# SEX RATIOS AND CRIME: EVIDENCE FROM CHINA

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*Abstract*—Since the introduction of the one-child policy in China in 1979, many more boys than girls have been born, foreshadowing a sizable bride shortage. What do young men unable to find wives do? This paper focuses on criminality, an asocial activity that has seen a marked rise since the mid-1990s. Exploiting province-year level variation, we find an elasticity of crime with respect to the sex ratio of 16- to 25-year-olds of 3.4, suggesting that male sex ratios can account for one-seventh of the rise in crime. We hypothesize that adverse marriage market conditions drive this association.

## I. Introduction

IN 1979, China launched the so-called one-child policy, placing population control at the center of a raft of policies aimed at lifting the country out of poverty. Ultrasound B machines for prenatal screening and induced abortions came to feature prominently in the arsenal employed to reduce the number and "improve the quality" of births (Zeng et al., 1993). As is now well known, the policy also marked the start of steadily rising sex ratios (Zeng et al., 1993; Chu, 2001; PRC, 2002; Yang & Chen, 2004; Das Gupta, 2005).[1] By 2000, almost 120 boys were born for every 100 girls, and according to the 2010 census, the subsequent ten years offered little change. There are now an estimated 30 million "surplus" boys (Zhu, Li, & Hesketh, 2009).

The recent decades have also witnessed a drastic increase in crime. Between 1988 and 2004, criminal offenses rose at an annual rate of 13.6% (Hu, 2006), and arrest rates almost doubled (figure 1). Candidate explanations have included the economic reforms, rising inequality, and weakened social control (Liu, Zhang, & Messner, 2001; Bakken, 2005; Hu, 2006).

In this paper, we propose an additional explanation for the rise in crime: "surplus" men. Between 1988 and 2004, the sex ratio (males to females) of young adults (16 to 25 years old) rose from 1.02 to 1.06 (see figure 2), implying a tripling of surplus men, a trend that will continue until at least 2030 judging by the latest census. The potentially disruptive social consequences of millions of young men with few or no prospects of marriage have received increasing attention (Hudson & Den Boer, 2002). Crime is a case in point. That young unmarried men are the most crime prone is a truism in criminology, and China is no exception. In 2000, 90% of arrestees were men (Chinese Law Yearbook Editorial Office, 2001), and young adults are vastly overrepresented. Young men ages 16 through 25 account for more than two-thirds of violent and property crimes (Hu, 2006).

For our core analysis, we assemble a province-year panel data set covering the years 1988 to 2004, 1988 being the first year for which province-level crime statistics are available. We are interested in the impact of the increasing maleness of young adults and focus on the province sex ratio of 16- to 25-year-olds (henceforth, 16–25 sex ratio). For annual data, we rely on projections from the decennial censuses.

The 2000 census contains information on the provinces of both birth and residence. Thus, the 2000 census can be used to calculate sex ratios in a number of ways. Our preferred sex ratio is one that assumes that the province of birth constitutes the marriage market but that the province of residence is where crime is conducted. This sex ratio is a weighted average of the province-of-birth sex ratios of men residing in the province, where their shares in the resident population provide the weights. For instance, if male residents in Beijing are 90% Beijing natives and 10% Sichuan natives, then the weighted sex ratio combines the sex ratios by birth province for Beijing and Sichuan, weighted 0.90 and 0.10, respectively. We refer to this as the weighted sex ratio. Empirically, it tracks the sex ratio by province of birth quite closely, and we will refer to both sex ratios as being based on province-of-birth information.

Results using the weighted sex ratio are stronger than those using the sex ratio by birth province, but both point in the same direction. Using the weighted sex ratio and including a number of province-year varying controls (the sex ratios of other age groups and measures of the economic climate) and province-specific time trends, we estimate an elasticity of crime with respect to the 16–25 sex ratio of 3.4. Over the study period, the 16–25 sex ratio rose by 4% and crime rates by 82.4%, suggesting that higher sex ratios can account for one-seventh ($3.4 \times 4/82.4$) of the rise in criminality.

The positive relationship between the 16–25 weighted sex ratio and crime is evident in the bivariate correlation and robust to the inclusion of province fixed effects. However, once time-varying covariates are added, the sex ratio is

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[1] The premium placed on sons over daughters in Chinese culture is reflected in sayings such as, "Raising a daughter is like watering a plant in another man's garden," and, "A daughter is a thief."

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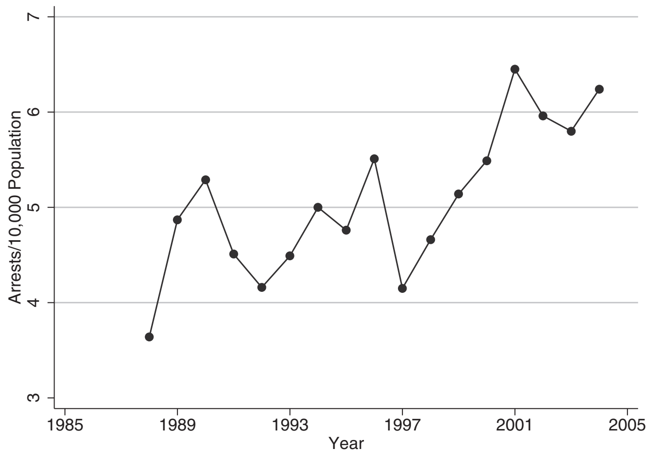
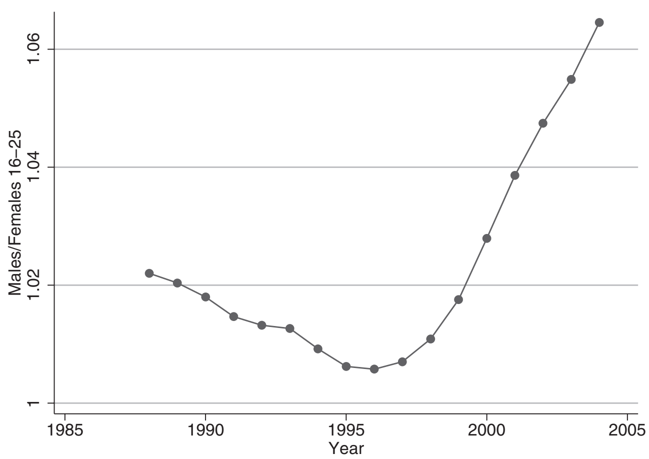
FIGURE 1.—ARRESTS PROPERTY AND VIOLENT CRIMES/10,000 POPULATION, 1988–2004




The drop in 1997 followed an amendment of the Criminal Law and Criminal Procedure Law.
*Source:* Chinese Supreme People's Procuratorate, *Procuratorial Yearbook of China* (Beijing, Publishing House of Law, 1986–2005).

FIGURE 2.—SEX RATIOS 16–25 YEAR OLDS, 1988–2004




Projected from the 2000 Chinese Population Census (1% sample).

estimated to raise crime only in a specification also allowing for a province-specific time trend.

Alternatively, we can calculate sex ratios by province of residence using either the 1990 or the 2000 census. The resident sex ratio is problematic, however. In China, as elsewhere, migration has a distinct economic and gender profile: women are more likely than men to leave economically depressed areas (Fan & Huang, 1998; Edlund, 2005). Thus, resident sex ratios and crime are likely to be linked through the economic climate. Migration was more limited in the 1990s. Therefore, we expect sex ratios based on the 1990 census to be more suitable than resident sex ratios using the 2000 census. Even so, the 1990 resident sex ratio yields only supportive results once the richest provinces (migration destinations) are excluded. Moreover, results are not robust to the inclusion of a province-specific time trend, further suggestive of the presence of a confounder, unmeasured economic climate being a prime candidate.

There are several reasons that higher sex ratios might affect crime. A higher sex ratio implies a more male population, which trivially leads to more crime if men are more crime prone than women—a maleness effect. A higher sex ratio also means that fewer men can be married, stiffening the competition for a wife and possibly raising the returns to risky activities, an incentive effect. Eventually, higher sex ratios means that fewer men can be married, and marriage may discipline men, a civilizing effect (Korenman & Neumark, 1991; Messner & Sampson, 1991; Barber, 2000; Sampson, Lavb, & Wimer, 2006).

Since the maleness effect can at most account for a 0.5 elasticity of crime with the respect to the sex ratio, effects beyond 0.5 might be attributed to men being more criminal. To see that 0.5 consists an upper bound, note that if the average male propensity to commit crime were unchanged, then a 1% increase in the fraction of men can raise crime at most by 1%. Since a 1% increase in the sex ratio maximally corresponds to a 0.5% increase in the fraction of men (assuming that men make up at least half of the population), more men alone can at most result in an elasticity of crime with respect to the sex ratio of 0.5 (see appendix A).

