Quantity-Quality: The Positive Effect of Family Size on School Enrollment in China (Incomplete)

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Abstract

Many policy makers in developing countries see restricting family size as a good strategy for increasing average human capital investment. This belief is consistent with the observed negative correlation between quantity and quality of children both across countries and across households within countries. However, because parents simultaneously choose the quantity and quality of their children, the observed correlation may reflect parental preferences rather than the causal relationship of quantity on quality. In addition, the decision to have a second child may be positively correlated with the quality of the first child. This paper exploits exogenous changes in family size caused by relaxations in China's One Child Policy to estimate the effect of family size on school enrollment. Specifically, it uses the relaxation that allows a rural household to have a second child if the first is a girl. First, it shows that the "1-son-2-child" rule increased family size for first born girls. Second, it uses the exogenous increase in family size to find that an additional sibling significantly *increases* school enrollment of the first child. Furthermore, this paper shows that the One Child Policy dramatically increased sex-imbalance in certain areas. (*JEL I2, I20, I21, J13, J16, J24, O10, R2*)

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1 Introduction

The trade-off between quantity and quality of children has been a question of long standing interest in labor economics. Understanding this tradeoff is especially relevant to developing countries today as policy makers in these countries attempt to curb population growth as a way of increasing average human capital investment. Both China and India, the world's two most populous countries, have experimented with different family planning policies to limit family size. Examining the trade-off between quantity and quality of children is of first-order importance for evaluating the effects of past policies as well as for constructing effective future ones.

Empirical findings on the quantity-quality trade-off are conflicted. On one hand, the effect of family size on education has been found to be negative by Rosenzweig and Wolpin (1980) in India; by Goux and Maurin (2004) in France; by Conley (2004), Berhman et. al. (1989) and Stafford (1987) in the U.S. On the other hand, studies by Lee (2003) in Korea, Kessler (1991) and Guo and VanWey (1991) in the U.S. have found that family size has no effect on education while Gomes (1984) found that family size was positively correlated with education attainment for first born children in Kenya. Adding to the controversy, in studies using data from Norway and the U.K., Black et. al. (2004) and Iacavou (2004) use dummy variables for family size instead of the traditional continuous variable and found that while family size and education outcomes are negatively correlated for children from households with two or more children, children from one-child families perform worse than children from two-child and three-child families. Black et. al. (2004) estimates that only-children, on average, attain 0.21 years less of schooling than other children.

The main empirical challenge in estimating the effect of family size on child outcomes is caused by two sources of endogeneity. The first source arises from parental heterogeneity. For example, if parents who value education more also prefer to have fewer children, the correlation between quantity and quality will be driven by parental preferences rather than by family size. To address this problem of joint determination, past studies have exploited the exogenous variation in family size caused by multiple births or the sex composition of the first two children (Rosenzweig and Wolpin, 1980; Conley, 2004; Lee, 2003). However, both instruments fail the exclusion restriction since they affect child outcomes other than family size. In a study of Indonesia, Duflo (1998) found that twin births of younger siblings were correlated with higher mortality rates of the first born child. She argued that short birth spacing may be a channel through which an increased number of children lower their average quality; the strain on resources is especially problematic if the household is credit constrained. The sibling-sex instrument is equally problematic: Dahl and Moretti (2004) and Ananat and Michaels (2004) find that sibling sex composition directly affect divorce rates. Using sibling sex composition has the additional limitation that it requires the sex of children to be randomly assigned, and consequently, it cannot be used in a country with sex-selection such as China (Qian, 2004).[1]

The second source of endogeneity arises from heterogeniety in the quality of the first child. For example, if parents are more likely to have a second child when the first child is of high quality, the OLS estimate of the family size effect will be biased upwards.

This principal contribution of this paper is to address the two sources of endogeneity by exploiting regional and time variation in relaxations of China's One Child Policy. Specifically, it uses the relaxation that allows families to have a second child if the first child is a girl to instrument for the family size of first born children born before the relaxation was announced. Three facts are exploited: first, an individual is only affected by the relaxation if she is born in a relaxed area; second, amongst first born children born in relaxed areas, only girls are affected; and third, amongst first born girls born in relaxed areas, the relaxation only affected girls whose family size were constrained by the initial One Child Policy (born in 1976 or after).[2] The instrument for family size is the triple interaction term of an individual's sex, date of birth and region of birth. The interaction between whether a girl was born in a relaxed area and whether

[1] Instead of looking for exogenous variation in family size, Guo and VanWey (1991) and Black et. al. (2004) attempt to control for the unobserved differences across households by controlling for household fixed effects in panel data. However, fixed effects estimates are biased if unobserved household-level heterogeneity are time varying.

[2] The One Child Policy began in 1978-1980. However, prior to that were policies which encouraged birth spacing of at least four years between children. I show that the One Child Policy was actually binding for the family size of cohorts born in 1976 and after.

she was born in 1976 or after estimates the effect of the relaxation of family size. The additional comparison with boys controls for region specific changes in school provision that affected boys and girls similarly.

There are two main benefits in setting this study in China. First, family planning policies provide a unique source of exogenous variation in family size. Second, this study can evaluate the effects of the One Child Policy, one of the most restrictive and large scale family planning policies ever undertaken. While demographers and sociologists have conducted descriptive studies of the policy's impact on fertility, the lack of local enforcement data has prevented an examination of the causal effect of the One Child Policy on child outcomes.

The empirical findings show that the One Child Policy *decreased* the fraction of girls amongst first born children in the surviving population by up to 10 percentage-points in certain regions. This increase in sex imbalance mostly reflects an increase in excess female mortality since pre-natal sex revealing technologies which enable selective abortion were not introduced to the population in this study until the end of this period (Zeng et. al., 1993). The results show that consistent with official reports, the 1-Son-2-Child relaxation was implemented in communities that experienced larger increases in boy-biased sex selection after the One Child Policy. The relaxation immediately decreased the level of sex selection although sex ratios did not return to their initial pre-One Child Policy levels. The difference in levels of sex selection between relaxed and un-relaxed regions suggests that parents who kept girls in relaxed regions are on average different from parents who kept girls in un-relaxed regions. Comparisons between these two types of regions will therefore suffer from selection bias. In particular, if parents of girls born in regions with the relaxation value education for girls more than parents of girls born in other regions, the two stage least squares estimate using the triple interaction term to instrument for family size will confound the family size effect with parental preferences and be biased upwards. To address this, I will estimate a lower-bound of the absolute value of the effect of family size on school enrollment.

The first stage results show that the relaxation increased family size for first born girls and had no effect on the family size of first born boys. The two stage least squares estimates show that an additional sibling *increases* school enrollment of the first born child by 18-20% on average. This finding can be explained by a model where there are fixed costs in education or a model where other children are complements in each child's production function. The plausibility of the latter hypothesis is strengthened by the finding that the family size effect varies by the age gap between the two children.

Although more research is needed to understand the effect of quantity on quality outside the one-child context, the empirical findings of this paper suggest that there is a strong only-child disadvantage and consequently reject the hypothesis that quality is monotonically decreasing with quantity. The results cast doubt on the idea that restricting family size will necessarily help to increase average human capital investment in developing countries. Policy makers should weigh the benefits of implementing restrictive family planning policies against the substantial "costs" related to sex selection and the long run consequences of the resulting sex imbalance. In addition, policy makers who wish to restrict family size to one child should consider implementing programs which increase interaction between children.

