# **One Child Policy and the Rise of Man-Made Twins**<sup>\*</sup>

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# Abstract

Population censuses of 1982, 1990 and 2000 census data and 2005 1% population survey reveal a rapid rise of twin births in China since the 1970s. Exploring provincial variations in the change of one-child policy regulatory fines over time, we find that the one child policy accounts for at least one third of the increase in twins. This result is robust to various checks including exploring the pre-post policy change and the difference in policy implementation between the ethnic Han and minorities. Furthermore, investigation in the China Health and Nutrition Survey reveals that an increase in policy fines by one year's income is associated with over a one centimeter height difference between twins of the same gender. These results are consistent with the hypothesis that the one-child policy has stimulated incentives to obtain twins through methods like taking fertility drugs.

JEL Codes: J08, J11, J13

Keywords: Twins, One-Child Policy, China

<sup>&</sup>lt;sup>\*</sup> We thank David Cutler, Richard Freeman, Edward Glaeser, Claudia Goldin, Lawrence Katz, Avraham Ebenstein, Adriana Lleras-Muney, Amanda Pallais, John Strauss, T. Paul Schultz, Fei Wang, and Yi Zhou for their helpful comments. All errors are ours.

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"One is too few,' said a woman (in China) waiting at the hospital..."By Elizabeth Grether, ABC News, Aug. 3rd, 2011

# I. Introduction

Over the past several decades, reported twin births in China have increased dramatically. As Figure 1a shows, twin births as reported in population censuses more than doubled between the late 1960s and the early 2000s, from 3.5 to 7.5 per thousand births. Although China is not unique in seeing rising twin births, in the U.S., for example, twinning rate went from 18.9 to 33.3 per thousand births from 1980 through 2009 (Martin et al., 2012), the reason for the rising twin rates is likely different in China. In the U.S., rising twin rates is mainly attributed to postponement of childbearing and/or in connection with fertility drug usage. In China, however, twin births occur mostly among young mothers. Because of China's draconian family planning policy, it is widely believed that fertility drugs are being used by young women in order to bypass the one-child rule.<sup>1</sup> Additionally, it has also been reported that couples register non-twin children as twins for the same purpose.<sup>2</sup> Although the medical literature is inconclusive on the impact of fertility drug use on maternal and child health, twin births have been known to be associated with high maternal risks and low birth weights, which in turn is negatively associated with later health and socio-economic outcomes. If, on the other hand, the twins are fake, then the older child's schooling will be delayed because the school entry age is determined according to one's registration.

#### [Figure 1a about here]

This paper examines the impact of one-child policy on twining in China. We first explore the temporal and geographical variation in one child policy fine to investigate the impact of one child policy on twin births. Then we use the start of the one-child policy (in 1979) and

<sup>&</sup>lt;sup>1</sup> http://abcnews.go.com/Health/chinese-women-fertility-drugs-bypass-child-policy/story?id=14219173 , accessed March 10th, 2013.

<sup>&</sup>lt;sup>2</sup> An example (in Chinese): http://www.wuhan.jcy.gov.cn/yasf/200905/t20090519\_221490.html , accessed March 10, 2013

Han-ethnic minority difference to provide supplementary evidence. Both approaches yield a large impact of the one-child policy on rising twinning rate.

Because the impact thus identified may come from either real twinning from fertility drugs or fake twins from reporting non-twin siblings as twins, and welfare implications are different in the two cases, we further distinguish between these two hypotheses. Since fertility drug use results in more DZ twins and same sex DZ twins are more different than MZ twins, while fake twins are different regardless of sex composition, if one-child policy leads to more fertility drug use rather than reporting fake twins, then we are likely to see greater difference among same-sex twins. We test this hypothesis by looking at height differences within twins identified from the China Health and Nutrition Survey (CHNS).

The structure of the paper is as follows. Section 2 introduces the data used in this study and provides background for the one-child policy. Section 3 examines the impact of one-child policy on twining in China using census data. Section 4 looks at the impact of one-child policy on height difference within twins using CHNS data. Section 5 concludes with a discussion of findings and the policy implications.

# II. Data

#### 2.1 Census Data and Twins in China

The data used in this study include a 1% sample of the 1982, 1990 Population Census, 0.1% sample of the 2000 Census, and 2005 Population Study sample. All the data sets contain variables including birth year, region of residence, type of residence,<sup>3</sup> sex, ethnicity, education, and relation to the head of the household. The datasets after 1982 also include month of birth. For women older than 15, the data provide information about their fertility history, including number of children ever born and number of children alive.

 $<sup>^{3}</sup>$  The 1990 Census does not provide the type of current residence. Instead, it provides whether the respondents lived in the same place 5 years ago and what the type of residence was then. Therefore, we constructed a variable to indicator the type of current residence. First, we keep those who were in the same place. Then, we calculate the proportion of people in this sample for each residence type 5 years ago. We use the type which has highest proportion as the type of current residence in the same area.

For analysis of twin births and family background, we restrict our sample using the following steps. We first keep only those households with at least one child and with mother's information available. We then restrict the sample to those households whose reported number of children alive equals the number of children born, which also equals to number of children observed in the survey. This is to make sure what we examine are all the children of the household, as the survey only has detailed information on children living in the household. For the same reason, in case we miss children who have moved out of the household, we further drop households with children over age 17 in the survey. We finally drop households where the mother's age when giving birth is either younger than 15 or older than 50, as this sample may be too special or may just be due to recording errors.<sup>4</sup>

We use samples of the children after the above restrictions, and twins are defined as children in the same household who have the same birth year and birth month.<sup>5</sup> We treat each birth as one observation, so two twins are only counted as one observation. For each birth, we also construct several variables including whether this birth is a twin, and if so whether this twin birth is male, female, or mixed gender. Figure 1a plots the rate of twins by year of birth between 1965 and 2005. We can see that the variation before 1970 and after 1990 is larger because of smaller sample size. Sampling weights are applied throughout this analysis. In Figure 2a, we divide twins into same-gender twins and mixed-gender twins, and plot again by year of birth. It shows that proportion of same-gender twins is higher than that of mixed-gender twins, which is consistent with statistics from the U.S. and Japan. In addition, the period between 1980 and 2000 also witnesses a significant increase in both types of twin birthrates. Figure 1c divides same-gender twins into male twins and female twins, but we do not find significant differences between them. As a result, we do not separate the two types of twins any more in further analyses.

<sup>&</sup>lt;sup>4</sup> We also drop Tibet data.

<sup>&</sup>lt;sup>5</sup> Because 1982 Census data do not have information on birth month, we just define twins in that year as those children born in the same household with the same birth year only. Actually, the results are almost the same if we drop the 1982 Census or define twins only using the year of birth in all data sets.

