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ABSTRACT

More Men, More Crime: Evidence from China's One-Child Policy*

Crime rates almost doubled in China between 1992 and 2004. Over the same period, sex ratios (males to females) in the crime-prone ages of 16-25 years rose sharply, from 1.053 to 1.093. Although scarcity of females is commonly believed to be a source of male antisocial behavior, a causal link has been difficult to establish. Sex-ratio variation is typically either small or related to social conditions liable to also affect crime rates. This paper exploits two unique features of the Chinese experience: the change in the sex ratio was both large and mainly in response to the implementation of the one-child policy. Using annual province-level data covering the years 1988-2004, we find that a 0.01 increase in the sex ratio raised the violent and property crime rates by some 5-6%, suggesting that the increasing maleness of the young adult population may account for as much as a third of the overall rise in crime.

JEL Classification: J12, J13, K42

Keywords: male-biased sex ratios, crime, one-child policy, China

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1 Introduction

The past three decades have seen steadily rising sex ratios (males to females) at birth in China. For most of the 1970s, sex ratios stayed within the biologically normal range of around 1.05, albeit trending upward. Following the one-child policy (which was initiated in 1979), sex ratios have seen a steep ascent (see Figure 1a). By 2000, 120 boys were born for every 100 girls.[1]

The implied surplus of young men has been worrying observers for a number of reasons, the future dearth of brides being a prominent concern (e.g., Hesketh & Xing (2006) and references therein).[2] In particular, unmarried men may not be "happy campers" and may thus engage in various types of anti-social behavior, including criminality.[3] Also, as elsewhere (Steffensmeier & Allan (1996); Hurwitz & Smithey (1998)), men, and especially young men, are the most crime prone. In 2000, males made up 90% of arrestees (*Law Yearbook of China* (2001)).

Between 1988 and 2004, criminal offences rose at an annual rate of 13.6% (or 12.5%, population adjusted) (Hu (2006)), and arrest rates were up by

[1]The sex ratio at birth of 1.20 was calculated from the 2000 census data pertaining to any new births (and their sex) in the past year (November 1, 1999 to October 31, 2000). In Figure 1a, the sex ratio for 2000 was 1.19. The slight discrepancy occurs because the sex ratio in Figure 1a pertains to the sex ratio of those enumerated in November 2000, by birth year.

[2]Also see "Missing sisters", *The Economist*, April 17, 2003 and "China: Too Many Men", *CBS 60 Minutes*, April 16, 2006,
http://www.cbsnews.com/stories/2006/04/13/60minutes/main1496589.shtml.

[3]The impact on females born has received less attention (exceptions include Goodkind (1996); Edlund (1999); Becker (2007)).

82.4% (Figure 2).[4] The overwhelming majority (70%) of perpetrators of violent and property crimes are between 16 and 25 years old. Since the mid-1990s, sex ratios for this age group have risen steeply.

The empirical literature most closely related to our work has focused on India, a country with comparably male-biased sex ratios. This literature has found a positive correlation between high sex ratios and crime (Dreze & Khera (2000)).[5] Another literature has argued that rising sex ratios, when there is a surplus of women, reduce crime (Messner & Sampson (1991); Barber (2000); Posner (1992)).[6] Although these studies suggest unbalanced sex ratios to be correlated with male criminality, and the possibility of a causal link between male-biased sex ratios and anti-social behavior (including crime) has been noted (e.g., Hudson & Den Boer (2004) and references therein), the establishment of empirical causality has remained elusive.

The pitfalls are many. For instance, crime may be viewed as a high-risk occupation that disproportionately employs (and victimizes) men, thus affecting the sex ratio. Moreover, violent and property crimes are low-skill crimes and may be more prevalent among economically disadvantaged

[4] As we do not observe province-level offences, our empirical analysis focuses on arrests.

[5] Based on cross-sectional data, Dreze & Khera (2000) found the elasticity of the murder rate with respect to the sex ratio to be greater than eight. They conjectured that cultural variations across regions may account for the pattern that was found. Specifically, a patriarchal culture may result in both higher sex ratios and more violent crime. Another possibility, advanced by Oldenburg (1992), is that high crime rates drive the demand for sons for physical protection. For a critique, see Mitra (1993).

[6] In the U.S., homicides and incarceration (of males) have resulted in female-biased sex ratios among the underclass. In other developed countries, the upward mobility of females means that male-biased or balanced sex ratios prevail among the underclass. For a link between female-biased sex ratios and unstable marriages, see Willis (1999).

groups, which face conditions (other than crime) that affect both marriage rates (Wilson (1987)) and sex ratios (through, e.g., gender differential migration, out-marriage, and mortality). Finally, although longitudinal data on individuals may seemingly offer a solution, the decision to enter into marriage is likely to be influenced by idiosyncratic labor market conditions unobservable to the econometrician, such as better employment or promotion prospects.

The Chinese experience offers a solution. It is by now well accepted that the Chinese one-child policy has been an important factor in the rise in the sex ratio – chiefly through the combination of a strong preference for sons and the arrival of prenatal sex determination technology in the early 1980s (Zeng *et al.* (1993); Miller (2001); Chu (2001); Li (2002); Yang & Chen (2004); Das Gupta (2005)).

There was substantial variation across provinces in the timing of the implementation of the various programs that make up China's one-child policy. Provinces with higher fertility, higher population density, and leaders who were eager to comply with central policies implemented the programs earlier. For example, although the first Family Planning Science and Technology Research Institute was established in Jiangsu in 1976, the national roll-out was not complete until 1992.

We exploit this provincial variation to identify the causal impact of male-biased sex ratios on crime. In particular, we identify the impact of sex ratios on the crime rate among the crime-prone cohort (16-25 years old) by using the lagged (17- to 26-year) family planning programs as instrumental variables (IVs). We believe the timing of these programs, by province, to be valid instruments. First, these programs are both *a priori* plausible instruments for the sex ratio at birth, and, empirically, highly correlated with the sex ratio. Second, the fact that these programs were implemented 17-26 years ago reduces the likelihood that they are correlated with the contemporaneous error term.[7]

The fact that the higher sex ratios were in response to family planning programs is attractive for our purposes for several reasons. First, it seems reasonable that children born as a result of these programs were positively selected and could therefore be expected to have better adult outcomes (Kane & Staiger (1996); Donohue & Levitt (2001); Pop-Eleches (2006)). Second, sex-selective abortion (Edlund (1999)) and the incentive structure of the one-child policy (Short & Zhai (1998)) are likely to have strengthened the positive selection of sons (that is naturally occurring; see Trivers & Willard (1973) and Almond & Edlund (2007)). Third, falling fertility meant that younger cohorts were also smaller, which should tend to reduce crime rates directly, assuming that younger cohorts are more crime prone, and indirectly, via more favorable labor market conditions and the quantity-quality tradeoff (Li *et al.* (Forthcoming)).

Employing province-level panel data (from 27 provinces) from the period 1988-2004 (mainly from the Statistical Year Books and the 1990 census), we estimate the effect of the sex ratio on criminality using both OLS and IV methods. For the IV estimation, we exploit provincial variations in the roll-out of the one-child policy. Both methods suggest that higher sex ratios had an economically and statistically significant impact on criminality. The IV estimates indicate that a 0.01 increase in the sex ratio raised the crime

[7]We control for pre-existing time-invariant differences between provinces by including province dummies in our estimations.

rate by about 5-6%. Since the sex ratio increased by some 0.04 (from 1.053 to 1.095), this suggests that higher sex ratios raised crime by 20-24%, thus accounting for about a quarter to a third of the overall rise (of 82.2%).

To further interpret our findings, we present a simple decomposition of the crime rate. To fix ideas, assume that males are more crime prone than females are, and that single men are more crime prone than married men are. Then, a higher sex ratio could translate into more crime through three avenues. First, a higher sex ratio implies more males and thus a higher crime rate. Second, assuming male-biased sex ratios, higher sex ratios imply more unmarried men. Finally, a third possibility is that the sex ratio affects the propensity of males to commit crimes. In particular, a lower likelihood of marriage may make single men more crime prone. We derive an upper bound for the first two effects and attribute additional increases to the third effect.

According to this decomposition, the magnitude of the empirical estimates suggests that the higher crime rate was not simply the result of more men, but also of more unmarried men. Moreover, the estimates are in the upper range of what could be expected from a pure composition effect, thus suggesting the possibility that the higher sex ratio also increased the propensity to commit crime, conditional on marital status. However, this conclusion is sensitive to various assumptions and therefore remains speculative.

The remainder of the paper proceeds as follows. Section 2 presents a simple decomposition of the crime rate. Sections 3-5 present the data, the empirical model and results. Section 6 concludes.

2 Male-Biased Sex Ratios and Crime

This section decomposes the crime rate into changes due to more men, changes due to more unmarried men, and the rest. We interpret the residual as being due to changes in the criminal propensity of men, conditional on marital status.

