THE QUANTITY-QUALITY TRADE-OFF OF CHILDREN IN A DEVELOPING COUNTRY: IDENTIFICATION USING CHINESE TWINS*

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Testing the trade-off between child quantity and quality within a family is complicated by the endogeneity of family size. Using data from the Chinese Population Census, we examine the effect of family size on child educational attainment in China. We find a negative correlation between family size and child outcome, even after we control for the birth order effect. We then instrument family size by the exogenous variation that is induced by a twin birth and find a negative effect of family size on children's education. We also find that the effect of family size is more evident in rural China, where the public education system is poor. Given that our estimates of the effect of having twins on nontwins at least provide the lower bound of the true effect of family size, these findings suggest a quantity-quality trade-off for children in developing countries.

The relationship between family size and outcomes for children has fascinated social scientists for decades, particularly since the emergence of the theory of the quantity-quality trade-off that was developed by Gary Becker and his associates (Becker 1960; Becker and Lewis 1973; Becker and Tomes 1976; Willis 1973).[1] According to this model, an increasing marginal cost of quality (child outcome) with respect to quantity (number of children) leads to a trade-off between quantity and quality. Numerous empirical studies have attempted to test the quantity-quality trade-off and either confirmed the prediction by observing a negative correlation between family size and child quality or found no such correlation (Anh et al. 1998; Blake 1981; Knodel, Havanon, and Sittitrai 1990; Knodel and Wongsith 1991; Sudha 1997).[2] However, most studies simply treat family size as an exogenous variable and thus cannot establish causality. Both child quantity and child quality are endogenous variables because childbearing and child outcome are jointly chosen by parents (Browning 1992; Haveman and Wolfe 1995), which means that they are both affected by unobservable parental preferences and household characteristics.

One important method for tackling endogeneity is to use the exogenous variations in family size that are caused by the natural occurrence of twins to isolate the causal effect of family size on child quality.[3] Rosenzweig and Wolpin (1980b), in a pioneering study

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1. Many aspects of household behavior have been shown to be associated with family size. For example, researchers have thoroughly documented evidence for the relationship between fertility and parental labor supply (Angrist and Evans 1998; Rosenzweig and Wolpin 1980a), maternal economic outcome (Bronars and Grogger 1994), stability of marriage (Jacobsen, Pearce, and Rosenbloom 2001; Koo and Janowitz 1983), and children's educational and economic attainments (Haveman and Wolfe 1995; King 1987).

2. Also see King (1987) and Blake (1989) for surveys of early studies. Education and health are usually used as measures of child quality in the literature.

3. Some researchers have also used the gender of the first child (Lee 2004) or the gender composition of the first two children (Angrist, Lavy, and Schlosser 2005; Conley 2004b) as the instrument for family size. The former instrument is based on the prevailing preference for sons that is observed in Asian countries; the idea behind the

that used twins as a means of identification, found that family size (as induced by the birth of twins) has a negative effect on children's educational attainment in a small sample (25 twins in approximately 1,600 children) from India. However, a recent study by Black, Devereux, and Salvanes (2005) that also used twins as the exogenous variation, but with a large sample of the entire population of Norway, found that the effect of family size is reduced to almost zero after controlling for birth order, and that there is a monotonic decline in educational attainment by birth order.[4] These new findings suggest that the omission of the birth order effect may lead to biased estimates of the effect of family size on child quality. Another recent study by Angrist et al. (2005) that used both twin births and gender composition as the instrumental variables found no evidence for a quantity-quality trade-off of children in Israel.

Black et al. (2005) and Angrist et al. (2005) raised a provocative question: Is there a quantity-quality trade-off as formulated by Becker? These studies made many improvements on the earlier study of Rosenzweig and Wolpin (1980b), particularly in terms of data quality and empirical specifications, and thus their evidence should be more robust. However, in addition to having larger samples and improved model specification, another important difference between these more recent studies and that of Rosenzweig and Wolpin is that the latter used data from a developing country, whereas the former used data from developed countries. In a developed country with a comprehensive welfare system, such as Norway, where there is both a good public education system (even college is free) and generous government support for childbearing and childcare, the cost of children, and particularly the educational expenditure, accounts for just a small proportion of the budget of parents. Thus, the quantity-quality trade-off may not be obvious in such countries. In contrast, in a developing country, such as India, where there is neither a well-functioning public education system nor generous support for childbearing and childcare, the cost of child quality is mostly borne by the parents. Thus, the quantity-quality trade-off is more likely to occur in a developing country.[5] Therefore, it is important to use good data from developing countries to verify whether the findings of Black et al. (2005) can be replicated.

In this analysis, we test the quantity-quality trade-off by using mainly the 1% sample of the 1990 Chinese Population Census. China has a poorly functioning education system, especially in rural areas, where poverty is the main reason that children drop out of primary and high school (Brown and Park 2002). Using educational level and school enrollment as measures for child quality, we find a negative correlation between family size and child quality under various specifications, even after controlling for the birth order effect. We identify the negative effect of family size on child education through two-stage least squares (2SLS) estimations using twin births as the instrumental variable (IV) for family size. Our findings strongly support the prediction of Becker and his associates on the quantity-quality trade-off of children but differ from those of Black et al. (2005).

Using twin births as the IV is not without caveats. Twinning may affect sibling outcomes through mechanisms other than family size, such as the reallocation of family resources from twins toward nontwin children and closer spacing between twins (Rosenzweig

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latter instrument is that parents of same-gender siblings are more likely to have an additional child (Angrist and Evans 1998).

4. Sociologists and psychologists have documented the effect of birth order on child outcomes. See, for example, the summary of the findings by King (1987) and Conley (2004a). Several earlier empirical studies were conducted by economists. For example, Hauser and Sewell (1985) found no significant effect of birth order, Behrman and Taubman (1986) showed that children born later tend to have an educational disadvantage, and Hanushek (1992) reported a U-shaped pattern of education by birth order for large families.

5. There is also some evidence from developing countries in studies of epidemiology and public health, although the methods used in these studies are usually different from those used by economists. See, for example, Karmaus and Botezan (2002).

and Zhang 2006). Thus, twinning is not a perfect IV. However, given that both the reinforcing intrafamily resource allocation (i.e., parents invest more in nontwin children who have greater endowments) and the potential correlation between sibling outcome and closer spacing between twins may bias the 2SLS estimates toward zero, our finding of a negative effect of family size implies that the true effect should be more negative after removing the bias, thus supporting the quantity-quality theory.