Thus, the estimated elasticity from our preferred specification exceeds what can be expected from simply more men, lending weight to the argument that higher sex ratios contributed to the rise in crime. A larger fraction of unmarried men (the civilization effect) could account for the found effect, but the fact that the average age at first marriage for men was close to 25 during the study period points toward a change in behavior, the incentive effect.[2]

To formalize the mechanics of the incentive effect, we present an adaptation of the lottery model proposed by Clark and Riis (1996). Men partake in a multiprize lottery, where the prize is one wife, by buying "tickets." As sex ratios rise above 1, the equilibrium level of male ticket buying initially increases.[3]

Crime is but one of a number of possible end points. To begin, it behooves us to ask whether marriage market outcomes vary with sex ratios in the assumed direction. Using the 2000 census, we find that counties with more men (to women) had a lower share of their men currently married. The pattern for women was the reverse. These findings are consistent with the basic premise of our paper: male sex ratios at birth present a challenge to men in those cohorts as they reach adulthood.

Would a future marriage squeeze not prompt men (or their parents) to invest more in education and training? To investigate this possibility, we use the annual Chinese Urban Household Surveys for the years 1988 to 2006. These surveys provide possibly the best household data for China. We focus on adults 18 to 45 years old, cohorts born in 1943 to 1988.

[2] The singulate mean age at marriage for men was 24.8 years in 1999. The UNDP World Fertility and Marriage Data Base 2003.

[3] Clearly, fewer potential wives will not have men increasing their investments ad infinitum. At some point, more men to women reduces the value of the lottery sufficiently as to make men turn to other pursuits.

For the bulk of these cohorts, sex ratios were roughly normal. To generate sex-ratio variation, we exploit the fact that men tend to marry women younger than themselves. As a result, cohort-size differences create sex-ratio variation. While we find higher sex ratios to increase the educational gender gap (in favor of men), lower female, rather than higher male, education accounts for the finding.

We also look at labor market outcomes, and the results are mixed. This is not entirely surprising. The incentive effect suggests a widening of the educational and labor market gender gap (in favor of men). The civilizing effect, however, works in the opposite direction.

Turning to marital bargaining, we use the Chinese Health and Nutrition Surveys (CHNS) and the information on intrahousehold allocation in the 1989, 1991, and 1993 waves. Because of temporal proximity, we use the 1990 census to calculate the sex ratios (again exploiting spousal age gaps) and find that higher sex ratios raise male time in household work and female participation in decision making, consistent with the findings of Porter (2007), who used the CHNS and the 1982 census.

Our paper builds on a long tradition in the sociobiological literature linking male behavior—risk taking, dominance, male-to-male aggression—to competition for partners. Trivers wrote (1972, p. 153), "Females compete among themselves for such resources as food but not for members of the opposite sex, whereas males ultimately compete only for members of the opposite sex, all other forms of competition being important only insofar as they affect this ultimate competition."

The possibility of a link between sex ratios and education and labor market investments has received some attention in economics (Angrist, 2002; Lafortune, 2008). The paper perhaps most closely related to ours is Wei and Zhang (2011) who argued that male sex ratios have contributed to China's high savings rate.

The remainder of the paper is organized as follows. Section II presents a simple model formalizing the incentive effect of a higher sex ratio. Section III discusses the different sex ratios. Section IV describes our data. Section V presents our main result: the effect of rising sex ratios on crime. Section VI presents auxiliary evidence that marriage market sex ratios have a gender differential effect on education, labor, and marital-bargaining outcomes. Section VII concludes.

## II. Incentive Effects of Male-Biased Sex Ratios: A Model

Consider a population of initially homogeneous men and women. Marriage is monogamous, and women are literally scarce. Men are assumed to invest in activities that increase their probability of obtaining a wife.

This formulation may serve as shorthand for a process where wives are allocated according to earnings power (his wage) and marriage matching is preceded by a phase where human capital investments are made. Since the purpose of these investments is partner competition, they can be characterized as premarital investments. For simplicity, we assume that only men make these investments.[4]

Sex ratio effects on premarital investments under monogamy may be couched in terms of a lottery model with multiple prizes where men compete to obtain one wife. Clark and Riis (1996) noted that the fact that winners are eliminated (a man can win only one wife) implies that the problem has a nested structure. In the special case that the valuation of the prize is invariant to the number of prizes (a reasonable approximation for the application at hand), Edlund (1996) showed that a decrease in the number of prizes can increase premarital investments, replicated below.

We are interested in how a decrease in the number of prizes (women) relative to the number of contenders (men) affects the level of male premarital investments (such as, crime). We focus on a symmetric equilibrium, that is, where men make identical investment decisions.[5]

Let there be $M$ men and $F$ women, $M > F$. Denote the vector of premarital investments by $\bar{h} = \{h_1, h_2, \ldots, h_M\}$, and let $h$ denote the average outlay. Each winner obtains one wife, valued at $z$. We can write the probability that player $i$ wins one of the $F$ prizes as the sum of probabilities of winning the $s$th prize conditional on not having won in the previous rounds:

$$P_i = p_i^1 + \sum_{s=2}^{F} \prod_{k=1}^{s-1} (1 - p_i^k) p_i^s.$$

Furthermore, we assume the following version of independence of irrelevant alternatives:

$$\begin{aligned} &p_i(h_1, h_2, \ldots, h_{j-1}, 0, h_{j+1}, \ldots, h_M) \\ &\qquad = \frac{p_i(h_1, h_2, \ldots, h_{j-1}, h_j, h_{j+1}, \ldots, h_M)}{1 - p_j(h_1, h_2, \ldots, h_{j-1}, h_j, h_{j+1}, \ldots, h_M)}, \end{aligned}$$

that is, the probability of man $i$ winning, when $j$ is not participating, equals the probability of winning conditional on $j$ not winning, which together with

$$p_i^1 = \frac{h_i}{Mh}$$

yields the following expression for the probability of winning in the $s$th round:

$$p_i^s = \frac{p_i^1}{(1 - (s-1)p_j^1)}, \text{ for } s > 1.$$

[4] Allowing for female investments would not change the qualitative results but would add considerable complexity. It would introduce female heterogeneity, which would mean that the stable matching would have to be described in terms of who marries whom, not just which men marry. Also, the associated transfer payment would be couple specific, which could reasonably have an impact on the premarital investment decision (e.g., Cole, Mailath, & Postlewait, 1992). Finally, to the extent that investments are made by parents of children of both genders, the decision to invest in one child's human capital might affect the other child's through a common budget constraint.

[5] The unique equilibrium in this game (Clark & Riis, 1996).

The expected payoff for player $i$ is

$$\pi_i(\bar{h}) = P_i z - h_i.$$

Hence, man $i$ sets $h_i$ to satisfy the first-order condition

$$\frac{\partial P_i}{\partial h_i} z - 1 = 0,$$

which, evaluated at a symmetric equilibrium, yields

$$h_i = h = \frac{z}{M} \frac{(M-1)}{M} \left( F - \sum_{k=1}^{F-1} \frac{F-k}{M-k} \right). \tag{1}$$

In order to study how male premarital investments are affected by a change in the sex ratio, we differentiate the expression for $h$ in equation (1), which for large $M$ implies that

$$\frac{dh}{d\frac{M}{F}} > 0 \text{ for } \frac{M}{F} < c$$
$$< 0 \text{ for } \frac{M}{F} > c,$$

where $c \approx \frac{e}{e-1}$.[6] In other words, there is a range for which higher sex ratios raise male premarital investments. In our setup, if sex ratios balanced, all men would be guaranteed a wife and there would be no marriage market return to premarital investments. As women become more scarce, the returns to such investments increase initially. However, fewer women lowers the value of the lottery, and eventually this effect dominates and men reduce their investments.

## III. Sex Ratios

To test our hypothesis that male sex ratios have contributed to the rise in crime in China, we need to decide on a sex ratio. On a conceptual level, what sex ratio do we think is important for criminal behavior? Is it the sex ratio at birth, during adolescence, or some later time? Furthermore, assuming the partner market is the reason sex ratios have an impact on behavior such as crime, what constitutes the relevant partner market?

We focus on the province-level sex ratio for a number of reasons. First, our outcome variable, crime, is reported at the provincial level. Second, the implementation of the birth planning policies was delegated to the provinces. In fact, by the late 1990s, it was the only major policy area lacking national enabling legislation: the national Law on Population and Birth Planning came into effect in 2002 (PRC, 2002). Third, until the early 1990s, interprovince migration was strictly regulated, arguably rendering marriage markets more provincial than national.

The mechanism enabling control over population mobility was the household registration (*hukou*) system put in place by the Communist party in the early 1950s. Under the system, everybody is registered in the place of parental *hukou* and obtains a *hukou* certificate. Changes to *hukou* are rare. Initially, access to everyday necessities such as food, cooking oil, and clothing was effectively regulated by the *hukou*. As free-market activities were increasingly tolerated through the 1980s, the de facto hold of the *hukou* system began to wane, and by the early 1990s, its role in the allocation of consumer goods was abolished. *Hukou* status still governs activities such as land distribution, identity card issuance, school registration, and marriage registration.