Consider a population of measure 1 with m men and $f = 1 - m$ women. Men can be unmarried, m^0, or married, m^1, and $m = m^0 + m^1$. Similarly for women, $f = f^0 + f^1$. We denote each demographic group's crime rate as c_m^0, c_f^0, c_m^1, and c_f^1, respectively. The crime rate, c, can then be expressed as

$$c = c_m^0 m^0 + c_m^1 m^1 + c_f^0 f^0 + c_f^1 f^1.$$

We restrict our attention to the case of male-biased sex ratios, and assume that men marry with probability $f/m (< 1)$ and women marry with certainty. Furthermore, we assume that

$$c_m^0 \geq c_m^1 > 0,$$

$$c_f^0 = c_f^1 = c_f, c_f \in [0, c_m^1).$$

To assume that married and unmarried females are equally crime prone is innocuous, as all women will be married by assumption.

2.1 Demographic Composition Effect

A higher sex ratio increases the fraction of males and the fraction of males who are unmarried, both of which may raise the crime rate. To focus on the first mechanism, we assume for now that

$$c_m^0 = c_m^1 (= c_m). \tag{1}$$

In this case, the crime rate, c, is simply

$$c = c_m m + c_f(1 - m).$$

Let r denote the sex ratio:

$$r = \frac{m}{f} = \frac{m}{1-m}. \tag{2}$$

We can write the elasticity of the crime rate with respect to the sex ratio, conditional on Eq. 1, as

$$\begin{aligned}
\epsilon(c, r)|_{c_m^0 = c_m^1 = c_m} &= (\frac{dc}{dm} \cdot \frac{m}{c}) \cdot (\frac{dm}{dr} \cdot \frac{r}{m}) \\
&= \frac{c_m m - c_f m}{c_m m + c_f(1-m)} \cdot f \\
&\leq f. \qquad (3)
\end{aligned}$$

That is, if the estimated elasticity of crime with respect to the sex ratio is greater than the fraction of females, then higher criminality cannot be due simply to more males.

Next, we consider the possibility that marital status has an impact on the propensity to commit crime. Specifically, an increase in the sex ratio, r, not only increases m, but also raises the fraction of unmarried males, which, assuming that $c_m^0 > c_m^1$, further raises criminality.

In this case, the ratio of all males to unmarried males gives an upper bound on the elasticity of the crime rate with respect to the sex ratio, because

$$\begin{aligned}\epsilon(c,r) &< (\frac{dc}{dm^0}\frac{m^0}{c})\cdot(\frac{dm^0}{dr}\frac{r}{m^0}) \\ &< (\frac{dc}{dm^0}\frac{m^0}{c})\cdot\frac{m}{m^o} \\ &\leq \frac{m}{m^o}. \end{aligned} \tag{4}$$

The first inequality in expression 4 follows from the fact that higher sex ratios result in an increase in unmarried males and a decrease in married males (and females). For the second inequality, we obviously need

$$\begin{aligned}\frac{dm^0}{dr}\frac{r}{m^0} &< \frac{m}{m^0}, \\ &\Leftrightarrow \\ \frac{dm^0}{dr} &< f.\end{aligned}$$

To see that this is the case, note that $f = (1 - m^0)/2$ and that

$$\frac{dm^0}{dr} = \frac{1}{dr/dm^0} = \frac{(1-m^0)^2}{2} < \frac{1-m^0}{2}, m^0 \in (0,1).$$

For the last inequality in expression 4, we note that $\frac{dc}{dm^0}\frac{m^0}{c} < 1$, because a percentage change in the fraction of unmarried men can at most bring about an equal change in the crime rate.

While expression 4 sets an upper bound for the elasticity of the crime rate with respect to the sex ratio, the empirical calculation of m/m^0 is sensitive to what ages are considered. Because the share of unmarried men decreases with age, m/m^0 rises with the age cut-off. Although older men are potential participants in the marriage market, their propensity to commit crime decreases with age. Thus, it may be that c_m^0 approaches c_m^1 as men age, which would suggest an age cap for the calculation of m/m^0.

Figure 3 shows the values for m/m^0 for ages x-40, where $x \in [16, 40]$ (using the 2000 Chinese population census). As expected, m/m^0 rises (almost) monotonically with the age cut-off, ranging from 2.64 to 16, where a cut-off age of 20 would return an upper bound on the elasticity of 3.7.[8] To anticipate events, our estimate of the elasticity is around 6%, which, depending on the age cutoff considered, could be consistent with the possibility that a lower probability of marriage raises the crime propensity of men, conditional on marital status (that is, $dc_m^i/dr \geq 0, i = 0, 1$, where the inequality is strict for at least one i).

2.2 Marriage Market Effect

The impact of the sex ratio on the crime rate could well exceed the upper bounds derived above. A high sex ratio could result in fiercer competition among males in the marriage market, and thus change men's labor market behavior (Chiappori *et al.* (2002)).[9] Considering crime as a high-risk occupation, a similar logic would suggest the possibility of a link between more competitive marriage markets and male criminality. To improve the ability to find/retain a partner in a more competitive marriage market, unskilled men may boost their incomes by turning to illegal activities. Thus, the marriage market could raise the effect of the sex ratio on crime rates beyond that attributable to a pure demographic composition effect.[10]

[8] The legal marriage age for men is 22, although it is not enforced. For instance, 5% of 21-year-old males were married in the 2000 census.

[9] That male-to-male competition for partners may lead to male aggression and risk taking was already proposed by Trivers (1972).

[10] Higher sex ratios imply a higher fraction of males being unmarried. If criminality among this group is influenced by peers, a peer effect may be at work. The existence of

3 Data Description

We analyze annual, province-level data from China's 27 provinces, covering the period 1988-2004.[11] Further variable descriptions and summary statistics are in Table 1.

Crime The crime rate is defined as the number of arrests per 10,000 population. The crime rate almost doubled in the study period, from 3.71 in 1988 to 6.77 in 2004 (Figure 2), and there were considerable variations across provinces. These ranged from 0.81 (in Tibet in 1988) to 13.1 (in Zhejiang in 2004).

We focus on violent and property crimes, because these are low-skill crimes (relative to, for example, corruption), and high sex ratios are likely to compromise the marriage prospects of low-skilled men disproportionately.

Data on crime are from the *China Law Yearbook* (Supreme People's court, 1989-2005) and the *Procuratorial Yearbook of China* (Supreme People's Procuratorate, 1989-2005). Due to the centralized judicial system, crime is consistently defined across provinces (Wang & Mo (1999); Chow (2003)).

Unfortunately, we cannot distinguish between violent crimes and property crimes. However, both types of crime may be similarly motivated (e.g., robbery). Moreover, property crimes made up between 77.3%

the peer effect would amplify the demographic composition and marriage market effects.

[11] The four municipalities directly under the central government – Beijing, Shanghai, Tianjin, and Chongqing – are excluded, as they are governed by a different judicial system and are not comparable to other provinces.

and 90.7% of all criminal cases between 1981 and 2004.[12] Among property crimes, larceny is by far the most common (86.7% in the same period) (Hu (2006)).

Sex ratio We focus on the 16-25 age group for three reasons. First, 16 years is the age of full criminal responsibility, and 25 years is the upper age for "juvenile crime." Second, these are the most crime-prone ages, accounting for more than 70% of the total number of criminal offenders since the mid-1980s. For example, the share of homicides, rapes, robberies, and larcenies committed by this age group in 1993 was 46.73%, 55.31%, 78.77%, and 66.16%, respectively (Hu (2006)). Third, these are also the ages that mark early adulthood and therefore partner search.

We use the 1990 census to predict the sex ratio for the 16- to 25-year-old cohort by province and year (henceforth, the 16-25 sex ratio). For instance, for Zhejiang and 1989, we use the sex ratio in Zhejiang for the 1964-1973 birth cohorts. However, these projections may deviate from the actual sex ratios for a number of reasons, e.g., gender-differential migration, mortality, and misreporting.

The 16-25 sex ratio has risen dramatically since the initiation of the one-child policy, from 1.053 in 1988 to 1.095 in 2004, a rise of four percentage points in 16 years (Figure 1b). The increase is even more dramatic if we consider the sex ratio at birth (an important early indicator of future 16-25 sex ratios): this sex ratio rose from about

[12] The national crime rate became available to the public in 1981, whereas the province-level crime rate was first made available in 1988.

1.04 for the birth cohort 1970 to 1.19 for the birth cohort 2000, a rise of 0.15 in three decades (Figure 1a). This implies that the 16-25 sex ratio will continue to rise in the next two decades. Another noteworthy feature is the substantial provincial variation. Guangdong and Hainan had the highest 16-25 sex ratio of 1.13 in 2003 and 2004, respectively, whereas Tibet (which has more liberal population control policies due to its higher fraction of minorities) had the lowest, 0.98, in 1991.[13]

One-Child Policy Instruments The one-child policy was formally launched in 1979. At the outset, each household was allowed only one child. Various exceptions were later introduced, and, most notably, many rural areas allowed a second child if the first were a girl (or handicapped). Households were given birth quotas, and "above-quota births" were penalized. Although there were variations at the sub-provincial level, e.g., the county, the main variation was at the provincial level.