We are among the first to draw on twins data from a developing country to test the theory of the quantity-quality trade-off of children. Given that the quantity-quality trade-off is expected to be more pronounced in developing countries, it is surprising that few previous studies have drawn on twins data from developing countries, although this is probably due to difficulty in obtaining data. We are also among the first to explicitly examine the trade-off in the context of China. Most of the previous related studies explored the determinants of Chinese children's educational attainment and emphasized the rural-urban gap (Connelly and Zheng 2003; Hannum 1999; Knight and Li 1993, 1996), gender inequality (Broaded and Liu 1996; Hannum 2002, 2003; Tsui and Rich 2002), or poverty and credit constraints (Brown and Park 2002). However, these studies either ignored the effect of family size or merely treated it as an exogenous control variable. To the best of our knowledge, the only exception is a paper by Qian (2005), who attempted to use China's birth control policy as an identification to test the quantity-quality trade-off.

Knowing the true effect of family size on child quality has important policy implications for developing countries, and in particular for China. Our findings suggest that the birth control policy in China has the potential positive effect of increasing the quality of children. If, as we find, a smaller family size is generally associated with a better average educational outcome for children, then the one-child (or two-child) policy has improved child quality by reducing the number of children in a household. In particular, we find that the trade-off between quantity and quality is more pronounced in rural areas, where the least well-off people live. This implies that the birth control policy, if it is as effective as expected by policy-makers, actually does enhance the quality of rural children and ultimately spurs economic growth (Li and Zhang 2007).

In the following sections, we specify our empirical strategy, describe our sample, present our estimates of the effect of family size on children's educational outcomes, and finally offer our conclusions based on the analyses.

EMPIRICAL METHOD

We follow the recent empirical literature and specify our general estimation as follows,

$$EDU = \beta_0 + \beta_1 SIZE + \mathbf{X}\beta_2 + \mathbf{Z}\beta_3 + \varepsilon, \tag{1}$$

where *EDU* is the educational attainment of the child as measured by the two educational outcome variables of educational level and school enrollment. The variable *SIZE* is the number of children in the family, and the coefficient β_1, which reveals the quantity-quality trade-off, is what interests us. **X** is a vector of child characteristics, including age, gender, ethnic group, birth order, and place of residence; **Z** stands for a set of parental attributes, including age and educational level. We also run separate regressions for the rural and urban samples to allow the effect of family size to interact with residence areas.

The coefficient β_1 as estimated by the ordinary least squares (OLS) method may merely suggest a correlation, rather than a causal effect, because family size is likely to be endogenous. Following Rosenzweig and Wolpin (1980b) and Black et al. (2005), we use the birth of twins as an identifying instrument for family size. The first stage of the two-stage least squares (2SLS) estimation is given by

$$SIZE = \alpha_0 + \alpha_1 TWIN + \mathbf{X}\alpha_2 + \mathbf{Z}\alpha_3 + \nu, \tag{2}$$

and Eq. (1) becomes the second stage. In Eq. (2), *TWIN* is a dummy variable that equals 1 if the *n*th delivery is a multiple delivery, and 0 otherwise; all of the other variables are the same as specified in Eq. (1).

As noted by Rosenzweig and Wolpin (2000), the presence of any twin birth in a family makes for an inappropriate instrument because its probability increases with the number of deliveries. To avoid this problem in estimating the 2SLS models, we restrict the sample to families with at least *n* births so that we can be fairly confident that the families with twins at the *n*th delivery have the same preference for the number of children as those with singleton births. If the occurrence of multiple births is randomly assigned by nature, then twin births should have little or no effect on children's education except through family size. Thus, the 2SLS estimate of β_1 would consistently measure the causal effect of family size on child quality. We further discuss the validity of the twins instrument in a later section.

DATA

We mainly use the 1% sample of the 1990 Chinese Population Census that was collected by the Chinese National Bureau of Statistics (formerly the State Statistic Bureau). It is the fourth of its kind, following the three censuses that were conducted in 1953, 1964, and 1982.[6] The 1% sample covers 11,475,104 individuals from 2,832,103 households. The data set contains a record for each household and includes variables that describe the location, type, and composition of the households. Each household record is followed by a record for each individual residing in the household. The individual variables include demographic characteristics, occupation, industry, educational level, ethnicity, marital status, and fertility.

We use the relation identifier to match children to their parents within the households. Specifically, we identify individuals who are labeled "child" as the primary observation, and obtain the family size by counting the number of children in the household. We then attach the data of the parents—that is, those who are labeled "household head" or "spouse"—to all of the children in the household. For each mother, we also have data on the total number of children born and the number of children still alive, which helps us identify whether the family size is complete.

To facilitate our analysis, we use a subsample of the census data. First, we use only children of the household head because we can match the parental information and count the number of children of a couple only for such children. Second, we drop households with no children or with a family size that exceeds the total number of surviving children, the latter of which is likely to be the result of data error.[7] Third, we restrict the sample to children who were between 6 and 17 years old and whose mothers were no older than 35 in the census year. We use 6 as the lower bound for the age of children because it is the minimum age of school enrollment in China, and no education information was recorded for children younger than 6 in the census. Restricting the mother's age to 35 or younger makes it fairly certain that no adult children have moved out of a household. We impose such a restriction because we are unable to track children who had already left the household by the time of the survey.[8] Finally, we exclude some households with missing information on fathers,[9] and a small number of families with a birth that occurred before the mother was 16.

6. The two earlier censuses are not available to researchers. The 1982 census is less useful for our purposes because of the lack of school enrollment information and explicit rural identifiers, although we perform some sensitivity analyses using the 1982 sample. The latest census, which was conducted in 2000, will be available soon.

7. This discrepancy may arise as a result of adopted children, but there is no information in the data to distinguish between adoption and birth.

8. With this restriction, only about 1% of the households have children who live outside of the household. We also conducted regressions using a sample excluding these families and obtained the same results.

9. Data on fathers were missing for 7% of all cases. In addition to dropping these observations, we also performed the estimations by creating categories of missing father variables, and the results were the same.

With these restrictions, we are left with a sample of 675,492 children from 447,159 households. Because the census does not include an explicit identifier for twins, we define twins as children who were reported to be born in the same year and month to the same woman. One percent of our sample comprises twin births. The first two columns of Table 1 report the summary statistics for the whole sample and for the sample excluding twins. No significant differences can be observed between the two columns; the statistics remain almost the same for each variable.