Following the liberalization of the economy and the relaxation of the *hukou* system, internal mobility increased substantially, giving rise to a large so-called floating population, a pejorative term connoting human flotsam. These unofficial migrants tend to be young, unskilled, and, contrary to conventional wisdom, more female than male. Still, a large number of largely unskilled male migrants have moved from the hinterland to the economically more vibrant coastal provinces, where they form an underclass. Most migrants expect to eventually return to their province of birth and marry someone of their own *hukou* status: this is particularly true of male migrants who are less likely to "marry up" than their female counterpart (Yusuf & Saich, 2008; Fan, 2008).

Another challenge is how to calculate the sex ratio without natality data. China has long delegated the registration of births, deaths, and marriages to the family, and nationally comprehensive vital registration data do not exist (Yang et al., 2005). The lack of natality data forces us to lean on projections from the decennial population censuses. The ages used in the 1990 and the 2000 censuses for the sex-ratio projections are tabulated in tables I and II in appendix B. (Appendixes B to D are in the online supplement.)

Since our study period is 1988 to 2004, the 1990 and the 2000 censuses are those best suited for our analysis. The 2000 census contains information on the provinces of birth and enumeration. The 1990 census contains information on the province of enumeration only.

Against this backdrop, three candidate sex ratios exploiting the availability of province of birth and residence information in the 2000 census stand out:

- Birth: The ratio of males to females born in the province. For province $i$, it is

$$r_i \equiv \frac{m_i}{f_i} = \frac{\sum_j m_{ij}}{\sum_j f_{ij}}, \tag{2}$$

  where $m_{ij}$ is the number of men born in province $i$ and residing in province $j$, and analogously for women, $f$. We refer to this sex ratio as the *birth* sex ratio.
- Resident: The ratio of males to females enumerated in the province. For province $j$, it is

$$r_j^r \equiv \frac{M_j}{F_j} = \frac{\sum_i m_{ij}}{\sum_i f_{ij}}. \tag{3}$$

  We refer to this sex ratio as the *resident* sex ratio.

[6] For smaller $M$, the same qualitative results will hold. However, the particular cutoff point may vary. Also, symmetry as to differentiation with respect to $F$ or $M$ is a limit result. $e$ is the base to the natural logarithm.

- Weighted: The weighted average of the birth sex ratio of male residents. For province $j$, it is

$$R_j = \sum_i \frac{m_{ij}}{M_j} \times r_i, \quad (4)$$

where $r_i$ is the birth sex ratio in province $i$ and $m_{ij}/M_j$ is the share of the male population from that province. We refer to this sex ratio as the *weighted* sex ratio.

For example, consider the age 20 sex ratio for Beijing and the year 2000. The birth sex ratio is the sex ratio of those born in Beijing in 1980. The sex ratio among residents is that of 20-year-olds residing in Beijing in the 2000 census. Finally, the weighted sex ratio is the weighted average of the birth sex ratio, where the weights are given by the composition of 20-year-old Beijing residents in the 2000 census. For instance, if Beijing's male population consisted of 10% men from Shanghai and 90% men from Beijing, the weighted sex ratio for Beijing is the Shanghai and Beijing birth sex ratios, weighted by 0.1 and 0.9, respectively.

### A. *Sex Ratio Comparison*

The birth sex ratio would be the most pertinent to crime if there were little migration or the province of birth constituted the marriage market regardless of province of residence. This sex ratio can be computed using the 2000 census.

For the resident sex ratio to bear on crime, we need to assume that the province of residence constitutes the marriage market. This is unlikely given that, especially in the more recent years, the bulk of migrants are unofficial (no change to *hukou* status) and therefore have difficulty settling permanently. An advantage of this sex ratio, however, is that it can be computed for both the 1990 and the 2000 censuses.

The weighted sex ratio is our favored sex ratio. It assumes that the province of birth is viewed as the relevant marriage market but crime is committed in the province of residence. For instance, young men from Sichuan residing in Beijing would be influenced by the Sichuanese marriage market, while any criminal behavior would show in Beijing's crime statistics. This sex ratio can be computed using the 2000 census, but the fact that we need to rely on projections and a strong age profile to migration means that the weights attached to the various sex ratios are likely to be inaccurate, especially for projections relying on younger ages. For instance, for the weighted sex ratio for 18-year-old men in Beijing in 2004, we would use the 2000 census composition of 14-year-old males residing in Beijing. Since migration is very limited before age 16 (migrants often leave children behind with other family members), this method gives Beijing's sex ratio too high a weight.

To help the reader assess the differences between the sex ratios, appendix C plots the four sex ratios by province. The provinces are presented in descending order of percapita income in the 2000 census.

There are drawbacks to using the 1990 census to project sex ratios, a census for which only province of residence is available. (The graphs in appendix C illustrate these problems.) Contrary to the pattern suggested by the birth sex ratio projected by the 2000 census and our expectations based on the national trend of rising sex ratios, the sex ratios for the three province-level municipalities—Beijing, Tianjin, and Shanghai—show a clear decline. Not only is the decline remarkable, the sex ratios are exceedingly male at the beginning of the period. For instance, the 16–25 sex ratio for Shanghai is 170 males per 100 females in 1988. This pattern is likely an artifact of the age and migration pattern in the 1990 census. Recall that the 16–25 sex ratio for 1988 is calculated using the sex ratio of 18- to 27-year-olds in 1990. These cohorts were born from 1963 to 1972, a time of no prenatal sex selection and limited postnatal selection. Therefore, the large deviations from normal sex ratios are likely migration driven. Male sex ratios from migration would also fit with the observation that at the time, there was little unofficial migration and therefore in-migration was likely limited to skilled labor (and therefore male). Migration might also be behind the negative, albeit less pronounced, trend for Guangdong province. While not a province-level municipality, Guangdong is a rich province (its per capita income is just below that of Beijing, Tianjin, and Shanghai). Because Guangdong is not a special administrative division, migration into this province might have been less regulated and thus more gender balanced. Recall that for projections for later years, we use successively younger ages in the 1990 census. For the last year, 2004, we use 2- to 11-year-olds in 1990, ages with little interprovince migration (as evidenced in the 2000 census) and we see that the projected sex ratios for 2004 show considerably less variation. For poor provinces, a negative trend could also result from young women leaving for marriage (Fan & Huang, 1998).

Second, the weighted sex ratio, $R$, and the birth sex ratio, $r$, track each other closely. For provinces with little in-migration, this is what we would expect. Absent migration, the two sex ratios coincide. Richer provinces have attracted more migrants, and the two sex ratios diverge somewhat. Thus, our preferred sex ratio is quite similar to the birth sex ratio. The high degree of congruence suggests that the mismeasurement in the weighted sex ratio that results from reliance on projections and migration having a distinct age profile may be modest. (The reason we do not favor the birth sex ratio is that it assumes that crime is committed in the province of birth, an assumption at variance with the common perception of unofficial migrants eking out a living at the margins of society.)

Third, there is a fair amount of separation between the resident and the weighted sex ratios (both based on the 2000 census). This is because, say, an inflow of men moves the resident sex ratio one-to-one (in percentage terms) but has a much more muted impact on the weighted sex ratio since the inflow gives a higher weight only to the source-province birth sex ratio.

Fourth, these graphs show a considerable gap between the projected sex ratios using the two censuses. Although we expect a small decline in cohort sex ratios from male mortality, the gap is often outside the expected range.[7]

We will return to the issue of the discrepancy between the 1990 and the 2000 sex ratios, but for now we conclude that the resident sex ratio using the 1990 census appears unsuitable, especially for the four richest provinces. Among the sex ratios using the 2000 census, we favor the weighted one, although in practic, the birth sex ratio is a close substitute. The resident sex ratio appears to be the least suitable.

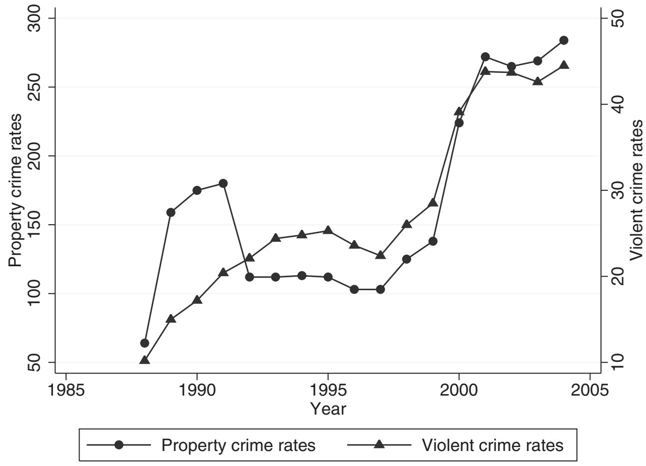
## IV. Data

Our province-year panel data set covers the years 1988 to 2004 and 30 province-level administrative divisions, assembled from statistic yearbooks and the 1% microsamples of the 1990 and 2000 Chinese censuses.[8] The crime statistics are from the *China Law Yearbooks* (Supreme People's Court, 1989–2005) and the *Procuratorial Yearbooks of China* (Supreme People's Procuratorate, 1989–2005), which provide annual aggregate statistics for all provinces (but unfortunately no age breakdowns). We define the crime rate as arrests for violent and property crimes per 10,000 inhabitans.[9] Thus measured, criminality almost doubled in the study period: from 3.71 in 1988 to 6.77 in 2004 (figure 1); and there was considerable variation, ranging from 0.81 (Tibet, 1988) to 13.1 (Zhejiang, 2004).