The paper is organized as follows. Section 2 discusses family planning policies, education in rural China and the conceptual framework. Section 3 describes the data. Section 4 presents the empirical results. Section 5 offers interpretation of the results and concluding remarks.

2 Background

2.1 Family Planning Policies

In the 1970s, after two decades of explicitly encouraging population growth, policy makers in China enacted a series of measures to curb population growth. The policies applied to individuals of Han ethnicity, who comprise 92% of China's population. Beginning around 1972, the policy "Later [age], longer [the spacing of births], fewer [number of children]" gave economic incentives to parents to space the birth of their children at least four years apart. The One Child Policy was formally announced in 1979. Actual implementation began in certain regions as early as 1978 and enforcement hardened across the country until the policy was firmly in place in 1980 (Croll et. al., 1985; Banister, 1987). Past studies generally consider the One Child Policy to have only affected the family size of cohorts born after 1978-1980. However, this paper will show that because of the previous four year birth spacing rule, the One Child Policy affected cohorts born in 1976 and after.

Policy tightened gradually and second births became forbidden except under extenuating circumstances. Local cadres were given economic incentives to suppress fertility rates. In the early 1980s, parts of the country were swept by campaigns of forced abortion and sterilization and reports of female infanticide became widespread (Greenlaugh, 1986; Banister, 1987).

Local governments began issuing permits for a second child as early as 1982. However, permits for a second child were not made widespread until the Central Party Committee issued "Document 7" on April 13, 1984. The two main purposes of the document were to: 1) curb female infanticide, forced abortion and forced sterilization; and 2) devolve responsibility from the central government to the local and provincial government so that local conditions can be better addressed. It asked cadres to deal with each case individually and move away from inflexible, uniform enforcement. The document allowed for second births for rural couples with "practical" difficulties, and strictly prohibited coercive methods (Greenlaugh, 1986). The main relaxation following Document 7 is called the "1-son-2-child" rule. It allows rural couples to have a second child if the first child was a girl (Greenlaugh, 1986).[3] The explicit purpose of this relaxation was to decrease female infanticide of the first born child.

White (1992) found that 5% of rural households were allotted second child permits in 1982. These permits were generally granted to regions with extremely high levels of infanticide. After Document 7, the permits expanded to 10% of the rural population in 1984, 20% in 1985 and 50% by 1986.

Document 7 made provincial governments responsible for both maintaining low fertility rates and decreasing infanticide. While the exact process of granting permits is unclear, I use county level data on family planning policy to show in

[3] Practical difficulties included households where a parent or first born child was handicapped, or if a parent was engaged in a dangerous industry (e.g. mining).

the next section that the probability for a county to obtain the 1-son-2-child relaxation is positively correlated with the rate of pre-relaxation sex selection, and both are positively correlated with distance from the provincial capital. These facts most likely reflect that in order to maintain low aggregate fertility rates and decrease female infanticide, provincial governments granted relaxations to regions that were distant to the administrative capital, and where female infanticide was more prevalent. The higher prevalence of sex selection in rural areas can be due to both more boy-preference in distant rural areas and the fact that geographic distance increases the provincial government's difficulty of preventing infanticide.[4] Issues of identification that arise from the correlation of obtaining a relaxation and sex selection will be addressed explicitly in section 4.

2.2 Rural Education

Inequality in education provision greatly increased during the 1980s both across provinces and across counties within a province. Inequality between school finance increased as changes in the fiscal system reduced subsidies from rich regions to poor regions. The system of "eating from separate pots" (*fen zou chi fang*) devolved expenditure responsibilities from the central and provincial governments onto local governments in order to give the latter stronger incentives to generate revenue. The ratio of the per capita schooling expenditure in the highest spending province to the lowest spending province doubled in one decade.

Many rural schools were closed; rural enrollment rates dropped dramatically and did not recover until the mid to late 1990s (Hannum and Park, mimeo). Using spending data from Gansu, Hannum and Park (mimeo) found that per capita school expenditure was positively correlated with income and significant variation in school quality existed across counties. They found little variation within counties, suggesting that studies examining education outcomes should focus on variation at the county level.

[4] Levels of income between counties with some relaxation and counties with no relaxation are comparable in the CHNS data. This is consistent with the findings of Qian's (2004) study of rural China, where she finds that sex selection was driven by the female-to-male income ratio and not by total household income.

Hannum (1992) show that difference in school provision between rich and poor areas are much greater for middle school and high school than primary school. This is consistent with the CHNS data used in this study, where primary school enrollment remained stable while middle school and high school dropout rates increased for poor areas (Hannum and Park, mimeo).

The CHNS data show that counties with some relaxation and counties with no relaxation have similar geographic access to schooling in 1989. However, the data does not reveal quality of schooling or the changes in school availability during the early 1980s. Because relaxed areas tend to be more rural, it is likely that the quality of schools declined in relaxed areas during the same time that the 1-son-2-child relaxation took effect. To control for this, I will compare outcomes for girls to boys within counties. The strategy is robust as long as the changes in school quality and the economic conditions that determine school quality in relaxed areas have the same impact on boys and girls.

2.3 Conceptual Framework

There are two models in the economics and sociology literature that predict an interaction between the quantity and quality of children. The quantity-quality model, known in sociology as the "resource dilution" model, dates backs to Becker (1960), Becker and Lewis (1973) and Becker and Tomes (1979). They theorized that when income increases, parents who prefer that children within a household have equal quality will want to increase the average quality of their children. Their model predicts that quality monotonically decreases with quantity. An alternative model is the "confluence model", which to date, has not been explored in the economics literature. Psychologists Zajonc and Gregory (1985) argue that children benefit from interacting with adults and teaching younger children. The quantity and quality of children are inversely related because increasing the number of children decreases the adult-to-children ratio within a household. At the same time, children from one-child families and the youngest child from a multi-child family are worse off because they cannot take advantage of the learning which comes from teaching younger children. This model, therefore, predicts an inverse "U" shape for the relationship between quantity and quality of children. This is consistent with findings from Iacavou's (2004) study of children in the U.K. She finds that although general family size is negatively correlated with measures of school performance, first born children from one-child families perform worse than first born children from two-child families. Moreover, the only-child effect decreases for children who interact more with other children outside of school.

3 Data

This paper matches data from the 0.1% 1990 *Population Census* with data from the 1989 *China Health and Nutritional Survey* (CHNS) at the county level. The 1990 *Population Census* contains 52 variables including birth year, region of residence, whether an individual currently lives in his/her region of birth, sex and relationship to the head of the household. The data allows children to be linked to parents. Thus, family size and birth order of children within a household can be calculated. Because the identification is partially derived from the region of birth, the sample is restricted to individuals who reported living in their birth place in 1990. The CHNS uses a random cluster process to draw a sample of approximately 3,800 households with a total of 16,000 individuals in eight provinces that vary substantially in geography, economic development, public resources, and health indicators. Most importantly, the survey provides detailed village and township level information on family policy enforcement. Since ethnic minorities were exempt from all family planning policies, I restrict the analysis to four provinces which are mostly composed of individuals of Han ethnicity. The matched dataset contains 21 counties in four provinces.[5] These provinces exclude rich coastal provinces or poor interior provinces.

For the analysis of family size and education, the sample is restricted to first born children in cohorts born during 1972-1981. This has two main advantages. First, all children in the sample have access to public schooling in 1990. Second, including children born after the relaxation may induce bias in the 2SLS estimate. After the relaxation, parents who prefer larger families may choose to keep girls. This means that the 2SLS estimate will show that girls with

[5] Liaoning, Jiangsu, Shandong and Henan.

larger family size are better off. But the estimate will be partially driven by parental preferences. Exclusion of first born children born after 1981 removes this possibility.