It is worthwhile to outline the institutional background of nontertiary education in China before we offer the definitions of the education variables. In 1986, the Law of Compulsory Education officially declared the implementation of nine compulsory years of schooling (six years of primary school and three years of junior high school) throughout China. However, the policy of compulsory education was not implemented uniformly across the country. The Resolution on Educational System Reform, which was initiated in 1985, devolved the total responsibility of implementing compulsory education to local governments, and thus the provision of basic education depends on the local budget or

Table 1. Descriptive Statistics of the 1% Sample of the 1990 Chinese Population Census

	Full Sample		By Area	
Variables	Including Twins (1)	Excluding Twins (2)	Rural (3)	Urban (4)
Observations of Children	675,492	665,738	595,729	79,763
Age	8.71 (2.39)	8.72 (2.39)	8.78 (2.42)	8.27 (2.08)
Male	0.52 (0.50)	0.52 (0.50)	0.52 (0.50)	0.52 (0.50)
Han	0.91 (0.28)	0.91 (0.28)	0.91 (0.29)	0.93 (0.26)
Rural	0.88 (0.32)	0.88 (0.32)	—	—
Education (aged 6 and above)				
Enrolled in school	0.70 (0.46)	0.71 (0.46)	0.71 (0.46)	0.70 (0.46)
Illiterate	0.28 (0.45)	0.28 (0.45)	0.28 (0.45)	0.30 (0.46)
Primary school	0.69 (0.46)	0.69 (0.46)	0.69 (0.46)	0.67 (0.47)
Junior high school and above	0.03 (0.15)	0.03 (0.15)	0.03 (0.15)	0.03 (0.17)
Education (aged 8 and above)				
Enrolled in school	0.91 (0.28)	0.91 (0.28)	0.91 (0.29)	0.97 (0.16)
Illiterate	0.07 (0.43)	0.07 (0.26)	0.08 (0.27)	0.03 (0.16)
Primary school	0.89 (0.31)	0.89 (0.31)	0.89 (0.31)	0.92 (0.27)
Junior high school and above	0.04 (0.19)	0.04 (0.19)	0.03 (0.18)	0.05 (0.22)
Observations of Families	447,159	442,423	376,680	70,479
Number of Children	2.10 (0.90)	2.09 (0.89)	2.26 (0.87)	1.27 (0.57)
Having Two or More Children	0.74 (0.43)	0.75 (0.43)	0.85 (0.36)	0.23 (0.42)
Having Three or More Children	0.27 (0.45)	0.27 (0.45)	0.32 (0.47)	0.04 (0.19)
Having a Multiple Birth	0.01 (0.10)	—	0.01 (0.10)	0.01 (0.09)
Mother's Age	31.60 (2.80)	31.60 (2.80)	31.50 (2.90)	32.50 (2.20)
Father's Age	34.20 (3.80)	34.20 (3.80)	34.10 (3.90)	34.80 (3.20)

Notes: Standard deviations are shown in parentheses. All sampled children were at least age 6 in 1990, with nonmissing information on both mothers and fathers. Mother's age is restricted to be 35 or younger in the census year.

level of economic development (He 1996). As a result, access to education in rural areas is much worse than in urban areas because rural citizens and governments are much poorer. In the poor rural areas, public schools are not widely available, and even in regions where the schooling system is publicly provided to all children, it is not totally free; parents still need to pay tuition and fees. Such a financial burden is one of the main reasons why poor families, who are often unable to borrow funds to finance their children's education, pull their children out of school (Brown and Park 2002).

In this analysis, we employ two education variables that are reported in the census: educational level and school enrollment. *Educational level* is defined as an ordered discrete variable that indicates three educational levels: illiterate, primary school, and junior high school and above.[10] *School enrollment* is defined as a binary indicator that equals 1 if a child was enrolled in school or had graduated, and 0 if a child had dropped out of school or never enrolled. Previous research has shown that school enrollment is a good indicator of educational attainment in developing countries (Alderman et al. 2001; Glewwe and Jacoby 1995; Glewwe, Jacoby, and King 2001). Table 1 shows that the average enrollment rate is 70% for the full sample and that children at the three educational levels account for 28%, 69%, and 3% of the sample, respectively. For children who were at least 8 years old, the enrollment rate rises to 91%, and the educational level also improves.

An important aspect of the data is that there is a large rural-urban difference in both education and fertility. In columns 3 and 4 of Table 1, we report the attributes of the rural and urban subsamples separately. Of all of the children, 88% were from rural areas. Note that although there is no rural-urban difference for the education variables for the whole sample of children, there is a large difference among children 8 years old and older. The reason for the lack of difference in the whole sample is that rural children went to school earlier. In urban areas, the enrollment age was normally 7 or 8 for the generation of children in our sample, and that age requirement has been strictly enforced. However, children in rural areas were able to go to school as early as age 6. Note also that the fertility of rural families is much higher than that of urban families, with the rural-urban gap in the number of children being as large as 1. Over four-fifths of the rural households had more than one child, in comparison to only one-fifth of the urban households.[11] These rural-urban differences make it important to analyze the rural and urban subsamples separately.

To gain a picture of how education may vary with family size, we present in Table 2 children's educational level by family size for both the rural and urban subsamples. To control for the age effect, we report the proportion among young children (aged 13 or below) who have at least primary school education and the proportion among older children (aged over 13) who have at least a junior high school education. Several aspects are worth noting. First, a greater family size is clearly associated with lower average education. Although among children younger than 13 years old, only children appear to have a lower education than children in two-child families, there is a monotonically decreasing trend for family sizes of two to six and above. Moreover, the advantage of two-child families over single-child families disappears for children who are older than 13. Second, urban children seem to have, on average, a higher educational level than rural children.

10. The census codes the educational level into seven categories: illiterate, primary school, junior high school, senior high school, technical school, junior college, and university. Because the proportion of respondents with an educational level of senior high school or above is very small in the sample (less than 0.01%), we classify all of these observations into the third level of junior high school and above. Having more categories for educational levels does not change our results.

11. Although the one-child policy had been in force for 10 years by the time of the census in 1990, there is empirical evidence that the policy was more effective in deterring second births in urban areas than in rural areas (Ahn 1994; Zhang and Spencer 1992).

Table 2. Descriptive Statistics of Educational Level, by Family Size: 1990 Chinese Population Census

	Family Size					
	One Child	Two Children	Three Children	Four Children	Five Children	Six or More Children
Full Sample	116,766	296,082	183,606	59,846	15,046	4,146
Primary school and above (age ≤ 13)						
All	0.68	0.73	0.68	0.64	0.61	0.57
Male	0.69	0.73	0.69	0.66	0.63	0.59
Female	0.64	0.72	0.66	0.63	0.60	0.56
Junior high school and above (age > 13)						
All	0.52	0.42	0.33	0.24	0.17	0.17
Male	0.51	0.43	0.36	0.28	0.22	0.21
Female	0.54	0.40	0.29	0.21	0.14	0.15
Rural Sample	61,784	277,474	179,236	58,579	14,639	4,017
Primary school and above (age ≤ 13)						
All	0.70	0.73	0.67	0.64	0.61	0.57
Male	0.72	0.73	0.69	0.66	0.63	0.59
Female	0.64	0.72	0.66	0.62	0.60	0.56
Junior high school and above (age > 13)						
All	0.42	0.37	0.31	0.22	0.16	0.15
Male	0.43	0.39	0.35	0.27	0.21	0.19
Female	0.40	0.34	0.27	0.19	0.13	0.13
Urban Sample	54,982	18,608	4,370	1,267	407	129
Primary school and above (age ≤ 13)						
All	0.65	0.78	0.77	0.72	0.67	0.59
Male	0.65	0.78	0.76	0.72	0.66	0.65
Female	0.65	0.77	0.77	0.72	0.67	0.55
Junior high school and above (age > 13)						
All	0.78	0.78	0.71	0.73	0.59	0.56
Male	0.75	0.76	0.68	0.73	0.54	0.56
Female	0.83	0.80	0.73	0.73	0.63	0.56

Notes: All sampled children were at least age 6 in 1990, with nonmissing information on both mothers and fathers. Mother's age is restricted to be 35 or younger in the census year.