# [Figures 1b and 1c about here]

#### 2.2 One Child Policy

In the 1970s, after two decades of explicitly encouraging population growth, policy makers in China enacted a series of measures to curb population growth, especially to the Han ethnicity. Beginning in around 1972, the policy "Later [age], longer [spacing of births], fewer [number of children]" offered economic incentives for parents to space their births over four years apart (Qian, 2009). The one child policy was formally conceived in 1979, but actual implementation had started in some regions in 1978, and enforcement gradually tightened until it was firmly established across the country in 1980 (Croll et. al., 1985; Banister, 1987)

The one-child policy, as the name suggests, restricted a couple to having only one child. The strictness of the one-child policy was partly reflected in its enforcement. Before 1978, Family Planning Policy (FPP) was mainly driven by political and administrative forces rather than law. In 1978, FPP appeared in the Constitution for the first time, and came up with more details in the 1982 amended Constitution (Wang, 2012). Legal measures, such as monetary penalties and subsidies, also ensured the effective enforcement of the one-child policy since 1979. Due to various development levels in different regions in China, the Central Party Committee "Document 7" devolved responsibility from central government to the local and provincial governments so that local conditions could be better addressed. It actually allowed for regional variation in family planning policies, like the amounts for monetary penalties or subsidies (Greenlaugh, 1986). As the population was still growing rapidly in the 1980s, Chinese policymakers felt more compelled to limit fertility, and emphasized the importance of this policy in several different documents. Afterwards, local governments tightened enforcement by reducing land allotments, denying public services, increasing unauthorized births, and so on.

However, it is worthwhile to note here that the one child policy mainly focused on the Han ethnicity. Though local governments might enact different regulations for a local minority of people, these regulations were less restrictive compared to those for Han. For example,

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regulations in Gansu province allow minority parents a second birth, which is not allowed for those of Han ethnicity.<sup>6</sup>

The measure of the one child policy in this study is the average monetary penalty rate for one unauthorized birth in the province-year panel from 1979 to 2000, which are from Ebenstein (2008).<sup>7</sup> The one child policy regulatory fine (policy fine) is formulated in multiples of annual income (Ebenstein, 2008; Wei, 2010). Though the monetary penalty is only one aspect of this policy and the government may take other administrative actions, it is still a good proxy for this policy because an increase in fines is usually associated with other stricter policies. Figure 2 shows the pattern of policy fines from 1980-2000 in each province.

# [Figure 2 about here]

It is obvious that fines and bonuses in different provinces generally have different patterns, both in timing and in magnitude. In looking at fine levels, Liaoning Province has increased it from one year's income to five in 1992, while Guizhou Province raised it from two years' income to five in 1998, and Hunan increased it from one to two in 1989. The average level of the fine is higher in the 1990s than in the 1980s, which is consistent with stricter policy enforcement in the 1990s. The geographical and temporal variance helps us identify the effects of the one-child policy in our empirical analysis.

Since approximately nine months are needed from the beginning of pregnancy unil the birth of the baby, the parents' decision to have a baby, if any, should be made close to a year in advance. We therefore match the sample from the census with the policy fine of the previous year before their birth in the corresponding province,<sup>8</sup> <sup>9</sup> and those not matched are dropped. Table 1 presents the summary statistics. Column 1 is the full sample, which is composed of over four million observations. We then divide the sample by the parents' ethnicity.<sup>10</sup> Column 2

<sup>&</sup>lt;sup>6</sup> http://www.mjrkj.gov.cn/html/gs-law/17\_31\_12\_360.html

<sup>&</sup>lt;sup>7</sup> The details about the construction of this variable can be found in Ebenstein (2008),

<sup>&</sup>lt;sup>8</sup> For example, a birth in 1986 in Hunan province will be matched with Hunan's fines in 1985. Births before 2001 are matched with fine data.

<sup>&</sup>lt;sup>9</sup> We assume fines are zero before 1979 because there was no formal one child policy then.

<sup>&</sup>lt;sup>10</sup> For simplicity, Han ethnicity mentioned below means both parents are Han ethnicity.

reports the results for whole Han sample, and Columns 3 and 4 present the statistics in this sample for those living in urban and rural areas, respectively. The last column reports the statistics for minorities. The twin birthrate in the full sample is 5.7 per thousand births, and that for the Han ethnicity is 5.8, which is higher than in the minority sample. Within the Han sample, the twin rate in urban areas, 2.3, is higher than in rural areas, 1.7. In addition, the mother's age when giving birth or the mother's education level is higher in urban areas, and mothers in urban areas give fewer births. The mean age of the sample is close to 8.

## [Table 1 about here]

Table 2 compares twin birthrates before and after the one child policy. For the Han ethnicity sample, twin birthrates increased by 0.26 percentage points, and both same-gender twins and different-gender twins contributed half. On the other hand, the twin birthrate increased by only 0.11 percentage points in the minority group, which is less than half of that in Han ethnicity sample. This simple comparison indicates the one child policy potentially accounts for a large proportion of the twin rate increase.<sup>11</sup>

[Table 2 about here]

#### **III. Empirical Results**

#### 3.1 Results Using One-Child Policy Fines

To estimate the effect of the one child policy on the twin birthrate, we estimate an equation as below:

(1) 
$$Twins_{ijkt} = \alpha + \beta Fine_{j,t-1} + \delta X_{ijkt} + \gamma_t YOB_t + \gamma_j Prov_j + \theta_j Prov_j \times Trend + \gamma_k wave_k + \varepsilon_i$$

where the dependent variable,  $Twins_{ijkt}$ , denotes whether birth *i* in year *t* is a twin in province *j* and wave *k*.  $Fine_{j,t-1}$  is the one child policy penalty in province *j* in year t-1. The coefficient  $\beta$  is the main coefficient of interest and gives the fine's association with the incidence of twins, and is interpreted as the impact of the one child policy.  $X_{ijkt}$  is a set of

<sup>&</sup>lt;sup>11</sup> More careful analysis is in Section 3.

covariates, which include dummies for residence type (urban/rural), parents' ethnicity (both Han/either a minority), birth order, mother's education level, and mother's age when giving birth.  $YOB_t$  is the set of birth year dummies and  $Prov_j$  denotes a set of province dummies. In addition, we also control for province-specific linear time trends, by adding the interactions between time trend (year of birth minus 1965) and province dummies.  $wave_k$  is a set of survey year dummies, and  $\varepsilon_i$  is the error term.

The coefficient estimates from equation (1) are presented in Table 3. Only the coefficients  $\beta$ 's and standard errors are reported. All coefficients should be interpreted as percentage points because all dependent variables in each panel have been multiplied by 100. The estimates for full samples are presented in Column 1, and those for Han ethnicity and minority samples are reported in Columns 2 and 3, respectively. The results indicate that a one year's income increase in the policy fine is associated with an approximately 0.06 percentage point increase in the twin birthrate. Corresponding estimates for different-gender twins and same-gender twins are 0.032 and 0.027, respectively. Therefore, the policy-associated twins are distributed almost equally in the different-gender and same-gender group, which is somehow consistent with what shown in Figure 1b. All these estimates are at least significant at the 10 percent significance level.

