

# **The Evolution of China's One-child Policy and Its Effects on Family Outcomes\* Online Appendix**

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## I. Empirical Framework

The major difficulty in estimating the effect of the one-child policy is that its introduction in 1979 coincided with that of the open-door policy and economic reforms. Following Li and Zhang (2016), my empirical identification is to explore the heterogeneity of the intensity of the one-child policy implementation across provinces/prefectures. Specifically, I construct a measure based on the excess births of each province/prefecture conditional on initial births and other variables. The important variable in my analysis is the excess fertility rate (EFR), which is defined as below:

$$EFR_j = \frac{\sum_i (Birth_{ij} \cdot 1(NSC_{ij} \geq 2) \cdot 1(25 \leq Age_{ij} \leq 44))}{\sum_i (1(NSC_{ij} \geq 1) \cdot 1(25 \leq Age_{ij} \leq 44)) - \sum_i (Birth_{ij} \cdot 1(NSC_{ij} = 1) \cdot 1(25 \leq Age_{ij} \leq 44))}$$

where  $Birth_{ij}$  is a dummy indicator for woman  $i$ , within an age range of 25-44 years old, in prefecture  $j$  giving a birth in 1981, and  $NSC_{ij}$  is the number of surviving children of women  $i$  in prefecture  $j$  by the end of 1981. In other words, the EFR in a prefecture here is defined as the percentage of Han mothers (i.e. women with at least one surviving child) aged 25-44 in the 1982 census who gave a higher order birth in 1981.

Using the EFR to represent the extent of violation of the one-child policy in prefecture  $j$ , I can further examine the effect of the policy on various family outcome variables. An outcome variable ( $y$ ) for an individual  $i$  at time  $t$  in prefecture  $j$  is related to the EFR in prefecture  $j$  as follows:

$$y_{ijt} = (EFR_j * T_i) * \alpha_1 + X_{ijt} * \gamma_1 + (C_j * T_i) * \delta_1 + \phi_j + \lambda_t + u_{ijt} \quad (1)$$

where  $T_i$  is a time dummy that equals to 0 (1) if the survey time is 1982 (1990);  $C_j$  is the set of prefectural level control variables;  $X_{ijt}$  is the set of individual/household level control variables;  $\phi_j$  and  $\lambda_t$  are respectively prefectural fixed effects and time fixed effect.

The specification is akin to a difference-in-differences (DiD) estimation.<sup>1</sup> The heterogeneity in the policy intensity across regions is one source of the variations (e.g. policy-strict areas vs. policy-loose areas). I take 1982 as the benchmark (it would have been ideal if I had data for 1979 or 1980). Suppose  $y$  is fertility or family size. The one-child policy affected family size (or number of young children) for only 2 years by 1982 but already affected family size for about 10 years by 1990. Family size is a stock measure, and the interaction term of  $EFR * T$  reflects the differential policy effect on the family size from 1982 to 1990. The coefficient on  $EFR * T$  largely corresponds to a DiD estimate.  $EFR$  and  $T$  each do not appear separately in the equation as usual because of the presence of the fixed effects  $\phi_j$  and  $\lambda_t$ .

The set of prefecture-level control variables,  $C_j$ , includes the average total number of births of females aged 45-54 years old; the shares of females aged 25-44 with 1, 2, 3, and 4+ births; the shares of females aged 25-29, 30-34, 35-39, and 40-44; the agricultural sector's employment share among adults aged 25-49 by gender; the shares of each education level category among adults aged 25-49 by gender. The set of household control variables,  $X_{ij}$ , includes the mother's age at first birth, the first child's age, the mother/father's education level, and the mother/father's employment sector.

Corresponding to the above individual-level regression, there are two “macro-level”

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<sup>1</sup> As noted in Li and Zhang (2016), this specification is analogous to Duflo (2001).

regressions at the prefecture level.

The fixed effect regression is given by

$$E_{jt}(y_{ijt}) = (EFR_j * T_i) * \alpha_1 + E_{jt}(X_{ijt}) * \gamma_1 + (C_j * T_i) * \delta_1 + \phi_j + \lambda_t + E_j(u_{ijt}) \quad (2)$$

where E stands for the expectation or average of the variable in prefecture j.

And the first difference regression is given by

$$E_{j90}(y_{ijt}) - E_{j82}(y_{ijt}) = (EFR_j) * \alpha_1 + (E_{j90}(X_{ijt}) - E_{j82}(X_{ijt})) * \gamma_1 + (C_j) * \delta_1 + v_j \quad (3)$$

Both of the macro-level regressions are weighted by the number of observations in each prefecture in 1982.

A key concept of my analysis is the EFR residual. Using the Frisch-Waugh-Lovell Theorem to obtain an equivalent regression in a residual form:

$$E_{j90}(y_{ijt}) - E_{j82}(y_{ijt}) = \theta_1 + (E_{j90}(X_{ijt}) - E_{j82}(X_{ijt})) * \beta_1 + C_j * \eta_1 + \mu_j \quad (4)$$

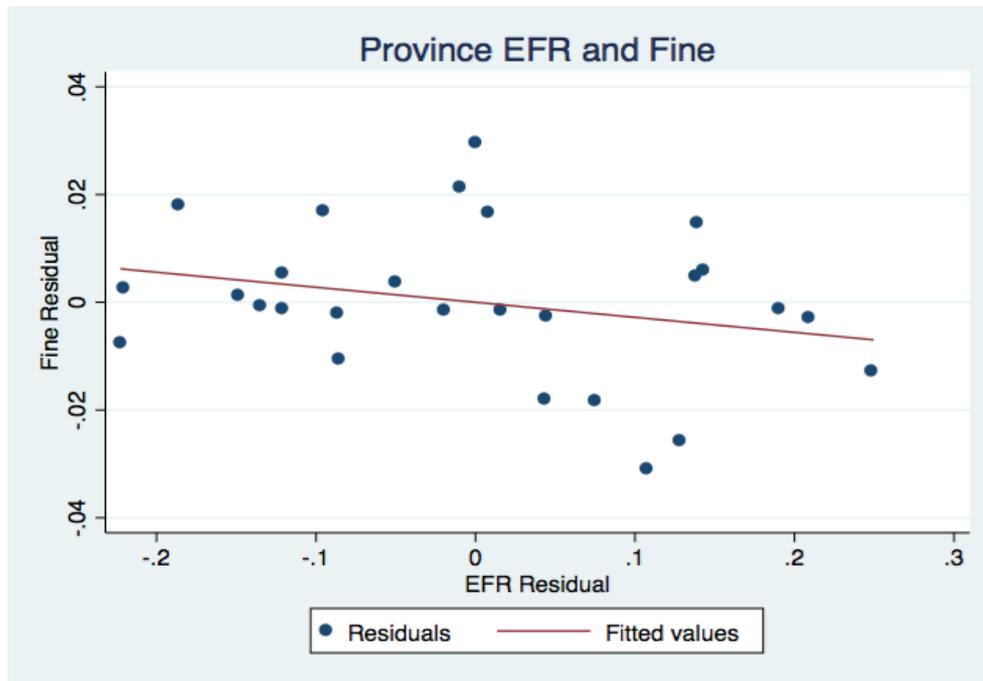
$$EFR_j = \theta_2 + (E_{j90}(X_{ijt}) - E_{j82}(X_{ijt})) * \beta_2 + C_j * \eta_2 + \varepsilon_j \quad (5)$$

Here I regress the family outcome variable (in its first difference form) and the  $EFR_j$  on the control variables to net out their influences. In Regression (4), I obtain  $\mu_j$  as the residual for the difference in the outcome variable. In Regression (5), I obtain  $\varepsilon_j$  as the EFR residual. Then I regress  $\mu_j$  on  $\varepsilon_j$ , and obtain the same estimate on  $\alpha_1$  as in Regressions (2) or (3). Controlling for the differences in pre-existing fertility preferences and socio-economic characteristics, which can remove both the demand for children and the possible policy responses to the demand, the EFR residual can proxy for regional differences in the one-child policy enforcement intensity. The measure takes into account both the harshness of the local fertility policy and the stringency of policy compliance in practice. A larger EFR residual represents a more relaxed policy.