Except for young children in the only-child group, urban children fare better in terms of education regardless of family size and gender. Finally, male children consistently have better education than female children in the rural sample, but the gender-based difference is less explicit in the urban sample.

THE EFFECT OF FAMILY SIZE ON CHILDREN'S EDUCATION

In this section, we present the results of OLS and 2SLS regressions designed to systematically test whether family size has a negative effect on children's educational attainment in China. We first discuss several issues regarding the validity of using twins as the IV. Then we use twins at the nth delivery (n = 1, 2, 3) to instrument family size, and perform estimations as specified by Eqs. (1) and (2). We also examine whether the effect of family size is different in rural versus urban areas and check the heterogeneity of the effect under other sample stratifications as well. For all of the estimations, we control for a full set of child and parent attributes that comprises the cubic form of child age, gender, indicator of being *Han* Chinese, birth order, parents' age and educational level, and rural (if applicable) and provincial dummy variables. Due to space constraints, the estimates for these control variables are not reported.

Twins Instrument

Unobserved family preferences. Before reporting the estimation results, we first discuss the validity of using twin births as our IV. A good instrument should be highly correlated with the number of children in a family but should not affect the child outcome except through family size. That is to say, a valid IV should not be correlated with unobserved parental and household characteristics that are captured by the error term in Eq. (1). The birth of twins is an important source of exogenous variation in fertility that has been used in previous research (Rosenzweig and Wolpin 2000) and is believed to be unlikely to depend on family background.[12] Although the correlation between twin births and unobserved household attributes is untestable by design, we follow Black et al. (2005) and examine whether the occurrence of twins is associated with certain observed characteristics, such as the educational level of parents. Similar to their findings, the F tests based on linear probability models suggest that the probability of having a twin birth is uncorrelated with the educational level of either mothers or fathers in our sample.

Birth spacing. Another concern is that a twin birth may affect child outcome through birth spacing. There are two possible ways that twin births may affect sibling outcome via spacing. In both cases, the 2SLS estimates of the effect of family size could be biased. First, if the space between the two following siblings has a significant effect on the quality of previous children, then the birth of twins may influence the outcome of early children by effectively reducing the space toward zero. In other words, twin births may affect the quality of prior children through both increased family size and narrowed spacing, which are inherently indistinguishable.

To address this possibility, we follow Black et al. (2005) and use samples of families without twins to check whether child education is correlated with the age gap (spacing) between the two immediately following siblings. Specifically, we examine the first children in families with at least three births, and first and second children in families with at least four births. As shown in Appendix Table A1, almost all of the OLS coefficients on spacing appear to be significantly negative, which means that a child is better educated if the following births have a closer spacing.[13] If this can be arguably extended to the case of

12. The existence of sex-selective abortion in China might undermine the validity of twinning instrument because the access to ultrasound use and abortion services allows parents to "choose" which birth to give. This became a more serious issue after China implemented the one-child policy in 1979. However, our analysis using the 1982 census data suggests that this does not seem to be a big concern.

13. One explanation for this result is that parents tend to give more equal treatment to children having closer spacing. More equal treatment would make increasing their average quality more expensive and drive parents to move more resources away from children in the following births and to older siblings, hence increasing the siblings' quality. Rosenzweig and Zhang (2006) used this logic to argue for the intrafamily resource reallocation from twins to nontwin siblings, as we show later.

twins, then twinning should improve sibling outcome because the spacing between twins is zero, and thus the spacing effect of twinning should bias our estimate of the quantity-quality trade-off toward small or no negative effect. Given this potential bias, if we still find a large negative effect of family size, we can be fairly certain that a quantity-quality trade-off exists.

A second way that a twin birth can affect child quality through spacing is that the probability of twins increases with maternal age at birth (Bronars and Grogger 1994). Thus, a mother is more likely to give a twin birth if that birth is spaced farther from the previous birth, conditional on her age at the previous birth. If such spacing similarly affects the outcome of prior children, then a twin birth will be (negatively) correlated with sibling outcome beyond the effect of twins through family size, leading to a negative bias in the 2SLS estimates of the family size effect.

However, this potential bias can be tackled by including the spacing between the potential twin birth and the previous birth as a control in our estimation. In practice, when we use twins at the nth delivery to instrument family size, we add a spacing variable that measures the age difference between the nth and $(n - 1)$th deliveries.[14] Unless there is a serious bias, the estimates will not be much changed by the additional control. As we will show later in this section, controlling for spacing immediately prior to the potential twin birth has very little effect on our estimates.

Interchild reallocation. Finally, we discuss the concern that twin births may directly affect child quality by changing the intrafamily resource allocation. This point, raised by Rosenzweig and Zhang (2006), argues that for parents who reinforce endowment differences across children (i.e., invest more in children with greater endowments), twinning will result in the allocation of resources toward nontwin siblings because (1) per-child investments in twins are more costly compared with nontwins due to closer spacing, and (2) twins tend to have inferior birth endowments, such as lower birth weight, compared with nontwin siblings (Behrman and Rosenzweig 2004). Moreover, consistent with the findings of Behrman, Rosenzweig, and Taubman (1994), Rosenzweig and Zhang (2006) found some empirical evidence of the reinforcing behavior of parents, using a sample of Chinese twins.

Therefore, without taking account of such a reinforcing effect on nontwins, which is a positive bias, the negative effect of increased family size on the average child outcome will be underestimated (i.e., biased toward zero) if researchers look only at the impact of twin births on nontwin children. As Rosenzweig and Zhang (2006) put it, the estimates of the effects of twinning on twins and nontwin siblings bound the true quantity-quality trade-off for an average child, with the latter estimates always giving the lower bound, which may be zero or even positive, as found in some recent studies. Although controlling for birth weight may help tighten the range of the upper and lower bounds, the census data that we use do not contain such information. However, to the extent that our estimates can be interpreted as the lower bound of the effect of family size, if we still find negative estimates, the true effect should be more negative; thus the findings would support the quantity-quality theory. The important point is that since we know the direction of possible bias in the IV estimate (i.e., biased toward zero), the IV bias is not a problem for us in inferring the direction of the quantity-quality trade-off if our IV estimate is negative. On the other hand, if we find a positive IV estimate, we would be unable to draw any conclusions about the trade-off.