# [Table 3 about here]

There is concern that policy fines may be endogenous due to reverse causality, so we test the endogeneity of policy fines by using prior twin birthrates to predict future policy fines. If the association is significant, then the endogeneity problem is worthy of concern. More specifically, we use twins or different types of twins as independent variables, and take the policy fine of the next year, two years later or five years later as dependent variables. Because the policy fine is assumed to be zero before 1979 and we do not want this assumption to drive the results, we restrict the sample to those born after 1979, run the same regressions, and report the results in Table 4. Obviously, conditional on other covariates, neither twins nor different types of twins can predict the one policy fine over the next one, two or five years. This indicates that the endogeneity problem of policy fines due to reverse causality may not be serious.

#### [Table 4 about here]

# 3.2 Results Using Starting Time and Ethnicity

Results in Column 2 of Table 3 present strong evidence for the positive association between the policy fine and twin birthrate for the Han ethnicity. According to the statistics in Table 2, coefficients in the second column indicate that the one child policy find explains the approximately 35 percent twin increase. The effects of the one child policy are almost equal on different-gender and same-gender twins. Expectedly, we do not find any evidence that the policy fine is associated with the twin birthrate in the minority group. This is additional evidence that the result is driven by the one child policy as it is only applies to the Han ethnicity.

As Li et al. (2011) argue, the spatial and temporal variations of the one-child policy may be endogenous. As Li and Zhang (2008) find, the level of the fine increases with the community wealth level and the local government's birth-control incentives, but decreases with the local government's revenue incentives. In particular, the spatial and temporal variations in the one-child policy have been documented as being affected by the fertility rate, which, in turn, may be correlated with twin incidence. To avoid this endogeneity problem, we follow the methodology in Li et al. (2011) by estimating the following equation:

(2) 
$$Twins_{ijkt} = \alpha + \beta(Policy_{t \ge 1980} \times Han) + \delta X_{ijkt} + \gamma_t YOB_t + \gamma_j Prov_j + \theta_j Prov_j \times Trend + \gamma_k wave_k + e_i$$

in which  $Policy_{t\geq 1980}$  denotes an indicator of whether the twins are born after 1980, and *Han* is an indicator for *Han* ethnicity. The sample is the same as the one used to estimate Equation (1). The estimates for coefficient  $\beta$  in Equation (2) are reported in Table 5, which also provide strong evidence for the positive association of the one child policy with twin incidence. This Difference-in-differences estimate indicates that the one-child policy explains about 44 percent of the twins increase, indicating our estimates in Table 3 may have underestimated the policy impact. This may be reasonable since the policy fine is only one dimension for the regulations, which may miss many other useful variations.

# [Table 5 about here]

Equation (2) is a standard Difference-in-Differences estimation, which requires the trends in the treated group (Han ethnicity) and control group (Minority group) to be similar. To test this, we plot the twin birthrate by the parents' ethnicity group in Figure 3. Though we do not have counterfactual data to see how the twins rate changes over time in the Han group without the one child policy, it is still meaningful to see how twins rates evolved before 1979. As shown in the figure, the trends in the two groups are almost identical between 1970 and 1979.<sup>12</sup> Though we derived similar results in this DID model, it is worth noting that the minority group may not be a perfect control group because they are also under the control of the one-child policy. This is why we use temporal and geographical variations in fines to estimate the effect.

[Figure 3 about here]

#### 3.3 Heterogeneous Effects

#### 3.3.1 Heterogeneous Results by Residency

The one child policy is enforced differently in urban and rural areas. For example, the policy in many rural areas allows a household to give a second birth if the first is a girl, called the "one and a half" child policy. This varying enforcement, together with other differences between these two areas, may result in heterogeneous effects. Those living in rural areas are generally more poorly educated and may have stronger children or boy preferences, so they may have stronger incentives to obtain more children and more sons. We test this possibility by the same set of regressions for the urban and rural residents separately and the results are reported in Panel A of Table 6. As expected, policy fine is positively associated with twins in both urban and rural areas, and the effect of the one child policy in urban areas is larger.

[Table 6 about here]

# 3.3.2 Heterogeneous Results by Birth Order

<sup>&</sup>lt;sup>12</sup> Data before 1970 is not taken into account because of its small sample size.

In most rural areas a couple is allowed to have an additional child if the first is a girl. There may therefore be heterogeneous effects of birth order in rural and urban areas. We interact policy fines with birth order dummies, and report the coefficients of the interactions in Panel B of Table 6. Estimates in the full sample show that twins in the second birth order are mostly policy-associated, and this also holds in both urban and rural areas. What is interesting is that the coefficient on the interaction between the first birth order and the policy fine is also significant in urban areas, but that in the rural areas is much smaller and insignificant, which is consistent with the fact that a large part of rural areas are governed by the "one and a half" child policy.

#### 3.3.3 Results by Maternal Education Levels

Finally, policy may have heterogeneous effects on parents of different education levels due to the following facts: First, those with more education have less preference in children or in having a boy. Second, the illiterate may not know that taking fertility drugs can increase the probability of a twin birth, and have no idea how to get access to them. Third, the more educated are more likely to occupy positions in state-owned firms, where the one-child policy is enforced more strictly. Again, the direction is uncertain without an empirical test. We thus interact policy fines with the mother's education level dummies and run the same regressions from Equation (1). The results are shown in Panel C of Table 6, and provide evidence for heterogeneity. In the full sample, the coefficients on the interactions show an inverse-U pattern against mother's education. However, the patterns in urban and rural areas are different. As mentioned above, methods for producing twins are easier to access in urban areas, and thus it is the boy or children preference of parents that drives the results of the urban areas. The situation in rural areas is complex, however. It is possible that less educated individuals in rural areas do not know how to make twins, and the one child policy is not binding for them because these people are very difficult to administrate or follow. Thus, access to twin-making methods, boy or children preferences, and knowledge all contribute to the heterogeneity between and within different areas.

# IV. Differentiating Real "Man-Made" Twins and Fake Twins

The previous section showed significant effects of the one-child policy on the twin birthrate in China. As discussed earlier, this can be achieved either by producing real twins through technologies like fertility drugs, or by registering siblings of different ages as twins. In this section, we hope do provide some evidence on which one of these two cases is more likely to be true.

To keep it concise, the detailed derivations of how to test these two hypotheses are put in the appendix. Summarizing these derivations, we can use height differences of the two children in same-sex twins and mixed-sex twins to test. If we find the height difference within same-gender twins increases with policy fines and that within mixed-gender twins does not, then it is most likely that the increase in twin births are real "man-made" twins; if we find, however, that height differences within mixed-gender twins increase with policy fines, then they are most likely to be fake twins.. We then turn to the China Health and Nutrition Survey (CHNS) data to do this work.

#### 4.1 CHNS data

The China Health and Nutrition Survey (CHNS) included 26,000 individuals in nine provinces that vary substantially in geography, economic development, public resources, and health indicators.<sup>13</sup> The nine provinces contain approximately 56% of the population of mainland China. Data collection began in 1989 and has been implemented every 2 to 4 years since (Jones-Smith and Popkin, 2010). We chose CHNS to conduct this study because it is currently the largest and longest household panel data of China for public use and contains detailed information on family background, demographic variables, and biological measures, including height, for both adults and children. For children under 18, height is measured in millimeters in each wave.