We focus on violent and property crimes because these crimes require little skill and the perpetrators are predominantly young men. We cannot distinguish between violent crimes and property crimes but they are highly correlated (figure 3) and may be similarly motivated (such as, robbery). Property crimes made up between 77.3% and 90.7% of all criminal cases between 1981 and 2004.[10] Among property crimes, larceny was by far the most common (86.7% in the same period) (Hu, 2006).[11]

From various yearbooks, *Comprehensive Statistical Data and Materials on 55 Years of New China* (National Bureau of Statistics, 2005) and the *China Statistical Yearbooks*, 1989–2005, we obtain information on a number of province-year variables that likely bear directly on the economic climate and thus crime: population size, per capita income, employment rate, income inequality (urban over rural household income), urbanization rate, welfare expenditures, openness to trade (exports, imports and foreign direct investments as a share of GDP), share of (out-of-province) immigrants, construction (square meters), and police expenditures (as a share of provincial government expenditures).[12]

FIGURE 3.—PROPERTY CRIME AND VIOLENT CRIME RATES BY YEAR




The number of cases registered by the police per 10,000 population.
*Source:* Chinese Supreme People's Court, *Law Yearbook of China* (Beijing: Publishing House of Law, 1985–2005).

Summary statistics for the variables used in the analysis of sex ratios and crime are in table 1. Data sources are detailed in table 2.

## V. Analysis

The posited link between sex ratios and crime turns on sex ratios affecting the marriage market and we start by showing that marital status varies with sex ratios in the hypothesized direction. We use county-level tabulations based on the full sample of the 2000 census.[13] This data set contains marriage rates and divorce rates for 15- to 45-year-olds. There are 2,871 counties, a substantial gain in sample size over province-level data. However, since mobility across counties is much higher than between provinces, the problem of internal migration is aggravated and therefore a causal interpretation is not appropriate (see Edlund, 2005).

We estimate by OLS a regression model of the form

$$y_j = \beta r_j^r + X_j + \rho + \varepsilon_j, \tag{5}$$

where $y_j$ is the outcome of interest (for example, the fraction of men never married) in county $j$. The sex ratio, $r^r$, is the number of men to women, 15 to 45 years old, enumerated in the county. $X_j$ is a vector of county characteristics: county-type indicator variables (urban district, county-level city, minority autonomous county, or county), urbanization

[7] Take year 2000, the cohort ages 16 to 25 was born between 1975 and 1984, and was 6 to 15 years old in 1990. Absent sex discrimination, sex ratios reach parity in the early 20s, a drop of 5 percentage points. For the cohorts at hand, we would expect a smaller drop because because of both male bias and the fact that the cohort under 5 years old, the ages of the highest mortality, contributes little to the cohort sex ratio.

[8] The is the first year for which crime statistics at this level of disaggregation are available is 1988. Of the thirtyone province-level divisions, only Chongqing is excluded because it was part of Sichuan until 1998.

[9] We focus on arrests rates for want of province-level offense rates. Conviction rates are very high in China.

[10] At the national level, publicly available data are available from 1981 onward.

[11] We do not have annual data on age-specific criminality or its breakdown by gender.

[12] An earlier version of the paper also included the secondary school enrollment rate; however, on closer inspection this variable seemed noisy. Its removal does not affect our results.

[13] "The Tabulation of the 2000 Population Census at the County Level" by the National Bureau of Statistics.

TABLE 1.—DESCRIPTIVE STATISTICS: PROVINCE-YEAR VARIABLES FOR TABLES 5–7

| Variable | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|
| Variable | | | | |
| Crime rate (arrests/10,000 population)[a] | 5.78 | 2.12 | 0.82 | 16.76 |
| Corruption rate (arrests/10,000 population) | 0.48 | 0.26 | 0.09 | 2.05 |
| Sex ratios (males/females) | | | | |
| R1015 (ages 10–15) | 1.06 | 0.05 | 0.94 | 1.23 |
| R1625 (ages 16–25) | 1.02 | 0.03 | 0.95 | 1.15 |
| R2645 (ages 26–45) | 1.03 | 0.02 | 0.97 | 1.11 |
| R4665 (ages 46–65) | 1.04 | 0.05 | 0.82 | 1.14 |
| Basic covariates | | | | |
| Population (10,000) | 3,928 | 2,574 | 212 | 9,717 |
| Income, per capital (RMB 1,000 at 2000 prices) | 3.14 | 1.83 | 0.99 | 14.16 |
| Employment (% employed, 16–65 year olds) | 69.08 | 9.87 | 46.82 | 97.96 |
| Inequality (urban/rural per capita income) | 2.63 | 0.72 | 1.24 | 5.16 |
| Urbanization (% living in urban areas) | 33.49 | 14.98 | 13.11 | 81.16 |
| Additional covariates | | | | |
| Welfare (% of government expenditures) | 2.36 | 0.80 | 0.66 | 11.41 |
| Economic openness: $\frac{\text{Export+Import+FDI}}{\text{GDP}}$ | 0.27 | 0.34 | 0.03 | 2.02 |
| Immigration rate (% born outside province)[b] | 2.86 | 2.46 | 0.31 | 16.34 |
| Construction (10,000 $m^2$) | 4,126 | 5,713 | 22 | 50,982 |
| Police (% of government expenditures) | 5.46 | 1.45 | 1.67 | 11.48 |

There are 510 observations: annual data from thirty province-level administrative divisions, 1988–2004. For data sources, see table 2.
[a] Violent and property crimes. A nonexhaustive list includes homicide, assault, robbery, rape, abduction of women and children, larceny, fraud, and smuggling. [b] Change to official household-registration status.

TABLE 2.—DATA SOURCES

| | |
|---|---|
| Tables 1, 4–7 | |
| Sex ratios | 2000 Census, 1-in-100 sample |
| Crime rate, corruption rate | *Law Yearbook of China, 1989–2005, Procuratorial Yearbook of China, 1989–2005.* |
| Per capital income, inequality, | *Comprehensive Statistical Data and Materials on 55 Years of* |
| urbanization rate | *New China; China Population Statistical Yearbook, 1989–2005.* |
| construction, employment (Export + Import + FDI)/GDP | |
| Welfare expenditures, police expenditures | *China Statistical Yearbook, 1989–2005* |
| Immigration rate | 1988–1991: *China Population Statistical Yearbook: 1992–2004: China Population Statistical Data and Material by Provinces and Cities* |
| Table 5 | |
| Sex ratios | 1990 and 2000 Censuses, 1-in-100 sample |
| All other variables | Urban Household Survey, 1988–2006 (annual). |
| Table 6 | |
| Sex ratios | 1990 Census, 1-in-100 sample |
| All other variables | Chinese Health and Nutrition Survey, 1989, 1991, 1993 |
| Table 7 | |
| All variables | *Tabulations of the 2000 Census at the County Level* |

rate, percent minorities, percent illiterate, fraction of households with potable tap-water, and percent immigrants (change to *hukuo* status). $\rho$ is a vector of province dummy variables. We expect higher sex ratios to raise the fraction of never married or currently divorced men and the reverse for women.

As expected, we find that the sex ratio correlates positively with both the fraction of men never married or currently divorced (table 3). For women, the pattern is reversed (although somewhat mixed for the fraction of women divorced).

## A. Sex Ratios and Crime

To investigate the hypothesis that the rising male surplus among 16- to 25-year-olds has materially contributed to the rise in crime in the past decades, we estimate a regression model of the following form:

$$\begin{aligned}\ln c_{jt} &= \beta_{1625} \ln \mathtt{R1625}_{jt} \\ &+ \beta_{1015} \ln \mathtt{R1015}_{jt} + \beta_{2645} \ln \mathtt{R2645}_{jt} + \beta_{4665} \ln \mathtt{R4665}_{jt} \\ &+ X_{jt} + \rho + \tau + t \times \rho + \varepsilon_{jt}, \quad (6)\end{aligned}$$

where $c_{jt}$ is the crime rate in province $j$, year $t$; $\mathtt{R1625}_{jt}$, $\mathtt{R1015}_{jt}$, $\mathtt{R2645}_{jt}$, and $\mathtt{R4665}_{jt}$ are the weighted sex ratios (see Equation (4)) for the 16–25, 10–15, 26–45, and 46–65 year-old cohorts, respectively. We hypothesize that $\beta_k > 0$ for $k = 1625$, and 0 otherwise.

A positive effect on crime could follow simply from there being more men. As discussed in section I (and shown in appendix A), an estimate of $\beta_{1625}$ greater than the share of females (roughly 0.5) supports the more interesting hypothesis that higher sex ratios make men more crime prone.