Local governments were given incentive contracts that specified the birth target according to the local population density and fertility rate. Financial rewards and penalties were linked to these birth targets (Peng (1996); Short & Zhai (1998); Li & Zhang (2007)). In addition, local government officials were held personally responsible for meeting the targets, and above-quota births could result in demotions, a serious penalty in the Chinese context, where government positions carry significant perks.

Our instrument for the sex ratio considers the differential roll-out of

[13] Minority-dominated areas such as Tibet may also have different crime rates. We include province fixed effects to control for such differences in our estimations.

family planning programs across provinces. The variation in the timing (see Figure A1) reflects the strictness or urgency of the local government in implementing the one-child policy.[14]

We focus on the following three programs established between 1976 and 1992: the establishment of *Family Planning Science and Technology Research Institutes*, *Family Planning Education Centers*, and *Family Planning Associations*.[15]

1. *Family Planning Science and Technology Research Institutes* are medical and research facilities that are available at all government levels. These institutes provide birth control services and collect data on the cost effectiveness of various birth control methods (chiefly intra uterine devices [IUDs], sterilizations, and abortions). They also distribute free contraceptives to households.

2. *Family Planning Education Centers* promote family planning through movies, pictures, books, and posters. These centers educate people on the use and usefulness of contraceptives and extol the virtues of a small family.

3. *Family Planning Associations* hold academic exchange activities

[14] In this paper, the one-child policy is defined in a broad sense and includes actions and measures implemented before 1979. For an excellent overview, see Scharping (2003).

[15] *Family Planning Commissions* were established at all government levels in China. They are in charge of implementing the one-child policy and oversee the process of implementation. All the three programs are subsidiaries of the *Family Planning Commissions*. We cannot use the roll-out of *Family Planning Commissions* as an IV because there was little variation in the timing of these establishments, which was more directly controlled by the central government.

on local population issues, conduct socioeconomic research on family planning, and make recommendations concerning population policy and planning. This program has more variation in the roll-out across provinces than the other two programs (see Figure A1).

Additional controls Province level per capita income, employment rate, income inequality, urbanization rate, population density, age structure, secondary school enrolment rate, welfare expenditures, police expenditures (as a share of provincial government expenditures), and share of (out-of-province) immigrants are included as additional controls. These data are from the *Comprehensive Statistical Data and Materials on 55 Years of New China* (National Bureau of Statistics (2005)), and the *China Statistical Yearbooks*, 1989-2005.

Income inequality is measured as urban over rural household income. The employment rate is defined as the fraction of the working age population who are employed. Welfare expenditure is measured as the ratio of pension and social welfare expenditure to total government expenditure.[16] Detailed definitions and summary statistics are reported in Table 1.

[16]It would be ideal to have additional control variables, such as the marriage rate and deterrence measures, and urban/rural breakdown in crimes. However, such data are not available at the province level. Nonetheless, to the extent that the marriage rate and deterrence measures are not highly correlated with the sex ratios, the estimates for the sex ratio impact would be little affected.

4 Empirical Model

We specify the log crime rate equation in the following linear form,

$$\ln c_{it} = \alpha r_{it} + X_{it}\beta + \delta_i + \tau_t + \varepsilon_{it}, \tag{5}$$

where c_{it} is the crime rate in province i, year t, r_{it} is the corresponding sex ratio among 16-25 year olds (males to females), and X_{it} is a vector of province-year level controls. δ_i and τ_t are vectors of the province and year dummies, which control for pre-existing differences between provinces and year-specific effects that are common to all provinces. The residual ε_{it} is assumed to be $i.i.d$ with zero mean. We first estimate Eq. 5 by OLS and then by IV. We expect a higher sex ratio to raise the crime rate, i.e., $\alpha > 0$.

Income inequality proxies for the difference between the potential gains from crime and the associated opportunity cost (Ehrlich (1973); Kelly (2000); Bourguignon (1999)). Hence, a rise in income inequality would unambiguously increase crime rates. Similarly, per capita income may also have a positive effect on crime rates, because it may magnify the difference in returns between illegal and legal activities. However, per capita income could also proxy for a pure income effect, in which case we would expect a negative effect on crime.

We expect urbanization (Glaeser & Sacerdote (1999)) and population density to be positively related with crime rates, and employment rate, secondary school enrollment, and welfare expenditure to be negatively correlated with crime rates (Zhang (1997)). Moreover, we expect police expenditures to have a negative effect on crime rates (Becker (1968)). However, the estimates may be biased by simultaneity between police expenditures and crime rates (Levitt (1997, 1998)). For want of suitable instruments for police expenditures, we estimate Eq. 5 with and without police expenditures, thus allowing us to gauge the sensitivity of our results to the inclusion of this variable.

5 Results

The correlation between the sex ratio and the crime rate is evident from Figure 4, which plots the crime rate against the sex ratio (both variables are demeaned of province and year fixed effects). The scatter plot shows a general upward trend. The fitted line has a slope of 1.896, which is precisely estimated with a t-statistic of 3.52, suggesting that an increase in the sex ratio by 0.01 raises the crime rate by 1.90%.

5.1 OLS Estimates

We first estimate Eq. 5 by OLS, using heteroskedasticity robust standard errors. We start by including only the sex ratio (in addition to the province and year dummies) in Column 1 of Table 2 (equivalent to the fitted line in Figure 4). Additional controls are added sequentially in Columns 2-5. Throughout, the estimated coefficient on the sex ratio variable is positive and statistically significant at the 5% level, and neither the point estimates nor the standard errors are much affected by the inclusion of the additional controls.

The estimated coefficient on log per capita income is consistently positive, which suggests that the positive scale effect dominates the negative income effect. The employment rate has the expected negative effect, albeit borderline significant. Also, as expected, inequality and the urbanization rate have positive and statistically significant effects, suggesting that the rise in crime rates may be partly attributable to rapid urbanization and rising inequality. Population density has the expected positive effect, albeit not consistently significant. Neither secondary school enrollment nor welfare expenditure seem to have an effect.

We also include log police expenditure (Columns 3-6) to capture its potential impact on crime. However, reverse causality – rising police expenditure as a result of rising crime – is a distinct concern (Levitt (1997)). Because we do not have a suitable instrument to identify the causal effect of police expenditure, its inclusion may be viewed as a sensitivity test. Interestingly, and consistent with the reverse causality argument, the coefficient on the police expenditure variable is positive. Throughout, however, the coefficient on the sex ratio variable remains positive and significant.

As another robustness test, we include (out-of-province) immigration in the regression equation (Columns 4-6). The estimated coefficient on the share of immigrants is positive in all of the specifications, which is consistent with the theoretical prediction of the standard crime model (Becker (1968); Ehrlich (1973)).[17] Still, the magnitude and significance of the estimated coefficient on the sex ratio remains essentially unchanged, which suggests that variations in migration patterns are largely orthogonal to changes in the sex ratio.

[17] High population mobility presumably reduces the probability of being caught. Alternatively, a high immigrant share may reflect the economic climate (not captured by the income measures).

Finally, we include some variables that capture the age composition of the population. In Column 5, we estimate a regression in which the proportion of the population that is between 16 and 25 is included. This is to control for the mechanical impact of the relative size of the crime-prone cohort on the crime rate. As expected, the estimated coefficient is positive and statistically significant (1%). Still, the estimated effect of the sex ratio remains largely unchanged, further supporting the hypothesis that the sex ratio has an independent effect on crime. The last column includes the fraction of the population that is young (0-14 years old) and old (65 years or above). Although both have a significantly positive effect, this finding goes away once we instrument for the sex ratio (Table 4, Column 6).

5.1.1 Endogeneity

Although province fixed effects can control for time-invariant differences between provinces, sex ratios may be endogenous to crime rates, and the resulting bias is *a priori* ambiguous. To the extent that crime is a high-risk occupation that disproportionately employs and victimizes men, crime would depress the sex ratio and, consequently, bias the OLS estimates downward. An additional factor working in this direction is the aforementioned Trivers-Willard effect: poor social conditions may depress the sex ratio (Trivers & Willard (1973)). In contrast, crime may raise the sex ratio. For instance, women may disproportionately leave high crime areas. To address these potential biases, we use the one-child policy variables to instrument the sex ratio.

5.2 IV estimates

We use lagged (one year before the birth year for the cohorts aged 16-25) province-level, family-planning policy variables as IVs. These IVs should be suitable for addressing the problem of endogeneity, as they are highly correlated with the sex ratio, and it is unlikely that these lagged variables are contemporaneously correlated with the crime rate (other than through the posited channel).