OLS and 2SLS Estimations

Table 3 presents the OLS and 2SLS estimates of the effect of family size on children's education for the full 1990 sample, along with the first-stage relationship between family size and

14. Implicitly, this is equivalent to controlling for mother's age at the nth delivery because we already include children and their mother's age in the regression.

Table 3. Ordinary Least Squares (OLS) and Two-Stage Least Squares (2SLS) Estimates of the Effect of Family Size on Children's Educational Outcomes: 1990 Chinese Population Census

	Educational Level[a]			Whether Enrolled in School[b]		
Independent Variable	OLS (1)	First Stage (2)	2SLS (3)	OLS (4)	First Stage (5)	2SLS (6)
Twins at the First Delivery						
All nontwin children (N = 672,207)						
Number of children	–0.028**	0.555**	–0.040**	–0.027**	0.555**	–0.030*
	(–42.59)	(42.38)	(–2.83)	(–43.42)	(42.38)	(–2.19)
Twins at the Second Delivery						
Nontwin children in families with two or more births (N = 553,438)						
Number of children	–0.038**	0.696**	–0.011	–0.036**	0.696**	–0.009
	(–48.07)	(57.62)	(–1.11)	(–47.27)	(57.62)	(–0.94)
First children in families with two or more births (N = 327,363)						
Number of children	–0.031**	0.780**	0.002	–0.027**	0.780**	0.002
	(–29.58)	(56.55)	(0.18)	(–28.64)	(56.55)	(0.21)
Number of children (control for spacing)	–0.033**	0.833**	0.002	–0.028**	0.833**	0.002
	(–31.09)	(61.84)	(0.27)	(–29.44)	(61.84)	(0.19)
Twins at the Third Delivery						
Nontwin children in families with three or more births (N = 256,487)						
Number of children	–0.044**	0.821**	–0.027†	–0.040**	0.821**	–0.025†
	(–29.19)	(51.25)	(–1.95)	(–27.94)	(51.25)	(–1.87)
First and second children in families with three or more births (N = 204,901)						
Number of children	–0.038**	0.857**	–0.024†	–0.032**	0.857**	–0.025†
	(–21.42)	(51.75)	(–1.70)	(–19.85)	(51.75)	(–1.82)
Second child	–0.029**		–0.031**	–0.021**		–0.022**
	(–16.26)		(–11.15)	(–12.55)		(–8.42)
Number of children (control for spacing)	–0.040**	0.884**	–0.023†	–0.035**	0.884**	–0.023†
	(–22.35)	(54.29)	(–1.65)	(–21.05)	(54.29)	(–1.73)
Second child (control for spacing)	–0.027**		–0.030**	–0.019**		–0.021**
	(–15.16)		(–10.14)	(–11.31)		(–7.49)

Notes: Robust t statistics, which allow for correlation of errors within family, are shown in parentheses. All regressions include age, age squared, age cubed, indicators for male and Han, parents' age and age squared, parents' educational level, and rural and provincial dummy variables.

[a] 1 = illiterate; 2 = primary school; and 3 = junior high school or above.

[b] 1 = yes; 0 = no.

† $p < .10$; * $p < .05$; ** $p < .01$

twins at the nth delivery.[15] The results with educational level as the dependent variable are reported in the first three columns, and the results with school enrollment as the dependent variable are reported in the last three columns. From top to bottom, we list in three panels the estimates for families with at least n births in increasing order of n from 1 to 3. For n = 2 or 3, we examine children of all parities and children prior to parity n, respectively.

Similar to the pattern we observe in Table 2, the OLS estimates in columns 1 and 4 consistently show a significantly negative correlation between family size and children's education, regardless of the dependent variable and sample used. For example, the OLS coefficient in the top panel (column 4) suggests that, everything else being constant, having one more child in the family reduces a child's probability of enrollment by approximately 3 percentage points.

Using twin births as the IV, the 2SLS estimates in columns 3 and 6 continue to suggest a negative effect of family size on child outcome except for the middle panel of families with two or more births, and the results are qualitatively the same for both education outcomes. In particular, the 2SLS coefficients on family size are significant at the 1% level for families with one or more births (top panel), and significant at the 10% level for those with three or more births (bottom panel). Note that given previous discussions, our 2SLS estimates may be subject to positive biases induced by not taking into account the closer space between twins or resource allocation from twins to nontwin siblings. Hence, that our negative estimates understate the true effect of family size indeed implies the existence of a quantity-quality trade-off. Moreover, as shown here, controlling for the space between parity n and parity $n - 1$ only marginally changes the estimates, suggesting that the bias from omitting this variable is negligible. Finally, it is worth noting that the first-stage relationship is significant for all of the specifications, with t ratios of well above 40. Consistent with the previous literature, the effect of a twin birth on family size increases with a higher parity, which ranges from 0.6 to 0.9 in our sample.

Although not shown in the table, the control variables have the expected signs. In general, male or Han children have an educational advantage over female or minority children, and rural children tend to have inferior education outcomes compared with urban peers. We also add a vector of birth order indicators to examine whether the effect of family size is partially driven by birth order. In fact, the addition of birth order controls has very little effect on both the OLS and 2SLS coefficients on family size. This result is in stark contrast to that of Black et al. (2005), who found that the effect of family size becomes trivial once the birth order effect is controlled. We also find little evidence of a monotonic decline of child quality by birth order, as distinct from Black et al. (2005). Rather, although the coefficients on second child have a negative sign in Table 3, we find that the coefficients on higher birth orders are positive in some cases, which indicates that children who are born later in large families are more likely to have an advantage over children who are born earlier (conditional on family size).

Effects in Rural and Urban Areas

As discussed in the Data section, there is a considerable rural-urban gap in access to and completion of schooling in China. This gap is the result of both supply- and demand-side factors. On the supply side, the average school quality is much better in urban China than in rural China. While urban public schools receive substantial subsidies from local governments, many rural schools are badly funded and thus short of well-trained teachers. The lack of government funding compels many rural schools to become self-financed, which forces many rural children out of school because their parents cannot afford to pay the school fees (Brown and Park 2002). On the demand side, rural parents may have lower

15. The t statistics that are reported here, as in all of the regressions in this analysis, allow for the correlation of errors for any two children in the same family.

educational aspirations for their children than urban parents. This is probably due to the lower return and higher opportunity cost of sending children to school for rural families, because rural children can contribute to the household income by carrying out farm and house work even at very young ages.[16]

Because of the rural-urban education gap, we expect the effect of family size on child quality to be different in rural and urban areas. Given that public education is more prevalent and children's education is held to be more important in urban China, having an additional child in the family may result in a smaller adverse impact on the average child education compared with the effect in a rural family. In this sense, the rural-urban difference within China to some extent resembles the difference between China and Norway. To allow for disparity in the effect of family size between rural and urban areas, we present in Table 4 the results of the same regressions as in Table 3 using the rural and urban subsamples, respectively. We skip reporting the estimates for enrollment because they are very similar to those for educational level.