We include the sex ratios for ages 10 to 15, 26 to 45 and 46 to 65 as a check against spurious correlation. The 10–15 age group is too young to be criminally responsible. The older age groups are too old to be criminally active, at least going by conventional wisdom. Were we to find higher sex ratios among these age groups to raise crime as well, that would be worrisome. The falsification test is particularly strong for the 10–15 age group. For the older age groups, the case is less clear cut, the 26–45 age group not having aged out of crime entirely.

TABLE 3.—SEX RATIOS AND MARITAL STATUS.
County-level data from 2000 Census

| | Dependent Variable | | | |
|---|---|---|---|---|
| | Fraction Men | | Fraction Women | |
| | Never Married (1) | Divorced[a] (2) | Never Married (3) | Divorced (4) |
| Mean | 0.240 | 0.013 | 0.160 | 0.009 |
| Covariates | | UnWeighted | | |
| No | 0.157*** | 0.005 | −0.103** | 0.001 |
| | (0.037) | (0.003) | (0.049) | (0.004) |
| Yes | 0.140*** | 0.004 | −0.130*** | −0.003 |
| | (0.033) | (0.003) | (0.042) | (0.004) |
| | | Population Weighted | | |
| No | 0.135 | 0.008*** | −0.212 | 0.007** |
| | (0.110) | (0.003) | (0.188) | (0.003) |
| Yes | 0.153** | 0.004* | −0.174* | 0.001 |
| | (0.056) | (0.002) | (0.089) | (0.002) |

[a]Currently divorced. There are 2,871 observations (counties). Each entry is from a separate regression and shows the estimated coefficient on the (logged) sex ratio defined over the county population ages 15 to 45. All regressions include province fixed effects and county-type indicator variables (urban district, county-level city, minority autonomous county, or county). Additional covariates (included as indicated) are urbanization rate (%), minority (%), illiterate (%), fraction of households with potable tapwater (%), and immigrants (%). Weights are total population by county. Robust standard errors clustered at the province level in parentheses. Significant at *10%, **5%, ***1%.

TABLE 4.—SEX RATIOS AND CRIME

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Dependent variable: ln(Crime) | | | | | | |
| ln(R1625) | 1.829** | 2.293*** | −0.847 | −1.000 | −1.502 | 3.713** |
| | (0.823) | (0.755) | (1.103) | (0.998) | (0.996) | (1.812) |
| ln(R1015) | 1.030* | 1.567*** | −0.303 | −0.523 | −0.687 | 0.570 |
| | (0.526) | (0.452) | (0.827) | (0.830) | (0.766) | (0.585) |
| ln(R2645) | −4.090** | 1.984 | 0.495 | −0.571 | −0.644 | 6.551 |
| | (1.749) | (2.915) | (2.800) | (2.393) | (2.291) | (4.534) |
| ln(R4665) | 0.570 | 3.347** | 2.699** | 2.203** | 2.048** | 3.080 |
| | (1.053) | (1.609) | (1.204) | (0.987) | (0.891) | (2.080) |
| Observations | 510 | 510 | 510 | 510 | 510 | 510 |
| Adjusted $R^2$ | 0.284 | 0.284 | 0.516 | 0.544 | 0.565 | 0.732 |
| Population weighted | | | | | | |
| ln(R1625) | 2.807*** | 2.621*** | −0.966 | −1.199 | −1.672 | 3.363* |
| | (0.795) | (0.722) | (1.276) | (1.001) | (1.129) | (1.764) |
| ln(R1015) | 0.928* | 1.054** | −0.884 | −0.673 | −0.946 | 1.186 |
| | (0.474) | (0.476) | (0.750) | (0.672) | (0.608) | (0.839) |
| ln(R2645) | −4.591** | −0.336 | −3.513 | −2.514 | −2.720 | 6.346 |
| | (1.727) | (2.327) | (2.153) | (2.171) | (2.141) | (4.492) |
| ln(R4665) | 0.939 | 3.059*** | 2.844** | 1.468* | 0.616 | 4.265 |
| | (1.339) | (1.003) | (1.194) | (0.733) | (0.858) | (2.636) |
| Observations | 2,003,494 | 2,003,494 | 2,003,494 | 2,003,494 | 2,003,494 | 2,003,494 |
| Adjusted $R^2$ | 0.224 | 0.750 | 0.845 | 0.864 | 0.871 | 0.920 |
| Province FE | No | Yes | Yes | Yes | Yes | Yes |
| Year FE | No | No | Yes | Yes | Yes | Yes |
| Basic | No | No | No | Yes | Yes | Yes |
| Additional | No | No | No | No | Yes | Yes |
| Province time trend | No | No | No | No | No | Yes |

The sex ratio is the weighted sex ratio: see equation (4). Each regression has 510 observations (30 provinces, 17 years). Basic covariates are ln(population size), ln(per capita income), employment, inequality, and urbanization. Additional covariates are welfare expenditures, (export + import + FDI)/GDP, immigration, ln(construction $m^2$), and police expenditures. Weights are total population by province and year. Robust standard errors clustered at the province level in parentheses. Significant at *10%, **5%, ***1%.

$X_{jt}$ is a vector of controls. Our basic controls are population size, income, employment, inequality, and urbanization. Additional controls are welfare (percent of government expenditures), export and import and foreign direct investments (FDIs) as a share of income (GDP), construction (m$^2$), immigrants, and police expenditures (percent of government expenditures). $\rho$ and $\tau$ are province and year fixed effects, and $t \times \rho$ is a province-specific time trend.

We estimate equation (6) by OLS, clustering standard errors at the province level. Table 4 reports results using the weighted sex ratio (our favored sex ratio). We report both unweighted results (top panel) and population-weighted results (bottom panel). Column 1 shows the results from regressing crime on the sex ratios, no covariates. Column 2 adds province fixed effects, and the estimated elasticity of crime with respect to the 16–25 sex ratio is 2.3. Column 3 adds year fixed effects, and the coefficient turns negative. Column 4 adds basic covariates and column 5 additional covariates, rendering the coefficient on the 16–25 sex ratio negative and borderline significant. Finally, column 6 adds province-specific time trends, and the positive and significant relationship between the 16–25 sex ratio and crime

TABLE 5.—SEX RATIOS AND CRIME: ALTERNATIVE SEX RATIO MEASURES
Dependent Variable: ln(Crime); Independent Variable Log 16–25 Sex Ratio

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | | A. 2000 Census | | | |
| Birth Sex Ratio, ($n = 510$) | | | | | | |
| | 1.934** | 2.299*** | −0.680 | −0.925 | −1.365 | 2.965* |
| | (0.819) | (0.700) | (1.063) | (0.955) | (0.972) | (1.628) |
| | | | Population weighted | | | |
| | 2.767*** | 2.509*** | −0.934 | −1.299 | −1.653 | 2.462 |
| | (0.762) | (0.682) | (1.197) | (0.937) | (1.041) | (1.477) |
| Resident Sex Ratio, ($n = 510$) | | | | | | |
| | 0.593 | 1.149 | −1.620** | −1.640** | −1.587** | 0.938 |
| | (0.751) | (0.698) | (0.649) | (0.659) | (0.690) | (1.461) |
| Population weighted | | | | | | |
| | 0.164 | 0.417 | −2.675*** | −2.755*** | −2.668*** | −1.167 |
| | (0.870) | (0.722) | (0.484) | (0.522) | (0.490) | (0.962) |
| | | | B. 1990 Census, Resident Sex Ratio | | | |
| All Provinces ($n = 510$) | | | | | | |
| | 0.794 | 0.260 | 1.171 | 0.968 | 1.377* | 1.089 |
| | (1.056) | (1.489) | (0.778) | (0.652) | (0.762) | (0.931) |
| Population weighted | | | | | | |
| | 1.966* | 0.675 | 0.812 | 0.544 | 0.593 | 1.435 |
| | (1.157) | (1.698) | (0.605) | (0.605) | (0.597) | (1.096) |
| Excluding Beijing, Tianjin, Shanghai, and Guangdong ($n = 442$) | | | | | | |
| | 3.652*** | 6.311*** | 4.811*** | 4.542*** | 4.182*** | 2.041 |
| | (1.060) | (2.194) | (1.558) | (1.331) | (1.240) | (1.726) |
| | | | Population weighted | | | |
| | 3.520*** | 5.222*** | 3.627*** | 3.529*** | 3.367*** | 2.007 |
| | (1.074) | (1.108) | (0.980) | (0.989) | (0.993) | (1.478) |
| Province FE | No | Yes | Yes | Yes | Yes | Yes |
| Year FE | No | No | Yes | Yes | Yes | Yes |
| Basic | No | No | No | Yes | Yes | Yes |
| Additional | No | No | No | No | Yes | Yes |
| Province time trend | No | No | No | No | No | Yes |

The table shows the regression coefficient on the 16–25 sex ratio. Regressions in panel A include the 10–15, 26–45 and 46–65 sex ratios. Regressions in panel B include the 26–45 and 46–65 sex ratios. The 10–15 sex ratio is excluded because the 1990 census cannot be used to project this sex ratio beyond year 2000. Basic covariates are ln(population size), ln(per capita income), employment, inequality, and urbanization. Additional covariates are welfare expenditures, (export + import + FDI)/GDP, immigration, ln(construction $m^2$), and police expenditures. Weights are total population by province and year. Robust standard errors clustered at the province level in parentheses. Significant at *10%, **5%, ***1%.

reappears. Moreover, the 16–25 sex ratio is the only statistically significant one. In the unweighted regression, the estimated elasticity is 3.7, and in the population-weighted one, it is 3.36.