To match the policies to the crime-prone cohorts, we use the cohort-size-weighted policy variables as our instruments. Specifically, we compute the weighted policy variable P_t^* in year t and province i as follows (the province subscripts are suppressed for clarity of exposition):

$$P_t^* = \frac{\sum_{g=16}^{25} C_g P_{t-g-1}}{\sum_{g=16}^{25} C_g}, \tag{6}$$

where C_g is the number of people who are g-years old in year t (and province i), and P_{t-g-1} is a policy dummy $g+1$ years ago, i.e., the year preceding the g-year-old's birth. In other words, P^* is the proportion of the 16-25 age cohort whose mothers were subject to the one-child policy the year before they were born. We construct our IVs by calculating P_t^* for our three policy variables: *Family Planning Science and Technology Research Institutes, -Education Centers and -Associations.*

The first-stage estimates are presented in Table 3. In addition to the three IVs, we also include all of the other control variables. Note that the IVs are jointly significant at the 1% level in all six cases, with F-statistics exceeding 10. These IVs are also individually significant in most cases.

Table 4 reports the IV estimates with heteroskedasticity-robust standard errors. Because we have three instruments for the sex ratio variable, we can conduct the Sargan-Hansen overidentification restriction test.[18] The p-values for the Hansen J-statistics reported in all of the specifications in Table 4 are very large, and we therefore cannot reject the validity of our instruments (conditional on a correctly specified model and on at least one of the instruments being valid).

The results from the IV estimations confirm our OLS results: increasingly male-biased sex ratios have had a large impact on crime. First, we report the results of a regression that only includes the (instrumented) 16-25 sex ratio (in addition to the province and year dummies) (Table 4, column 1). The 16-25 sex ratio has a positive and significant (1%) effect on the crime rate. As before, our key result is robust to the inclusion of other variables (Columns 2-6). The signs of the estimated coefficients for the control variables remain mainly the same as in the OLS estimations.

The estimated coefficient on the 16-25 sex ratio variable ranges from 4.68 to 6.18, that is, a rise in the sex ratio of 0.01 increased the crime rate by 4.68-6.18%. The fact that the sex ratio (ages 16-25) rose by 0.04 (from 1.053 in 1988 to 1.095 in 2004) suggests that the rise in the sex ratios raised crime rates by some 20-25%, accounting for a quarter to a third of the overall rise

[18] We first regress the residuals from the second-stage regression on all exogenous variables, including the IVs, and obtain the unadjusted R^2. We then compute Sargan's statistic, $N \times R^2$, where N is the number of observations. Under the joint null hypothesis that the instruments are valid, and that the instruments are correctly excluded from the estimated equation, the test statistic is χ^2-distributed with the number of degrees of freedom equal to the number of overidentifying restrictions (two in our case). In the presence of heteroskedasticity, Sargan's statistic becomes the Hansen J-statistic (Hayashi (2000), pp. 227-8, 407, 417).

during the sample period (82.4% from 1988 to 2004).

Overall, the IV estimates of the effect of the sex ratio are larger than the OLS estimates. In addition to the possibility of the endogeneity of the sex ratio to the crime rate already mentioned (specifically, the possibility that crime lowers the sex ratio), sex-differential migration may be another reason for the OLS estimates to understate the effect. In industrialized countries, women of marriageable age migrate to richer (and more urban) areas, which offer two (better) sources of income: potential partners (Fan & Huang (1998)) and work. Although, arguably, men also migrate, their incentives are less pronounced, as only the labor market is a source of income (marriage being on the expenditure side (Edlund (2005))). As a result, the projected sex ratio may understate the 16-25 sex ratio in poor and rural areas and overstate it in richer and urban areas. Because the latter also have higher crime rates, this would result in the projected sex ratios underestimating the effect on crime.

5.2.1 Robustness

Another concern is misreporting of births (Zeng *et al.* (1993); Croll (2000)).[19] To avoid or at least defer heavy fines for above-quota births, many households hid births. As these children have no legal identity and thus are not

[19] Gender-differential mortality is another potential source of measurement error. However, gender differences in mortality are most pronounced during infancy and old age. We focus on the 16-25 sex ratios for the period of 1988-2004, which means that the cohorts of interest were born between 1963 (1988-25) and 1988 (2004-16). In 1990, these cohorts were between the ages of two and 27, the mortality rates of which were not very sex-biased.

entitled to social benefits, daughters are more likely to be unreported. However, most scholars agree that these girls are eventually reported, usually upon reaching school age (and their parents will probably have paid the fine by then), because unregistered children are not allowed to go to school (Cai & Lavely (2003); Banister (2004)). Therefore, we can use the different censuses to test the seriousness of the measurement error that stems from misreporting.

Compared to 1990, people were eight years younger in 1982, the estimate using this census should be subject to more serious measurement error. Therefore, it would be reassuring if the regression results using the 1982 census were similar to those using the 1990 census. This is largely what we find (Column 1, Table 5).[20]

5.2.2 Falsification Tests

A possible objection is that the sex ratio may be correlated with policies or socioeconomic changes that are liable to exert an independent effect on the crime rate. To address this issue, we conduct two falsification tests. In the first test, we regress the corruption rate on the 16-25 sex ratio. If our posited mechanism holds, then we would expect lower or little impact on high-skill or white-collar crime, such as corruption, because the men who are liable to engage in this type of crime have largely been insulated from the shortage

[20] The number of observations is lower, because the latest year for which the 1982 census can provide a projected sex ratio for the 16-25 age group is 1998. In addition to fewer years, we only have 26 provinces, as Hainan was part of Guangdong in the 1982 census. Thus, the total number of observations for the 26 provinces over the period 1988-1998 is 286.

of women. If the 16-25 sex ratio is insignificant, then this would support our hypothesis that the surplus of young men had an impact on low-skill crime and is not mirroring an overall increase in criminality.

Figure 5 plots the corruption rate against the sex ratio (both variables are demeaned of province and year effects). The fitted line has a slightly negative slope and is not significant (p-value 0.273). The IV estimates confirm this conclusion (Table 5, Column 2). That is, we find no evidence that the 16-25 sex ratio has raised corruption rates, which is what we would expect were our model correctly specified.

Our second falsification test is to examine whether the sex ratios of the 10- to 15-year-old cohort (too young to commit crimes) have an impact on the crime rate. If they do, then this suggests that our results could be driven by unobservable determinants of crimes that are correlated with the sex ratios (and with the IVs).

The IV estimates (Table 5, Columns 3-4) show that the 10-15 sex ratio has no impact on the crime rate, whether we control for the 16-25 sex ratio or not. The coefficient on the 10-15 sex ratio is negative (Column 3), but is not significantly different from zero. When we also include the 16-25 sex ratio, the coefficient on the 10-15 sex ratio remains negative and insignificant. Moreover, the 16-25 sex ratio continues to have a large and significant impact on the crime rate (Column 4). These results suggest that our main findings are not likely to be driven by unobserved factors.

5.2.3 Is the Rise in Crime Rates Driven by Sex-Related Crimes?

One possibility is that the rise in violent and property crimes is driven by "sex related" crimes: rape and abduction of women and children.[21] While a breakdown does not exist at the province level, we can examine the patterns at the national level. These data do not indicate the rise in violent and property crimes to be driven by sex related crimes. Figure 6a shows that the rape rate rose between 1985 and 1992. However, this was probably not related to the sex ratio, as the 16-25 sex ratio during this period (Figure 1b) was rather flat (the first cohort of the one-child policy were still young in 1992). More interestingly, the rape rate began to drop in 1992 (as did abduction of women and children, not shown), which is in stark contrast to the overall rise in property and violent crime rates in the same period (Figure 6b).[22] Thus, our finding on the large impact of the sex ratio on property and violent crimes is likely driven by "non-sex" related types of crimes.

[21] There is only one "rape" category (attempted rape and sexual assault are either not crimes or subsumed under rape). Abduction of women and children is to a large extent for the purpose of selling as brides or prostitutes.

[22] A possible explanation for the decline in sex related crimes is that prostitution (illegal, but not part of violent and property crimes) has increased rapidly in the last two decades (Jeffreys (2004)). In his article "Surfeit of boys could spread AIDS in China's cities" published in *Nature* 434, 425 (24 March 2005), David Cyranoski said "...it [China] has admitted that there are now between 4 million and 6 million [prostitutes], compared with only 25,000 in 1985." doi:10.1038/434425b; Published online 23 March 2005. http://www.nature.com/nature/journal/v434/n7032/full/434425b.html

6 Summary and Discussion

A preference for sons has resulted in Chinese births being increasingly male. In 2000, 120 boys were born for every 100 girls, a development that has raised a number of concerns, ranging from human rights issues surrounding the fate of the "missing females" to the social impact of surplus men. Who are these men going to marry, and, if they do not, what are the consequences? High sex ratios are not unique to China, but, unlike India, where population growth has buffeted some of the impact, the shortage of brides is likely to be felt more acutely. Also, whereas South Korean men have been able to turn to poorer neighbors to source brides, China's bachelors, likely to be poorer in a still poorest country, are unlikely to be in a position to do so.