Interestingly, the OLS estimates show that the effect of family size is smaller in urban areas than in rural areas. As shown in column 1, the estimates for the rural sample are very close to those for the full sample. In contrast, the OLS coefficients on family size for the urban sample, listed in column 4, are smaller in magnitude, and some are not statistically different from zero. It is also worth noting that ethnic- and gender-based differences are less explicit among urban children (not shown). Although there is a clear educational advantage for male or Han children in rural areas, the evidence from urban children shows an insignificant ethnic effect and even a negative male effect.[17]

Not surprisingly, the quantity-quality trade-off appears to exist only for rural families, as suggested by the 2SLS estimates in column 3. As with the full sample estimates, we find an effect of family size significant at the 1% level for the $n = 1$ case and an effect significant at the 10% level for the $n = 3$ case in the rural sample, although the estimates for first and second children in the latter case are marginally insignificant. Nevertheless, for the urban sample (column 6), none of the 2SLS estimates are statistically different from zero at the 10% level, which implies the absence of quantity-quality trade-off in urban families. Note that because our estimates are potentially biased upward, the zero effects for urban families are still likely to be consistent with the quantity-quality trade-off, and the consistency is more evident for the rural sample, for which the negative effects are detected.

So far, we find that family size is negatively correlated with children's educational attainment in China when we measure education both by discrete levels and by the probability of being enrolled in school. The negative effect is not sensitive to the inclusion of controls for birth order and spacing. By examining the rural and urban subsamples separately, we find that the adverse impact of family size is smaller in urban China. We also observe some evidence of a negative second-child order effect but do not identify a significant negative effect of higher birth orders in large families.

The One-Child Policy

One concern about the previous rural-urban differences in the effect of family size on child outcome is to what extent such disparities can be attributed to the variation in birth control policy between rural and urban China. China introduced its unique one-child policy in 1979. Under this policy, each couple is allowed to have only one child.[18] Households are

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16. See Becker (1991), Dasgupta (1995), Johnson (1994), and Ray (1998) for arguments on the benefits of children in developing countries.

17. The absence of an educational advantage for boys in urban China has also been observed in recent literature (e.g., Connelly and Zheng 2003; Tsui and Rich 2002).

18. This policy applied only to the Han Chinese during most of the 1980s; minorities were normally allowed to have two children. In some regions, such as Xinjiang and Tibet, minorities can even have more than two children.

Table 4. Ordinary Least Squares (OLS) and Two-Stage Least Squares (2SLS) Estimates of the Effect of Family Size on Children's Educational Level: 1990 Chinese Population Census (rural vs. urban)

	Dependent Variable: Educational Level[a]					
	Rural			Urban		
Independent Variable	OLS (1)	First Stage (2)	2SLS (3)	OLS (4)	First Stage (5)	2SLS (6)
Twins at the First Delivery						
All nontwin children						
Number of observations		593,186			79,021	
Number of children	–0.030** (–43.69)	0.505** (33.61)	–0.042* (–2.46)	–0.003 (–1.37)	0.785** (37.81)	–0.024 (–1.13)
Twins at the Second Delivery						
Nontwin children in families with two or more births						
Number of observations		529,511			23,927	
Number of children	–0.038** (–47.22)	0.689** (54.93)	–0.013 (–1.21)	–0.022 (–4.96)	0.849** (21.04)	–0.008 (–0.25)
First children in families with two or more births						
Number of observations		312,378			14,985	
Number of children	–0.030** (–27.91)	0.771** (53.64)	0.000 (0.04)	–0.027** (–4.16)	0.922** (22.52)	0.007 (0.23)
Number of children (control for spacing)	–0.031** (–29.27)	0.826** (58.85)	0.001 (0.11)	–0.026** (–3.93)	0.947** (23.52)	0.006 (0.18)
Twins at the Third Delivery						
Nontwin children in families with three or more births						
Number of observations		250,646			5,841	
Number of children	–0.044** (–29.06)	0.824** (50.51)	–0.026† (–1.85)	–0.023** (–2.31)	0.680** (7.89)	–0.050 (–0.50)
First and second children in families with three or more births						
Number of observations		200,538			4,403	
Number of children	–0.038** (–21.38)	0.856** (50.81)	–0.021 (–1.49)	–0.012** (–0.93)	0.803** (8.88)	–0.049 (–0.55)
Second child	–0.030** (–16.78)		–0.032** (–11.70)	0.004 (0.39)		0.009 (0.58)
Number of children (control for spacing)	–0.040** (–22.26)	0.885** (53.38)	–0.020 (–1.44)	–0.012** (–0.96)	0.790** (8.95)	–0.049 (–0.56)
Second child (control for spacing)	–0.028** (–15.73)		–0.032** (–10.71)	0.005 (0.42)		0.011 (0.59)

Notes: Robust t statistics, which allow for correlation of errors within family, are shown in parentheses. All regressions include age, age squared, age cubed, indicators for male and Han, parents' age and age squared, parents' educational level, and provincial dummy variables.

[a]1 = illiterate; 2 = primary school; and 3 = junior high school or above.

†$p < .10$; *$p < .05$; **$p < .01$

given birth quotas, and they are penalized for "above-quota births." Parents with above-quota children are forced to pay for each additional birth and may be subject to other punishment or criticism. In contrast, parents who comply with the one-child policy receive cash subsidies from the government, and their children can receive free health care, such as immunizations.

However, the local implementations of this policy demonstrate great heterogeneity, especially between rural and urban areas. In general, the penalties for above-quota births are much more severe in the urban areas than in the rural areas (Banister 1987). Urban citizens who violate the policy have to pay fines that are proportional to their monthly salaries, sometimes as high as 70%. They are demoted or rendered ineligible for promotion forever if they work in state-owned enterprises or institutions, which were the major urban employers in the 1980s. In contrast, the only severe punishment in rural areas is a one-shot payment for above-quota births, and the payment may not be very effective in rural areas because many poor farmers cannot afford to pay it (Li and Zhang 2004). Because of the difficulty in implementations and the potential for social unrest, in some rural areas and in certain years, the policy was relaxed to allow people to have second children if the first was female (Chow 2002; Qian 1997).

Given that the one-child policy has been enforced more strictly in urban China, one may argue that parents who have above-quota children are inherently different from those who comply with the birth control policy by having fewer children. This may explain why the quantity-quality trade-off is not observed in our urban sample. For example, richer families who are able and willing to pay fines to have additional children can invest more per child anyway. Likewise, parents may choose to have fewer children not because they want to trade quantity for quality but because they are not allowed to have more.