An argument for favoring the results from the specification including province-specific time trends is that they control for trends in unmeasured, province-specific factors. We will return to this issue below, but taking the elasticity estimate of 3.36 at face value (column 6), the rise in sex ratios can account for 16% of the overall rise in crime. (The 16–25 sex ratio rose by 4%, from 1.02 in 1988 to 1.06 in 2004, and crime rates rose by 82.4% over the same period, $0.16 = 3.36 \times 4/82.4$.) As for the question as to whether criminality rose because men became more crime prone (not just more men), the answer is a qualified yes. The fraction of females constitutes an upper bound on the effect higher sex ratios through merely more men. The fraction of females was about 0.485. The result from the unweighted regressions supports this interpretation, whereas the result from the weighted regressions makes for a weaker case.

To sum up, we find a positive and significant correlation between the 16–25 weighted sex ratio and crime. The positive relationship between the 16–25 weighted sex ratio and crime is evident in the bivariate correlation and robust to the inclusion of province fixed effects. However, once time-varying covariates are added, the sex ratio is estimated to raise crime only if province-specific time trends are also controlled for. We now turn to the role of the province-specific time trends.

### B. *Alternative Sex Ratios*

The results from using the alternative sex ratios may cast additional light on the forces at hand. (For brevity, 2000 and 1990 refer to the respective census unless otherwise noted. For instance, the 2000 resident sex ratio refers to resident sex ratios using the 2000 census.)

The results from alternative sex ratios for the 16- to 25-year-old group are reported in table 5. From the high degree of congruence between the weighed and the birth sex ratio, we expect the 2000 birth sex ratio to deliver similar results, and that is what we find. The use of the birth sex ratio assumes rather unrealistically that crime is committed in the province of birth, and in view of this, it is perhaps not surprising that the estimates are noisier than those obtained using the weighted sex ratio (the weighted sex ratio is easier to defend on conceptual grounds but is likely measured with error, especially for the later years). As expected, the 2000 resident sex ratio does quite poorly (table 5, panel A). As for the resident sex ratio using the 1990 census, the results are only suggestive of a positive effect of sex ratios, but the point estimate of the elasticity is only 1.4, with an even lower $t$-ratio when estimated on the whole sample (table 5, panel B). As noted, migration renders the resident sex ratio problematic (although less so for the 1990 census). Recall that the richest provinces exhibited very male sex ratios for the early years of our study period. This pattern likely reflects the age and gender pattern of migrants in 1990, not the marriage market prospects of young males. In view of this, we also present results when Beijing, Tianjin, Shanghai, and Guangdong are excluded (table 5, panel B). Restricting the sample brings the results for the 16–25 sex ratio more in line with our hypothesis. However, results are not robust to the inclusion of a province-specific time trend, in contrast with the results using the 2000 birth and weighted sex ratios.

The estimated coefficients on the other age groups, 10–15, 26–45, and 46–65, using alternative measures of sex ratios, are reported in table III.A–D in appendix B. With the 2000 birth sex ratio, most of the coefficients are imprecisely measured and statistically insignificant. Surprisingly, when we use the 2000 residential sex ratio, some of the estimates for these age groups are negative and statistically significant. We believe this result to be driven by confounders such as gender-biased temporary migration and unmeasured economic conditions. When we use the 1990 residential sex ratio, the results are in the opposite direction with respect to those using the 2000 residential sex ratios. This difference can partly be attributable to sex ratios being projected. For example, the cohorts used to calculate the sex ratio for the 46–65 age group in year 1988 were 58–77 years old in 2000. These results support the case for using sex ratios based on province-of-birth information available in the 2000 census.

Although we will not be able to bring conclusive evidence to bear on this issue, the different roles of the province-specific time trend suggest the presence of a confounder that affects crime and is not reflected in the sex ratios calculated using the 2000 census information on province of birth but is reflected in the sex ratio using the 1990 census information on province of residence. To investigate this hypothesis further, we examine the difference between the 1990 census (resident) sex ratio and the 2000 census birth sex ratio. We choose the birth sex ratio since it easier to interpret and has a high level of congruence with the weighted sex ratio. This difference, by province, is graphed in appendix D.

We would expect the 1990 resident sex ratio of a cohort to reflect the cohort's sex ratio at birth and any migration (up to 1990). The 2000 birth sex ratio in turn would reflect the cohort's sex ratio at birth and gender differential mortality. Because of the ages we study (briefly elaborated on below), mortality alone suggests a positive difference that increases with the study year, in the ranged 1 to 3 percentage points.

Normally males suffer higher mortality than females at every age. Thus, absent sex discrimination, cohort sex ratios decline from about 1.05 at birth to 1 by age 20. Since mortality before midlife is concentrated in the infant and juvenile period, a 5% gap constitutes an upper bound on the gap we expect measuring sex ratios ten years apart. This age profile to mortality also means that we expect the gap to be greater for the cohorts born later. They were younger in 1990 and therefore would have suffered higher mortality between 1990 and 2000. For instance, the 15–24 sex ratio for the year 2004 is calculated using the 2–11 age group in the 1990 census and the 12–21 age group in the 2000 census, whereas for 1994, the corresponding ages are 12 to 21 and 22 to 31, respectively. Add to this picture biased preferences for sons and we would expect the decline in sex ratios as the cohort ages to be smaller than the 5% given by naturally occurring mortality. Thus, if the 1990 sex ratio did not reflect migration, we would expect the difference to be in the range of 1 to 3 percentage points, toward the top of the ranger for more recent years. For most provinces, that is not what we see.

Instead, a striking pattern with respect to income appears. As mentioned, the four richest provinces (Beijing, Tianjin, Shanghai, and Guangdong) are outliers. Among the next four richest provinces, there is an upward trend, as would be expected from mortality alone, but the magnitude of the gap is often well outside the expected 1–3 percentage point band. By contrast, in the four poorest provinces, the trend is negative. Moreover, for the more recent years, the difference is negative. In other words, the number of girls living in poor provinces (in 1990) was higher than the number of girls born in these provinces (as measured by the 2000 census information on province of birth). Gender differential mortality cannot account for this pattern. Instead, the pattern is consistent with the 1990 census reflecting economically motivated migration. If so, the economic climate is not fully captured by our controls and suggests the need to include province-specific time trends when using the sex ratio based on the 2000 census information on province of birth.

Recall that for the early years, 1988 to 1990, sex ratios are projected using 16- to 27-year-olds in 1990. Among this age group we would expect marriage migration of women from poor provinces (see Fan & Huang, 1998). This could account for the positive gap at the beginning of the study period among poor provinces (the resident sex ratio in 1990 being above the sex ratio at birth) and the negative, or close to negative, gap for many rich provinces.

The pattern then reverses, with positive gaps in rich provinces and negative gaps in poor provinces. Preferential treatment of sons might be driving this pattern.

For the years toward the end of our study period, say 2000 and onward, the 16–25 cohort was 2–15 years old in 1990 and therefore unlikely to have migrated by themselves. Instead, preferential treatment of sons may account for more young males residing in richer provinces (than born there). First, if immigrants were more likely to bring their sons (than their daughters) and immigration was predominantly to richer provinces, we would expect more boys in the 1990 census than in the 2000 census, which could account for a positive gap (outside the range of 1% to 3% expected from mortality) in rich provinces. Girls being left behind could similarly account for relatively more girls residing in poor provinces than born there, thus accounting for the negative gap seen in the four poorest provinces toward the end of our study period. Preference for sons might be a reason for such gender differential treatment of sons and daughters with parents being more attached to their sons.

Second, if parents were more likely to send away a daughter in provinces that had a stricter one-child policy and these provinces were also richer, this could also account for the pattern we see. Generally the one-child policy was stricter in richer provinces, being more urbanized and having smaller minority populations. An exception is Sichuan, which ranks toward the bottom in terms of per capita income (twenty-fifth at of 30) but had a strict one-child policy (possibly for idiosyncratic political reasons; Sichuan was Deng Xiaoping's birthplace).

For the period 1995 to 2000, the 16–25 sex ratio is generated by the 1970–1984 birth cohorts, ages 6 to 20 in 1990. For these years, a mix of the above two mechanism might have been at play.