The same period has also witnessed a dramatic increase in crime. Although the notion that unbalanced sex ratios may raise crime is a long-standing hypothesis, such causality has been difficult to establish. This paper has documented a causal link between surplus young men and criminality by exploiting the natural experiment created by the one-child policy, which, in combination with a traditional preference for sons, raised sex ratios substantially above the naturally occurring 1.05 (males to females at birth). We find that the sex ratio among those 16-25 years old has had a large and statistically significant impact on crime. Using cross-province variation in the roll-out of three key features of the one-child policy, we estimate that male-biased sex ratios may account for a quarter to a third of the overall rise in violent and property crime during the period 1988-2004.

Granted, China has seen other dramatic changes during the study period. Although the state tightened its grip on fertility, its influence over virtually all other spheres of life diminished markedly. The introduction of a *de facto* market economy, rapid economic growth, and reduced state control over all facets of life (except fertility) are factors that may have contributed to the rise in crime rates. However, the fact that our results are robust to the inclusion of many province-level variables that proxy for these developments (including province-fixed effects) and that the sex ratio for a younger cohort (10-15 years old, young to commit crimes) has no impact on the crime rate suggest that our estimate is unlikely to be caused by unobserved time-varying factors. Furthermore, the effect of the sex ratio is not evident in a type of crime that we would not expect to be influenced by the sex ratio – corruption – further supports our hypothesis that the surplus of young men has had a causal and economically important impact on low-skill crimes.

To interpret our quantitative results, we have also presented a simple decomposition illustrating how we would expect crime rates to change in response to changing demographics. A higher sex ratio raises crime through three channels. First, (and trivially) a higher sex ratio implies more men, which by itself would raise crime rates as men are more prone to crime than women are. Second, a higher sex ratio raises the fraction of single men and lowers the fraction of married men (and women), which raises crime rates if unmarried men are more crime prone than married men. We derive upper bounds on the elasticity of the crime rate with respect to the sex ratio for these two effects, and attribute residual increases to a third mechanism: higher sex ratios may impact the propensity to commit crime, *conditional* on marital status. One reason for this is that a surplus of males is likely to result in more intense male competition in the marriage market. This is particularly the case for low-skilled men, as they are likely to be the ones left without brides. As a consequence, these men may resort to more risky activities to improve their marriage prospects, with crime being a prime candidate.

According to our decomposition, the estimated size is well above what we would expect if the rise in crime were simply due to more men. Moreover, the magnitude is suggestive of the possibility that higher sex ratios (and the reduced probability of marriage or partnership) have raised male propensity to commit crime (conditional on marital status). The implied elasticity of the crime rate with respect to the sex ratio ranges from 4.95 to 6.59 when the sex ratio is evaluated at its sample mean of 1.06 (IV estimates). This finding should be compared to the upper bound on the elasticity from the first two effects, derived to be the ratio of all men to unmarried men. Although straightforward, the calculation of this number is sensitive to which age groups are considered. We believe that an upper bound of 40 years old makes sense, as there is a pronounced drop in (low skill) criminality with age. As for the lower bound on the ages to include, the late teens and early 20s mark the entry into adulthood (and the search for a partner), and it may therefore be reasonable to include 20-year-old. Such cut-offs would yield an upper bound on the elasticity from the first two effects of 3.7, using the 2000 Census (Figure 3). As a point of reference, we note that the conclusion that higher sex ratios have raised male criminality across the board remains when ages 22 to 40 are considered, but becomes more tenuous as the lower age bracket is moved up.

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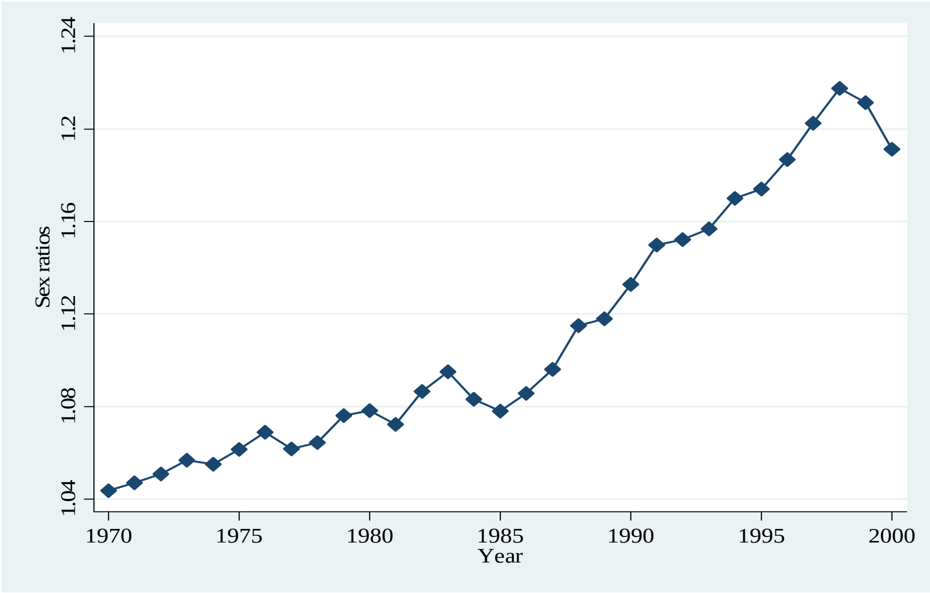
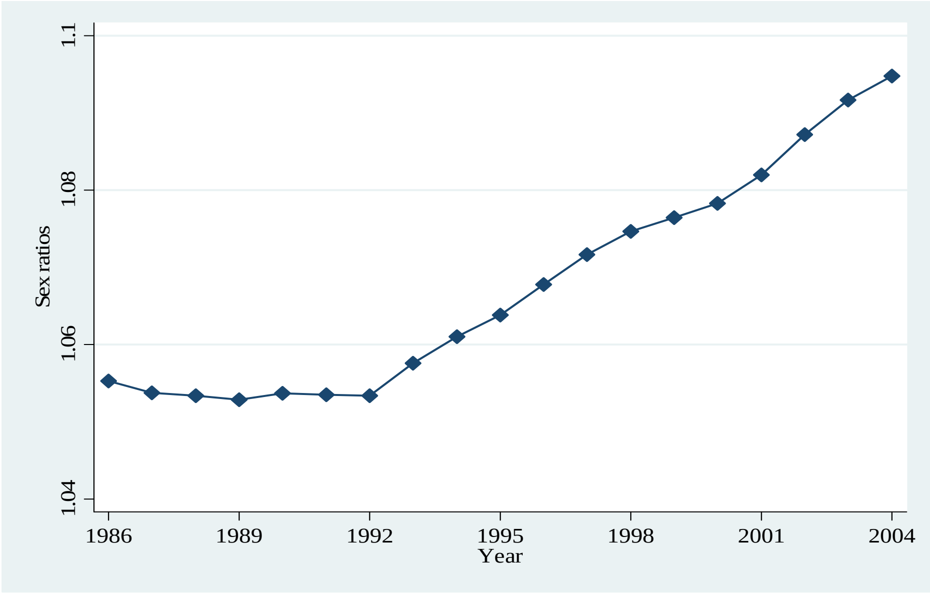
Figure 1a: Sex ratios by birth year



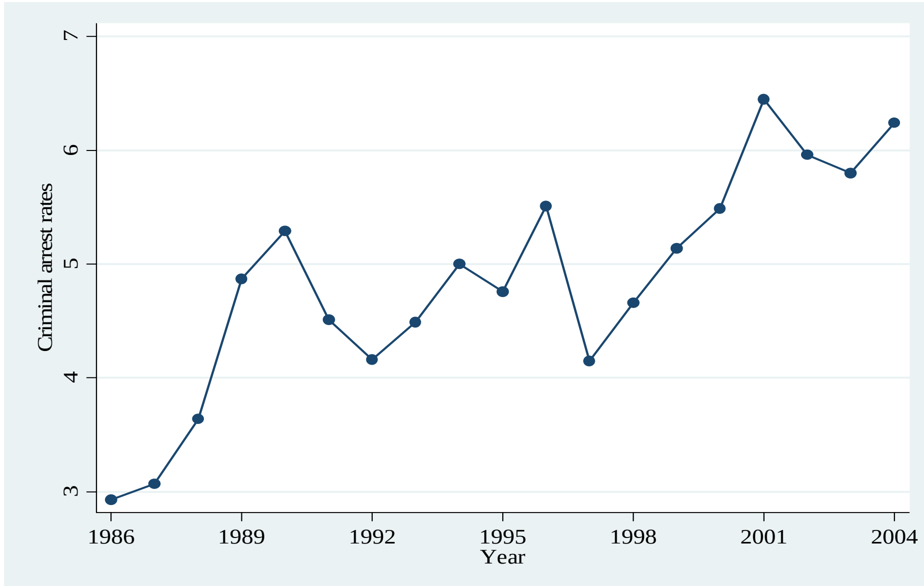
Note: Calculations for the 1970-1990 birth years are based on the 1990 Chinese population census (1% sample), and calculations for the 1991-2000 birth years on the 2000 census (1% sample).