Same as before, twins are defined as children less than 18 years old with exactly the same

<sup>&</sup>lt;sup>13</sup> The nine provinces are Liaoning, Heilongjiang, Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi and Guizhou.

birth year and birth month within the same household. There are 85 pairs in total,<sup>14</sup> including 60 same-gender pairs and 25 mixed-gender pairs. We then define the height gap of one pair as the height of the taller individual less that of the shorter one. Considering that the height gap may change as the children grow up, we also introduced ratio of height gap to the mean height in the pair (Gap/Height) as another measure of height difference. We also match the twins sample to the fines data, and drop those born after 2001, as we did for the census sample.

## 4.2 Summary Statistics and Descriptive Patterns

Summary statistics are provided in Table 7. The mean height gap is 2.1 centimeters and gap/height ratio is 1.71 percent. In addition, the fine is about 1.5 years' income on average, with the standard deviation being 0.6.

# [Table 7 about here]

Figures 4a and 4b plot same-gender twins' height gap and gap/height ratios against fine levels respectively, and both include a linear fit relation. The shadowed area shows 95% CI of the linear fit. The two figures show that height difference presents a strongly positive correlation with policy fine level.<sup>15</sup>

#### [Figures 4a and 4b about here]

Similarly, Figures 5a and 5b show the relation between height difference and fine level within different-gender twins. However, the linear fit is flat, thus no clear relation between height difference and fine level in the different-gender twins sample.<sup>16</sup>

<sup>&</sup>lt;sup>14</sup> Because CHNS data are panel, some particular twins may be interviewed for several waves. We also dropped 3 outliers with abnormal heights.

<sup>&</sup>lt;sup>15</sup> Specifically, the regression function in Figure 4a is Height gap = 1.13 (0.365) Fine - 0.46 (0.613), and that in Figure 4b is Gap/Height = 0.92 (0.254) Fine - 0.41(0.426).

<sup>&</sup>lt;sup>16</sup> The linear regression functions in 5a and 5b are Height gap = -.85 (2.70) Fine + 5.58 (3.92) and Gap/Height =

[Figures 5a and 5b about here]

However, the results are not sufficient to conclude that the one child policy has an impact on the height difference in twins, due to confounding factors: the fine and the height difference may both be correlated with the time trend, and the time trend is correlated with some unobservables related to twins' height difference.<sup>17</sup> In addition, the policy fine and twins' height differences may be also correlated to geographic factors, like gene agglomeration, climate, and so on. The relation between height difference and policy penalties may also be driven by these factors. Therefore, we rely on regressions to control for these confounding factors.

# 4.3 Regression Results on Height Gap

The regression equation we use is:

(3) 
$$HD_{ijkt} = \alpha + \beta Fine_{j,t-1} + \gamma Same_i + \delta X_{ijkt} + \gamma_t YOB_t + \gamma_j Prov_j + \theta_j Prov_j \times Trend + \gamma_k wave_k + \mu_i$$

where the dependent variable,  $HD_{ijkt}$  denotes the height difference within twin pair *i* in province j born in year t of wave k. It can be either the height gap or gap/height ratio as defined above. The coefficient  $\beta$  gives the association between the policy fine and the height difference of twins. *Same<sub>i</sub>* is an indicator of whether pair *i* is a same-gender twin or a different-gender one. We control for a full set of covariates that are potentially correlated with height difference, including indicator variables such as urban residence, same-gender pair, the boy being taller in different-gender twins, and also average height of pair i, the mother's age when giving birth, age, and age squared of the twin. Because of the small sample size, we combine the birth year into four groups (every five years as one group), *YOB<sub>t</sub>*, to capture the time effects. Finally, province

<sup>-0.05 (2.16)</sup> Fine + 3.80 (3.12), respectively.

<sup>&</sup>lt;sup>17</sup> For example, as GDP grows and more foods became available, composition of the Chinese diet becomes richer and people are better nourished. This is likely to influence the twin birthrate and DZ ones' proportion as well. Therefore, the positive association between fines and height differences may be driven by the time trend since both are correlated with the trend.

dummies and province-specific time trends are also controlled for.

Table 8 reports the estimates of equation (3) and some of its extensions, with standard errors clustered at province-year levels. The first column shows that increasing the fine by one year's income suggests an increase in twins' height gap by 1.4 centimeters on average. This result rules out the possibility that policy-associated twins are MZ, and provides evidence that people may have employed methods to create twins by themselves (by taking drugs or false reporting). We then add interactions of the fines variable with same-gender and different-gender twins in the same regression to identify the one-child policy's impact on the height difference of different types of twins. As shown in Column 2, policy fines are significantly and positively associated with same-gender twins' height differences, and the policy has no significant relation with that of different-gender pairs, given that the coefficients of the two interactions are 1.6 (0.5) and 0.5(1.3), respectively. Furthermore, the third column reports that the regression result using only the same-gender sample, with an additional indicator of whether the pair is female controlled for.<sup>18</sup> The coefficient of the fine becomes more significant, and a one standard deviation increase in the fine is associated with a 0.44 standard deviation increase in the height gap of the same gender twins. The results in the first three columns seem to support the man-made twins hypothesis rather than the fake twins hypothesis because there is no evidence showing the one-child policy has a positive effect on different-gender twins' height differences, which is predicted in A.2.

Columns 4 through 6 present regression results with the dependent variable being the Gap/Height ratio, and the results are very consistent with those in the first three columns. Specifically, a one year income's increase in fines is significantly associated with the gap/height ratio increasing by 1.2 percentage points (Column 4), and the effect only exists significantly in same-gender twins (Columns 5 and 6).

[Table 8 about here]

#### 4.3 Robustness Checks

<sup>&</sup>lt;sup>18</sup> The regression for different-gender twins is not implemented due to a small sample.

Since CHNS is a panel data and one certain pair of twins may be surveyed in different waves, the height difference may change as the twins grow up because the individuals in one pair would have different environmental influences throughout their lives. To avoid potential econometric problems, we only keep one observation for each pair and run the same regressions as the first two columns in Table 8, and the results are reported in Table 9. In the first four columns of Table 9, we only keep the latest wave for the duplicated twins, and keep the earliest wave in the last four columns. Though such a small sample, the results in the first four columns still show a robust pattern. The estimates show that when the policy fine increases by one unit, the height gap of same gender twins will increase by about 2.5 centimeters, and their gap/height ratio will increase by 1.8 accordingly. The coefficients on fine or its interactions in the last four columns are not significant, and the reason may be larger measurement error when the twins are younger.

[Table 9 about here]

#### **IV. Conclusions**

Using 1982, 1990, 2000, and 2005 census data, this paper documents that the birthrate of twins in China has more than doubled from the late 1960s through the early 2000s, from 3.5 to 7.5 per thousand births. Though a similar phenomenon also exists in other countries like the U.S. and Japan, China may be a special case because of the one-child policy. To those who prefer more children or those with a strong preference for boys, twins are favorable because individuals are able to have more children, and have a higher probability of having a boy without going against regulations. Given such a strong incentive, individuals are likely to take measures to increase the probability of giving birth to twins, which may include taking fertility drugs, asking for help from private hospitals, and even hiding older children and registering them with younger ones as twins.