The discussion has focused on migration accounting for the difference between the 2000 census birth sex ratio and the 1990 resident sex ratio. Another possibility is that the province of birth is misreported. However, misreporting would have had to be of boys being underreported as born in rich provinces (or girls overreported) and the reverse pattern to hold in poor provinces. As far as we are aware, there is no basis for such a pattern to reporting bias. On the contrary, theory suggests that those residing in rich provinces will have less trouble marrying their sons than their poorer compatriots and therefore can indulge their son preference (Edlund, 1999). Anecdotal evidence also suggests such a direction: daughters born to parents in richer areas are sent to relatives in poorer provinces so as to make room for a son.

In sum, the differences between the 1990 resident and the 2000 birth sex ratios appear systematic and consistent with a scenario where the 2000 birth sex ratio is closer to the true sex ratio at birth and the 1990 resident sex ratio is a composite of the sex ratio at birth and the province's relative economic standing in 1990. This observation begs the question why the 1990 and the 2000 resident sex ratios do not return more similar result. A possible explanation lies in the strong age profile to migration that appeared after the relaxation of the *hukou* system in the early 1990s and the fact that we are relying on projections from the censuses.

### C. *Falsification Tests and Specification Checks*

Appendix B, table IV presents further falsification and robustness checks. One concern is that violent and property crime statistics may reflect various province-specific "fight crime" drives rather than underlying criminal activities. To address this concern, we replace violent and property crime with corruption as the dependent variable. The idea is that a general law-and-order zeal would affect corruption statistics as well, whereas there is little reason for the sex ratio of young adults to bear on corruption rates. Relative to violent and property crimes, corruption is high skilled and the perpetrators tend to be older and, tautologically, in a position of some influence. Age and economic standing would largely insulate perpetrators of white-collar crime from the demographics of younger adults. Consistent with this argument, we find no significant effect of the 16–25 sex ratio on corruption rates (column 1).

Next, we test model specification by including the sex ratio of each age group alone in a regression. Results are largely supportive of the hypothesis that it is the 16–25 sex ratio that matters most for crime, columns 2 to 5. We see that only the sex ratio of the 16 to 25 year age group is statistically significant (this result is weaker in the population weighted regressions).

Finally, we estimate specifications with the number of 16- to 25-year-old males and females in (logged) levels rather than as a ratio. Since the weighted sex ratio does not have an obvious analogy in levels, we use the number of men and women born in the province (the level version of the birth sex ratio). The results using only the 16–25 age group are in column 6, and those including all age groups are in column 7. The results are in line with expectations in terms of magnitude and sign but are only borderline significant. The elasticity estimate is in the 2–3 range for the number of men and similarly for the number of women (with the opposite sign).

## VI. Other Outcomes

The primary focus of this paper is on the role of male sex ratios for crime. However, crime is but one of a number of end points. We now turn to education, labor market, and intrahousehold bargaining. For education and labor market outcomes, we use the Urban Household Surveys. For intrahousehold bargaining, we use the China Health and Nutrition Surveys, years 1989, 1991, and 1993.

For the most part, the cohorts in question were born before the one-child policy and abnormal sex ratios at birth. However, in China, as elsewhere, men marry younger women. This age gap implies that cohort-size variations can have an impact on the marriage market. An event of particular significance is the Great Leap Forward famine, which sharply reduced fertility in the years around 1960 and resulted in a post-famine baby boom.

### A. *Education and Labor Market Outcomes*

For many of the same reasons crime might be affected, high sex ratios may have an impact on education and labor market outcomes. For education, we would expect the incentive effect to raise male education and lower that of women. For labor market outcomes, we would expect higher sex ratios to allow women to perform worse. Marriage is all but universal for women in China, but strengthened intrahousehold bargaining is a possible channel. For men, the prediction is less clear. The incentive effect would work toward better outcomes, while the civilizing effect would have the opposite effect (higher sex ratios imply fewer married men).

TABLE 6.—SEX RATIOS AND EDUCATION AND LABOR MARKET OUTCOMES, URBAN HOUSEHOLD SURVEYS, 1988–2006

| | Dependent Variable | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Education | | | | Labor Market | | | |
| | Years (1) | College and above (2) | High School and Above (3) | Middle School and above (4) | Employed (5) | Professional (6) | ln(Wage) (7) | ln(Income) (8) |
| Mean | 12.070 | 0.273 | 0.709 | 0.973 | 0.777 | 0.089 | 8.020 | 8.05 |
| Without province-specific time trend | | | | | | | | |
| ln(sex ratio) | −1.316 | −0.191 | −0.103 | −0.112** | −0.268* | −0.0279 | −0.168 | −0.502 |
| | (0.558) | (0.118) | (0.0521) | (0.0184) | (0.0752) | (0.0423) | (0.236) | (0.233) |
| Male×ln(sex ratio) | 1.290 | 0.220 | 0.118 | 0.0732*** | −0.0318 | −0.0176 | −0.289* | −0.203 |
| | (0.582) | (0.125) | (0.0482) | (0.00975) | (0.0457) | (0.0264) | (0.0817) | (0.0904) |
| With province-specific time trend | | | | | | | | |
| ln(sex ratio) | −1.567* | −0.236* | −0.148 | −0.0867** | −0.274** | −0.0200 | −0.307 | −0.692** |
| | (0.446) | (0.0908) | (0.0610) | (0.0175) | (0.0584) | (0.0241) | (0.150) | (0.152) |
| Male×ln(sex ratio) | 1.340 | 0.228 | 0.126* | 0.0709** | −0.0336 | −0.0169 | −0.272* | −0.177 |
| | (0.570) | (0.123) | (0.0466) | (0.0109) | (0.0453) | (0.0257) | (0.0839) | (0.0971) |
| *N* | 157,132 | 157,133 | 157,133 | 157,133 | 157,133 | 157,133 | 121,831 | 134,816 |

Annual microdata are from the six provinces: Beijing, Liaoning, Zhejiang, Sichuan, Guangdong, and Shaanxi. The sample is restricted to adults 18–45. All regressions include cohort, year and province fixed effects, and a gender dummy. The sex ratio is the resident sex ratio projected by the urban population in the census closest to the survey year (for 1988 to 1995, the 1990 census is used; and for 1996 to 2006, the 2000 census is used). The sex ratio is defined by province, age, and gender and assumes an age difference of two years between spouses. Table III in appendix B presents results using the analogous birth sex ratio (2000 census). Robust standard errors clustered at the province level in parentheses. Significant at *10%, **5%, ***1%.

The Urban Household Surveys have been conducted yearly since 1988 by the National Bureau of Statistics. Our data cover six provinces—Beijing, Liaoning, Zhejiang, Sichuan, Guangdong, and Shaanxi—and the period of 1988 to 2006 (nineteeth rounds of repeated cross-sections). We restrict the sample to respondents ages 18 to 45.[14]

We estimate by OLS a regression model of the form

$$y_{ijt} = \alpha r^r_{ijt} + \mathtt{male}_i + \beta \mathtt{male}_i \times r^r_{ijt} + X_i + \rho + \tau + \varepsilon_{ijt}, \quad (7)$$

where $y_{ijt}$ is the outcome of interest (educational attainment, labor market outcome) for individual $i$ in province $j$ and year $t$. For educational outcomes, we expect $\hat{\alpha} < 0, \hat{\beta} > 0$ and $\hat{\alpha} + \hat{\beta} > 0$—a higher sex ratio to lead to lower human capital investments by women, greater human capital investments by men, and a widening of the gender gap. For labor market outcomes, we expect $\hat{\alpha} < 0$ but are agnostic with respect to $\hat{\beta}$. The vector $X_i$ contains cohort, province, and year fixed effects and, if indicated, a province-specific trend. We control for birth cohort rather than age since there are likely strong cohort effects for those born between 1943 and 1988 (famine, cultural revolution, and so on).

Individuals in the Urban Household Surveys (UHS) have the coveted urban *hukou* status, and the vast majority have not migrated. Therefore, we believe the sex ratio most pertinent to these individuals is the local urban sex ratio, and we use the resident sex ratio projected from the 1990 and the 2000 censuses (employing the census nearest to the survey year and restricting the sample to individuals with urban registration status).

We define the sex ratio over cohort, province, and gender. For instance, for a 30-year-old man in Beijing in 2000, we use the number of men 28 to 32 years old of urban status in Beijing to the number of women 26 to 30 years old of urban status in Beijing in the 2000 census. Because relative cohort sizes play a role in the calculation of the sex ratio, there is more variation than when calculated within cohort. The mean for this sex ratio is 1.104, ranging from a low of 0.851 to a high of 1.899.

[14] For further description of these surveys, see, Zhang et al. (2005).

The results are presented in table 6. The first four columns show education outcomes: number of years of education and whether at least college, high school, or middle school, respectively. While more male sex ratios improve men's relative to women's educational status, this result is driven by the sex ratio reducing women's educational attainment.