Figure 1b: Sex ratios of the 16-25 age cohort, by year



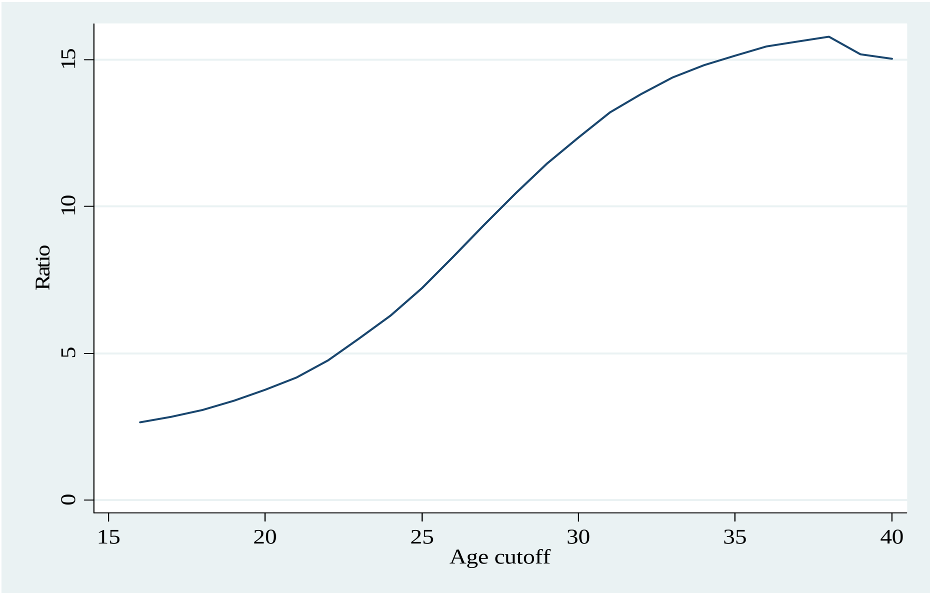
Note: Calculation is based on the 1990 Chinese population census (1% sample).

Figure 2: Criminal arrest rates (property and violent crimes): 1986-2004



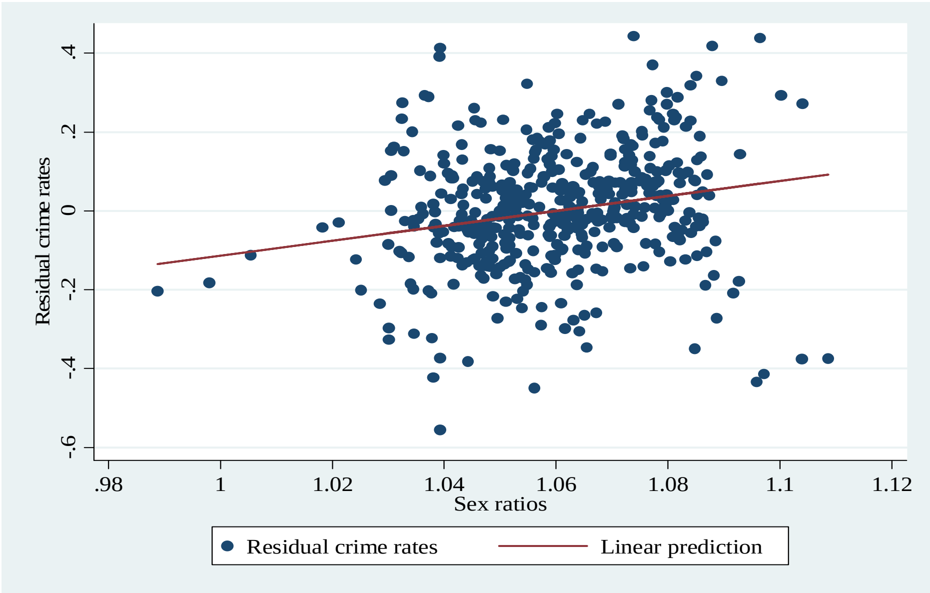
Notes: (i) data sources: Chinese Supreme People's Procuratorate (1986-2005), *Procuratorial yearbook of China*, Beijing, The Publishing House of Law; (ii) two structural breaks of the criminal arrest rates in China deserve noting. First, in 1989-1990 there is a marked increase in violent and property offences, possibly due to weakened government and social control in the aftermath of the *Tiananmen Square Protests of 1989*. Second, in 1997 the 5th session of the 8th NPC passed the new amended *Criminal Law* and *Criminal Procedure Law*, which considerably modified the preceding 1979 *Criminal Law* and *Criminal Procedure Law*, leading to a structural break in 1997. (The regression analysis will control for year fixed effects.)

Figure 3: Ratio of males to unmarried males for age group (between age cutoff and 40 years old), by age cutoff



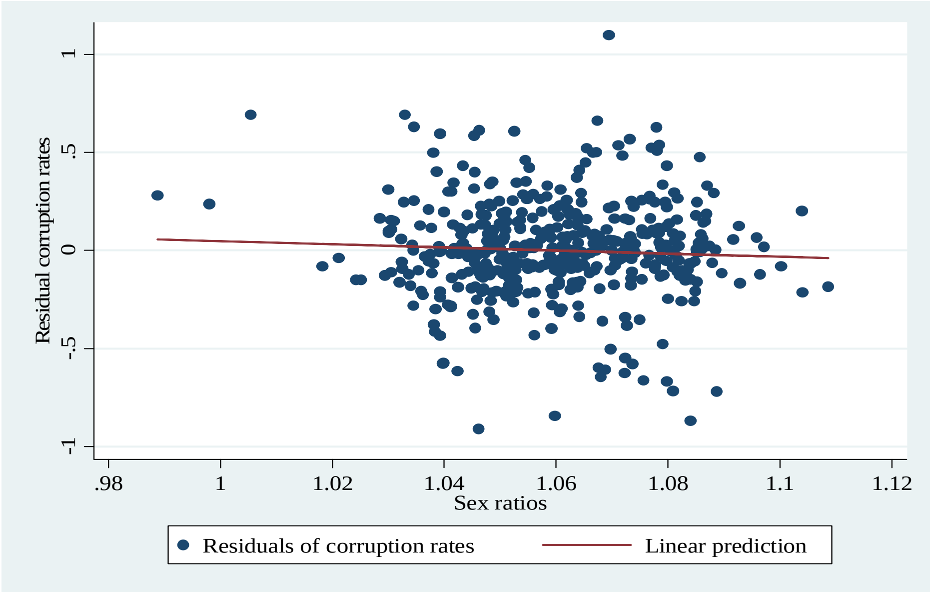
Note: Calculation is based on the Chinese 2000 population census (1% sample).

Figure 4: (Residual) crime rates by sex ratios, 1998-2004



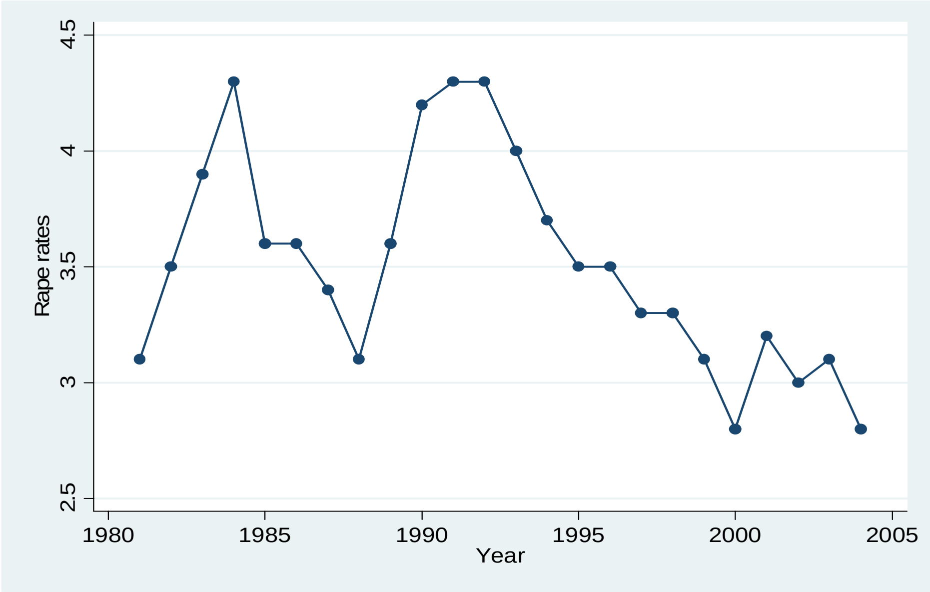
Note: Crime refers to property and violent crimes. Crime rates and sex ratios are purged of province and year fixed effects; coefficient: 1.896; *t*-statistic: 3.52, *p*-value: 0.001. Crime rates are in log form.