However, there is no previous research on this issue. This paper sheds some light on the question whether and how the one child policy and twins are related. Above all, through matching census data to one-child policy regulatory fine data by province and year, we find that

an increase in policy fines by one year's income is associated with an approximately 0.06 per thousand births increase in the rate of twins. The estimates in our preferred model indicates that about one third of the twins rate increase since 1970s can be explained by the one child policy. By dividing twins into same-gender and different-gender groups and doing a similar analysis, we find policy-associated twins are distributed almost equally between the two groups.

We then divide the full sample into urban and rural groups and add interacts of fines with birth order and mother's education level in the regressions. We find that the impact of the policy is larger in urban areas, and that the policy increases the twin birthrate in the first two births in urban areas, but higher orders in rural areas. We also find the heterogeneous effect of the policy on mothers with different education levels. These results indicate that twins may not be randomly distributed throughout the population, conditional on the age of the mothers giving birth.

Finally, we try to provide evidence on what ways those increase in twins comes from. Estimates show that height difference within twins is positively associated with policy fines and this association only exists in same-gender twins. This finding is consistent with the hypothesis that the one-child policy may have stimulated people's incentives to have twins by taking fertility drugs. We cannot, however, rule out the possibility of the fake twin hypothesis due to the small sample of different-gender twins in the CHNS sample.

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# Appendix: The One Child Policy and Height Differences between Twins

Though evidence for policy-associated twins has been provided in the census data, it is still unknown whether policy-associated twins are real "man-made" twins or fake twins.

In this section, different consequences of real "man-made" twins and fake twins hypothesis are derived, and empirical analysis is presented afterward. For simplicity, we only use financial penalties (fine) to measure the one-child policy, and assume it equals one when the financial penalty policy is established in the local province, and zero if otherwise.

# A.1 Real "Man-made" Twins Hypothesis

The man-made twins hypothesis indicates that individuals are motivated to take technologies, like fertility drugs, to give birth to twins, under the one-child policy.<sup>19</sup> The mechanisms of these drugs are usually to induce ovulation, which is usually used to treat infertility. When normal women take them, the possibility of multiple ovulation is increased, and thus they are more likely to have twins. Though these drugs are classified as prescription medicines, it is still possible for individuals to purchase them in certain private hospitals<sup>20</sup> and get prescriptions from certain doctors.<sup>21</sup>

Under the Man-Made Twins Hypothesis, individuals are more likely to take fertility drugs (Take = 1) to have more children while avoid being punished under the one-child policy, that is, Pr(Take = 1|Fine = 1) > Pr(Take = 1 | Fine = 0). According to the medical literature and stylized facts, we know that taking certain fertility drugs or using technologies really increases the probability of giving birth to twins: Pr(Twins = 1|Take = 1) > Pr(Twins = 1 | Take = 0). We also assume that, conditional on individuals' behaviors (Take), giving birth to twins is independent of the one-child policy. In addition, the biological results of fertility drugs or embryo technologies are to develop multiple zygotes in the uterus at the same time, rather than to stimulate a single fertilized egg in the mother's body to divide into two or more embryos. Thus,

<sup>&</sup>lt;sup>19</sup> Taking drugs should be more reasonable because embryo technologies did not appear in China until the late 1990s.

<sup>&</sup>lt;sup>20</sup> Some even have advertisements on Google to "ensure" twins.

<sup>&</sup>lt;sup>21</sup> They may bribe the doctors or tell lies.

these actions only raise the probability of dizygotic (DZ) twins rather than that of monozygotic (MZ) twins, <sup>22</sup> that is, Pr(DZ = 1 | Take = 1) > Pr(DZ = 1 | Take = 0) and Pr(MZ = 1 | Take = 1) = Pr(MZ = 1 | Take = 0).

According to the medical literature, MZ twins are genetically nearly identical and they are always of the same sex unless there has been a mutation during development. But it is possible that same-gender twins are DZ. Certain characteristics of MZ twins become more alike as twins age, such as IQ and personality (Segal, 1999). DZ twins, however, like any other siblings, have an extremely small chance of having the same chromosome profile. DZ twins may look very different from each other, and may be of different sexes or the same sex. The above also holds for brothers and sisters from the same parents, meaning that DZ twins can be viewed as siblings who happen to be of the same age. Therefore, we have Pr(DZ = 1|DG = 1) = 1, Pr(SG = 1 | MZ = 1) = 1, 0 < Pr(DZ = 1|SG) < 1, and 0 < Pr(MZ = 1|SG) < 1.

An established strand of literature has proved that DZ twins tend to have more differences than MZ ones as they grow up because of genetic disparity, including height (Fischbein, 1977; Smith et al., 1973), weight (Stunkard et al., 1986), mental ability profiles (Segal, 1985), bone mass (Smith et al., 1973) and so on. Specifically, Fischbein (1977) found that MZ twins have a significantly higher concordance in height than for DZ pairs during puberty, for both boys and girls, and yearly height increments are also more similar for MZ pairs, indicating that the height spurt occurs more simultaneously for MZ twins in comparison to DZ twins. Thus, we presume that the height difference (HD) within a DZ (same-gender) pair<sup>23</sup> should be larger than that of a MZ pair if other factors are equalized. Thus, E(HD|SG = 1, DZ = 1) > E(HD|SG = 1, MZ = 1).

Additionally, we also assume that the actions people may take do not influence twins' height differences conditional on these twins' type (MZ or DZ). Based on the facts or assumptions above, it can be shown that

<sup>&</sup>lt;sup>22</sup> There are only two types of twins, MZ and DZ.

<sup>&</sup>lt;sup>23</sup> Because MZ twins must be of the same sex, Pr(SG=1|MZ=1)=1

(1) E(HD|Fine=1,Twins=1) > E(HD|Fine=0,Twins=1)

(2) E(HD|Fine=1,SG=1) > E(HD|Fine=0,SG=1)

# and (3) E(HD|Fine=1,DG=1)=E(HD|Fine=0,DG=1)

Equation (1) tells us that twins' height differences would be larger under the one-child policy because there will be more DZ twins due to the methods the individuals take in response to the one-child policy. This response will increase DZ proportions in same-gender twins, and thus the height difference in this group will become larger (Equation 2). However, different-gender twins' height differences will not change because they themselves are DZ (Equation 3). Because we do not know whether a woman took fertility drugs or not.<sup>24</sup> equations (1) through (3) are important because of height difference, one-child policy fine and twins' gender composition are all observables, and thus allow for empirical tests.

# A.2 Fake Twins Hypothesis

The fake twins hypothesis means that parents report siblings as twins in order to avoid punishment under the one-child policy. It is somehow feasible under some special circumstances in earlier China. First, many pregnant women gave birth at home in the 1980s, and the population administration would not notice until parents report their infants, though they were required to do so. Second, birth certificates did not launch until 1997,<sup>25</sup> and the children's birthdates were easy to revise before that. Third, children-especially siblings-look alike, in particular when the age difference is not large. Though parents will face harsher punishment once they are found to have reported fake twins, many parents may still choose to do so because of strong children or boy preferences.