As for labor market outcomes, the results are mixed (table 6, columns 5–8). As expected, higher sex ratios lower women's employment, professional employment, wage, and income. In view of this and the finding that relative to women, men had improved educational outcomes, it is somewhat surprising that the interaction effects between the male dummy and the sex ratio are negative. These findings are, however, consistent with the observation that the civilizing effect (of marriage) suggests that higher sex ratios would result in men doing worse in the labor market.

We also calculate the corresponding birth sex ratio using the 2000 census. This sex ratio had lower variance, consistent with its giving no role to migration (barring systematic misreporting). It averaged 1.069 and ranged from 0.939 to 1.61. The regression results are similar, albeit with a lower point estimate possibly stemming from the lower variance of the sex ratio variable (see appendix B, table V). That the birth and resident sex ratios would yield similar results is perhaps not surprising given the rarity of rural to urban *hukou* conversions.

## B. *Household Bargaining*

To investigate the effect of sex ratios on intrahousehold bargaining, we use the China Health and Nutrition Surveys

TABLE 7.—SEX RATIOS AND HOUSEHOLD BARGAINING, CHINESE HEALTH, AND NUTRITION SURVEYS: 1989, 1991, 1993

| | Dependent Variable | | | | | |
|---|---|---|---|---|---|---|
| | Household Chores (hours/week) | | | | Major Purchases[a] | |
| | Food Preparation (1) | Wash Clothes (2) | Child Care (3) | Total (4) | Participate (5) | Decide (6) |
| Mean | | | | | | |
| Women | 10.19 | 3.92 | 10.25 | 24.81 | 0.73 | 0.09 |
| Men | 2.63 | 0.48 | 4.23 | 7.89 | 0.91 | 0.31 |
| Independent variable | | | | | | |
| Male×ln(Sex ratio) | 8.099*** | 2.269* | 11.98** | 23.73*** | −0.158** | −0.214** |
| | (2.293) | (1.123) | (4.758) | (4.274) | (0.062) | (0.083) |
| $N$ | 2,778 | 2,765 | 2,398 | 2,269 | 4,483 | 4,483 |

[a]Electric fan, TV, and radio are the three major appliances. Each entry is from a separate regression and shows the estimated coefficient on the interaction term between (logged) sex ratios and male. Full regression results are in appendix B, table VI. The following nine provinces were surveyed in the CHNS: Guangxi, Guizhou, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong. The sample is restricted to married individuals 18 to 45 years old. All regressions control for sex ratio, male, ln(population), age, education years, rural, province, and year fixed effects. The sex ratio is calculated using the 1990 census, and is defined over province, age, gender, and rural/urban status, and assumes a spousal age difference of two years and a five-year window. For example, a 25-year-old man's sex ratio would be calculated using the number of men ages 23 to 27 to the number of women ages 21 to 25. Robust standard errors clustered at the province level in parentheses. Significant at *10%, **5%, ***1%.

(CHNS) carried out in nine provinces: Guangxi, Guizhou, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong.[15] We utilize the information in the first three waves (1989, 1991, and 1993) on household chores and decision making.

As for the sex ratio, we favor the 1990 census. Temporally, this census is the closest. Moreover, for the provinces at hand, migration was limited. Thus, the resident sex ratio using the 1990 census is possibly the most accurate measure of marriage market conditions.

We restrict the sample to married individuals, ages 18 to 45. The sex ratio is defined over province, age, gender, and rural/urban status (urban households were oversampled and make up more than half of the sample). As above, we assume a spousal age gap of two years and a five-year age window (a 25-year-old man's sex ratio is calculated as men ages 23 to 27 over women ages 21 to 25). This sex ratio averaged 1.07, with a standard deviation of 0.20, ranging from a minimum of 0.71 to a maximum of 1.73.

We examine two sets of outcomes pertaining to bargaining power: time spent on household chores, cooking, laundry, and child care broken out and who decides major household-appliance purchases (electric fan, TV, radio).

We estimate by OLS a regression model of the form

$$y_{ij} = r^r_{ij} + \mathtt{male}_i + \beta\mathtt{male}_i \times r^r_{ij} + X_{ij} + \rho + \tau + \varepsilon_{ij}, \tag{8}$$

where $y_{ij}$ is the outcome of interest for individual $i$ in province $j$, $r^r_{ij}$ is the sex ratio and $X_{ij}$ is a vector of controls (the individual's education in years, the province population size, whether rural, as well as province and cohort dummies). We hypothesize that higher sex ratios lead to men spending more time doing household chores and lower their decision making.

Results are in table 7 (full results are in table VI, appendix B). As expected, higher sex ratios raise the amount of time men spend on household chores and reduce their decision making. Our estimates suggest that a 10% (half a standard deviation) increase in the sex ratio reduces the gender gap in time spent cooking by 0.8 hours per week; washing clothes by 0.2 hours per week; and child care by 2 hours. Overall, the gender gap in household chores is reduced 2.3 hours, or 13%. Similarly, the gap in participation in deciding whether to purchase a major household appliance is reduced by 1.6 percentage points, or 9%. As for the final decision, a 10% increase in the sex ratio is estimated to reduce the gender gap by 2.1 percentage points, or some 9%.

These results are in the expected direction and mesh with those of Porter (2007), who, also using the CHNS, found higher sex ratios (computed using the 1982 census and same year fertility survey) to lead to fewer but healthier and better-educated children, findings she interpreted as evidence of greater bargaining power of the wife. The results are very similar when we use the 2000 birth sex ratio (table VII, appendix B).

## VII. Summary and Discussion

In 2000, almost 120 boys were born for every 100 girls in China, and the latest census revealed a similarly high number ten years on. Abnormally male sex ratios raise a number of concerns, ranging from the human rights issues surrounding the fate of the "missing girls" to the social impact of surplus men. Are these men going to be able to marry? What will happen if they do not? High sex ratios are not unique to China, but, unlike India, where population growth has buffered some of the impact, or South Korea, where the level of economic development allows the import of brides, the shortage of brides is likely to be felt acutely in China.

The rise in the sex ratios has coincided with a dramatic increase in crime. Although the notion that male sex ratios may raise crime is longstanding, a causal link has been difficult to establish. This paper exploits the rise in sex ratios at birth following the introduction of the one-child policy and the resulting rise in the sex ratio of young adults starting in the mid-1990s.

[15] The survey was conducted by the Carolina Population Center, the Institute of Nutrition and Food Hygiene, and the Chinese Academy of Preventive Medicine.

Results are sensitive to the sex ratio used and the inclusion of province-specific time trends. We favor a sex ratio based on the province of birth—information available in the 2000 but not the 1990 census. Our rationale is as follows: A sex ratio based on province of residence reflects migration patterns. Richer provinces not only attract more migrants, but there is a distinct gender pattern. Young women are more likely to migrate. Moreover, most migration since the early 1990s is unofficial (entailing no change to *hukou* status) and therefore temporary. If migrants plan to return to their province of birth to marry, then the resident sex ratio is misleading.

Taking the results from our preferred specification at face value, we estimate an elasticity of crime with respect to the 16–25 sex ratio of 3.4, implying that higher sex ratios may account for up to one-seventh of the overall rise in violent and property crime during our study period (1988–2004). As a point of reference, the next two decades will likely see another 10% increase in the sex ratio of young adults.

The positive relationship between the 16–25 weighted sex ratio and crime is evident in the bivariate correlation and robust to the inclusion of province fixed effects. However, once time-varying covariates are added, the sex ratio is estimated to raise crime only if province-specific time trends are also controlled for. The inclusion of province-specific time trends is a common strategy employed to guard against unmeasured confounders and one that is appropriate in the context at hand, we would argue.

The social consequences of male sex ratios extend beyond crime. Still, crime is an end point of particular interest. It is a plausible outcome. The low barriers to entry and the possibility of some upside may give it particular appeal to low-skilled men looking for a break, the men likely to be at the receiving end of the looming wife shortage foreshadowed by years of abnormal sex ratios. Crime also serves as a reminder that the social cost of male sex ratios is not limited to the men who will not be able to marry or the daughters who have gone "missing."

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## APPENDIX A

### More Men, More Crime

A higher sex ratio implies more men and more unmarried men. In this section, we decompose the elasticity of crime and show that the fraction females $f \in [0, 0.5]$ is an upper bound on the elasticity of crime with respect to the sex ratio if the only source of the increase in crime is that there are more men.

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.

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 that

$$c_m^0 = c_m^1 (= c_m). \tag{A1}$$

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

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

Let $r$ denote the sex ratio:

$$r = \frac{m}{f} = \frac{m}{1 - m}. \tag{A2}$$

We can write the elasticity of the crime rate with respect to the sex ratio, conditional on equation (9), as

$$\begin{aligned} \epsilon(c, r)|_{c_m^0 = c_m^1 = c_m} &= \left(\frac{dc}{dm} \times \frac{m}{c}\right) \cdot \left(\frac{dm}{dr} \times \frac{r}{m}\right) \\ &= \frac{c_m m - c_f m}{c_m m + c_f (1 - m)} \cdot f \\ &\leq f. \end{aligned} \tag{A3}$$

Thus, 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 alone.

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