Figure 5: (Residual) corruption rates by sex ratios, 1988-2004



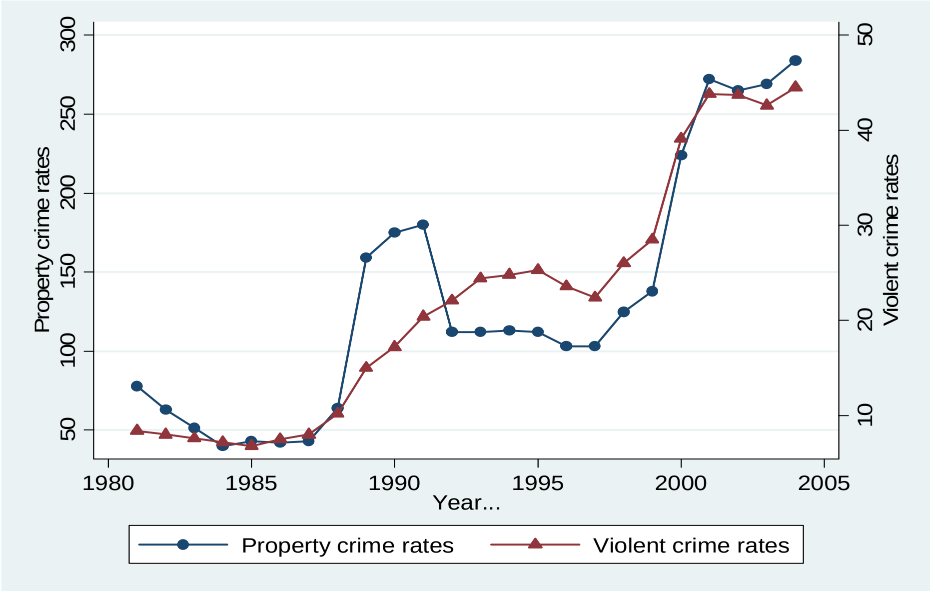
Note: Corruption rates and sex ratios are purged of province and year fixed effects estimation; coefficient: -0.799; *t*-statistic: -1.10, *p*-value: 0.273. Corruption rates are in log form.

Figure 6a: Rape rates by year



Notes: (i) data sources: Chinese Supreme People's Court (1985-2005), *Law Yearbook of China*, Beijing, The Publishing House of Law; (2) Rape rates are defined as the number of cases registered by the police per 10,000 population.

Figure 6b: Property crime rates and violent crime rates by year



Notes: (i) data sources: Chinese Supreme People's Court (1985-2005), *Law Yearbook of China*, Beijing, The Publishing House of Law; (2) both the property and violent crime rates are defined as the number of offence cases registered by the police per 10,000 population

Table 1: Summary Statistics of Variables

Variables	Mean	Standard deviation	Min	Max
Crime rate (number of arrests per 10,000 population)*	5.406	1.711	0.816	13.042
Corruption rate (number of corruption cases per 10,000 population)	0.454	0.214	0.086	1.515
Sex ratio (the ratio of males to females 16-25 projected by the 1990 census)	1.060	0.030	0.984	1.137
Real per capital income (RMB 1,000 at 2000 prices)	2.778	1.194	0.993	8.773
Employment rate (% employed of 16-65 year olds)	67.623	8.961	46.816	97.961
Inequality (the ratio of urban per capita income to rural per capita income)	2.717	0.695	1.528	5.159
Urbanization rate (% of population living in urban areas)	29.973	10.970	13.113	56.011
Population density (number of inhabitants per square km)	236.277	178.891	1.728	724.415
Secondary school enrollment rate (%)	87.059	11.058	39.6	100
Welfare expenditures share (% of government expenditures)	2.418	0.781	1.032	11.412
Police expenditures share (% of government expenditures)	5.436	1.404	2.588	10.273
Immigration rate (‰ of cross provincial immigrants)	2.237	1.291	0.310	8.680
Age 16-25 years old (% of population)	20.373	3.888	13.948	33.811
Age 0-14 years old (% of population)	26.037	4.629	13.892	36.739
Age 65 and above (% of population)	6.377	1.503	2.967	11.495

*Violent and property crimes. A non-exhaustive list includes: homicide, assault, robbery, rape, abduction of women and children, larceny, fraud, smuggling.

Data sources: *China Population Statistical Yearbooks, 1989-2005; China Statistical Yearbooks, 1989-2005; China Population Statistical Data and Material By Provinces and Cities, 1992-2004; Comprehensive Statistical Data and Materials on 55 Years of New China; Chinese Population Census (1982, 1990), 1% Sample; Law Yearbook of China, 1989-200; Procuratorial Yearbook of China, 1989-2005*

Table 2: The Effect of Sex Ratios on Crime Rates (province level data, 1988-2004, OLS)

	Dependent variable: ln(crime rate)					
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio (16-25)	1.896***	1.666**	1.796***	1.654**	1.663**	1.422**
	(0.565)	(0.686)	(0.672)	(0.663)	(0.672)	(0.631)
ln(per capita income)		0.129	0.118	0.111	0.106	0.141
		(0.094)	(0.093)	(0.093)	(0.092)	(0.093)
Employment rate (%)		-0.004*	-0.003	-0.003	-0.001	-0.003
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Inequality		0.155***	0.170***	0.164***	0.143***	0.141***
		(0.058)	(0.059)	(0.059)	(0.053)	(0.051)
Urbanization rate (%)		0.008***	0.007**	0.006**	0.003	0.003
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Population density		0.003*	0.002	0.002	0.002	0.002
		(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Secondary school		0.001	0.001	0.001	0.001	0.001
enrollment (%)		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
ln(welfare expenditures)		-0.033	-0.032	-0.040	-0.025	-0.001
		(0.042)	(0.042)	(0.042)	(0.042)	(0.045)
ln(police expenditures)			0.248***	0.239***	0.261***	0.234***
			(0.087)	(0.086)	(0.088)	(0.086)
Immigration rate (‰)				0.023	0.024*	0.025*
				(0.014)	(0.014)	(0.014)
Age 16-25 (%)					0.023***	0.032***
					(0.008)	(0.010)
Age 0-14 (%)						0.020**
						(0.009)
Age 65- (%)						0.039*
						(0.021)
Observations	459	459	459	459	459	459
Number of provinces	27	27	27	27	27	27
R-squared	0.50	0.56	0.57	0.57	0.59	0.59

Note: Robust standard errors in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. The sex ratio is defined as the ratio of males to females aged 16-25 with a sample mean of 1.06. Year- and province-fixed effects are included in all specifications.

Table 3: First Stage Regressions of the IV-Estimations

	Dependent variable: Sex ratio (16-25)					
	(1)	(2)	(3)	(4)	(5)	(6)
Instrumental variables (lagged 17-26 years)						
Family planning science& technology research institute	0.014*	0.005	0.006	0.013*	0.013*	0.011
	(0.008)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Family planning education center	0.031**	0.013	0.012	0.005	0.005	0.008
	(0.014)	(0.016)	(0.017)	(0.017)	(0.017)	(0.017)
Family planning association	0.049***	0.051***	0.050***	0.053***	0.055***	0.055***
	(0.008)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
Control variables						
ln(per capita income)		0.002	0.003	0.001	0.001	0.004
		(0.011)	(0.010)	(0.010)	(0.010)	(0.011)
Employment rate (%)		-0.001***	-0.001***	-0.001***	-0.001***	-0.001***
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Inequality		-0.001	-0.001	-0.003	-0.002	-0.002
		(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Urbanization rate (%)		-0.001**	-0.001**	-0.001**	-0.001**	-0.001**
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Population density		0.001**	0.001***	0.001***	0.001***	0.001***
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Secondary school enrollment (%)		-0.001	-0.001	-0.001	-0.001	-0.001
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
ln(welfare expenditures)		-0.013***	-0.013***	-0.014***	-0.014***	-0.011**
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
ln(police expenditures)			-0.020*	-0.022**	-0.023**	-0.025**
			(0.011)	(0.011)	(0.011)	(0.011)
Immigration rate (‰)				0.005***	0.005***	0.005***
				(0.001)	(0.001)	(0.001)
Age 16-25 (%)					-0.001	0.000
					(0.001)	(0.001)
Age 0-14 (%)						0.003***
						(0.001)
Age 65- (%)						0.005**
						(0.002)
Joint F test of IVs						
F-statistics	24.83	11.05	11.83	15.97	15.93	15.34
p-values	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001
Observations	459	459	459	459	459	459
Number of provinces	27	27	27	27	27	27
R-squared	0.39	0.42	0.43	0.45	0.45	0.47

Note: Robust standard errors in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. The dependent variable in all specifications is the 16-25 sex ratio (males to females) with a sample mean of 1.06. Year- and province-fixed effects are included in all specifications. See Figure A1 for the distributions of the three instrumental variables.

