stage)	0.146	--	--	0.130	--	--
N	1911	1059	1059	3418	1711	1711

Notes: The exogenous regressions of years of school completed are estimated using OLS (reported in Table A1) and those of school attendance are estimated using probit models (reported in Table 4). Results are duplicated here in the first row for ease of comparison. The IV coefficients are estimated using 2SLS for years of school completed and maximum likelihood IV probit for school attendance. The instrument is described in the text. The full set of IV results is shown in Appendix Tables A.2 and A.3. *p < .05; † p<.10

Appendix Table A.1. OLS Models Predicting Completed Education by Cohort and Rural Residence, IFLS 1997, (Total N=10,573, robust standard errors in parentheses)

Cohort	Urban 1948-57		Urban 1958-67		Urban 1968-77		Rural 1948-57		Rural 1958-67		Rural 1968-77	
Model	3.1	3.2	3.3	3.4	3.5	3.6	3.7	3.8	3.9	3.10	3.11	3.12
Num. of sibs (linear, 0-14)	0.247*		0.071		-0.138*		0.070		0.034		-0.017	
	(0.088)		(0.064)		(0.050)		(0.064)		(0.050)		(0.040)	
Num. of sibs (categories):												
0-2		-1.461*		-0.039		0.920*		-0.482		-0.292		0.175
		(0.683)		(0.456)		(0.345)		(0.427)		(0.347)		(0.267)
3-4		-1.629*		-0.801*		0.584†		-0.642		-0.128		0.030
		(0.677)		(0.382)		(0.328)		(0.403)		(0.327)		(0.248)
5-6		-1.450*		-0.402		0.514		-0.588		-0.393		-0.075
		(0.545)		(0.380)		(0.325)		(0.438)		(0.325)		(0.255)
7 plus (reference)												
Mother’s education	0.245*	0.250*	0.345*	0.349*	0.331*	0.331*	0.481*	0.488*	0.432*	0.435*	0.354*	0.353*
	(0.076)	(0.076)	(0.040)	(0.040)	(0.031)	(0.031)	(0.078)	(0.079)	(0.049)	(0.050)	(0.031)	(0.031)
Father’s education	0.266*	0.275*	0.339*	0.340*	0.238*	0.239*	0.450*	0.444*	0.376*	0.376*	0.315*	0.316*
	(0.064)	(0.063)	(0.039)	(0.038)	(0.027)	(0.027)	(0.054)	(0.055)	(0.035)	(0.035)	(0.026)	(0.026)
Female	-1.786*	-1.807*	-0.733*	-0.780*	-0.183	-0.190	-1.564*	-1.571*	-1.216*	-1.206*	-0.502*	-0.503*
	(0.550)	(0.551)	(0.288)	(0.286)	(0.195)	(0.196)	(0.384)	(0.382)	(0.241)	(0.243)	(0.162)	(0.162)
Oldest boy	0.098	0.013	-0.022	-0.083	-0.381†	-0.354	0.357	0.315	-0.371	-0.385†	-0.056	-0.083
	(0.506)	(0.511)	(0.317)	(0.318)	(0.217)	(0.216)	(0.329)	(0.327)	(0.229)	(0.231)	(0.174)	(0.173)
Oldest girl	0.237	0.119	-0.561†	-0.589*	-0.483*	-0.447*	0.006	-0.031	-0.233	-0.254	-0.553*	-0.580*
	(0.429)	(0.444)	(0.307)	(0.297)	(0.216)	(0.217)	(0.303)	(0.303)	(0.217)	(0.216)	(0.154)	(0.153)
Constant	2.130	4.004	7.665*	8.685*	5.541*	4.430*	2.474	3.256	5.876*	6.241*	6.115*	6.024*
	(3.021)	(3.000)	(1.500)	(1.558)	(0.978)	(1.092)	(1.915)	(1.904)	(1.060)	(1.036)	(0.691)	(0.735)
R-squared	0.34	0.35	0.41	0.41	0.35	0.35	0.34	0.35	0.31	0.31	0.34	0.33
Number of Observations	559		1203		1911		1053		2429		3418	

Notes: All models control for child’s age, mother’s age, and province of residence (coefficients not shown). *p < .05; † p<.10

Appendix Table A.2. First Stage Regression Results for Models Shown in Table 5, IFLS 1997