While all these reduced-form regressions are reported in the fixed effects form, all graphs are drawn in the form of the residual of the outcome variable difference on the EFR residual (i.e. the first-difference form). This will give us a clear visual representation of a DiD estimate of the one-child policy effect.

*Figure 1*

**Province EFR and Fine**



*Source:* Data resource of the fine, Ebenstein (2010). "The “missing girls” of China and the unintended consequences of the one child policy." *Journal of Human Resources* 45(1): 87-115.

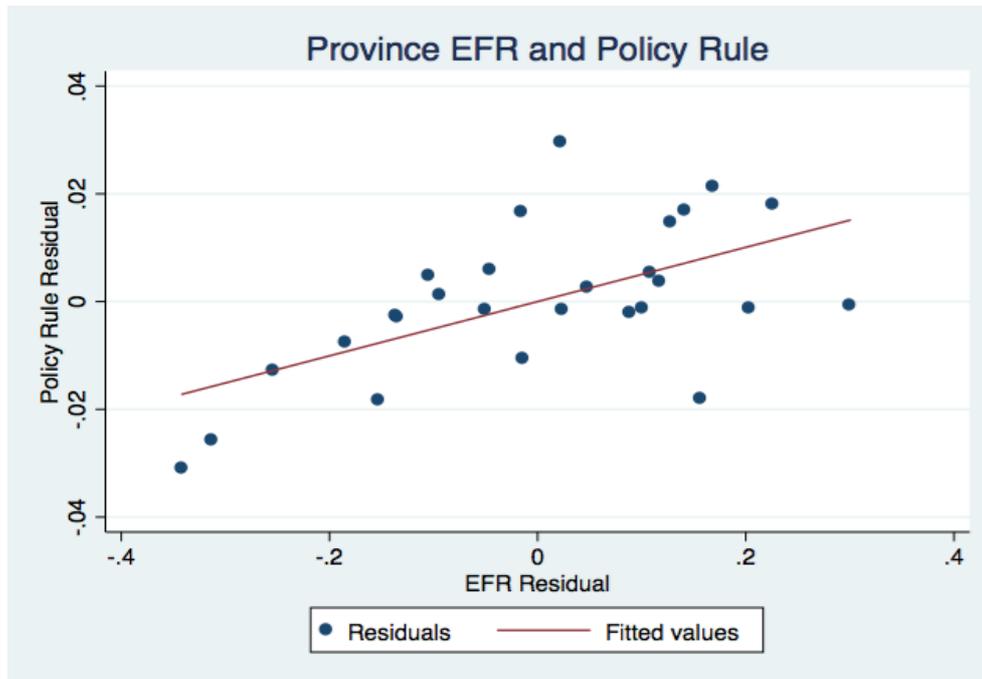
*Note:* The fine is defined as the fine rates in years of household income. For example, in 1980 Guangdong Province ratified a fine of approximately 1.21 years of household income.

To gauge the plausibility of the EFR residual as the policy intensity, I look at its correlation with two direct policy measures that are available at the province level. Figure 1 shows a negative correlation between the EFR residual and the level of fines imposed on above-quota births. The correlation of the EFR residual and the fine is -0.2648. Provinces with harsher policy have both higher fines on above-quota births and lower EFR residuals. Figure 2 shows a positive correlation between the EFR residual and the number of children allowed in rural areas of China. The correlation of

the EFR residual and Policy Rule is 0.5917. Provinces with more relaxed policy have both higher EFR residuals and higher numbers of births in the rural areas.

Figure 2

### Province EFR and Policy Rule



Source: Data resource of Policy Rule, Wang. (2005). "Can China afford to continue its one-child policy?"

Notes: The majority of Chinese reside in rural areas, and fertility policies covering them fall into three broad categories (Ebenstein (2010)). For the first category, Policy Rule equals 1 in seven provinces wherein a mother was allowed to have one child only. For the second category, Policy Rule equals 1.5 in nineteen provinces where the mother was allowed to have one additional birth following a first born daughter. And for the third category, Policy Rule equals 2 in five provinces where the mother was allowed to have two or more children.

In my main empirical analysis, I use the data from China's 1982 and 1990 censuses. Table 1 shows a detailed statistical summary of the two key variables, the EFR and EFR Residuals, for the regression on fertility between 1982 and 1990. The mean and standard deviation of the EFR are 7.39% and 4.36%, respectively. Moreover, the minimum and maximum values of the EFR equal to 0 and 20.85%, respectively. (A perfect compliance of the one-child policy would mean that an EFR values at 0.) The

25th percentile and 75th percentile of the EFR equal to 4.24% and 9.76%, respectively. These statistics suggest a considerable variation in the realization of the one-child policy across the around 290 prefectures in China

*Table 1*

**Statistical Summary of the EFR and EFR Residuals**

|           | EFR    | EFR Residual |
|-----------|--------|--------------|
| Mean      | 0.0739 | 0.0024       |
| Std. Dev. | 0.0436 | 0.0248       |
| Min       | 0.0000 | -0.0657      |
| Max       | 0.2085 | 0.0902       |
| P25       | 0.0424 | -0.0138      |
| P75       | 0.0976 | 0.0162       |

For the EFR residual, which is intended to measure the truly exogenous one-child policy intensity (net out the endogenous individual and community effects in the one-child policy by controlling for the pre-existing fertility preferences and socio-economic characteristics), the mean and standard deviation are 0 and 2.48%, respectively. Besides, the minimum and maximum values of the EFR residual equal to -6.57% and 9.02%, respectively, and the 25th percentile and 75th percentile of the EFR residual equal to -1.38% and 1.62% respectively. These numbers show a large variation in the one-child policy intensity across all prefectures.

By performing a crude comparison between the EFR and the EFR residual, I can gauge the extent of the policy intensity in the occurrence of the excess births. The difference between the maximum and minimum values of the EFR is about 21%. The difference between the maximum and minimum values of the EFR residuals is about 15%. This seems to indicate that overall, up to 71% of the maximum dispersion of the excess births could be attributed to the stringency of both local fertility policy and the

policy compliance in practice. However, if we look at the variances of the two measures, a simple calculation indicates that only about 29% of the EFR could be attributed to the policy stringency per se. Nevertheless, 29% is still a large fraction of the excess births occurred.

In my assessment of the policy effect later, I will look at the difference between the 25th and 75th percentiles of the EFR residuals to obtain a more generalizable effect. The 25th and 75th percentiles of the EFR are equal to 4.24% and 9.76%, respectively, and the 25th percentile and 75th percentile of the EFR residuals equals to -1.38% and 1.62%, respectively. The differences between the 25th percentile and 75th percentile of the EFR and the EFR residuals equals to 5.52% and 3%, respectively. Thus, looking at the issue this way, the EFR residuals account for 54.3% ( $0.03/0.0552$ ) of the EFR. In other words, approximately 54.3% of the EFR can reflect the regional differences in the one-child policy enforcement intensity which is slightly lower than the estimate based on the difference between the maximum and minimum values. The other 45.7% of the EFR has been controlled for by the pre-existing fertility preferences and socio-economic characteristics.

## **II. Micro Evidence**

In this section, I discuss the empirical results in the literature and from my own analysis using the empirical framework in Section I. I look at the overall effects of the one-child policy on all women aged 20-64 years and their family members. Table 2 reports my empirical results using the EFR approach. In the main text, I did not present the results on the effect of the one-child policy on fertility and education. Here I first show the empirical results for these two well-studied outcome variables using the EFR approach in subsection A. I then present and discuss the empirical results for the other family outcome variables in great details in subsection B. Additionally, I conduct a further analysis of all the family outcome variables as a robustness check in

subsection C by separately looking at the results for different age groups of women. In my basic analysis, I use the 1982 census data to calculate the EFR and examine the outcome variables in 1990 (except migration that uses the 2010 census data).

How does the one-child policy affect family outcomes? I attempt to analyze the effects on six factors which are closely related to the family. First, the one-child policy is aimed to control population and should have reduced fertility. The magnitude of the estimates of the impact varies across several studies. Second, the quantity-quality tradeoff theory implies that a reduction in fertility can enhance human capital investment per child. The one-child policy may help to increase children's education. Third, low fertility should reduce the youth dependency ratio. Fourth, low fertility may prompt an increase in the number of divorces. Fifth, lower fertility may increase parental labor supply, which is conducive to economic development. Sixth, low fertility may facilitate parental migration, thereby enabling the reallocation of the labor force from rural to urban areas.