The descriptive statistics in Table 1 Panel A show that counties with no relaxation are very similar along demographic characteristics to counties with some relaxation. Each has 52% boys on average and is mainly composed of ethnic Hans. Children in relaxed counties have on average one more sibling than children from counties without the relaxation. Approximately 65% of children are enrolled in school.

The data shows that counties with some relaxation are almost four times as far from the provincial capital as counties with no relaxation. Distance to school is similar between the two types of counties.

Panel B of Table 1 describes the data for first born children from one-child families and from families with two or more children. 47% of children in multi-child families are boys while 60% of one-child families are boys. Children without siblings are on average enrolled in school 12% more than children with at least one sibling.

4 Empirical Framework

4.1 Identification

Sex, date and region of birth jointly determine an individual's exposure to the 1-Son-2-Child relaxation. The relaxation allowed parents to have a second child only if the first born child was a girl. Therefore, family size should be positively correlated with being a girl. The One Child Policy, introduced around 1980, followed family planning polices which encouraged birth spacing of at least four years. Consequently, the relaxation should only affect girls born 1976 or after.

The interaction between whether a girl was born in a relaxed area and whether she was born 1976 or after estimates the effect of the relaxation on family size. The additional comparison with boys controls for education provision changes that affected both boys and girls similarly. The instrument for family size is the triple interaction of an individual's sex, date and region of birth. Only the combination of the three is exogenous. The exclusion restriction for the instrument is that it must be correlated with family size and have no direct effect on school enrollment or other right hand side variables.

Like simple differences-in-differences estimators, cohort-invariant differences across regions are differenced out by the comparison across cohorts. Changes across cohorts which affect different regions similarly are differenced out by the comparison across regions. The triple difference adds the advantage that cohort varying differences that affect boys and girls similarly across regions are also differenced out by the comparison between girls and boys within each cohort and region. The exclusion restriction is only violated if a change with differential impacts on relaxed and un-relaxed regions *and* on boys and girls occurs at the same time the relaxation took effect. In other words, the 2SLS estimate will be biased only if there is a sex-specific, region-specific change for the treated cohort.

I find in the next section that consistent with official reports, the extent of the relaxation is strongly correlated with the extent of sex selection for One Child Policy cohorts (1976-1982). The determinants of sex-selection may also affect education investment differentially for boys and girls. For example, Qian (2004) shows that increasing male-to-female earnings increase boy-biased sex-selection. She also shows that increasing male-to-female earnings has no effect on education investment for boys but decreases education investment for girls. This means that sex-selection is correlated with lower education investment for girls relative to boys. This will not bias the estimates as long as the correlation is time invariant, in which case it will be differenced out by the before and after comparison.[6]

A potential source of bias introduced by the One Child Policy is the selection of parents who choose to keep girls. Parents who choose to keep girls born during 1976-1982 in relaxed counties may have different preferences from parents who keep girls in counties without the relaxation. For example, if parents who decide

[6] The CHNS does not have accurate data on individual income within the household since much of rural production is conducted at the household level and income cannot be accurately assigned to individual members. Consequently, I cannot directly examine the role of relative earnings in this study.

to keep girls in relaxed counties also value education more than parents who keep girls in non-relaxed counties, the 2SLS estimate will overestimate the true effect of family size on school enrollment.

To address the problem of sample selection, I construct an alternative sample where the "extra" boys from relaxed counties in the actual sample are taken out and replaced with girls so that for each cohort, the sex ratio is equivalent between counties with some relaxation and counties without any relaxation. In order to estimate the lower bound of the absolute value of the effect of family size on school enrollment, I remove only boys who are not enrolled in school and add girls who are not in enrolled in school. This increases the average enrollment rate for boys born 1976-1982 in counties with the relaxation, and decreases average enrollment rate for girls in counties with the relaxation. 2SLS using this "stacked" sample will underestimate the true effect of family size on school enrollment. Thus, using the actual sample and the constructed sample, I will be able to estimate the upper and lower bounds of the absolute value of the family size effect.

4.2 The Effect of the 1-Son-2-Child Relaxation

4.2.1 Effect on Family Size

One benefit of this policy experiment is that it is possible to check whether the policy was enforced correctly by estimating the effect of the policy on family size for boys and girls separately. If the policy was correctly enforced, it should increase the number of siblings for girls born 1976 and after and have no effect on boys. The following equation separately estimates the effect of the relaxation on family size for boys and girls born during 1962-1981.

$$sibs_{itc} = \sum_{l=1973}^{1981} (relax_c \times d_{il})\beta_l + \gamma_t + \alpha + \psi_c + \upsilon_{itc} \tag{1}$$

The number of siblings for individual i, born in county c, birth year t, is a function of: the interaction term of $relax_c$, the extent of relaxation in county c and d_{il}, a dummy indicating whether the individual was born in year l; γ_t, birth year fixed effects and ψ_c, county fixed effects. The reference group is comprised of individuals born during 1962-1972. It and all of its interaction terms are dropped. For all regressions, standard errors are clustered at the county level.

β_l is the effect of being born in a relaxed county on family size for an individual born in year l. The estimates for girls and boys are shown in Table 2, columns (1) and (2). The estimates for girls are statistically significant at the 1% level for the affected cohorts (born 1976-1981). The estimates for boys are statistically insignificant. The coefficients are plotted in Figure 2A. It shows that family size for boys and girls were similar for cohorts born 1973-1976, after which the family size for girls increased and the family size for boys remained the same.

This difference in the effect of the relaxation on family size between boys and girls can be written as the interaction between sex, date of birth and region of birth.

$$
\begin{aligned}
sibs_{itc} = & \sum_{l=1973}^{1981} (relax_c \times girl_{itc} \times d_{il})\beta_l + \sum_{l=1973}^{1981} (relax_c \times d_{il})\delta_l \\
& + \sum_{l=1973}^{1981} (girl_{itc} \times d_{il})\zeta_l + (relax_c \times girl_{itc})\lambda + girl_{itc}\kappa \\
& + \alpha + \gamma_t + \psi_c + v_{itc}
\end{aligned} \tag{2}
$$

The number of siblings for individual i, born in county c, birth year t, is a function of: the triple interaction term of $relax_c$, the extent of relaxation in county c, $girl_{itc}$, a variable indicating whether a child is a girl and d_{il}, a dummy indicating whether the individual was born in year l; the interaction term of $relax_c$ and d_{il}; the interaction term between $girl_{itc}$, and d_{il}; the interaction term between $relax_c$ and $girl_{itc}$; $girl_{itc}$; γ_t, birth year fixed effects; and ψ_c, county fixed effects. As before, the reference group of cohorts born 1962-1972 and all its interactions are dropped. β_l is the difference in the effect of being born in a relaxed areas on family size between girls and boys. The estimates should be zero for cohorts who were not affected by the One Child Policy and relaxation (1973-1976) and positive for affected cohorts (1976-1981). The coefficients are shown in Table 2, column (5). They are statistically significant at the 5% level for the effected cohorts. Figure 2A plots the coefficients. It shows that the difference in the effect of being born in a relaxed area on family size is zero for unaffected cohorts and positive for the affected cohorts. The relaxation increased family size of first born girls by approximately 0.25 children on average.