Under the fake twins hypothesis, the one child policy stimulates people's incentives to report fake twins, that is,  $Pr(Twins^*|Fine=1) > Pr(Twins^*|Fine=0)$ , in which Twins^\* denotes the observed twins, including real ones and fake ones. For real twins (Twins), we assume all of them

<sup>&</sup>lt;sup>24</sup> Because CHNS does not have information and respondents may not tell the truth even they are asked. Whether a pair of twins is MZ or DZ is not provided by CHNS.<sup>25</sup> Though it may start in some areas earlier.

are reported, that is,  $Pr(Twins^*|Twins)=1$ .

For simplicity, we do not consider the gender factor in height differences in different-gender twins here,<sup>26</sup> then because of the age difference, the height difference within fake twins is supposed to be larger, so  $E(HD|Twins^*) > E(HD|Twins)$  if  $Pr(Twins|Twins^*) < 1$ .

The condition  $Pr(Twins|Twins^*) < 1$  ensures that fake twins do exist. If parents have strong children preference and do not care about the gender, then the gender composition of fake twins should be random,<sup>27</sup> so that height differences in both (observed) same-gender twins and different-gender ones should be larger. However, if parents have a strong boy preference, and they report siblings as twins only if the first baby is a girl and try to give birth to a boy in the next,<sup>28</sup> then the height difference within (observed) different-gender twins is expected to be larger. No matter which case it is, under the fake twins hypothesis, we must have

# (4) $E(HD|Twins^*,Fine=1) > E(HD|Twins^*,Fine=0)$ ,

# and (5) $E(HD|DG^*,Fine=1) > E(HD|DG^*,Fine=0)$ ,

in which  $DG^*$  denotes the observed different-gender twins. Same as before, equations (4) and (5) are all based on observables so that can be tested in empirical analysis. Same as the man-made twins hypothesis, the height difference also is expected to be larger under the one-child policy according to the fake twins hypothesis. However, the fake twins hypothesis predicts that the height difference within different-gender twins should be larger under the one child policy, which is different from the results in A.1. Such a difference provides us an identification strategy to differentiate the two hypotheses.

<sup>&</sup>lt;sup>26</sup> Conditional that boys may be taller than girls conditional on the same age, height gap, when defined as the taller sibling's height minus the shorter one's, is possible to be larger or narrower if the proportion of fake twins increase, when those with boy preferences are more likely to report fake twins if they first gave birth to a girl. In section 5.2.2, we try to replace it with the girl's height minus that of the boy's in the different-gender sample, and thus the following equation should be more reasonable, as well as equations (4) and (5).

<sup>&</sup>lt;sup>27</sup> Although the probability of boys is slightly higher in natural cases.

<sup>&</sup>lt;sup>28</sup> Use ultrasound technology to select.



Data source is 1982, 1990, 2000 and 2005 Census.



Data source is 1982, 1990, 2000 and 2005 Census.



Data source is 1982, 1990, 2000 and 2005 Census.



# Fig 2. One Child Policy Regulatory Fines, by Province

Data source is Ebenstein (2008).





NOTE: Same-gender twins sample is from China Health Nutrition Survey, and policy fine data are from Ebenstein (2008). The shadowed area represents the 95% Confidence Interval.



NOTE: Diffrent-gender twins sample is from China Health Nutrition Survey, and policy fine data are from Ebenstein (2008). The shadowed area represents the 95% Confidence Interval.

Table 1: Summary statistics

	(1)	(2)	(3)	(4)	(5)	
<b>X</b> 7 · 11	Г. Ц. – Ц. – – – – – – – – – – – – – – –		Parents are Han	Either parent		
Variables	Full sample	Full	Urban	Rural	minority	
Fine in years of income	1.03	1.01	1.27	0.91	1.30	
	(0.99)	(0.97)	(1.07)	(0.91)	(1.20)	
Twins (%)	0.57	0.58	0.66	0.55	0.46	
	(7.53)	(7.58)	(8.07)	(7.39)	(6.77)	
Male twins (%)	0.18	0.19	0.23	0.17	0.14	
	(4.27)	(4.31)	(4.77)	(4.12)	(3.68)	
Female twins (%)	0.18	0.18	0.21	0.17	0.17	
	(4.27)	(4.29)	(4.61)	(4.16)	(4.09)	
Same-gender twins (%)	0.37	0.37	0.44	0.34	0.30	
	(6.03)	(6.07)	(6.62)	(5.84)	(5.49)	
Different-gender twins (%)	0.20	0.21	0.21	0.21	0.16	
	(4.52)	(4.56)	(4.63)	(4.53)	(3.98)	
Rural area	0.73	0.72	0.00	1.00	0.81	
	(0.45)	(0.45)	0.00	0.00	(0.40)	
Both parents Han ethnicity	0.93	1.00	1.00	1.00	0.00	
	(0.26)	0.00	0.00	0.00	0.00	
Age	8.06	8.09	8.72	7.84	7.65	
	(4.67)	(4.67)	(4.69)	(4.64)	(4.63)	
Mother's age when giving birth	23.26	23.29	24.31	22.89	22.83	
	(2.97)	(2.95)	(3.05)	(2.82)	(3.18)	
Mother's education level						
Illiterate	0.21	0.20	0.07	0.25	0.29	
	(0.40)	(0.40)	(0.25)	(0.43)	(0.46)	
Primary	0.38	0.38	0.24	0.44	0.41	
	(0.49)	(0.49)	(0.43)	(0.50)	(0.49)	
Middle high	0.30	0.31	0.41	0.27	0.22	
	(0.46)	(0.46)	(0.49)	(0.44)	(0.41)	
Senior high or above	0.11	0.11	0.28	0.04	0.08	
	(0.31)	(0.31)	(0.45)	(0.20)	(0.28)	
<u>Birth order</u>						
First	0.57	0.57	0.72	0.51	0.51	
	(0.50)	(0.49)	(0.45)	(0.50)	(0.50)	
Second	0.30	0.29	0.22	0.33	0.32	
	(0.46)	(0.46)	(0.41)	(0.47)	(0.47)	
Third or above	0.14	0.13	0.07	0.16	0.18	
	(0.34)	(0.34)	(0.25)	(0.37)	(0.38)	
Observations	4310227	4034009	893084	3140925	276218	

Notes: Standard deviations in parentheses. The sample is the births before 2001.

	1 5		
	(1)	(2)	(3)
	Born no later than 1979	Born after 1979	Difference $(2) - (1)$
Panel A: Parents Han ethnicity			
Policy fine rate		1.42	
		(0.86)	
Twins	0.39	0.66	0.264
	(6.26)	(8.08)	
Same-gender twins	0.28	0.41	0.133
	(5.26)	(6.39)	
Different-gender twins	0.12	0.25	0.132
	(3.40)	(4.97)	
Observations	1866384	2198616	
Panel B: Either of Parents Mino	rit <u>y</u>		
Twins	0.38	0.48	0.106
	(6.12)	(6.93)	
Same-gender twins	0.27	0.31	0.036
	(5.20)	(5.53)	
Different-gender twins	0.11	0.18	0.070
-	(3.24)	(4.18)	
Observations	104496	175118	

# Table 2: Twins rate before versus after one child policy

Notes: Standard deviations are in parentheses, and standard errors are in brackets. Only the sample in which both parents belong to Han ethnicity is kept, and the births born later than 2001 are dropped.