Numerous studies have looked into the effect on fertility and children education, and some have investigated the effect on the dependency ratio, but only little or no research has been conducted on marital status, labor supply and migration in the context of the one-child policy in China. I attempt to review the literature (if available) and provide novel evidence. There has been no systematic analysis of the one-child policy on multiple family outcomes in the literature.

### **A. The Effect of the One-child Policy on Fertility and Children's Education**

In this subsection, I discuss the empirical results for the effects of the one-child policy on fertility and children's education using the framework in Section I. Both results are consistent with the outcomes in the literature.

Table 2

**Regression Results for Six Outcome Variables, 1982-1990**

| Dependent Variable                  | Coefficient | Std. Err. | Obs. |
|-------------------------------------|-------------|-----------|------|
| (1)Fertility                        | 1.205***    | 0.229     | 586  |
| (2)Children's Education Attainment  | 0.552       | 0.512     | 582  |
| (3)Youth Dependency Ratio           | 0.873***    | 0.116     | 586  |
| (4) Marital Status (Being Divorced) | -0.00538*** | 0.00139   | 586  |
| (5a)Female Labor Supply             | 0.037       | 0.0731    | 586  |
| (5b)Male Labor Supply               | -0.0387***  | 0.00942   | 584  |
| (6)Rural Migration                  | -0.279***   | 0.052     | 530  |

Notes: For all of these regressions, the excess fertility rate residual was calculated using 1982 census data, and all the regressions used prefectural-level data weighted by the number of observations in each prefecture in 1982. The regression (1)-(5) used 1982 and 1990 census data, while the regression (6) on migration used 1982 and 2000 census data.

First, Row 1 of Table 2 shows the result for fertility. The coefficient has the expected positive sign and is statistically significant. As indicated in Figure 3, the magnitude of the effect is rather small. By construction, the EFR residual's mean is zero and has a standard deviation of 2.48%. The interquartile range of the EFR residual equals 3%, which is somewhat larger than one standard deviation. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual (i.e. the policy intensity goes from 1.62% to -1.38%) would lead to 0.0362 less births per household in 1990 (i.e.  $0.0300 \times 1.205$ ), representing a decrease of only about 1.40% in fertility (i.e.  $0.0362/2.579$ , where 2.579 is the mean number of children in 1982).<sup>2</sup> I

<sup>2</sup> Alternatively, we can look at the maximum and minimum values of the EFR residuals, which are 9.02% and -6.57% respectively, and the difference between the two values is 15.59%. A decrease of 0.1559 in the EFR residual would lead to a decrease of 0.1878 in births per family or cause a decrease of 7.28% in fertility, which implies that in the prefecture where the one-child policy is most strictly implemented, the

interpret the small effect of the one-child policy on fertility to indicate that, generally, had China waited for 10 years, the fertility outcome would have been very similar without the one-child policy.

Figure 3

**The Overall Effect of the One-child Policy on Fertility**



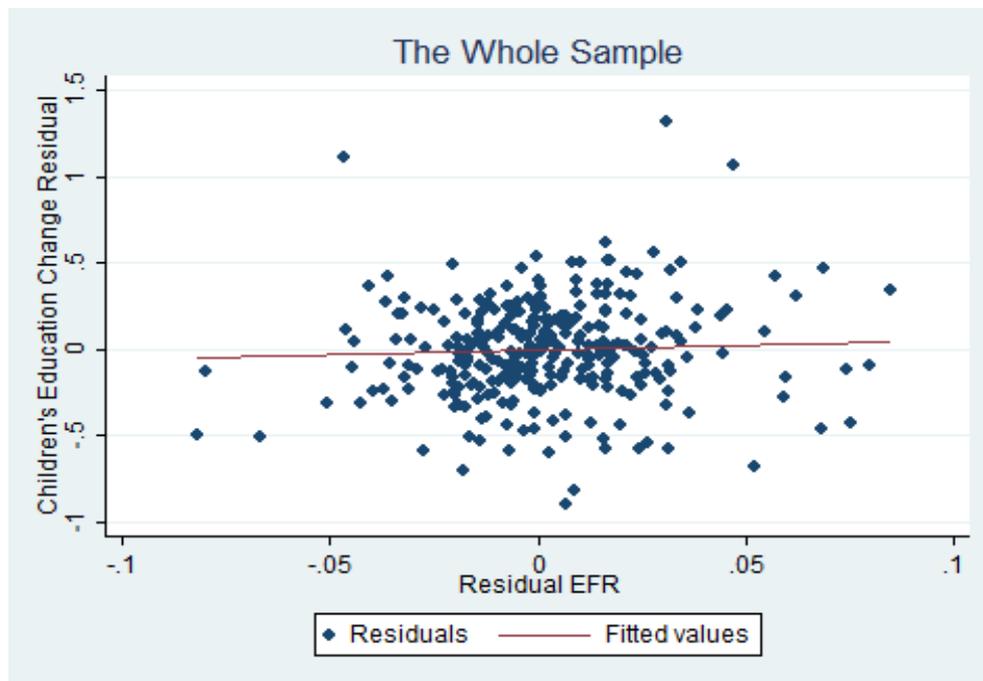
Second, Row 2 of Table 2 shows my estimation on children’s schooling years. Several rigorous studies mentioned in the main text indicate that fertility decline has either a modest or little effect on children education. My empirical estimation (0.552) is also not statistically significant. As indicated in Figure 4, the curve is almost horizontal, implying that the magnitude of the effect is almost zero. This finding is consistent with the small estimates in the literature. I will return to this issue when presenting results by separate age groups.

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fertility declines much more substantially than that in the prefecture where the one-child policy is most relaxed. Therefore, the effect of the one-child policy on fertility in these prefectures where the policy is strictly implemented is more considerable, but it is still not large.

Figure 4

### The Overall Effect of the One-child Policy on Child Education



## B. The Effect of the One-child Policy on Other Family Outcome

### Variables

In the main text, I only report the numerical estimation results for the other family outcome variables. In this subsection, I discuss the empirical results in the literature as well as my own estimation for the other family outcome variables (dependency ratio, marital status, labor supply and rural migration) in great details.

#### B.1 Dependence Ratio

Many studies have documented the falling youth dependence ratio in the short term and the rising elderly dependence ratio in the long term. For example, Poston (2000) investigated the factors leading to the huge current and projected number of the elderly population in China, which reflected the fertility transition happening in the

country since the 1970s. He found that the dependency burden on productive population in China would be increasing in the next few decades.<sup>3</sup> Hussain (2002) argued that the onset of the age structural change in 1970 predated the one-child policy. Meanwhile, until 2002, the most significant change occurred in the proportion of children in the population (decreasing by almost 17 percentage points from 40.4% in 1962 to 23.9% in 1998). Another large change was the rising share of working-age adults by 13.3 percentage points (largely due to the arrival of baby boomers in the 1960s). However, as a result of the demographic transition, the composition of the dependent population would change from a young to an elderly population.

Hesketh et al. (2005) pointed out that China is experiencing an increase in the proportion of the elderly population and an increase in the ratio between elderly parents and adult children. The percentage of the population over the age of 65 years was 5 percent in 1982 and increased to 7.5 percent in 2004 (and further increased to 10.5 percent in 2015), but was expected to rise to more than 15 percent by 2025. They argued that a lack of adequate pension coverage in China meant that financial dependence on offspring would still be necessary for approximately 70 percent of elderly people. This problem was called the “4:2:1” phenomenon in China, meaning that an increasing number of couples would be solely responsible for the care of one child and four parents. Zhang and Goza (2006) also argued that the trend of aging in China is producing profound social impact and the development of appropriate policies is required. They concentrated on the so-called sandwich generation, which means those who need to care for younger and older generations at the same time. Hu and Yang (2012) employed statistics from China’s Statistical Yearbook 2010 and found that the share of working age population began to decrease in 2015 in China because of the one-child policy and the longer life expectancy. The peak number of

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<sup>3</sup> In 1999, China had a modest old dependency ratio compared with the US and some other countries. However, the situation would be changed in 2050 when China is projected to turn into an “old” country. The old dependency ratio at that time in China would be higher than that in the US.

elderly people in China would be reached in 2038 as they projected, when the dependency ratio would increase to more than 50%. Zhang et al. (2015) examined a more detailed effect of the age structure on economic development and its possible channels. Using a panel data set composed of several demographic and economic variables for 28 provinces in China in four census year (1990, 1995, 2000 and 2005), this paper found that the ratio and the internal composition of the working-age population had significant effect on GDP per capital. The higher the working-age ratio and the prime-age share (the ratio of the age 35–54 years to the working age), the higher the GDP per capital. The authors argued that more than 19% of the GDP growth can be attributed to the favorable changes in the age structure during 1990-2005, especially through the influencing channel of the total factor productivity. However, owing to the strict one-child policy, the contribution of the advantageous demographic changes would gradually vanish as both the working-age ratio and the prime-age share would decrease.