4.2.2 Effect on Sex Ratios by Birth Parity

This section evaluates the effect of the relaxation on sex ratios by birth parity. To observe the effect of the relaxation on sex ratios, the sample must be expanded to include cohorts born after the relaxation. Past studies comparing hospital birth records and population census data, or by comparing sex ratios for the same cohort at different ages have found that sex selection mostly occurs at very young ages, which is consistent with the lack of prenatal gender revealing technology and tough government enforcement against infanticide (Qian, 2004; Zeng et. al., 1993). Hence, any sex selection caused by the One-Child Policy should be observed for cohorts born very close to 1980. I estimate the following equation using a sample of cohorts born between 1962 and 1989 by birth order. Because of widespread under reporting of children under one year of age, I exclude the 1990 cohort (Zeng, 1992). The reference cohort is composed of individuals born during 1962-1968.

$$male_{itc} = \sum_{l=1969}^{1989} (relax_c \times d_{il})\beta_l + \gamma_t + \alpha + \psi_c + \upsilon_{itc} \tag{3}$$

This equation is similar to (1). The dependent variable indicates whether an individual is male. Table 3 column (1) shows the estimates of β_l for first born children. They are statistically significant. Column (2) shows that the estimates are robust to the addition of a control for whether individuals are ethnically Han. Columns (3) and (4) show the estimates for second born children. Columns (5) and (6) show the estimates for children of higher birth parity. The coefficients for first, second and later born children from columns (1), (3) and (5) are plotted in figures 2A, 2B and 2C with their 95% confidence intervals. The solid vertical line in the figures indicates the beginning of the [initial] One Child Policy in 1978. The dashed line indicates the beginning of the relaxation in 1982. Figure 2A shows that in areas that received the relaxation, the fraction of males increased after the One Child Policy relative to other areas. It also shows that the relaxation decreased the fraction of males.

Figures 2B and 2C show that the One Child Policy and subsequent relaxations did not affect sex ratios of higher order births in relaxed counties differently from counties without relaxations. The relaxation did not change the sex composition of siblings for first born children born between 1972 and 1982. This is important because the exclusion restriction for using the triple difference as an instrument for family size requires that the instrument does not affect any right hand side variable other than family size. Dahl and Moretti (2004) and Ananat and Michaels (2004) show that the sex composition of children has a direct affect on the divorce rates of parents. Hence, if the relaxation also changed the sex composition of children in families of the affected cohort, the 2SLS estimate will be biased.

To estimate the effect of the relaxation on sex ratios, I estimate the following equation using the sample of first born children. The children are divided into three groups according to birth cohort. The reference group is comprised of individuals not affected by the One Child policy and the relaxation (born before 1978). The second group comprises of children born after the One Child Policy but before the relaxation (1978-1981). The third group comprises of children born after the relaxation (1982-1989).

$$male_{itc} = \sum_{l=2}^{3}(relax_c \times post_{il})\delta_l + \alpha + \gamma_t + \psi_c + \varepsilon_{itc} \tag{4}$$

The probability of being male for individual i, born in county c, birth year t is a function of: the interaction term between $relax_c$, and $post_{il}$, a variable indicating the individual's cohort group; ψ_c, county fixed effects and γ_t, cohort group fixed effects.

The estimate for δ_l is shown in column 1 of Table 3. It shows that first born children born in relaxed regions after the initial One Child Policy are 8% more likely to be male than children born in un-relaxed regions. After the relaxation, first born children born in relaxed areas are only 4% more likely to be male than children born in areas without the relaxation. Both estimates are statistically significant at the 1% level.

It is interesting to note that although the One Child Policy constrained the family size of individuals born as early as 1976, sex selection from the One Child Policy appears only in cohorts born after 1978. This is consistent with past findings that sex selection in China mostly occurs for very young children. In other words, once the policy is announced in 1978-1980, parents were unwilling (or unable) to kill girls that were more than 1 or 2 years of age in order to have a boy.

These results suggest that parents in counties which received relaxations reacted differently to family planning policies than parents in counties without relaxations.These differing reactions may reflect different preferences towards investment in children that is time, sex and region specific, which will confound the two stage least squares estimate for the effect of family size.. Excluding cohorts born after the relaxation (1982-1990) partially addresses this problem. It has the additional advantage of excluding households which kept girls in order to have a second child. The sample selection issue from the [initial] One Child Policy (1978-1981) will be addressed by estimating the absolute value of the lower bound effect of family size with the alternative sample.

4.2.3 Effect on Female Labor Supply

If the relaxation caused parents to have a second child and mothers to stay home to take care of the child, the 2SLS estimate will confound female labor supply effects with family size effects. To address this, I estimate the effect of the relaxation on mother's work status controlling for mother's age. The results are not reported in this paper. They show that mothers of affected girls were *less* likely to stay at home.

4.3 The Effect of Family Size on School Enrollment

4.3.1 OLS

The correlation between school enrollment and family size can be obtained by estimating the following equation for the sample of first born children.

$$enroll_{itc} = sibs_{itc}b + X_{ict}^{'}\kappa + \sum_{l=1973}^{1981}(urban_c \times d_{il})\delta_l + \alpha + \gamma_t + \psi_c + \varepsilon_{itc}$$

School enrollment for individual i, born in county c, birth year t, is a function of: $sibs_{itc}$, the number siblings he or she has; X_{ict}, individual characteristics; the interaction term between $urban_c$, distance to urban area, and d_l, a variable indicating whether an individual was born in year l; γ_t, birth year fixed effects; and ψ_c, county fixed effects. The estimate in Table 5 column (1) shows that on average, one additional sibling is correlated with 1.7 percentage point less of enrollment. The estimate is statistically significant at the 1% level. Columns (2)-(5) show that the OLS estimate is robust to controls for the full set of double interaction terms from equation (2), a variable indicating whether an individual is ethnically Han, distance to urban area and mother's education.[7] Panel B shows the OLS estimates using the constructed sample. The point estimates are similar to those of the original sample and statistically significant.

4.3.2 Reduced Form Estimates

To illustrate the identification strategy, I will first estimate the effect of the relaxation on enrollment separately for boys and girls. This can be characterized by the following equation.

$$enroll_{itc} = \sum_{l=1973}^{1981} (relax_c \times d_{il})\beta_l + \alpha + \gamma_t + \psi_c + \upsilon_{itc} \tag{5}$$

The reference group is comprised of individuals born during 1962-1972. The dummy variable for the reference group and all its interactions are dropped. The coefficients for girls and boys are shown in Table 2, columns (3) and (4). The estimates are statistically significant for girls. Figure 3A plots the estimates for boys and girls. Cohort to the right of the solid line are those affected by the relaxation. The plot of the reduced form shows that for the affected cohort, girls have higher education enrollment than boys, whereas for the unaffected cohort, girls had lower school enrollment rates than boys.