	Urban 1968-77	Urban 1978-81	Rural 1968-77	Rural 1978-81
Residual	-1.303*	-1.283*	-1.275*	-1.454*
	(0.182)	(0.183)	(0.178)	(0.112)
Residual^2	1.137*	1.036*	0.986*	1.555*
	(0.189)	(0.177)	(0.283)	(0.138)
Residual^3	-0.201*	-0.165*	-0.162*	-0.277*
	(0.043)	(0.031)	(0.064)	(0.034)
Mother's Education	-0.026	-0.050*	-0.027	-0.025
	(0.021)	(0.018)	(0.021)	(0.015)
Father's Education	0.013	-0.007	0.034*	0.008
	(0.018)	(0.017)	(0.016)	(0.013)
Child's Age (linear)	0.057*		0.011	
	(0.015)		(0.011)	
Age 16		0.106		-0.016
		(0.118)		(0.097)
Age 17 (reference)				
Age 18		0.293*		0.099
		(0.121)		(0.095)
Age 19		0.197		-0.007
		(0.110)		(0.091)
Female	0.087	-0.099	-0.132	-0.066
	(0.116)	(0.125)	(0.080)	(0.100)
Mother's Age	0.023*	0.076*	0.016*	0.082*
	(0.011)	(0.012)	(0.007)	(0.008)
Oldest boy	-1.138*	-1.096*	-1.083*	-1.057*
	(0.117)	(0.137)	(0.094)	(0.101)
Oldest girl	-1.296*	-0.823*	-0.994*	-1.013*
	(0.124)	(0.133)	(0.090)	(0.109)
Child <6 in household		1.371*		1.390*
		(0.160)		(0.100)
Constant	1.673*	0.319	3.324*	-0.175
	(0.608)	(0.544)	(0.449)	(0.403)
R-squared	0.34	0.46	0.30	0.56
Joint F test: Residual, Residual^2, Residual^3 X^2 (3)	18.18*	17.52*	24.78*	64.03*
Partial R^2	0.146	na	0.130	na
Observations	1911	1059	3418	1711

Notes: Dependent variable is number of living siblings. Robust standard errors are shown. Residual is estimated from Poisson regression of number of miscarriages on number of pregnancies, and number of pregnancies squared and cubed estimated by GLS. Models control for province of residence (coefficients not shown). IV strategy is described in the text. Models for the 1978-81 cohorts are IV probits thus there is no first stage. I have included these results in the table for descriptive purposes. *p < .05; † p<.10

Appendix Table A.3. Full IV Results for Models Shown in Table 5, IFLS 1997

	Urban 1968-77 Yrs of School Completed	Rural 1968-77 Yrs of School Completed	Urban 1978-81 Complete Jr. Sec.	Urban 1978-81 Enter Sr. Sec.	Rural 1978-81 Complete Jr. Sec.	Rural 1978-81 Enter Sr. Sec.
	2SLS		IV Probit			
No. of Siblings	-0.260†	-0.167	-0.110	-0.143†	-0.070	-0.054
	(0.144)	(0.17)	(0.085)	(0.079)	(0.049)	(0.055)
Mother's Education	0.325*	0.346*	0.133*	0.081*	0.117*	0.114*
	(0.031)	(0.031)	(0.021)	(0.017)	(0.015)	(0.015)
Father's Education	0.238*	0.321*	0.081*	0.086*	0.122*	0.117*
	(0.027)	(0.026)	(0.016)	(0.015)	(0.012)	(0.013)
Child's Age (20-49)	0.063*	-0.010				
	(0.030)	(0.021)				
Child age=16			-0.231†	-0.536*	-0.178†	-0.684*
			(0.134)	(0.125)	(0.095)	(0.115)
Child age=18			0.212	0.370*	-0.026	0.121
			(0.141)	(0.122)	(0.095)	(0.098)
Child age=19			0.157	0.394*	-0.139	0.079
			(0.136)	(0.118)	(0.094)	(0.096)
Female (1=yes)	-0.163	-0.523*	0.345*	0.212†	-0.064	-0.000
	(0.196)	(0.162)	(0.131)	(0.119)	(0.092)	(0.100)
Mother's Age	0.022	0.004	0.017	0.022†	0.011	0.007
	(0.017)	(0.011)	(0.014)	(0.012)	(0.008)	(0.008)
Oldest Boy (1=yes)	-0.538*	-0.247	0.055	0.153	0.060	-0.055
	(0.273)	(0.219)	(0.179)	(0.159)	(0.118)	(0.128)
Oldest Girl (1=yes)	-0.664*	-0.730*	-0.091	-0.015	-0.054	-0.248*
	(0.284)	(0.202)	(0.179)	(0.157)	(0.116)	(0.126)
Child <6 in household			0.210	0.261	-0.043	-0.097
			(0.211)	(0.185)	(0.131)	(0.139)
Constant	5.717*	6.635*	-0.956	-1.587*	-1.166*	-1.600*
	(1.025)	(0.801)	(0.492)	(0.445)	(0.311)	(0.340)
Observations	1911	3418	1059	1059	1711	1711
R-squared	0.35	0.33				
Log Likelihood			-2323	-2429	-3934	-3768