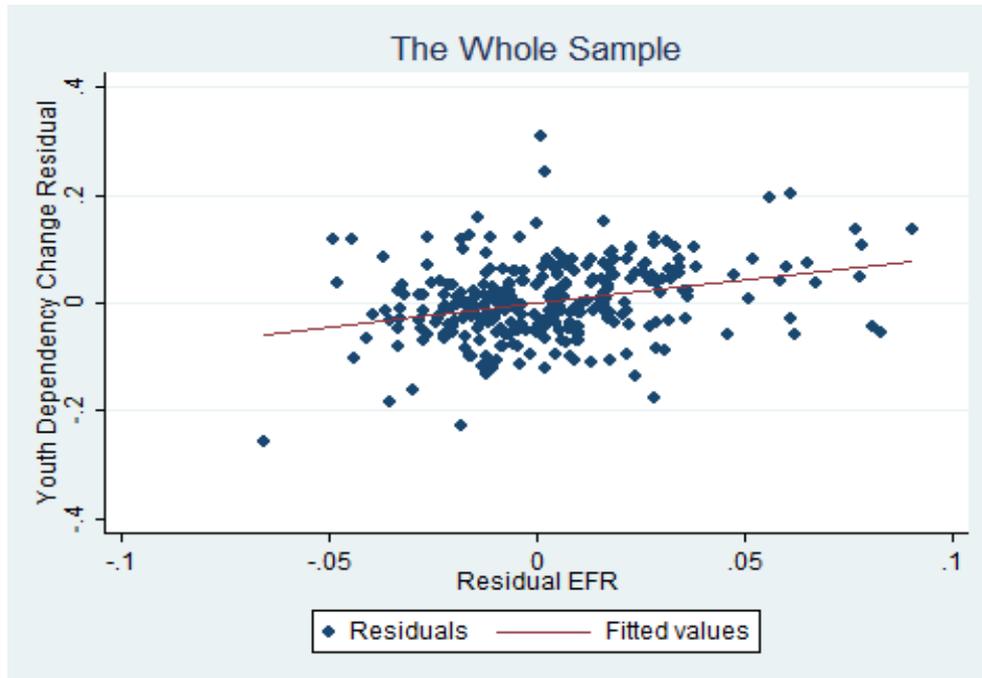
I have access to the data of the 1982, 1990 and 2000 censuses. For those born after 1979, they were only about 20 years old by 2000. Thus it is not meaningful to analyze the elderly dependence ratio using the data available because they are not old enough. Thus, I will focus on the youth dependence ratio.

Row 3 of Table 2 reports my results on the youth dependence ratio. The dependent variable is defined as the ratio of the number of individuals below 18 years old to the number of individuals between 18 and 64 in the household. The coefficient has the expected positive sign and is statistically significant. As indicated in Figure 5, the size of the effect is still rather small. The standard deviation of the EFR residual is 2.48% and the interquartile range of the EFR residual equals 3%. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual would decrease the youth dependency ratio in the household by 2.62 percentage points in 1990 (i.e.  $0.0300 \times 0.873$ ), representing a decrease of only about 2.46% in the youth dependency ratio (i.e.  $0.0262/1.063$ , where 1.063 is the mean of

the youth dependency ratio in 1982).

Figure 5

### Youth Dependency Ratio Regression



## B.2 Marital Status

The one-child policy may influence the marital outcomes through its impacts on fertility within marriage. As sources of joy or even future supporters, children may help stabilize marriage and increase marital benefits (Becker et al., 1977; Becker, 1991). Therefore, compulsory fertility restrictions because of the one-child policy may reduce potential marriage gains and consequently affect individuals' marital behaviors. Several studies have examined the possible effects of the one-child policy on marital outcomes, including inter-ethnic marriage and unmarried rates.

Huang and Zhou (2015a) employed Choo and Siow's (2006) transferable utility model of marriage and data from census 2000 and 1% mini-census in 2005 to identify the impact of the one-child policy on unmarried rates and Han-Minority marriages by using plausibly exogenous variations in the ethnicity-specific assigned birth quotas

and fertility penalties across Chinese provinces over time. They found that the one-child policy induced higher unmarried rates among the population, and the effect on unmarried status was greater and more significant for the Hans but insignificant and smaller for the minorities. They further found that the one-child policy resulted in more inter-ethnic marriages and allowed minorities in H-M marriages to marry Han spouses with higher education in the preferential-policy regions (that allowed a second birth for the minorities) by using non-preferential-policy regions as the control group. Another paper by Huang and Zhou (2015b) also reported that the likelihood of inter-ethnic marriage increased with fines in both rural and urban areas and the one-child policy accounted for one fourth of the rise in inter-ethnic marriage in the last two decades.

The one-child policy not only affects the marriage outcomes of the first generation by changing the expectations in the number of potential births and the relevant costs of having and raising a child, but also influences the marriage markets of the next generation through imbalanced sex ratios. Studies found that sex imbalance would increase males' unmarried rates, cause women to marry up and also affect marital satisfaction. For example, Du, Wang and Zhang (2015) studied the effects of sex-ratio imbalance on matching patterns in China's marriage markets by using data from China General Social Survey 2006 and Census 2000. They found that unbalanced sex ratios would lead to women's hypergamy and enhance the wife's relative bargaining power in the intra-household resource allocation. Cheng and Smyth (2015) also examined China's unbalanced sex ratio, marriage patterns and marital satisfaction.

Row 4 of Table 2 shows my estimation result for the likelihood of divorce for the marriageable population (females aged above 19 and males above 21). The dependent variable is a dummy indicating whether an individual is divorced or not. The coefficient has the expected negative sign and is statistically significant. As indicated in Figure 6, the magnitude of the effect is rather small. The standard deviation of the EFR residual is 2.55 percent and the interquartile range equals 2.85 percent. A

tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual is associated with an increase in the probability of divorce by 0.0153 percentage points (that is,  $0.0285 \times 0.00538$ ).