To address this problem, we attempt to control for family preferences to some extent by restricting the sample to families with at least *n* births in previous estimations. In another check, we redo the analysis using the 1982 census data. Because all the sampled children (aged 6 and above) in 1982 were born before 1979, the impact of the one-child policy, if any, should be minimal. Table 5 replicates the regressions in Table 4 using the rural and urban subsamples from the 1982 census.[19] Although the OLS coefficients on family size are closer between the two subsamples, none of the 2SLS estimates for the urban sample is significantly different from zero. However, for the rural families, the 2SLS estimates in the middle ($n = 2$) and bottom ($n = 3$) panels show some evidence of a negative effect of family size on children's educational level. The results in Table 5 suggest that, even in absence of a (potentially) large effect of birth control policy, we are still unable to find a quantity-quality trade-off in the urban sample. This implies that our results presented in the Effects in Rural and Urban Areas section are not largely driven by the birth control policy.

The Heterogeneous Effects of Family Size

In this section, we test the sensitivity of our estimates to more stratification of the sample. Specifically, we estimate the effect of family size by child gender and by mother's education. Because the effect has been shown to differ between rural and urban areas, we skip the estimations for the full sample and perform the sensitivity test for the rural and urban samples separately. The upper and lower panels in Table 6 report the OLS and 2SLS coefficients on family size for rural and urban samples, respectively.

In the first two columns, we break the samples down by gender to see whether the effect of family size differs between boys and girls. Although the OLS estimates show that

19. Because the 1982 census does not include an explicit rural identifier, we use the occupation code to define rural children as those whose parents were engaged in a broad range of agricultural business. Although this categorization may understate the rural population (77% in 1982 compared with 88% in 1990), it is the best approximation we can make. To see whether this would lead to a severe problem, we reestimated the 1990 sample using the occupation-based rural identifier, and our results were not significantly changed.

Table 5. Ordinary Least Squares (OLS) and Two-Stage Least Squares (2SLS) Estimates of the Effect of Family Size on Children's Educational Level: 1982 Chinese Population Census (rural vs. urban)

	Dependent Variable: Educational Level[a]					
	Rural			Urban		
Independent Variable	OLS (1)	First Stage (2)	2SLS (3)	OLS (4)	First Stage (5)	2SLS (6)
Twins at the First Delivery						
All nontwin children						
Number of observations		530,596			158,976	
Number of children	–0.040**	0.348**	–0.024	–0.033**	0.369**	0.014
	(–52.29)	(17.26)	(–0.70)	(–25.29)	(11.61)	(0.29)
Twins at the Second Delivery						
Nontwin children in families with two or more births						
Number of observations		521,180			151,276	
Number of children	–0.042**	0.567**	–0.062**	–0.038**	0.610**	–0.032
	(–53.20)	(29.51)	(–3.10)	(–27.25)	(20.74)	(–1.17)
First children in families with two or more births						
Number of observations		249,495			75,735	
Number of children	–0.037**	0.630**	–0.024	–0.039**	0.723**	–0.010
	(–29.78)	(24.56)	(–1.05)	(–17.17)	(19.46)	(–0.34)
Number of children (control for spacing)	–0.052**	0.681**	–0.014	–0.049**	0.773**	–0.004
	(–40.47)	(28.00)	(–0.69)	(–20.78)	(21.52)	(–0.12)
Twins at the Third Delivery						
Nontwin children in families with three or more births						
Number of observations		403,746			92,828	
Number of children	–0.049**	0.689**	–0.036*	–0.050**	0.924**	0.007
	(–44.63)	(38.80)	(–2.07)	(–21.31)	(31.04)	(0.32)
First and second children in families with three or more births						
Number of observations		301,237			68,285	
Number of children	–0.049**	0.742**	–0.028	–0.055**	0.906**	0.020
	(–34.30)	(40.17)	(–1.59)	(–18.09)	(28.03)	(0.76)
Second child	–0.028**		–0.035**	–0.022**		–0.040**
	(–16.24)		(–5.67)	(–6.49)		(–5.48)
Number of children (control for spacing)	–0.057**	0.785**	–0.023	–0.060**	0.922**	0.020
	(–38.82)	(44.05)	(–1.38)	(–19.30)	(29.39)	(0.78)
Second child (control for spacing)	–0.024**		–0.036**	–0.019**		–0.040**
	(–13.86)		(–5.92)	(–5.68)		(–5.28)

Notes: Robust *t* statistics, which allow for correlation of errors within family, are shown in parentheses. All regressions include age, age squared, age cubed, indicators for male and Han, parents' age and age squared, parents' educational level, and provincial dummy variables.

[a]1 = illiterate; 2 = primary school; and 3 = junior high school or above.

*$p < .05$; **$p < .01$

Table 6. Ordinary Least Squares (OLS) and Two-Stage Least Squares (2SLS) Estimates of the Effect of Family Size on Children's Educational Level by Gender and Mother's Education: 1990 Chinese Population Census (rural vs. urban)

	Dependent Variable: Educational Level[a]				
	Gender		Mother's Education		
Independent Variable	Male (1)	Female (2)	Low (3)	Median (4)	High (5)
Rural Sample					
Twin at first delivery					
OLS	–0.024**	–0.035**	–0.051**	–0.021**	–0.016**
	(–26.39)	(–37.86)	(–38.10)	(–22.88)	(–11.96)
2SLS	–0.022	–0.061**	0.021	–0.018	–0.141**
	(–0.93)	(–2.79)	(0.64)	(–0.73)	(–4.24)
Twin at second delivery					
OLS	–0.033**	–0.041**	–0.057**	–0.026**	–0.024**
	(–29.36)	(–39.54)	(–38.36)	(–24.20)	(–14.21)
2SLS	–0.025†	–0.001	–0.028	–0.009	–0.004
	(–1.87)	(–0.09)	(–1.04)	(–0.63)	(–0.24)
Twin at third delivery					
OLS	–0.038**	–0.046**	–0.061**	–0.029**	–0.024**
	(–17.08)	(–25.19)	(–23.57)	(–14.63)	(–6.81)
2SLS	–0.045*	–0.013	–0.061*	–0.011	–0.006
	(–2.24)	(–0.74)	(–2.09)	(–0.55)	(–0.26)
Urban Sample					
Twin at first delivery					
OLS	0.000	–0.006*	–0.017**	–0.001	0.013**
	(0.05)	(–2.08)	(–4.05)	(–0.19)	(2.88)
2SLS	–0.019	–0.027	–0.078	–0.038	0.007
	(–0.67)	(–1.02)	(–0.98)	(–1.08)	(0.28)
Twin at second delivery					
OLS	–0.018**	–0.026**	–0.026**	–0.019**	–0.009
	(–2.82)	(–4.65)	(–3.86)	(–2.79)	(–0.79)
2SLS	0.021	–0.026	–0.040	0.036	0.020
	(0.45)	(–0.69)	(–0.77)	(0.61)	(0.39)
Twin at third delivery					
OLS	–0.022	–0.026*	–0.024†	–0.018	–0.020
	(–1.57)	(–2.22)	(–1.74)	(–1.18)	(–0.73)
2SLS	–0.061	–0.034	–0.118	0.060	–0.164
	(–0.50)	(–0.25)	(–0.68)	(0.38)	(–0.81)

Notes: Robust *t* statistics, which allow for correlation of errors within family, are shown in parentheses. All regressions include age, age squared, age cubed, indicators for male and Han, parents' age and age squared, parents' educational level, and provincial dummy variables. The low, median, and high levels of mother's education refer, respectively, to illiterate, primary school, and above primary school for the rural sample; they refer to below junior high school, junior high school, and above junior high school for the urban sample.