The estimates in Figure 3A show that relative to areas without the relaxation, enrollment for both boys and girls decreases after primary school. This

[7] The double interactions include the interaction term of $relax_c$ and d_{il}; the interaction term between $girl_{itc}$ and d_{il};the interaction term between $relax_c$ and $girl_i$; and $girl_{itc}$. The reference group is comprised of cohorts born during 1962-1972. The dummy variable for the reference cohort and all its interactions are dropped.

is consistent with the hypothesis that school provision and quality in relaxed regions relative to regions without the relaxation declined during this period. I control for this by comparing the effect of the relaxation on enrollment for boys with the effect of the relaxation on enrollment for girls, which can be characterized by an equation similar to equation (2) with school enrollment as the dependent variable.

$$\begin{aligned} enroll_{itc} = & \sum_{l=1973}^{1981} (relax_c \times girl_{itc} \times d_{il})\beta_l + \sum_{l=1973}^{1981} (relax_c \times d_{il})\delta_l \\ & + \sum_{l=1973}^{1981} (girl_{itc} \times d_{il})\zeta_l + (relax_c \times girl_{itc})\lambda + girl_{itc}\kappa \\ & + \alpha + \gamma_t + \psi_c + v_{itc} \end{aligned} \tag{6}$$

The reference group is comprised of individuals born during 1962-1972. The dummy variable for the reference group and all its interaction terms are dropped. The coefficients are shown in Table 2, column (6). The estimates show that for older cohorts not affected by the relaxation, individuals born in relaxed areas have on average 1% to 17% less school enrollment than areas without the relaxation. However, for cohorts affected by the relaxation, individuals born in relaxed areas are on average enrolled in school 5% more than individuals born in areas without the relaxation. The estimates are statistically significant at the 1% level. Figure 3B plots the triple difference reduced form estimates. It shows that school enrollment in relaxed areas is higher for girls of the affected cohort than for boys.

4.3.3 Two Stage Least Squares

Using the predicted residuals from the first-stage equation (2), I estimate the following second stage:

$$\begin{aligned} enroll_{itc} &= sibs_{itc}b + \sum_{l=1973}^{1981} (relax_c \times girl_i \times d_{il})\beta_l \\ &+ \sum_{l=1973}^{1981} (relax_c \times d_{il})\delta_l + \sum_{l=1973}^{1981} (girl_i \times d_{il})\zeta_l \\ &+ \sum_{l=1973}^{1981} (urban_c \times d_{il})\delta_l + (relax_c \times girl_i)\lambda \\ &+ X_{ict}^{'}\kappa + \alpha + \gamma_t + \psi_c + v_{itc} \end{aligned}$$

$X_{ict}^{'}$ is a vector of individual controls (e.g. mother's education, ethnicity). $urban_i$ is the average distance to the nearest urban area. Column (7) in Panel A of Table 5 shows that contrary to the negative OLS estimate, an additional sibling increases school enrollment by 20.5% in the actual sample. The estimate is statistically significant at the 5% level. I repeat the estimation for the alternative constructed sample to estimate the lower bound effect of family size on school enrollment. The result is shown in Panel B of column (7). It shows that one additional sibling increases school enrollment of the first born child by 18.4%. The estimate is statistically significant at the 1% level. Columns (8)-(10) show that the 2SLS estimates of both the actual sample and the constructed sample are robust to individual and county level controls.

5 Conclusion

This paper has two purposes. It evaluates the effects of the One Child Policy and the subsequent 1-son-2-child relaxation. Then, it uses exogenous variation in family size caused by this relaxation to evaluate the causal effect of family size on school enrollment.

The One Child Policy is one of the most internationally controversial policies undertaken by the post-Mao Chinese government. It reportedly increased female infanticide and led to a generation of "spoiled children". However, the common misunderstanding that the One Child Policy is uniformly enforced across China and the lack of local enforcement data has, until recently, prevented researchers from measuring the causal effects of China's family planning policies. The lack of transparency in the policy enforcement decision process added to the difficulty of such studies.

This paper uses local enforcement data of the 1-son-2-child relaxation to evaluate the effects of the relaxation and the One Child Policy on sex ratio and family size. It shows that although the One Child Policy was enacted in 1978-1980, previous family planning laws which encouraged birth spacing meant that the former was actually binding for cohorts born as early as 1976. The results show that the 1-son-2-child relaxation was indeed implemented in regions where sex selection was more severe after the initial One Child Policy. The relaxation decreased sex selection from the levels immediately following the implementation of the One Child Policy, but sex ratios in these regions did not return to their initial pre-One Child Policy levels.

I use the exogenous increase in family size of girls born in relaxed regions to evaluate the causal effect of family size on school enrollment. The advantage of this method is that it addresses the endogenous relationships between family size and parental preferences over education and between family size and the quality of the first child. The results show that school enrollment for girls from one-child households *increased* by 18-20% when parents had an additional child. The findings reject models which predict that quality is monotonically decreasing in quantity. However, the results are consistent with evidence from previous studies which show that although family size is negatively correlated with education outcomes for children from households of two or more children, only-children are disadvantaged compared to children from two or three child families.Further research is needed to examine the family size effect beyond the two-child context.

There are several hypotheses that can explain the only child disadvantage. In a simple framework where parents and children have the same preferences (or where parents internalize child preferences), family size can increase school enrollment if children complement each other in their respective production functions. Iacavou's (2004) finding that the only child disadvantage decreases as the child interacts more with children outside of school suggests that there are complementarities in learning or development for children. In this study, I find that the family size effect varies according to the age gap of the two children. Similarly, psychologists Zajonc and Gregory (1982) hypothesized that children benefited from teaching younger children. This hypothesis predicts that the younger child will be worst off relative to the older child. Exploration of this hypothesis is prevented by the fact that the second child in this study is 6 years of age or younger.

There may also be economies of scale in schooling costs and learning (or other psychological responses). In the context of a developing country, text books and clothes can be considered as fixed costs for sending children to school. Then having a second child will lower the average and marginal cost of school attendance for the first child as long as the secondary market for these goods functions such that transferring the goods to children from other families is more costly than transferring them to children within the household. [8]

In summary, this paper presents strong empirical evidence that only-children are worst off compared to children from two or three child families. More research is needed to understand the family size effect beyond the one-child context

[8] If parents and children have different preferences, the results can also be explained by child behavioral responses in regard to the decrease in her share of tangible and intangible resources within the household. While there has been many studies about parents' decisions to reallocate resources in response to the decrease in average resources, the effect of the first born child's behavioral response and parental response to child behavior has been left unexplored. The results of this paper, however, suggest that this is worthwhile considering for there are several ways that child behavior can cause parents to increase the first child's school attendance. For example, if the first child dislikes sharing tangible goods and parental attention with the younger sibling, she may behave badly and therefore increase her parents' desire to send her to school, away from the sibling during the day. Simultaneously, being at school may have an added attraction relative to being at home for the first child as a place where her position is not affected by the birth of the latter. In addition, the first child may be more motivated to attend school because she feels that academic distinction will increase her stature in the household relative to the younger child.

It is important to note that in China, there is no schooling beyond high school in rural areas. Universities are highly concentrated in the largest cities. In fact, rural students with academic potential generally leave their homes during high school, or even middle school to attend better quality schools in urban areas. The lack of economic opportunities in rural areas means that such children do not return home after graduating from college. Therefore, if parents desire to keep at least one child near them, they are more likely to encourage the child to pursue higher education if they have a second child to keep near them. This will translate into lower drop out rates for the first child. Like the hypothesis proposed by Zajonc and Gregory (1982), this explanation implies that family size effects may be different for the children of different birth order.

and to understand the mechanisms underlying this effect. In the mean time, the results show that the relation ship between quantity and quality is not monotonically decreasing and a richer theoretic model is needed to understand the effect of family size on child quality. Policy makers should note that "one" is not the optimal number of children per household and consider administering aggressive family planning policies in conjunction with programs that increase interaction between children outside of school.