Figure 6

### Marital Status (Being Divorced) Regression



### B.3 Labor Supply

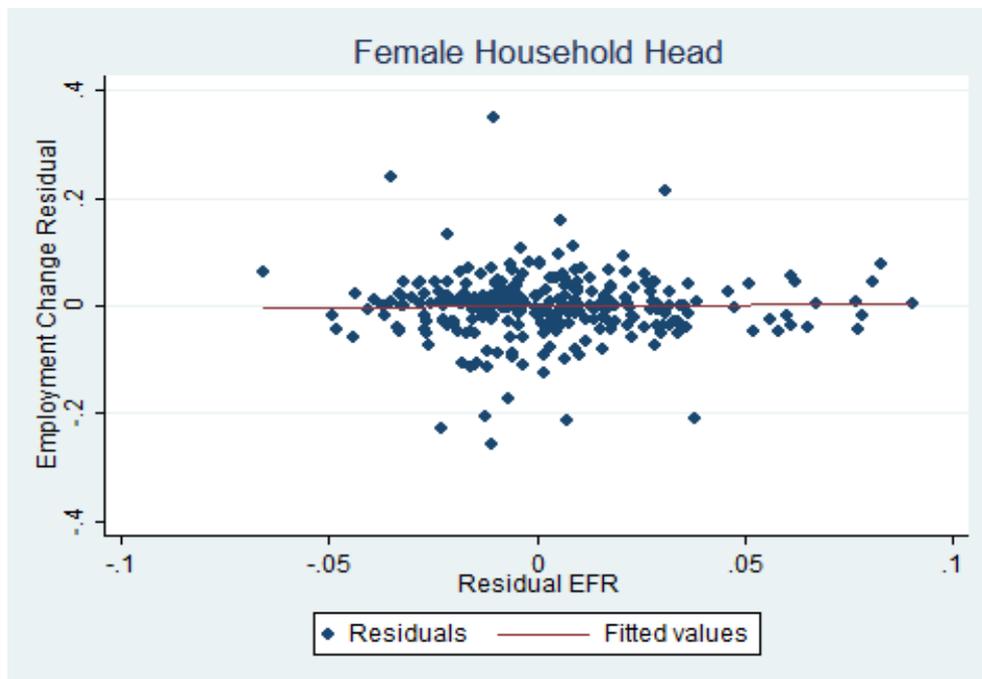
There seems to be no study on the effect of the one-child policy on parental labor supply. Two rather descriptive studies focused on other issues but delved on labor supply. Fong (2002) conducted a two-year fieldwork in Dalian during the late 1990s which included a survey covering 2273 students and some in-depth interviews, and found that urban daughters can benefit from the one-child policy. On the one hand, they enjoyed unprecedented parental support because they did not have to compete with brothers for parental investment. On the other hand, the one-child policy led to a low fertility and enabled mothers to acquire paid work, since women were freed from heavy child-rearing burdens. This fertility transition made women take part in education and labor force rather than in motherhood alone. As a result, mothers could

demonstrate their filiality by giving their parents financial support, which showed that daughters could provide their parents with support when they get old. Chen (1985) argued that the one-child policy would free women from the childrearing and reinforce the phenomenon that the labor participation rate of married Chinese women is pretty high.

Row 5 of Table 2 presents my estimation results on female labor supply. The dependent variable is a dummy indicating whether the female head (or the spouse of the household head) in a household participated in the labor force. The coefficient (0.037) is statistically insignificant. As indicated in Figure 7, the curve is almost horizontal, implying that the size of the effect is almost zero.

Figure 7

### Female Labor Supply Regression

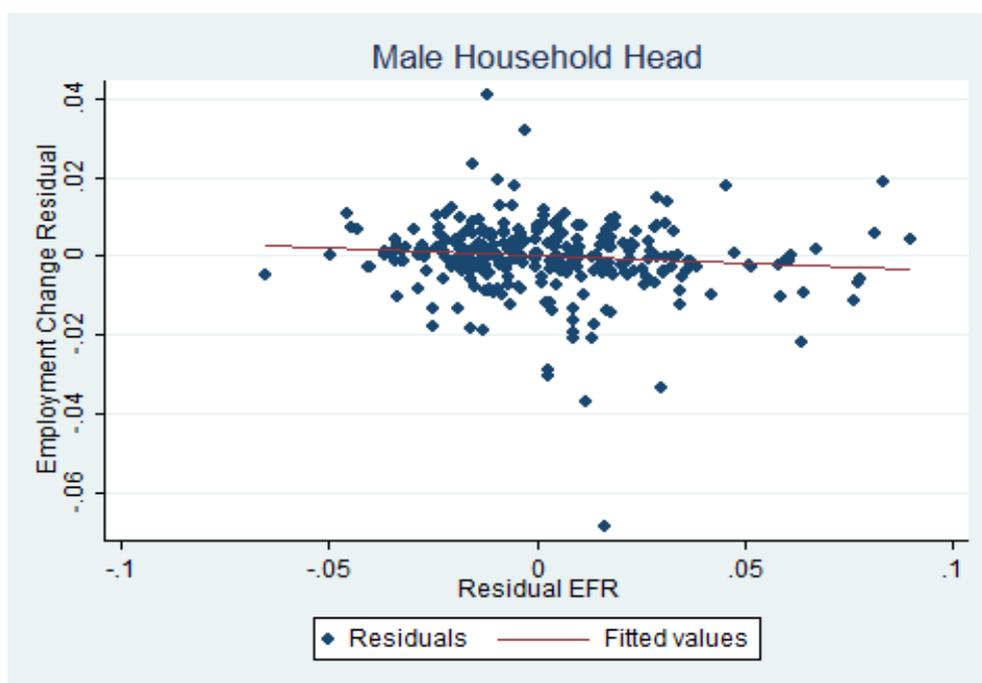


Row 6 of Table 2 shows the result for labor supply for the male head of the household. The dependent variable is a dummy indicating whether the male head in a household

participated in the labor force. The coefficient has the expected negative sign and is statistically significant. As indicated in Figure 8, the size of the effect is still rather small. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual is associated with an increase in the probability of labor force participation of a male household head by 0.117 percentage points (that is,  $0.0302 \times 0.0387$ ).

Figure 8

### Male Labor Supply Regression



### B.4 Rural Migration

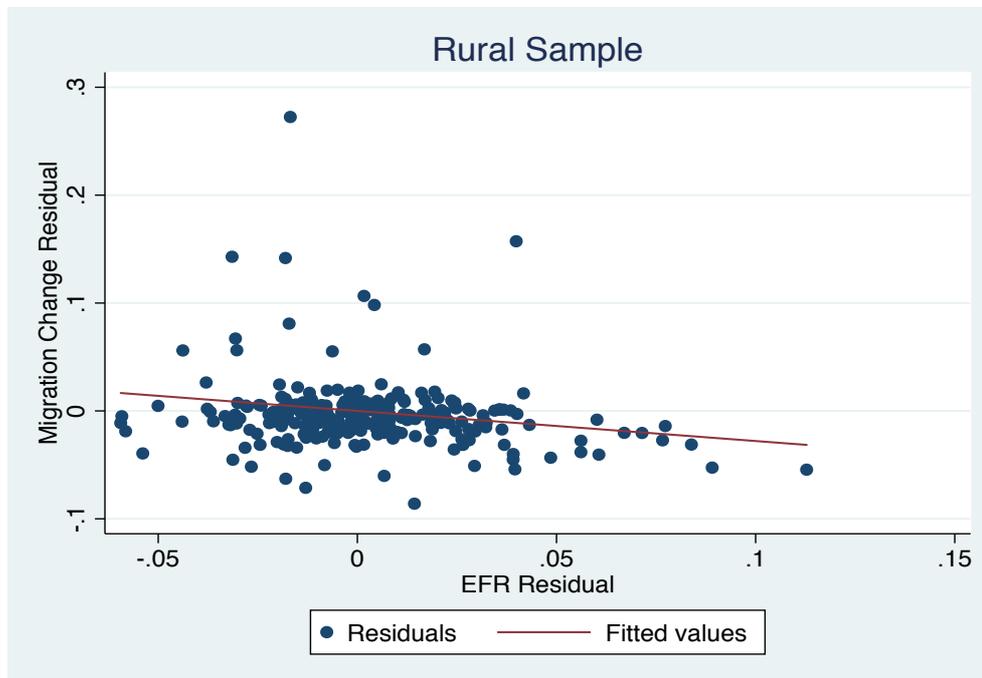
There is no study of the effect of the one-child policy on migration in China. The closest research is some work on the effect of children on migration in China. For example, Zhao (1999) found that the presence of the preschool children in the family would decrease the probability of the family members to migrate to an urban area for work (though the effect is not statistically significant) but more school-aged children would encourage family members to migrate. Yang and Guo (1999) reported that more children in the household would reduce the probability of both men and women

to migrate and the magnitude of the impact is larger for women but neither estimate was statistically significant.

Row 7 of Table 2 shows my estimation results for the rural migration. The dependent variable is defined as the ratio of the number of individuals leaving the household for more than one year to the number of the labor force in the household in rural areas. The coefficient has the expected negative sign and is statistically significant. that is, having fewer children tends to be associated with a higher level of migration. As indicated in Figure 9, the size of the effect is moderate. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual can increase the rural migration rate by 0.823 percentage points in 2000 (that is,  $0.0295 \times 0.279$ ).