Table 4: The Effect of Sex Ratios on Crime Rates (province level data, 1988-2004, IV estimations)

	Dependent variable: ln(crime rate)					
	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio (16-25)	4.676***	6.180***	6.121***	6.160***	5.385***	5.594***
	(1.297)	(1.923)	(1.874)	(1.707)	(1.529)	(1.598)
ln(per capita income)		0.053	0.042	0.039	0.047	0.050
		(0.105)	(0.102)	(0.099)	(0.095)	(0.103)
Employment rate (%)		-0.000	0.001	0.001	0.002	0.002
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Inequality		0.153***	0.173***	0.172***	0.150***	0.149***
		(0.053)	(0.055)	(0.056)	(0.050)	(0.050)
Urbanization rate (%)		0.013***	0.012***	0.011***	0.008**	0.008**
		(0.004)	(0.004)	(0.004)	(0.003)	(0.004)
Population density		0.000	-0.001	-0.001	-0.001	-0.001
		(0.003)	(0.003)	(0.003)	(0.002)	(0.002)
Secondary school		0.002	0.002	0.002	0.002	0.002
enrollment (%)		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
ln(welfare expenditures)		0.022	0.020	0.019	0.023	0.034
		(0.052)	(0.051)	(0.051)	(0.049)	(0.049)
ln(police expenditures)			0.333***	0.332***	0.338***	0.328***
			(0.106)	(0.107)	(0.104)	(0.105)
Immigration rate (‰)				0.005	0.009	0.009
				(0.017)	(0.016)	(0.016)
Age 16-25 (%)					0.023***	0.028***
					(0.008)	(0.009)
Age 0-14 (%)						0.007
						(0.010)
Age 65- (%)						0.026
						(0.023)
Overidentification test:						
Hansen *J*-statistics	1.625	0.052	0.176	0.133	0.065	0.015
p-values	0.444	0.975	0.916	0.935	0.968	0.982
Observations	459	459	459	459	459	459
Number of provinces	27	27	27	27	27	27

Note: Robust standard errors in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. The sex ratio variable is defined as the ratio of males to females aged 16-25 with a sample mean of 1.06. Year- and province-fixed effects are included in all specifications. See Figure A1 for the distributions of the three instrumental variables.

Table 5: Robustness Tests on The Effect of Sex Ratios on Crime Rates (province level data, 1988-2004, IV estimations)

	Dependent variables			
	ln(crime rate)	ln(corruption rate)	ln (crime rate)	ln (crime rate)
	(1)	(2)	(3)	(4)
Sex ratio (16-25)	5.705**	-1.977		3.766**
	(2.852)	(2.935)		(1.828)
Sex ratio (10-15)			-2.871	-0.314
			(2.031)	(1.953)
ln(per capita income)	0.303*	0.164	0.351***	0.215
	(0.159)	(0.184)	(0.134)	(0.143)
Employment rate (%)	-0.004	0.003	-0.011**	-0.005
	(0.005)	(0.006)	(0.005)	(0.005)
Inequality	0.170***	-0.235***	0.141**	0.188***
	(0.054)	(0.062)	(0.058)	(0.059)
Urbanization rate (%)	-0.002	0.001	-0.001	0.001
	(0.006)	(0.005)	(0.005)	(0.005)
Population density	-0.002	0.002	0.005**	0.001
	(0.003)	(0.002)	(0.002)	(0.003)
Secondary school	0.004	-0.003	0.003	0.002
enrollment (%)	(0.005)	(0.003)	(0.004)	(0.003)
ln(welfare expenditures)	0.025	0.137*	0.036	0.038
	(0.054)	(0.079)	(0.064)	(0.060)
ln(police expenditures)	0.235	-0.238	0.113	0.365**
	(0.170)	(0.150)	(0.126)	(0.172)
Immigration rate (‰)	0.011	-0.010	0.010	0.005
	(0.022)	(0.019)	(0.020)	(0.021)
Age 16-25 (%)	0.042***	0.015	0.038***	0.038***
	(0.013)	(0.012)	(0.013)	(0.012)
Age 0-14 (%)	0.019	0.009	0.028**	0.016
	(0.013)	(0.019)	(0.011)	(0.013)
Age 65- (%)	0.080**	0.041	0.079***	0.065**
	(0.036)	(0.031)	(0.031)	(0.031)
Overidentification test:				
Hansen J-statistics	2.989	1.852	6.411	9.495
p-values	0.224	0.396	0.041	0.054
Observations	286	459	351	351
Number of provinces	26	27	27	27

Note: Robust standard errors in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. The sex ratio (16-25) is defined as the ratio of males to females aged 16-25, and the sex ratio (10-15) is defined as the ratio of males to females aged 10-15. The sex ratio in column 1 is projected from the Chinese population census in 1982 with a sample mean of 1.05, while the sex ratio in columns 2-4 is projected from the Chinese population census in 1990, with sample means of 1.06 (for ages 16-25) and 1.08 (for ages 10-15) respectively. Since Hainan was still included in Guangdong province in the Chinese population census in 1982, the number of provinces in the 1982 census is 26. In addition, the latest year projected for the age cohort of 16-25 by the 1982 census is 1998. Thus the total number of observations for the 26 provinces over the period of 1988-1998 is 286 for column 1. Columns 3-4 also have fewer observations because of missing values for the 10-15 sex ratio. See Figure A1 for the distributions of the instrumental variables. Columns 1-2 use the three program variables (lagged 17-26 years) as IVs; column 3 uses the three program variables (lagged 11-16 years) as IVs; column 4 uses both sets of variables as IVs.

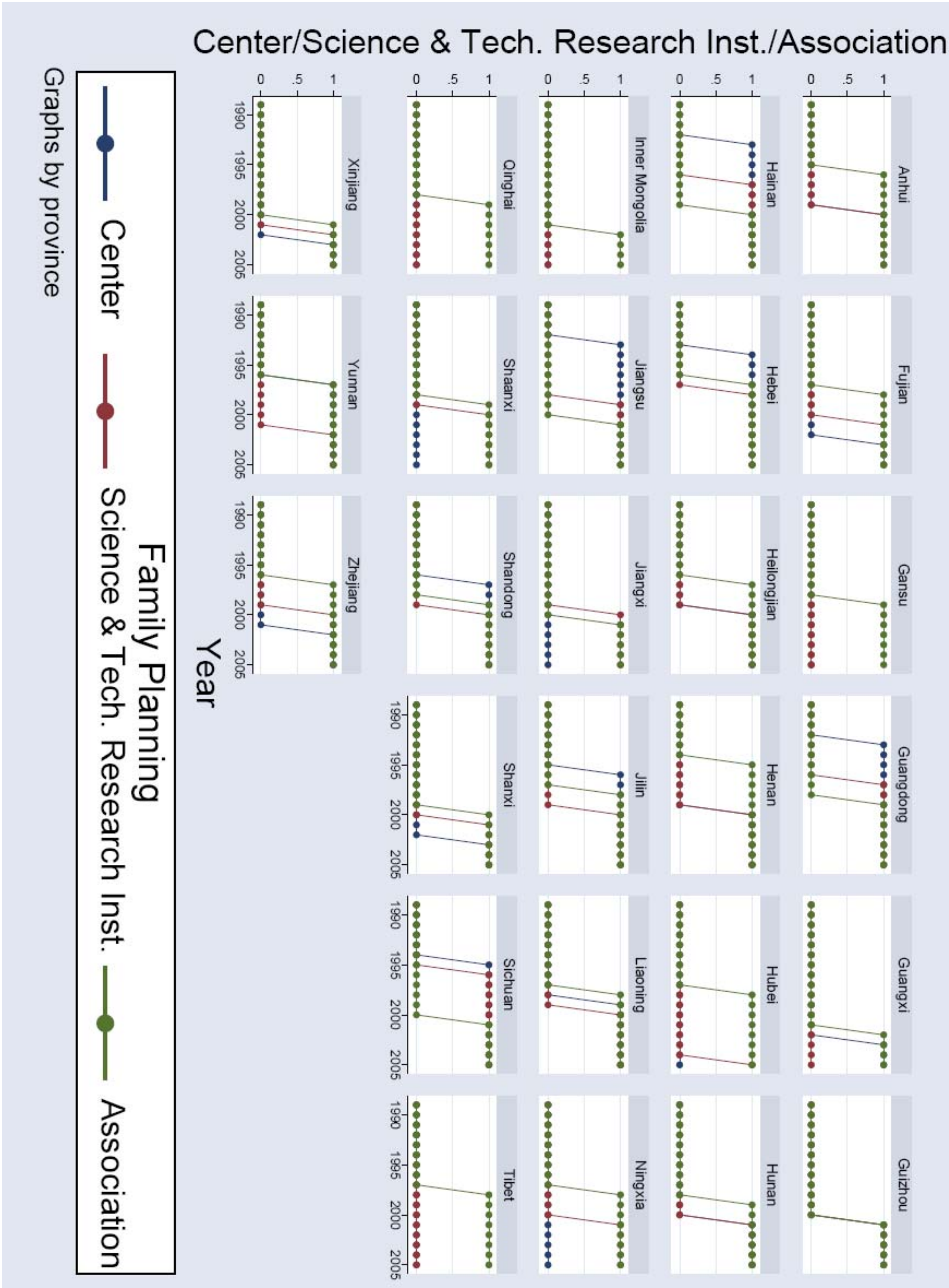


Figure A1: Program roll-out, by Province and Year of Implementation+17