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Figure 1A: The Effect of Relaxation on Family Size
Coefficients of the Interactions between
Born in a Relaxed Area * Birth Regions

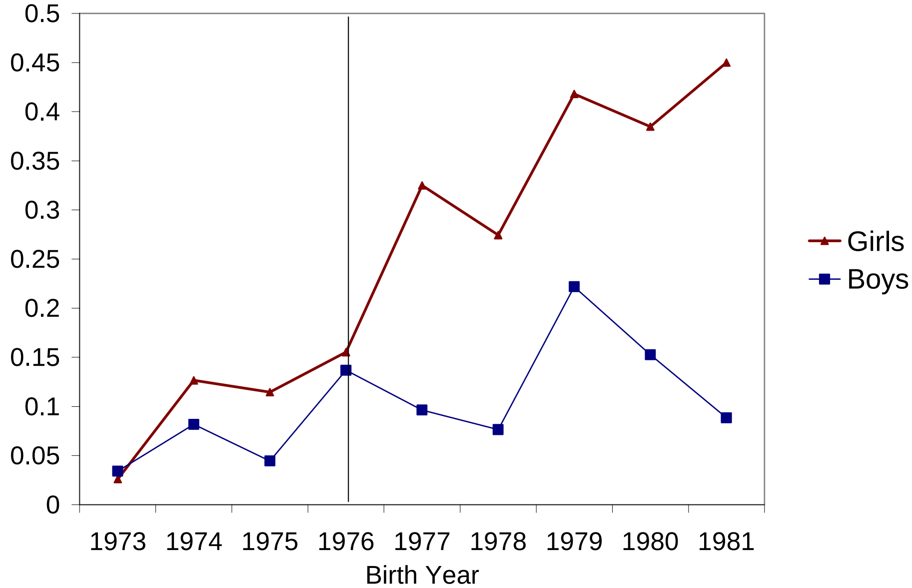

Figure 1B: The Effect of Relaxation on Family Size
Coefficients of the Interactions between
Dummy for Girl * Born in a Relaxed Region * Birth Year

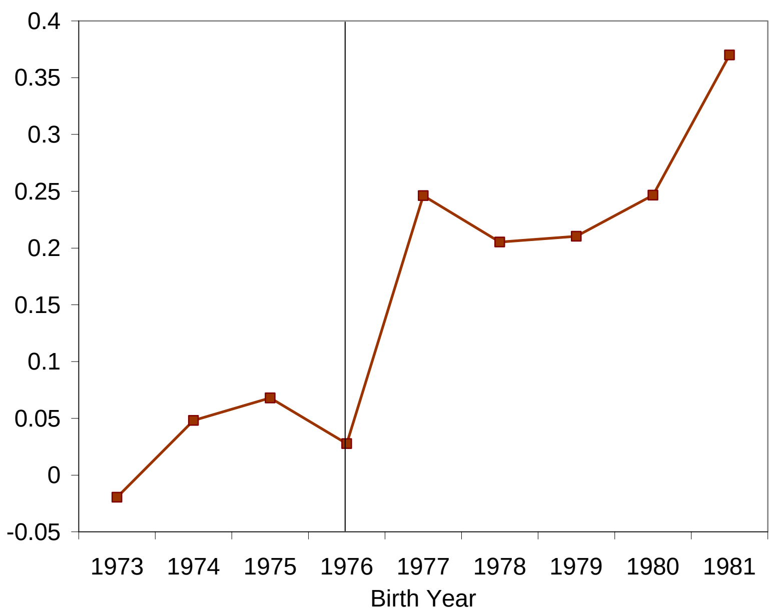

Figure 2A: The Effect of Relaxation on Sex Ratios of First Born Children and 95% Confidence Intervals
Coefficients of the Interactions between
Born in Relaxed Region * Birth Year

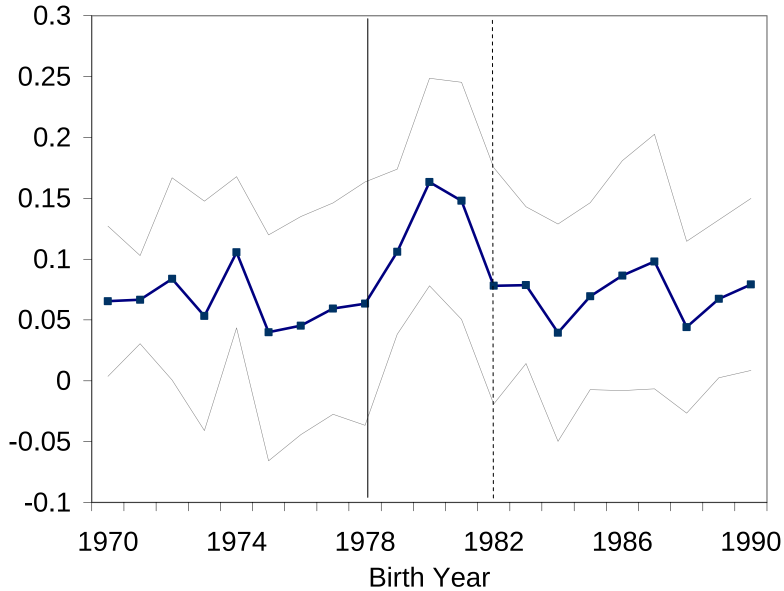

Figure 2B: The Effect of Relaxation on Sex Ratios of Second Born Children and 95% Confidence Intervals
Coefficients of the Interactions between
Born in Relaxed Region * Birth Year

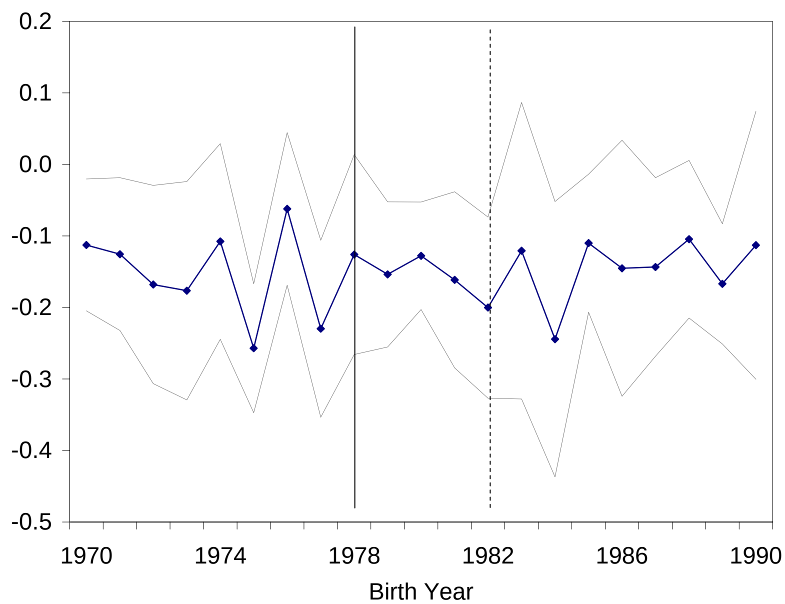

Figure 2C: The Effect of Relaxation on Sex Ratios of Later Born Children and 95% Confidence Intervals