Notes: Robust standard errors are shown. Number of siblings is instrumented as described in the text. Models control for province of residence (coefficients not shown). *p < .05; † p<.10

Appendix Table A.4. OLS Models Predicting Completed Education by Cohort and Rural Residence both with and without 1993 Log Per Capita Expenditure Controlled (Comparable to Table A.1 above), IFLS 1997

Model	3.1		3.3		3.5		3.7		3.9		3.11	
	Urban 1948-57		Urban 1958-67		Urban 1968-77		Rural 1948-57		Rural 1958-67		Rural 1968-77	
Num. of sibs (linear, 0-14)	0.297* (0.094)	0.288* (0.090)	0.061 (0.070)	0.097 (0.066)	-0.179* (0.060)	-0.082 (0.058)	0.084 (0.068)	0.086 (0.068)	0.034 (0.054)	0.038 (0.051)	-0.043 (0.046)	-0.0004 (0.044)
Log Per Capita Expenditure		0.778* (0.294)		1.040* (0.213)		1.102* (0.174)		0.666* (0.214)		1.185* (0.161)		1.087* (0.124)
Mother's education	0.232* (0.082)	0.218* (0.082)	0.356* (0.043)	0.324* (0.042)	0.355* (0.039)	0.288* (0.039)	0.509* (0.081)	0.478* (0.079)	0.437* (0.051)	0.400* (0.052)	0.368* (0.035)	0.317* (0.033)
Father's education	0.272* (0.068)	0.249* (0.069)	0.318* (0.040)	0.261* (0.043)	0.234* (0.034)	0.192* (0.032)	0.421* (0.056)	0.397* (0.056)	0.376* (0.037)	0.317* (0.036)	0.314* (0.029)	0.254* (0.028)
Female	-1.399* (0.583)	-1.271* (0.581)	-0.770* (0.315)	-0.921* (0.306)	-0.245 (0.251)	-0.178 (0.244)	-1.340* (0.410)	-1.384* (0.412)	-1.348* (0.266)	-1.284* (0.264)	-0.731* (0.190)	-0.594* (0.182)
Oldest boy	0.374 (0.536)	0.420 (0.532)	0.081 (0.351)	0.154 (0.343)	-0.551† (0.286)	-0.428 (0.280)	0.384 (0.341)	0.395 (0.335)	-0.360 (0.242)	-0.317 (0.237)	-0.177 (0.210)	-0.052 (0.201)
Oldest girl	0.421 (0.494)	0.247 (0.496)	-0.399 (0.338)	-0.268 (0.329)	-0.745* (0.295)	-0.668* (0.289)	0.018 (0.337)	0.085 (0.341)	-0.169 (0.240)	-0.129 (0.236)	-0.644* (0.177)	-0.562* (0.172)
Constant	1.300 (3.406)	-3.539 (3.708)	6.293* (1.635)	2.321 (1.754)	6.212* (1.219)	1.984 (1.263)	3.932 (2.052)	1.533 (2.172)	5.684* (1.146)	0.815 (1.367)	6.940* (0.796)	2.393* (0.900)
R-squared	0.33	0.35	0.40	0.43	0.39	0.43	0.34	0.35	0.32	0.36	0.37	0.40
Number of Observations	470		990		1230		923		2060		2502	

Notes: All models control for child's age, mother's age, and province in which household is located (coefficients not shown). For each model, first column shows results without control for wealth while second column shows same model with 1993 per capita expenditure controlled for those not residing with their mothers. The restriction of non-coresidence along with a small amount of missing expenditure data explains the differences in sample size between these models and those shown in Table A.1 above. *p < .05; † p<.10

Appendix Table A.5. Binary Probit Models Predicting School Enrollment and Continuation Beyond Primary School for the 1978-81 Cohort (age 16-19 in 1997) both with and without 1993 Log Per Capita Expenditure Controlled (Comparable to Table 4 above), IFLS 1997