Figure 9

### Rural-urban Migration Regression



Besides the basic empirical analysis mentioned above, I also perform additional analyses as robustness check. First, I used the 2000 census data and looked at the

longer-term effect in 2000. Second, since the one-child policy was modified around 1984 and then stabilized in 1990, thus I define a new EFR that is based on the 1990 census data and evaluate the outcome variables in 2000. Moreover, I separate the urban and rural samples for the 1990 and 2000 analysis. Finally, I defined the EFR according to the actual policy rule in the rural areas – some allowed only one, some allowed a second birth if the first child was a girl, and some allowed a second birth without qualifications. In all these alternative definitions and sample treatments, the results are very similar to what have been reported above, which states that the overall effects are either small or zero, especially by 2000. Having examined the overall effects, I turn now to the analysis of separate age groups.

### **C. A Further Analysis by Age Groups**

In the previous subsections, I identify the treatment effect by comparing the fertility of 20-64 years old women in 1982 and 1990 on the basis of the fact that women in 1982 were less affected by the one-child policy than women in 1990. However, such estimation strategy may under-estimate the policy effect. Evidently, not all 20-64 years old women's fertility were affected by the one-child policy during 1982 and 1990. For example, for a woman of 64 years old, she had finished her childbearing years before the one-child policy was introduced in 1979 and thus was not affected at all, whether in 1982 or 1990. The fertility levels of women of 64 years old, whether in 1982 and 1990, would not be affected by the one-child policy. While for a 20-year-old woman, she may have just been married but have not yet had any child, and thus would not be affected by the one-child policy either.

In brief, younger women are too young to be affected by the policy in the two census years, whereas elder women may be too old and may have finished their childbearing before the one-child policy was introduced. Therefore, the medium age group, say, 30 to 40 years old women, may be the most suitable cohorts to identify the policy effect during 1982 and 1990. Specifically, for women of 40 years old in 1982 sample, they

were less affected by the one-child policy because when the policy was introduced in 1979, they were already 37 years old and thus probably had two or more children already; whereas for 40-year-old women in the 1990 sample, they were more affected by the one-child policy because in 1979 they were only 29 years old and had more childbearing years under the one-child policy. Thus an alternative and perhaps more appropriate analysis is to divide women into different age groups to examine the effect of the one-child policy. Thus, I divide the women into six groups by women's age in the census year as follows: 20-24, 25-29, 30-34, 35-39, 40-44, and 45-49, and the empirical results for all the family outcome variables by separate groups are shown in Table 3.

*Table 3*

**Regression Results for Different Age Groups, 1982-1990**

| Age Group | Fertility           | Children's Education <sup>4</sup> | Youth Dependency Ratio | Labor Supply (Female) | Labor Supply (Male)  | Being Divorced          | Rural Migration      |
|-----------|---------------------|-----------------------------------|------------------------|-----------------------|----------------------|-------------------------|----------------------|
| 20-24     | -0.183<br>(0.164)   | -----                             | -0.105<br>(0.087)      | -0.0926<br>(0.088)    | 0.0132<br>(0.012)    | 0.00298<br>(0.0023)     | -0.344***<br>(0.092) |
| 25-29     | -0.104<br>(0.273)   | -6.535***<br>(1.012)              | -0.045<br>(0.132)      | 0.0227<br>(0.089)     | -0.039***<br>(0.006) | -0.00470***<br>(0.0014) | -0.400***<br>(0.086) |
| 30-34     | 2.559***<br>(0.331) | 0.685<br>(0.727)                  | 1.254***<br>(0.162)    | 0.022<br>(0.098)      | -0.0172<br>(0.009)   | -0.00131<br>(0.0015)    | -0.223***<br>(0.048) |
| 35-39     | 4.207***<br>(0.292) | 0.783<br>(0.742)                  | 1.824***<br>(0.148)    | 0.0631<br>(0.084)     | -0.029***<br>(0.009) | -0.00164<br>(0.0016)    | -0.229***<br>(0.046) |
| 40-44     | 3.392***<br>(0.317) | 0.0678<br>(0.780)                 | 0.989***<br>(0.145)    | 0.12<br>(0.087)       | -0.0296*<br>(0.012)  | -0.00356<br>(0.0027)    | -0.248**<br>(0.075)  |
| 45-49     | -0.515<br>(0.500)   | -1.418<br>(0.814)                 | 0.372*<br>(0.184)      | -0.176<br>(0.111)     | -0.0559<br>(0.029)   | -0.0166<br>(0.0093)     | -0.243***<br>(0.064) |

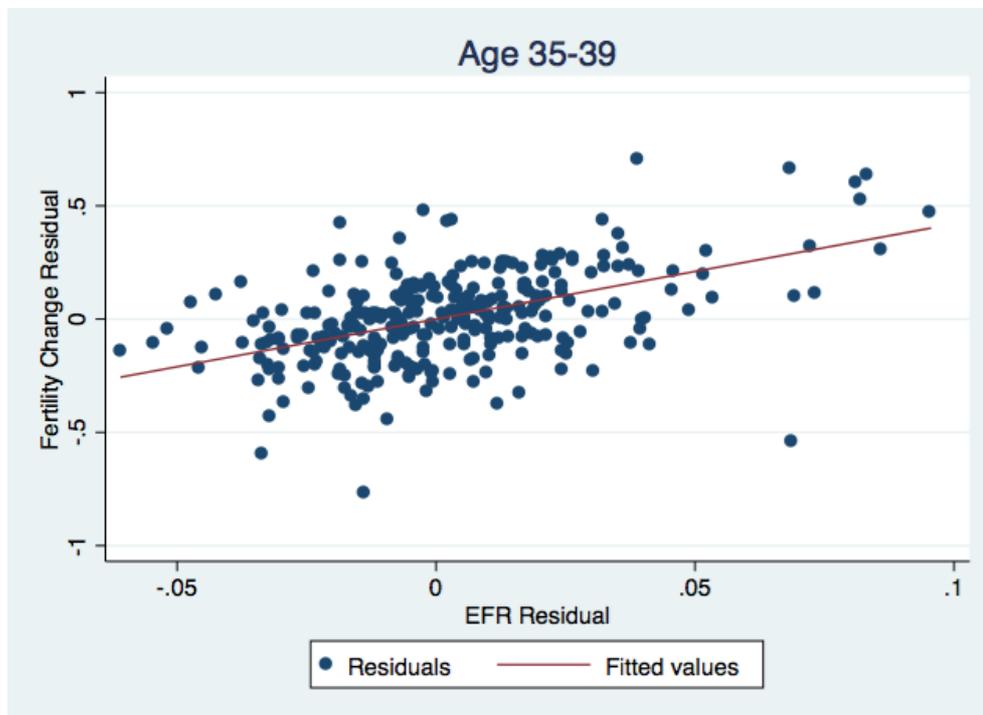
I expect that the policy effect will be small for the younger and elder women and most

<sup>4</sup> Here I did not report the estimate of the policy effect on the education level of children whose mothers are only 20-24 years old. Evidently, most of these children were actually very young and had not yet reached the primary school age. Thus, it is impossible to measure their education levels in the census year. Furthermore, even though there were very few observations who were elder than 6 years old and had an education record, the number of such observations in each prefecture (often less than 20) may be too small to be representative of the population.

significant for some intermediate groups. The empirical results in column 2 of Table 3 are completely consistent with my expectations. The policy effects on the younger cohorts (20-24, 25-29) and elder cohorts are generally small and statistically insignificant, whereas the effects on the intermediate cohorts are quite larger and significant at the 1% percent.