Coefficients of the Interactions between
Born in Relaxed Region * Birth Year

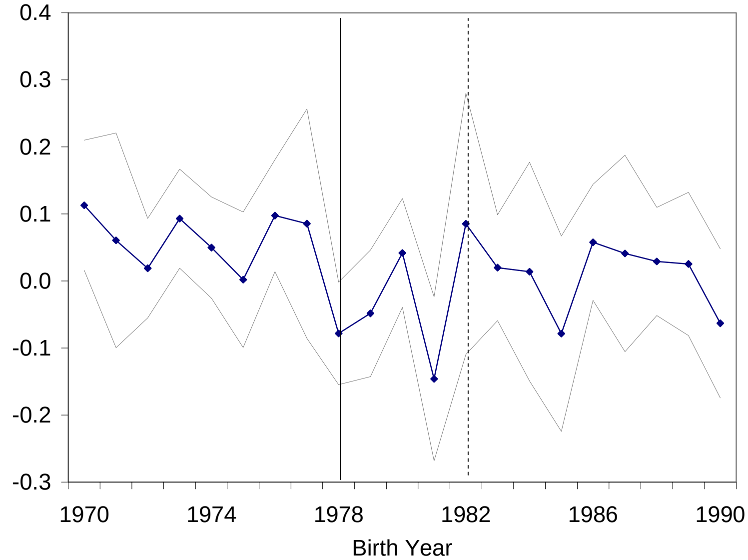

Figure 3A: The Effect of Relaxation on School Enrollment
Coefficients of Interactions between
Born in Relaxed Region * Birth Year

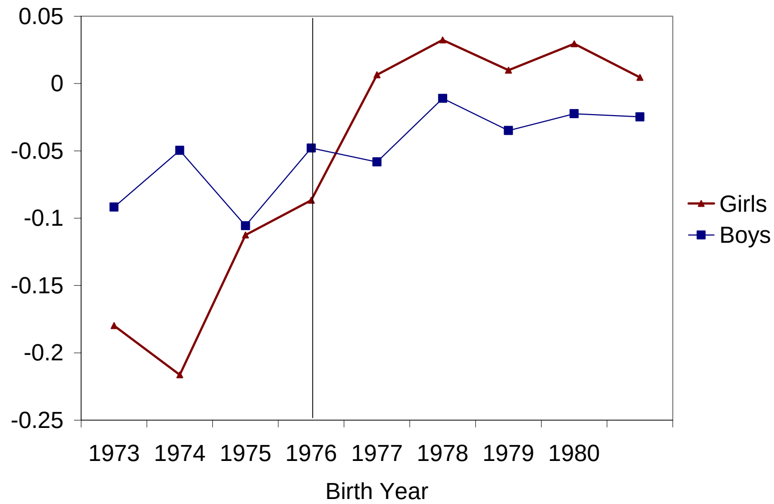

Figure 3B: The Effect of Relaxation on School Enrollment
Coefficients of Interactions between
Dummy for Girl * Born in Relaxed Region * Birth Year

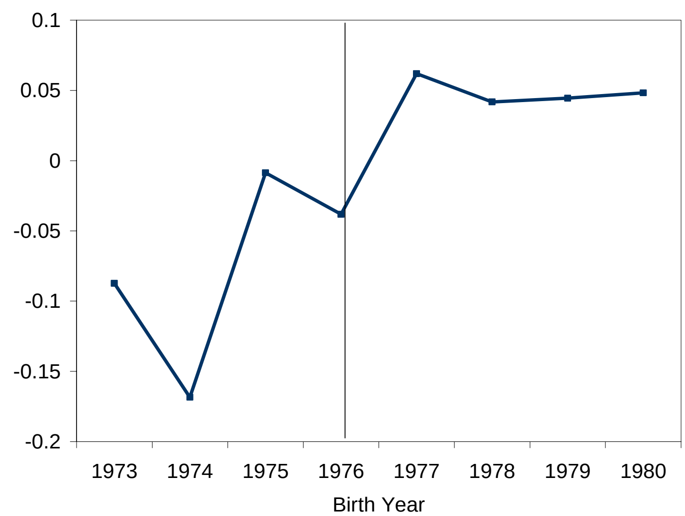

Table 1: Descriptive Statistics
CHNS 1989 and 0.1% Sample of China Population Census

	Obs	Mean	Std. Err.	Obs	Mean	Std. Err.
A. By Sex	**Female**			**Male**		
Han	13271	0.944	(0.002)	15500	0.949	(0.002)
Siblings	13271	1.153	(0.009)	15500	0.922	(0.008)
Sisters	13271	0.504	(0.006)	15500	0.511	(0.006)
Brothers	13271	0.649	(0.006)	15500	0.411	(0.005)
Enrollment	13271	0.473	(0.004)	15500	0.456	(0.004)
Mother's Education	12862	6.063	(0.037)	14890	5.668	(0.035)
Father's Education	12134	8.058	(0.034)	14239	7.628	(0.033)
Mother is Housewife	13271	0.119	(0.003)	15500	0.139	(0.003)
Relaxed Area	13271	0.254	(0.003)	15500	0.242	(0.003)
B. By Family Size	**Siblings**			**Only Child**		
Sex	19038	0.502	(0.004)	9733	0.611	(0.005)
Han	19038	0.941	(0.002)	9733	0.958	(0.002)
Enrollment	19038	0.399	(0.004)	9733	0.591	(0.005)
Mother's Education	18488	5.300	(0.029)	9264	6.952	(0.048)
Father's Education	17623	7.476	(0.027)	8750	8.530	(0.044)
Mother is Housewife	19038	0.134	(0.002)	9733	0.121	(0.003)
Relaxed Area	19038	0.279	(0.003)	9733	0.186	(0.003)
C. By Relaxation	**No Relaxation**			**Some Relaxation**		
Sex	10828	0.544	(0.005)	17943	0.535	(0.004)
Han	10828	0.968	(0.002)	17943	0.934	(0.002)
Siblings	10828	1.048	(0.010)	17943	1.016	(0.007)
Sisters	10828	0.512	(0.007)	17943	0.505	(0.005)
Brothers	10828	0.536	(0.007)	17943	0.511	(0.005)
Enrollment	10828	0.437	(0.005)	17943	0.480	(0.004)
Mother's Education	10454	5.034	(0.040)	17298	6.345	(0.033)
Father's Education	9914	7.439	(0.036)	16459	8.058	(0.031)
Mother is Housewife	10828	0.111	(0.003)	17943	0.141	(0.003)
Relaxed Area	10828	180.299	(1.256)	17943	147.335	(1.195)
Distance to Urban	9460	2.041	(0.017)	17943	11.849	(0.087)
Agriculture	10818	0.720	(0.004)	17903	0.569	(0.004)
Distance to Primary School	10828	0.230	(0.006)	16672	0.399	(0.004)
Distance to Middle School	10827	1.008	(0.009)	16672	1.584	(0.011)
Distance to High School	10827	4.920	(0.084)	16672	4.506	(0.067)

Sample of cohorts born 1962-1981

Table 2: The Effect of Relaxation on Family Size and School Enrollment by Size
Coefficients for columns (1)-(4) are the interaction terms between born in a relaxed region * year of birth.
Coefficients in columns (5)-(6) are the triple interactions between dummy for girl * born in a relaxed region * year of birth.