[a] 1 = illiterate; 2 = primary school; and 3 = junior high school or above.

† $p < .10$; * $p < .05$; ** $p < .01$

the effect of family size is more negative for girls than for boys, the picture from the 2SLS estimates is not as clear. The effect appears to be more pronounced for rural girls when we use the IV of twins at the first delivery but becomes larger for rural boys in families with at least three births. Despite the mixed results for the 2SLS estimations for the rural sample, we continue to identify a rural-urban gap that is independent of gender—namely, a smaller effect of family size in urban areas.

In the last three columns, we stratify our sample by mother's educational level. Household income is not observed in our sample, so we use mother's education as a control for financial constraints. If better-educated mothers are less financially constrained, then we should see a smaller effect of family size on the educational outcomes of their children. We categorize mother's education in the rural sample as illiterate, primary school, or all the other levels above primary school. Because urban women are generally better educated than rural women, we categorize urban women's education as below junior high school (illiterate and primary school), junior high school, and above junior high school to avoid a group with too few observations.

To some extent, the results by educational group are consistent with our expectations. With the OLS estimates, the effect of family size decreases in magnitude with the level of the mother's education for both rural and urban children, although a few OLS coefficients for the urban sample are not statistically significant. However, the evidence is less explicit when we look at the 2SLS estimates. Again, for the rural sample, the variation in effects across educational groups depends on the IV (and the sample) that we use. For the urban sample, we do not detect a tangible effect of family size for any subgroup; all the 2SLS estimates are statistically insignificant.

CONCLUSIONS

In this paper, we test the theory of quantity-quality trade-off of children by using a representative census data set from China. We find evidence that family size is negatively correlated with children's education. The negative effect of family size is robust to various specifications, including those that control for parental characteristics and birth order effect. We then instrument family size with twin births to explore the causal link between family size and child education and find supportive evidence. We further find that the effect of quantity on quality is not uniform between rural and urban areas. More precisely, the trade-off relationship is more evident in rural China, but the effect diminishes or even vanishes for urban China. We also find that the effect differs according to the gender of the child and the mother's educational level. Given that our estimates are probably upwardly biased toward zero due to the direct effects of twin births on child outcome through mechanisms other than family size, our results provide the lower bound of the negative effect and indeed suggest a quantity-quality trade-off.

Overall, our findings evidently support the prediction of Becker (1960) and Becker and Lewis (1973) of the quantity-quality trade-off of children, but differ from those of Black et al. (2005). The most important difference between our study and that of Black et al. (2005) is that they drew on data from Norway, which is a developed country, whereas we draw on data from China, which is a developing country. In a developed country like Norway, with a comprehensive welfare system and both a good public education system and generous government support for childbearing and childcare, the quantity-quality trade-off may not be obvious. However, in a developing country like China, where there is neither a good public education system nor generous support for childbearing and childcare, the cost of child quality is mostly borne by the parents. Thus, a quantity-quality trade-off is more likely in the Chinese case.

Although this study has limitations, it is among the first to explicitly measure the effect of family size on child outcome in China. Previous empirical tests were often limited by a small sample size or by the fact that they did not take into account the endogeneity of family size; we overcame both of these limitations in this paper. Given that public education is insufficiently funded in many areas of China, our findings suggest a plausible determinant of children's education in China that has not been well explored in the literature. Nonetheless, due to the data limitations, we are unable to examine more aspects of child quality (such as health and labor market outcomes) and are thus ill inclined to generalize our results to a broader extent. Future work may rely on more comprehensive and traceable household data that give researchers information on the completed education of children even if they have left the family.

This research may shed some light on other issues in China, such as the one-child policy. Since its inception in the late 1970s, China's one-child policy has been controversial and has drawn attention from politicians, the mass media, and academics alike. Although there is still no consensus on many of the positive or negative aspects of this forced birth-control policy, a recent study by Li and Zhang (2007) showed that the population reduction as a result of the dramatic population control policy has indeed helped the growth of the Chinese economy since the late 1970s. This study indicates that a possible effect may be that children are of better quality under the policy because the size of their families would have been larger if the policy had not existed. However, to better understand the long-term effect on child outcomes in adulthood, more work is badly needed in this area.

Appendix Table A1. Ordinary Least Squares Estimates of the Effect of Family Size and Birth Spacing on Children's Educational Level: 1982 and 1990 Chinese Population Census

	Dependent Variable: Educational Level[a]					
	1982			1990		
Independent Variable	Full Sample (1)	Rural Sample (2)	Urban Sample (3)	Full Sample (4)	Rural Sample (5)	Urban Sample (6)
First Children in Nontwin Families With Three or More Births						
Number of observations	202,295	165,819	36,476	120,291	117,885	2,406
Number of children	–0.058**	–0.054**	–0.066**	–0.043**	–0.042**	–0.029†
	(–31.99)	(–26.96)	(–15.01)	(–18.49)	(–18.21)	(–1.70)
Age gap (year) between the two following births	–0.010**	–0.011**	–0.007**	–0.004**	–0.004**	0.001
	(–13.62)	(–13.05)	(–4.31)	(–4.65)	(–4.45)	(0.26)
First and Second Children in Nontwin Families With Four or More Births						
Number of observations	143,930	122,602	21,328	51,977	50,916	1,061
Number of children	–0.061**	–0.057**	–0.071**	–0.038**	–0.037**	–0.076*
	(–21.86)	(–19.19)	(–9.49)	(–8.60)	(–8.38)	(–2.44)
Age gap (year) between the two following births	–0.015**	–0.015**	–0.011**	–0.008**	–0.008**	–0.008
	(–14.40)	(–13.60)	(–4.20)	(–5.12)	(–4.97)	(–0.91)
Second child	–0.038**	–0.040**	–0.030**	–0.033**	–0.034**	0.030
	(–13.77)	(–13.16)	(–4.41)	(–8.98)	(–9.40)	(1.17)

Notes: Robust t statistics, which allow for correlation of errors within family, are shown in parentheses. All regressions include age, age squared, age cubed, indicators for male and Han, parents' age and age squared, parents' educational level, and rural and provincial dummy variables.

[a]1 = illiterate; 2 = primary school; and 3 = junior high school or above.

†$p < .10$; *$p < .05$; **$p < .01$

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