	(1)	(2)	(5)
	Full sample	Parents Han	Either Parent Minority
Panel A: Twins			
Policy fine rate	0.0587***	0.0652***	0.0118
	(0.0218)	(0.0232)	(0.0399)
Panel B: Different-gender twins			
Policy fine rate	0.0321**	0.0316**	0.0199
	(0.0146)	(0.0151)	(0.0326)
Panel C: Same-gender twins			
Policy fine rate	0.0266*	0.0336**	-0.00801
	(0.0150)	(0.0163)	(0.0317)
Observations	4,310,227	4,034,009	276,218

NOTE: Robust standard errors in parentheses are clustered in province-year level. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1. One child policy fine is measured in years of local household income. Sampling weights are applied. Coefficients should be interpreted as percentage because all dependent variables have been multiplied by 100. Covariates include residency type, parents' ethnicity, province, birth year, birth order, survey year, mother's education level, mother giving birth age, if applicable.

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	Yea	r +1	Yea	r +2	Yea	r +5
			One Child	Policy Fine		
Twins	-0.00237		-0.00482		0.00881	
	(0.00668)		(0.00684)		(0.00636)	
Different-gender twins		-0.0135		-0.0182		0.0152
		(0.0110)		(0.0115)		(0.0106)
Same-gender twins		0.00420		0.00313		0.00501
-		(0.00837)		(0.00848)		(0.00794)
Observations	2,567,099	2,567,099	2,548,242	2,548,242	2,478,952	2,478,952
R-squared	0.752	0.752	0.761	0.761	0.800	0.800
F tests		0.875		1.320		1.229
P values		0.417		0.267		0.292

Table 4: One child policy fine predicted by prior twins rate in post-1979 births

NOTE: Robust standard errors in parentheses are all clustered in province-year level. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1. Only post-1979 births are used. Covariates include continuous variables like provincial time trend, and indicator variables, like residency type, parents' ethnicity, province, birth year, birth order, survey year, mother's education level mother giving birth age, if applicable.

	(1)	(2)	(3)
Dependent variables	Twins	Different-gender twins	Same-gender twins
Birth year >= 1980 interacting	0.117***	0.0433*	0.0740**
with Parents are Han	(0.0396)	(0.0221)	(0.0329)
Observations	4,310,227	4,310,227	4,310,228
R-squared	0.001	0.001	0.002

# Table 5: Twins, one child policy and difference-in-differences

NOTE: Robust standard errors in parentheses are clustered in household level. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1.Sampling weights are applied. Coefficients should be interpreted as percentage because all dependent variables have been multiplied by 100. Covariates include residency type, parents' ethnicity, province, birth year, birth order, survey year, mother's education level, mother giving birth age, if applicable.

	Ι	Dependent variable is Twin	S
-	Dononto Llon	Subsamples by t	type of residence
	Parents Han	Urban	Rural
Panel A: Policy Fine			
Policy fine rate	0.0652***	0.0827**	0.0570**
	(0.0232)	(0.0340)	(0.0273)
Panel B: Policy Fine interac	ting with birth order		
First birth order	0.0387	0.0752**	0.0135
	(0.0248)	(0.0354)	(0.0284)
Second birth order	0.131***	0.127**	0.126***
	(0.0283)	(0.0513)	(0.0335)
Third or higher	0.0601*	0.0495	0.0620
-	(0.0327)	(0.0669)	(0.0384)
Panel C: Policy Fine interac	eting with mother's education	on level	
Illiterate	0.0540	0.197**	0.0195
	(0.0332)	(0.0823)	(0.0354)
Primary school	0.0777***	0.138***	0.0512*
	(0.0259)	(0.0478)	(0.0287)
Middle high	0.0671***	0.0620*	0.0736**
	(0.0253)	(0.0369)	(0.0312)
Senior high or above	0.0397	0.0682*	0.0325
	(0.0310)	(0.0403)	(0.0627)
Observations	4,034,009	893,084	3,140,925

# Table 6: Heterogenous effects of policy fine

NOTE: Robust standard errors in parentheses are clustered in province-year level. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1. One child policy fine is measred in years of household income. Sampling weights are applied. Coefficients should be interpreted as percentage because all dependent variables have been multiplied by 100. Covariates include residency type, parents' ethnicity, province, birth year, birth order, survey year, mother's education level, mother giving birth age, if applicable.

Table 7: Summary	statistics in	CHNS
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	(1)	(2)	(1)
		By type	e of twins
Variable	All twins	Same gender	Different gender
Panel A: Biomarkers			
Height gap (cm)	2.18	1.30	4.65
	(2.74)	(1.59)	(3.68)
Mean of height (cm)	121.15	119.58	125.55
	(27.59)	(28.47)	(25.17)
Gap/Height Ratio	1.75	1.02	3.78
	(2.14)	(1.14)	(2.90)
Boy is taller in the pair $(1 = Yes)$			0.58
			(0.51)
Panel B: One-child policy			
Fine in years of income	1.47	1.55	1.26
	(0.66)	(0.68)	(0.56)
Panel C: Demographics			
Male twins $(1 = Yes)$	0.39	0.53	
	(0.49)	(0.50)	
Female twins $(1 = Yes)$	0.35	0.47	
	(0.48)	(0.50)	
Different gender twins $(1 = Yes)$	0.26		
	(0.44)		
Urban $(1 = Yes)$	0.26	0.19	0.47
	(0.44)	(0.40)	(0.51)
Mother's age when giving birth	27.57	26.79	29.86
	(3.64)	(3.10)	(4.24)
Age in years	8.48	8.42	8.63
	(4.64)	(4.82)	(4.20)
Observations	72	53	19

NOTES: Standard deviations are reported in brackets. Data source is China Health and Nutrition Survey. Twins are defined as children (aged below 18) born in the same household within the same month. For each pair of twins, height gap is defined as the height of the taller member minus that of the shorter one. Gap/Height ratio is defined as twins' height gap divided by the mean height of the pair, and the values reported are multiplied by 100.