	Completed Jr. Sec.				Entered Sr. Sec.			
	Urban		Rural		Urban		Rural	
Model	4.1		4.3		4.5		4.7	
No. of siblings (linear)	-0.088* (0.032)	-0.074* (0.034)	-0.015 (0.024)	0.014 (0.024)	-0.099* (0.031)	-0.079* (0.032)	-0.022 (0.026)	0.006 (0.028)
Log Per Capita Expenditure		0.422* (0.091)		0.492* (0.063)		0.415* (0.081)		0.399* (0.068)
Mother's education	0.129* (0.020)	0.117* (0.021)	0.121* (0.016)	0.106* (0.016)	0.087* (0.017)	0.071* (0.017)	0.116* (0.015)	0.103* (0.015)
Father's education	0.079* (0.016)	0.067* (0.017)	0.121* (0.012)	0.105* (0.012)	0.086* (0.015)	0.074* (0.015)	0.117* (0.013)	0.100* (0.013)
Female	0.318* (0.133)	0.326* (0.136)	-0.053 (0.094)	-0.018 (0.097)	0.221† (0.122)	0.244* (0.124)	0.019 (0.102)	0.056 (0.104)
Oldest boy in family	0.088 (0.156)	0.110 (0.159)	0.128 (0.107)	0.193† (0.110)	0.253† (0.137)	0.303* (0.138)	0.002 (0.115)	0.079 (0.119)
Oldest girl in family	-0.037 (0.159)	-0.025 (0.163)	0.024 (0.106)	0.037 (0.108)	0.047 (0.139)	0.053 (0.143)	-0.232* (0.118)	-0.205 (0.119)
Child under 6 in household	0.189 (0.161)	0.238 (0.159)	-0.121 (0.107)	-0.096 (0.108)	0.223 (0.152)	0.267 (0.151)	-0.102 (0.112)	-0.089 (0.114)
Constant	-0.847 (0.483)	-2.518* (0.629)	-1.217* (0.320)	-3.153* (0.421)	-1.566* (0.450)	-3.234* (0.553)	-1.633* (0.348)	-3.176* (0.466)
Log likelihood	-429	-416	-899	-855	-526	-511	-738	-712
N	1024		1647		1024		1647	

Notes: All models control for child's age (in single years), mother's age and province of residence (coefficients not shown). For each model, first column shows results without control for wealth while second column shows same model with 1993 per capita expenditure controlled. A small amount of missing expenditure data explains the differences in sample size between these models and those shown in Table 4 above. *p < .05; † p<.10

Appendix Table A.6. Instrumental Variables Results for Models Shown in Tables A.4 and A.5, both with and without 1993 Log Per Capita Expenditure Controlled (Comparable to Table 5 above), IFLS 1997

Cohort	Urban 1968-77		Urban 1978-81		Urban 1978-81		Rural 1968-77		Rural 1978-81		Rural 1978-81	
Dependent Variable	Yrs of school completed (linear)		Complete Jr. Sec. (0/1)		Enter Sr. Sec. (0/1)		Yrs of school completed (linear)		Complete Jr. Sec. (0/1)		Enter Sr. Sec. (0/1)	
	Original estimate	w/ PCE	Original estimate	w/ PCE	Original estimate	w/ PCE	Original estimate	w/ PCE	Original estimate	w/ PCE	Original estimate	w/ PCE
No. of Siblings exogenous	-0.138* (0.050)	-0.082 (0.058)	-0.087* (0.032)	-0.074* (0.034)	-0.089* (0.030)	-0.079* (0.032)	-0.017 (0.040)	-0.0004 (0.044)	-0.017 (0.023)	0.014 (0.024)	-0.026 (0.026)	0.006 (0.028)
No. of Siblings instrumented	-0.260† (0.144)	-0.274 (0.171) *p<.109*	-0.110 (0.085)	-0.110 (0.090)	-0.143† (0.079)	-0.149† (0.082)	-0.167 (0.117)	-0.191 (0.117)	-0.070 (0.049)	-0.046 (0.051)	-0.054 (0.055)	-0.030 (0.057)
F (IV first stage)	18.2	13.8	--	--	--	--	24.8	25.0	--	--	--	--
Partial R^2 (IV first stage)	0.146	0.152	--	--	--	--	0.130	0.142	--	--	--	--
N	1911	1230	1059	1024	1059	1024	3418	2502	1711	1647	1711	1647