	Dependent Variables						
	# Siblings		Enrollment			# Siblings	Enrollment
	(1)	(2)	(3)	(4)		(5)	(6)
Birth Year	Girls	Boys	Girls	Boys	Birth Year	1st	RF
1973	0.026	0.034	-0.180	-0.092	1973	-0.020	-0.087
	(0.110)	(0.116)	(0.082)	(0.071)		(0.099)	(0.037)
1974	0.127	0.082	-0.216	-0.050	1974	0.048	-0.168
	(0.115)	(0.107)	(0.098)	(0.078)		(0.073)	(0.070)
1975	0.115	0.045	-0.112	-0.106	1975	0.068	-0.009
	(0.071)	(0.139)	(0.078)	(0.048)		(0.132)	(0.056)
1976	0.155	0.137	-0.087	-0.048	1976	0.028	-0.038
	(0.128)	(0.157)	(0.062)	(0.030)		(0.170)	(0.074)
1977	0.325	0.096	0.007	-0.058	1977	0.246	0.062
	(0.136)	(0.101)	(0.055)	(0.037)		(0.116)	(0.061)
1978	0.274	0.076	0.032	-0.011	1978	0.205	0.042
	(0.152)	(0.161)	(0.028)	(0.027)		(0.171)	(0.022)
1979	0.418	0.222	0.010	-0.035	1979	0.210	0.044
	(0.158)	(0.159)	(0.034)	(0.025)		(0.183)	(0.022)
1980	0.385	0.153	0.030	-0.022	1980	0.247	0.048
	(0.180)	(0.128)	(0.033)	(0.028)		(0.168)	(0.018)
1981	0.450	0.088	0.005	-0.025	1981	0.370	0.029
	(0.186)	(0.154)	(0.035)	(0.030)		(0.194)	(0.014)
Observations	13271	15500	13271	15500	Observations	28771	28771
R-squared	0.26	0.23	0.70	0.69	R-squared	0.25	0.69

Regressions in columns (1)-(4) include county and birth year fixed effects.
Regressions in columns (5)-(6) inlude controls for relax*girl, relax*birthyear, girl*birthyear, girl, birthyear fixed effects and county fixed effects.
Standard errors clustered at county level.

Table 3: The Effect of Relaxation on Sex Ratios by Birth Parity

Coefficients are the interaction terms between born in a relaxed region * birth year.

	Dependent Variable: Dummy for Male											
	1st Borns				2nd Borns				Later Borns			
	(1)		(2)		(3)		(4)		(5)		(6)	
Birth Year	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
1970	0.065	(0.032)	0.064	(0.031)	-0.113	(0.047)	-0.110	(0.046)	0.113	(0.049)	0.112	(0.049)
1971	0.067	(0.018)	0.066	(0.018)	-0.126	(0.055)	-0.124	(0.054)	0.061	(0.082)	0.055	(0.081)
1972	0.084	(0.042)	0.086	(0.043)	-0.168	(0.071)	-0.170	(0.071)	0.019	(0.038)	0.018	(0.037)
1973	0.053	(0.048)	0.056	(0.049)	-0.177	(0.078)	-0.178	(0.077)	0.093	(0.038)	0.091	(0.037)
1974	0.106	(0.032)	0.103	(0.033)	-0.108	(0.070)	-0.111	(0.069)	0.050	(0.039)	0.047	(0.039)
1975	0.027	(0.047)	0.025	(0.049)	-0.257	(0.046)	-0.256	(0.047)	0.002	(0.052)	0.002	(0.052)
1976	0.045	(0.046)	0.050	(0.047)	-0.062	(0.054)	-0.066	(0.055)	0.097	(0.043)	0.094	(0.044)
1977	0.059	(0.044)	0.057	(0.045)	-0.230	(0.063)	-0.225	(0.063)	0.085	(0.087)	0.081	(0.087)
1978	0.063	(0.051)	0.056	(0.052)	-0.126	(0.071)	-0.122	(0.071)	-0.078	(0.039)	-0.080	(0.039)
1979	0.106	(0.035)	0.101	(0.036)	-0.154	(0.052)	-0.154	(0.051)	-0.048	(0.048)	-0.050	(0.049)
1980	0.163	(0.044)	0.162	(0.044)	-0.128	(0.038)	-0.126	(0.039)	0.042	(0.041)	0.041	(0.041)
1981	0.148	(0.050)	0.145	(0.051)	-0.161	(0.063)	-0.169	(0.064)	-0.146	(0.062)	-0.144	(0.063)
1982	0.078	(0.050)	0.077	(0.050)	-0.200	(0.065)	-0.200	(0.065)	0.085	(0.099)	0.073	(0.098)
1983	0.079	(0.033)	0.081	(0.034)	-0.121	(0.106)	-0.120	(0.105)	0.020	(0.040)	0.007	(0.037)
1984	0.040	(0.046)	0.038	(0.046)	-0.244	(0.098)	-0.241	(0.098)	0.014	(0.083)	0.003	(0.083)
1985	0.070	(0.039)	0.066	(0.039)	-0.110	(0.049)	-0.114	(0.048)	-0.079	(0.074)	-0.079	(0.076)
1986	0.086	(0.048)	0.089	(0.048)	-0.145	(0.091)	-0.158	(0.092)	0.058	(0.044)	0.047	(0.049)
1987	0.098	(0.053)	0.097	(0.053)	-0.144	(0.064)	-0.145	(0.062)	0.041	(0.075)	0.038	(0.075)
1988	0.044	(0.036)	0.038	(0.036)	-0.105	(0.056)	-0.108	(0.057)	0.029	(0.041)	0.028	(0.041)
1989	0.067	(0.033)	0.075	(0.033)	-0.167	(0.043)	-0.167	(0.044)	0.025	(0.055)	0.021	(0.055)
1990	0.079	(0.036)	0.077	(0.037)	-0.113	(0.096)	-0.118	(0.095)	-0.063	(0.057)	-0.064	(0.057)
Han	N		Y		N		Y		N		Y	
Observations	44754		44754		23306		23306		14495		14495	
R-squared	0.01		0.01		0.01		0.01		0.01		0.01	

Regressions include county and birth year fixed effects.

Standard errors clustered at the county level.

Table 4: The Effect of Relaxation on Sex Ratios for First Borns

Coefficients are the interaction terms between born in a relaxed region * born in the One-Child Policy Cohort and between born in a relaxed region * born in the relaxation cohort

The Effect of Relaxation on Sex Ratios for First Borns

Dependent Variable: Dummy for Male		
	(1)	(2)
Born in relaxed region * Born 1976-1981	0.106	0.098
	(0.024)	(0.025)
Born in relaxed region * Born 1982-1989	0.037	0.035
	(0.021)	(0.021)
Han, Han * Birth Cohort	N	Y
Observations	44234	44234
R-squared	0.00	0.00

Regressions include county and birth cohort fixed effects.

Standard errors clustered at the county level.

Table 5: OLS Estimate of the Effect of Family Size on School Enrollment
Coefficients are the number of siblings a child has

	Dependent Variable: School Enrollment								
	OLS					2SLS			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
A. Main Estimates									
# of Siblings	-0.017	-0.017	-0.017	-0.017	-0.017	0.205	0.206	0.205	0.193
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.103)	(0.103)	(0.103)	(0.090)
Observations	28771	28771	28771	28771	27752	28771	28771	28771	27752
R-squared	0.69	0.69	0.69	0.69	0.70	0.55	0.55	0.55	0.57
B. Lower-bound Estimates									
# of Siblings	-0.017	-0.016	-0.016	-0.016	-0.016	0.184	0.186	0.193	0.168
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.080)	(0.081)	(0.088)	(0.067)
Controls									
Mother's Education	No	No	No	No	Yes	No	No	No	Yes
Urban*Birthyear	No	No	No	Yes	No	No	No	Yes	No
Han*Birthyear	No	No	Yes	Yes	No	No	Yes	Yes	No
Han	No	No	Yes	Yes	No	No	Yes	Yes	No
All Interactions	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	28705	28705	28705	28705	27687	28705	28705	28705	27687
R-squared	0.70	0.70	0.70	0.70	0.71	0.58	0.58	0.57	0.61

All interactions includes relax*girl, relax*birth year, girl*birth year.

All regressions control for girl, county fixed effects and birth year fixed effects.

Standard errors clustered at county level.