Table 8: Height difference and one-child policy fine

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable		Height gap (cr	n)	Gap/Height	Ratio (Mutip	olied by 100)
	Full	sample	Same gender	Full sa	ample	Same gender
Fine in years of income	1.434**		0.944***	1.227**		0.777***
	(0.541)		(0.157)	(0.447)		(0.129)
Interactions						
Same gender pair & Fine		1.591***			1.240***	
		(0.516)			(0.415)	
Different gender pair & Fine		0.502			1.145	
		(1.342)			(1.093)	
Same gender pair $(1 = Yes)$	-3.746***	-5.297**		-3.093***	-3.229*	
	(1.357)	(2.197)		(1.112)	(1.718)	
Mean of height (cm)	-0.0233	-0.0260	0.0233	-0.0139	-0.0141	0.0218
	(0.0776)	(0.0787)	(0.0334)	(0.0557)	(0.0563)	(0.0179)
Boy is taller in the pair $(1 = Yes)$	1.548	1.264		1.252	1.227	
	(1.458)	(1.500)		(1.135)	(1.191)	
Female twins $(1 = Yes)$			-0.911			-0.418
			(0.532)			(0.356)
Urban ( $1 = Yes$ )	0.622	0.600	1.097**	0.786	0.784	0.660**
	(0.935)	(0.916)	(0.511)	(0.676)	(0.681)	(0.312)
Age	0.429	0.482	-0.446	0.275	0.279	-0.476
c	(0.600)	(0.617)	(0.622)	(0.421)	(0.427)	(0.356)
Age square /100	-0.125	-0.370	0.577	-0.0466	-0.0681	0.410
	(2.157)	(2.224)	(0.790)	(1.535)	(1.557)	(0.489)
Mother's age when giving birth	-0.171	-0.134	-0.249***	-0.181*	-0.178*	-0.208***
	(0.132)	(0.130)	(0.0518)	(0.102)	(0.103)	(0.0344)
Constant	4.887	5.579	6.682	5.589	5.650	6.319**
	(5.398)	(5.633)	(4.392)	(3.736)	(3.862)	(2.901)
Observations	72	72	53	72	72	53
R-squared	0.690	0.696	0.873	0.714	0.714	0.852
Year of birth group dummies	Yes	Yes	Yes	Yes	Yes	Yes
Wave dummies	Yes	Yes	Yes	Yes	Yes	Yes
Province dummies	Yes	Yes	Yes	Yes	Yes	Yes
Province interacting with time trend	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Standard errors in parentheses are clustered in province-year level. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1. Data source is CHNS. The twins sample used in this table are those born between 1979 and 2000. Height gap is defined as the height of the taller twin minus that of the shorter one, and gap/height ratio is defined as twins' height gap divided by the mean height of the pair. Each observation is derived from one pair of twins.

Table 9: Height difference and one-child policy fine

Ţ	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variable	Lat	test wave non-	-duplicated tw	vins	Earli	iest wave no	n-duplicated tv	vins
	Height g	gap (cm)	Gap/Hei	ght Ratio	Height g	ap (cm)	Gap/Heig	ht Ratio
Fine in years of income	2.536**		1.841**		1.055		0.837	
5	(1.031)		(0.686)		(0.931)		(0.640)	
Interactions					· · · ·			
Same gender pair & Fine		2.552**		1.837***		1.146		0.799
- *		(0.971)		(0.660)		(0.993)		(0.676)
Different gender pair & Fine		2.249		1.917		0.123		1.226
- *		(3.211)		(2.368)		(2.099)		(1.678)
Same gender pair $(1 = Yes)$	-1.928*	-2.364	-1.690**	-1.573	-4.321***	-5.759	-3.602***	-3.001
	(1.016)	(3.893)	(0.775)	(3.032)	(1.564)	(3.790)	(1.302)	(2.853)
Mean of height (cm)	-0.0706	-0.0736	-0.0825	-0.0817	0.0455	0.0461	0.0107	0.0105
	(0.0832)	(0.0880)	(0.0635)	(0.0659)	(0.143)	(0.144)	(0.122)	(0.129)
Boy is taller in the pair $(1 = Yes)$	1.205	1.150	1.089	1.104	1.093	0.873	1.130	1.222
• • • • •	(1.397)	(1.567)	(1.071)	(1.266)	(1.480)	(1.584)	(1.315)	(1.396)
Urban $(1 = Yes)$	1.017	1.001	0.616	0.620	0.418	0.400	0.674	0.682
	(1.353)	(1.346)	(1.014)	(1.026)	(1.228)	(1.257)	(0.901)	(0.948)
Age	1.315	1.355	1.219*	1.208*	-0.628	-0.576	-0.354	-0.376
-	(0.818)	(0.946)	(0.623)	(0.676)	(1.199)	(1.272)	(1.031)	(1.092)
Age square /100	-4.462	-4.596	-4.321*	-4.286*	3.954	3.683	2.352	2.465
	(3.024)	(3.518)	(2.210)	(2.441)	(3.432)	(3.771)	(2.882)	(3.054)
Mother's age when giving birth	-0.0142	-0.0102	-0.0564	-0.0575	-0.197	-0.178	-0.160	-0.168
	(0.219)	(0.206)	(0.166)	(0.160)	(0.173)	(0.204)	(0.175)	(0.184)
Constant	0.641	1.151	5.674	5.538	3.398	3.973	6.092	5.852
	(6.533)	(9.693)	(5.129)	(7.317)	(9.971)	(10.07)	(7.555)	(8.320)
Observations	33	33	33	33	33	33	33	33
R-squared	0.829	0.830	0.839	0.839	0.825	0.829	0.843	0.844
Province dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province interacting with time trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Standard errors in parentheses are clustered in province-year level. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1. Data source is CHNS. The twins sample used in this table are those born between 1979 and 2000. Height gap is defined as the height of the taller twin minus that of the shorter one, and gap/height ratio is defined as twins' height gap divided by the mean height of the pair. Each observation is derived from one pair of twins.

	1995 Cesus Survey								
Voor of hirth -	By different gender		Total	By MZ	Z or DZ	Twins r	ate (per 1000 de	eliveries)	
	Both male	Both female	Different gender	Total	MZ	DZ	MZ	DZ	Total
1980	3667	3749	1871	9287	5545	3742	3.46	2.34	5.80
1981	3718	3846	1807	9371	5757	3614	3.76	2.36	6.12
1982	3640	3711	1836	9187	5515	3672	3.65	2.43	6.08
1983	3735	3854	1771	9360	5818	3542	3.86	2.35	6.21
1984	3627	3758	1773	9158	5612	3546	3.77	2.38	6.15
1985	3491	3672	1735	8898	5428	3470	3.77	2.41	6.18
1986	3406	3306	1750	8462	4962	3500	3.60	2.54	6.14
1987	3417	3405	1651	8473	5171	3302	3.83	2.45	6.28
1988	3300	3317	1722	8339	4895	3444	3.73	2.63	6.36
1989	3279	3376	1925	8580	4730	3850	3.73	3.03	6.76
1990	3162	3266	1829	8257	4599	3658	3.74	2.97	6.71
1991	3177	3226	1881	8284	4522	3762	3.74	3.11	6.85
1992	3317	3165	2017	8499	4465	4034	3.70	3.34	7.04
1993	3235	3316	2120	8671	4431	4240	3.74	3.58	7.32
1994	3521	3489	2280	9290	4730	4560	3.94	3.80	7.74
1995	3377	3488	2373	9238	4492	4746	3.77	3.98	7.75
Total	55069	55944	30341	141354	80672	60682	3.73	2.81	6.54

Appendix Table 1: Twins in Japan

Note: From Journal of Population Problems, 1998: 13-35.