Notes: The exogenous regressions of years of school completed are estimated using OLS and those of school attendance are estimated using probit models. Results are duplicated here in the first row for ease of comparison. The IV coefficients are estimated using 2SLS for years of school completed and maximum likelihood IV probit for school attendance. The instrument is described in the text. Differences in sample sizes within each cohort result from restriction of non-coresidence in older cohort and a small amount of missing expenditure data across all groups. Original estimates are from Table 5, above. *p < .05; † p<.10

Figure 1. Educational Attainment by Family Size in Indonesia, 1997
Adults Ages 20-49, IFLS (N=10,573)

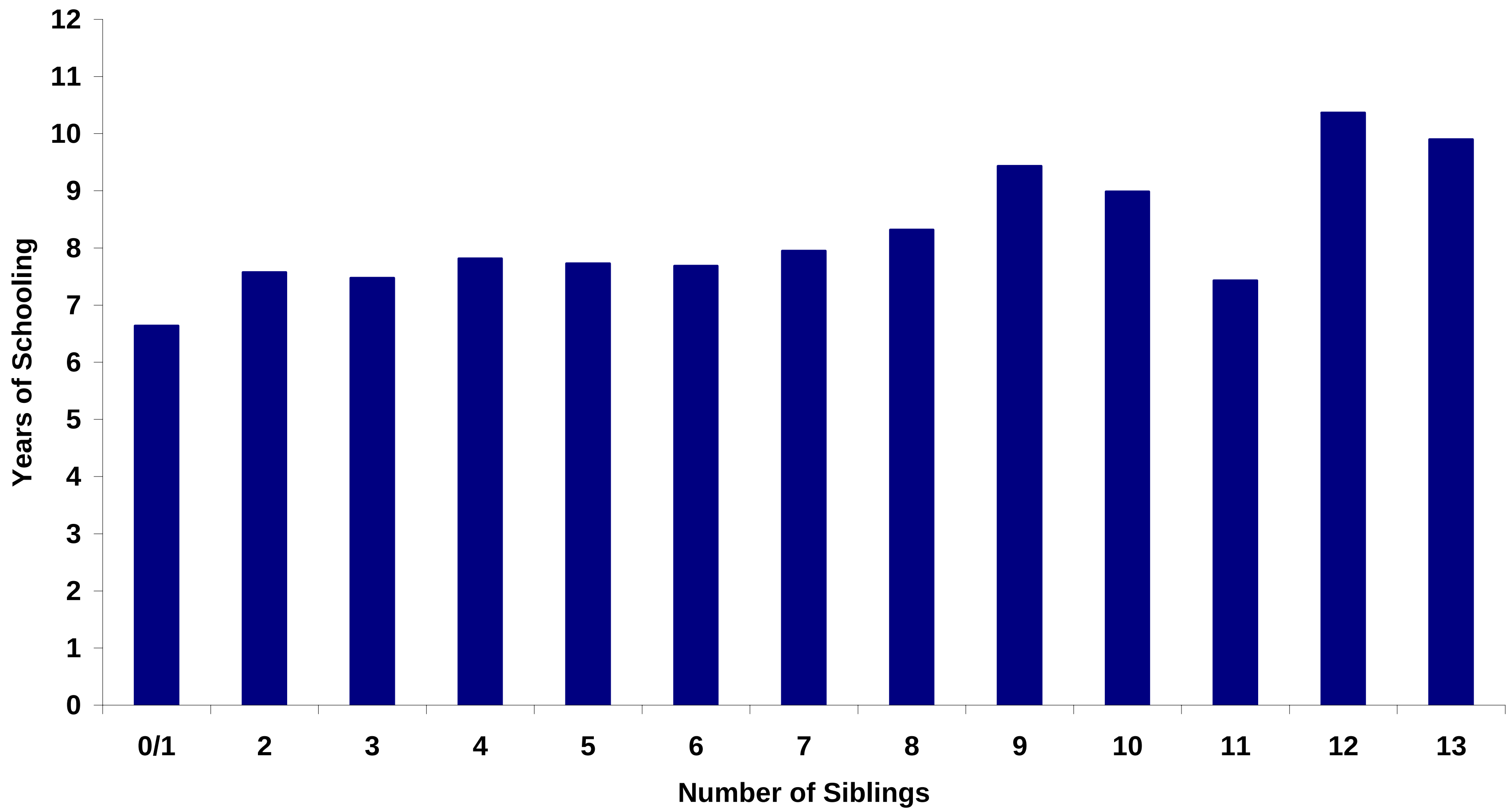

Figure 2. Educational Attainment by Family Size and Birth Cohort in 1997, IFLS, (N=10,573)