Figure 10

**Fertility Regression for Age Group 35-39**

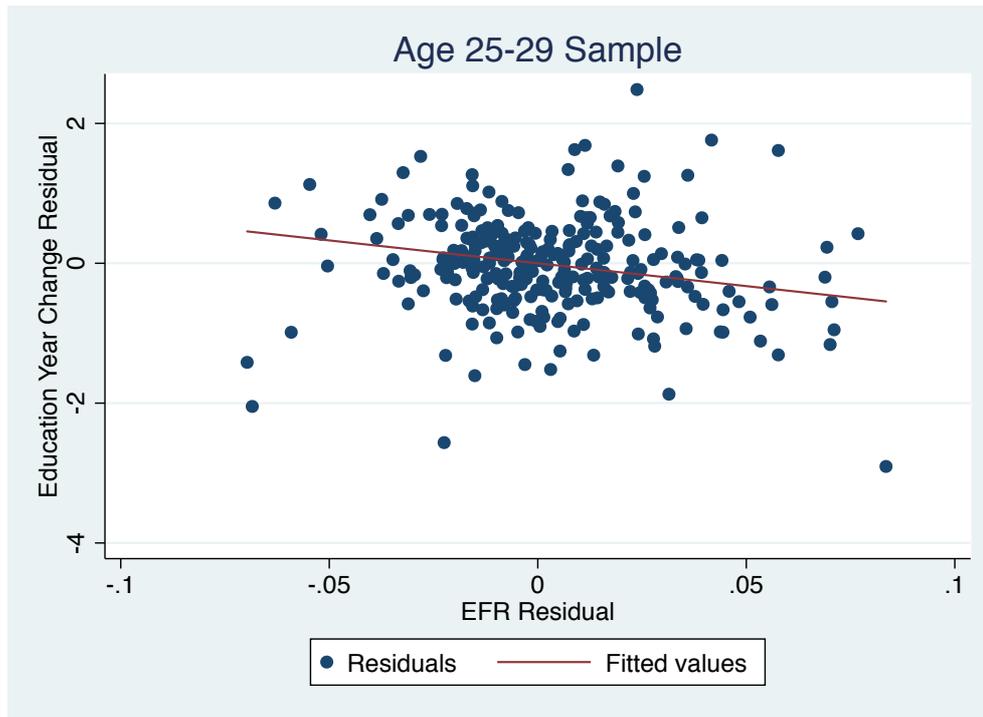


I evaluate the magnitude of the effect of the one-child policy on fertility according to the estimation results. As indicated in Figure 10, the size of the effect is rather small. For the most affected 35-39 age group, by construction, the mean of the EFR residual is zero with a standard deviation of 2.53%. The interquartile range of the EFR residual equals 3.03%. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual can lead to 0.1274 less births per household in 1990 (i.e.  $0.0303 \times 4.207$ ), representing a decrease of about 3.70% in fertility (i.e.  $0.1274 / 3.444$ , where 3.444 is the mean number of children for

35-39 age group in 1982). Therefore, column 2 shows evidence that the one-child policy has a small effect on China's fertility by 1990.<sup>5</sup>

Figure 11

**Children's Education Regression for Age Group 25-29**



Column 3 of Table 3 shows the estimates of the effects of the one-child policy on children's education by mothers' age group. The estimates are only significant for the children of youngest mothers (25-29 years old), which seems quite strange at first glance. The policy effects on this group is negative and statistically significant, which implies that a stricter policy induces a higher education level for their children. As indicated in Figure 11, the policy effect seems modest even for the 25-29 age group,

<sup>5</sup> Alternatively, if we look at the maximum and minimum values of the EFR residuals, which are 0.1128 and -0.0594 respectively, and the difference between the two values is 0.1722. A decrease of 0.1722 in the EFR residual would lead to a decrease of 0.724 in births per family or cause a decrease of 21% in fertility, which implies that in the prefecture where the one-child policy is most strictly implemented the fertility declines much more sharply than that in the prefecture where the one-child policy is most relaxed. Therefore, the effect of the one-child policy on fertility can be quite substantial in these prefectures where the policy is strictly implemented.

on which there was the largest effect. Specifically, for this age group, the standard deviation of the EFR residual is 2.50% and the interquartile range equals 2.94%. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual can lead to 0.192 more years of schooling for the child in 1990 (i.e.  $0.0294 \times 6.535$ ), which is an increase of about 6.59% in the schooling years of the child (i.e.  $0.192/2.912$ , where 2.912 is the mean schooling years of the children whose mothers are in this group in 1982).

However, it remains unclear as to why the policy effects become much smaller and insignificant for the older cohorts. Re-examining the youngest mothers more thoroughly, I find that most of them are from backward rural areas. In China rural women are generally married much earlier than urban women and the educational attainment of very young mothers are also relatively low. Previous studies found that the quantity-quality tradeoff was considerable in rural areas, but became much smaller or even vanished in urban areas, where resources constraint was less severe (Li et al., 2008; Rosenzweig and Zhang, 2009). Therefore, the significant quantity-quality tradeoff for the children of youngest mothers may imply that such tradeoff effect is largest for those less well-off rural families who usually face tight resources constraint.<sup>6</sup>

Column 4 of Table 3 shows the estimates of the policy effect on the youth dependency ratio, which are highly consistent with the effect on fertility in Column 2. As indicated in Figure 12, while the estimates for women of 30-44 years old are

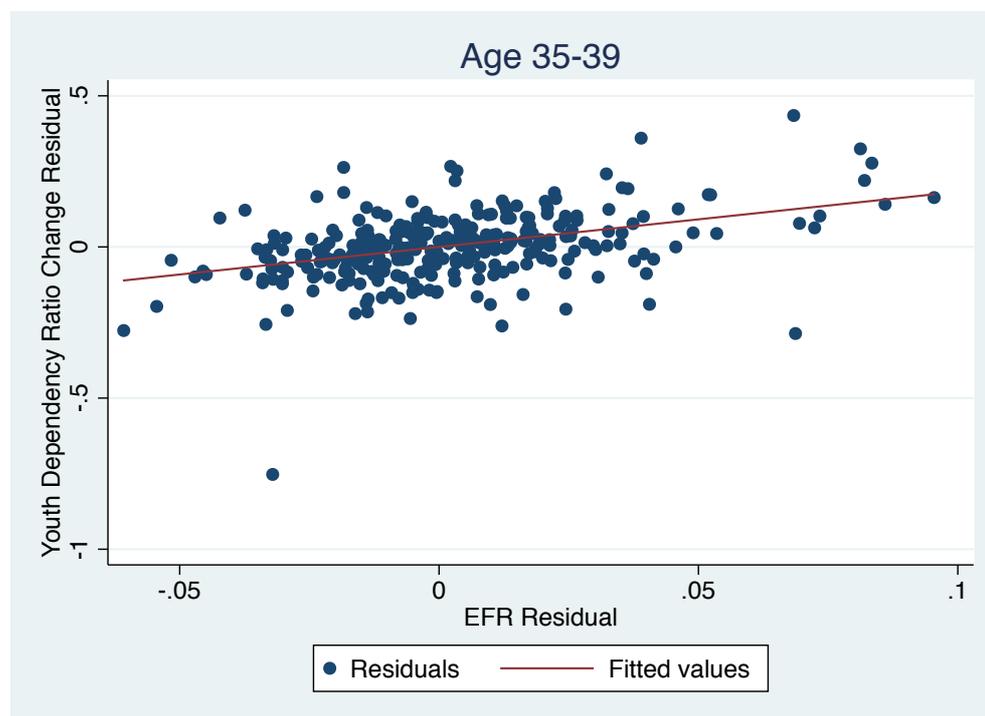
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<sup>6</sup> Note that Table 3 shows that in the census year the policy has no effect on the fertility of youngest mothers (20-29 years old). However, that does not mean that the quantity-quality tradeoff does not exist for their children. For example, in prefectures where the one-child policy is more strictly implemented, even though a mother only had one child in the census year, she may expect that this child would probably be the only child she could have and thus invested more resources in the only child. Therefore, for these youngest mothers, even the one-child policy has not affected their fertility up to the census year, the quantity-quality tradeoff still exists via their expectation of the number of children they could have that are obviously affected by the implementation intensity of the one-child policy.

positive and statistically significant, the magnitude of the effect is not very large. For 35-39 years old women on whom the effect is largest, a tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual would decrease the youth dependency ratio in the household by 5.53 percentage points in 1990 (i.e.  $0.0303 \times 1.824$ ).

*Figure 12*

**Youth Dependency Ratio Regression for Age Group 35-39**



Column 5 of Table 3 reports the estimates of the policy effect on female head’s labor supply. The estimates are small and not statistically significant for all age groups, which indicates that the one-child policy has no significant effect on female labor supply, as in the overall effect discussed earlier.

Column 6 of Table 3 shows the estimates of the policy effect on the labor supply of male heads of the households. We can find significant negative effect for some age groups but the size of the effect is rather small. As indicated in Figure 13, the effect is quite small, even for the age group 25-29, in which the effect is the largest. A tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual would increase the probability of male heads’ labor

participation by 0.119 percentage points in year 1990 (i.e.  $0.0306 \times 0.039$ ).

Figure 13

**Male Head Labor Supply Regression for Age Group 25-29**

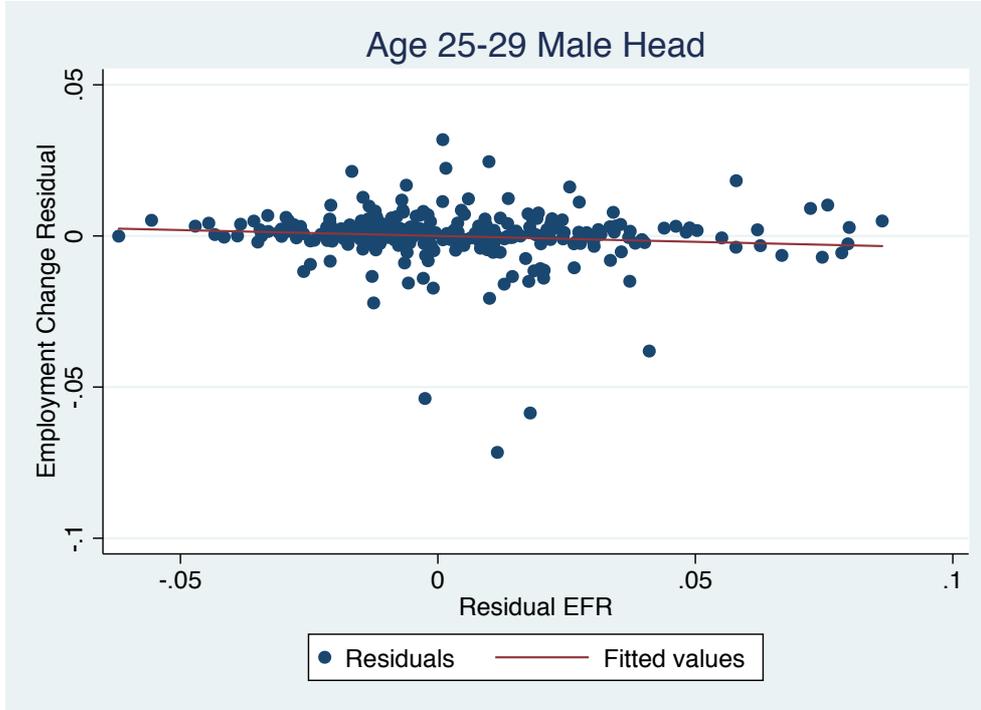
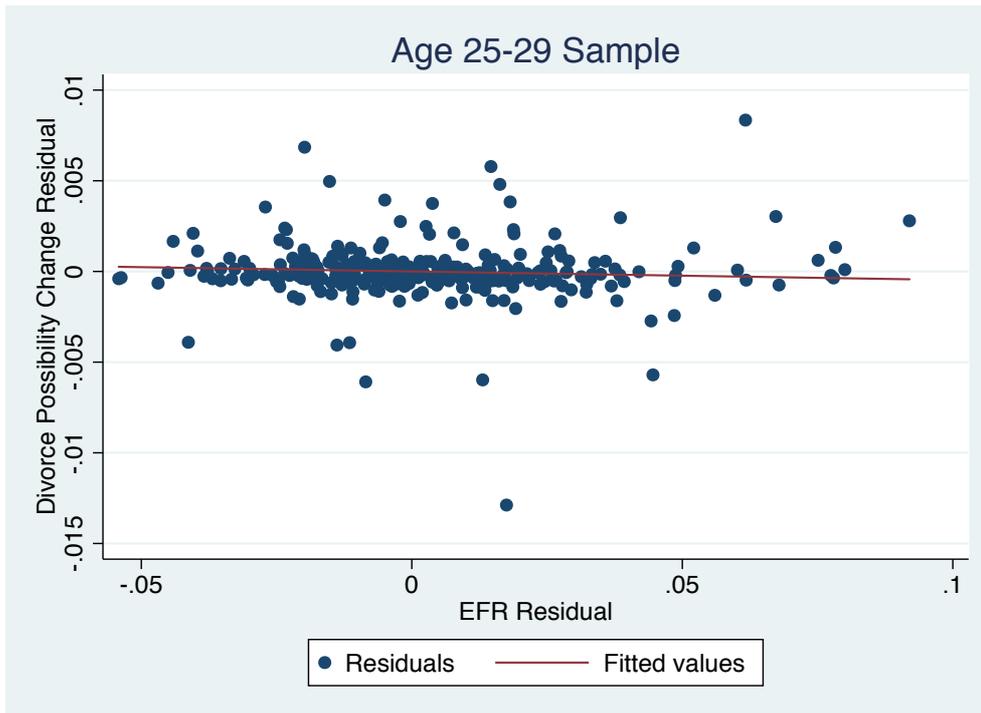


Figure 14

**Marital Status (Being Divorced) Regression for Age Group 25-29**

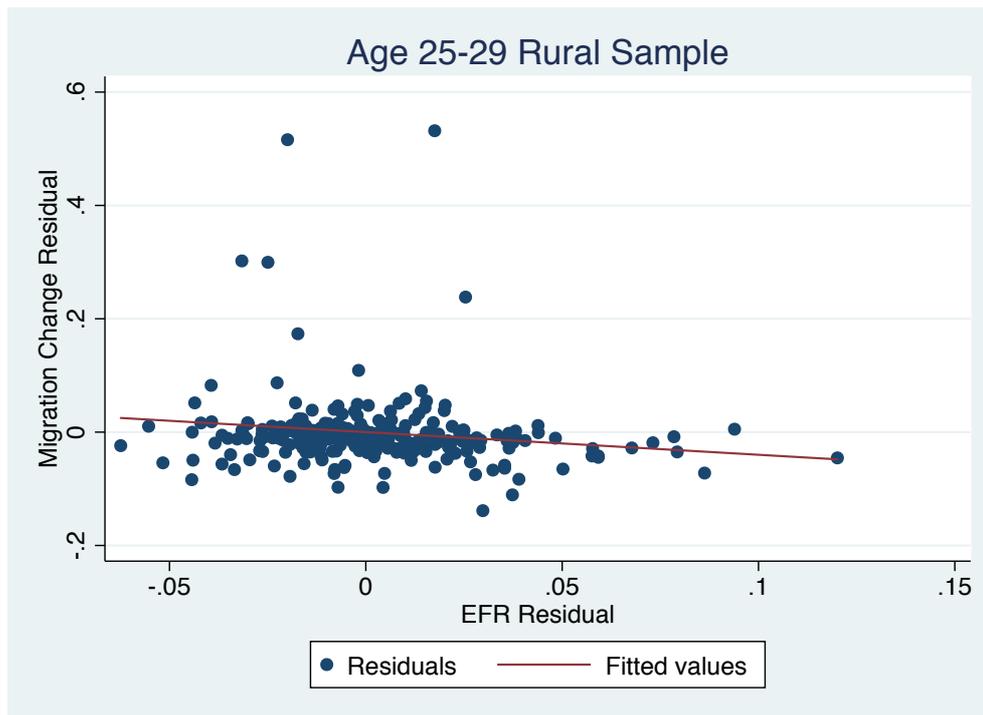


Column 7 of Table 3 shows the estimates of the policy effect on divorce. The only significant negative effect is for the age group of 25-29 years old. As indicated in Figure 14, the size of the effect is quite small.

Column 8 of Table 3 shows the estimates of the policy effect on rural migration. I found a negative and significant effect for all the age groups. As indicated in Figure 15, the size of the effect is small. For the age group 25-29 on whom the effect is the largest, a tightening of the one-child policy in terms of one interquartile range decrease of the excess fertility rate residual would increase the rural migration rate by 1.16 percentage points in 2000 (i.e.  $0.0291 \times 0.400$ ).

Figure 15

**Rural-urban Migration Regression for Age Group 25-29**



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