10-3-2009

Preliminary draft

The Sexual Foundations of Economic Growth: Evidence from China

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Abstract

A severe sex ratio imbalance is common in many Asian economies. The existing literature on its consequences has focused mostly on social ills such as crimes. In this paper, we examine the possibility that the imbalance may also stimulate economic growth by inducing more entrepreneurial activities and hard work. We provide evidence from China. First, new domestic private firms – an important engine of growth - are more likely to emerge from regions with a higher sex ratio imbalance (holding constant other determinants of firm growth). An increase in the sex ratio by one standard deviation accounts for about half of the growth in the number of private firms. Second, households with a son in regions with a more skewed sex ratio demonstrate a greater willingness to accept relatively dangerous or unpleasant jobs and supply more work days. In contrast, the labor supply pattern by households with a daughter is unrelated to the sex ratio. Third, the growth rates of per capita GDP across the provinces are systematically related to the local sex ratio. Since the imbalance for the pre-marital age cohort will become worse over the next two decades, this growth effect is likely to persist.

Key words: excess men, missing women, entrepreneurship, growth
JEL codes:

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The authors would like to thank Patrick Bolton, Qingyuan Du, Lena Edlund, Amit Khandelwal, Rick Mishkin, and seminar/conference participants at the NBER Beijing conference, National University of Singapore, and Carleton University for helpful discussions, and Kyle Gerry and Joy Glazener. All errors are the authors' responsibilities.

1. Introduction

A sex ratio imbalance at birth and in the marriageable age cohort – too many men relative to women – is a common demographic feature in many Asian economies, such as China, Korea, India, Vietnam, Singapore, Taiwan, and Hong Kong. In many such economies, parents voluntarily want to limit the number of children they wish to have. This, together with a strong preference for sons, and the availability of inexpensive technology to screen the gender of a fetus (most commonly by Ultrasound B) to abort the unwanted pregnancy, leads parents to engage in sex selective abortions in favor of sons. The imbalance is particularly acute in China where a strict family planning policy has restricted the number of children most families can have to one or two. The imbalance is still increasing in some economies. For example, the United Nations recently warned Vietnam to watch out for a sharp rise in the sex ratio imbalance[1]. In China, the sex ratio at birth in 1980, when the strict family planning policy was first introduced, was 1.07 boys per girl, which was moderately higher than the natural rate (about 1.06). The ratio deteriorated steadily to 1.12 in 1990, 1.18 in 2000, and 1.22 boys per girl in 2007 (Li, 2007; Zhu, Lu, and Hesketh, 2009).

The existing literature has identified several negative consequences of a serious sex ratio imbalance. First, a rising imbalance translates into an increasing pool of men who have no realistic hope to get married. The scale of the problem is frightening. For example, the number of excess Chinese men under age 20 exceeded 32 million in 2005 (Zhu, Lu, and Hesketh, 2009). This number is greater than the entire male population of Italy or Canada. Second, the imbalance may cause crimes. Using data across Chinese provinces, Edlund, Li, Yi, and Zhang (2007) estimate that every one basis point increase in the sex ratio (e.g., from 1.10 to 1.11 boys per girl) raises violent and property crime rates by 3%, and the rise in the sex ratio imbalance may account for up to one-seventh of the overall rise in crimes in China. Den Boer and Hudson (2004) boldly hypothesize that the sex ratio imbalance should generate security concerns for other countries since the one with a high sex ratio "might actively desire to send its surplus young males to give their lives in a national cause," although they provide no rigorous data analysis to back up their theory. Third, the imbalance may also trigger competitive savings among households – men and households with sons forego current consumption to accumulate wealth in

[1] "UN warns Vietnam of sex ratio imbalance," 11/9/2009, http://en.vietnamplus.vn/Home.

order to improve a young man's standing in the marriage market relative to other men. This increase in the savings rate is inefficient since it does not alter the number of unmarried men in the aggregate. Wei and Zhang (2009) estimate that about half the increase in the household savings rate in China during 1990-2005 can be attributed to the rise in the sex ratio. By raising the aggregate savings, the sex ratio imbalance contributes to China's current account surplus, and the global current account imbalances. To the extent that the global current account imbalances were a significant factor in the asset price bubbles during 2002-2007 (according to Greenspan, 2009, among others), the sex ratio imbalance may be a contributing factor to the onset of the 2007-2009 global financial crisis.

In this paper, we study a possibly positive effect of the sex ratio imbalance on economic growth. If the family wealth of a man relative to those of other men is a sorting variable for a man's relative standing in the marriage market, then a rise in the sex ratio would inspire men and parents with a son to find ways to accumulate more wealth. Working harder and longer, and becoming more entrepreneurial are ways to achieve this objective. As a result, the economy may grow faster than it would have otherwise. As far as we know, this effect has never been investigated.

We choose to conduct the empirical analysis using data from censuses of firms and household surveys from China. Several reasons make the country a particularly good candidate for this research topic. First, China presents one of the fastest increases in the sex ratio in the world due to its draconian family planning policy. As a result, there is a better chance to detect this growth effect if one exists. Second, a within-country study has advantages over cross-country studies as the legal system and other institutions can be more plausibly held constant across regions within a country than across countries. As a very large country, there are many sub-national geographic units in China that allow us to have sufficient statistical power when exploring regional variations. Third, while the Chinese economy is about half the size of the United States on a PPP-adjusted basis, the contribution of the Chinese growth to the incremental world GDP has been the largest in the world since 2002 (IMF 2009). Therefore, understanding the determinants of Chinese growth has intrinsic value for international macroeconomics due to its direct global implications.

The empirical results not only support the hypothesis qualitatively, but are significant quantitatively. First, based on the data from two censuses of firms in 1995 and 2004 (the two most recent censuses), we estimate that an increase in the sex ratio by one standard deviation can account for about half of the extensive margin of the private sector growth (i.e., the birth of new private firms) across regions. To address concerns about possible biases due to missing regressors, endogeneity, and measurement errors associated with the sex ratio measure, we employ both an instrumental variable approach and a placebo test, and use household data to check distinct implications of the hypothesis for families with a son versus those with a daughter. Second, from a survey of rural households, we estimate that households with a son respond to a rise in the sex ratio by a combination of working more days off farms (including as migrant workers) and becoming more willing to accept unpleasant or relatively dangerous jobs. In contrast (but consistent with our hypothesis), the labor supply pattern of daughter-households is not linked to the local sex ratio. Third, to capture the general equilibrium effect, we also directly check whether the growth of per capita GDP across regions is linked to the local sex ratio imbalance and find that the answer is affirmative. We estimate that about 15% of the growth rate of GDP per capita, or slightly over one percentage point per annum, can be attributed to the rise in the sex ratio. Since the sex ratio imbalance is projected to become worse in the next decade, this effect may become relatively more important over time.

The rest of the paper is organized in the following way. In Section 2, we develop the argument more fully and make connections to related literatures. In Section 3, we provide statistical evidence for our hypothesis. Finally, in Section 4, we conclude and discuss possible future research.

2. Developing the Argument and Connecting to the Literature

In this section, we first explain why a sex ratio imbalance could trigger a race to accumulate wealth, stimulating more hard work and more entrepreneurial activities. We then connect this discussion to related literature.

From an unbalanced sex ratio to a desire to accumulate wealth

Most men have a desire to get married. It may be safer to say that, in an Asian society, parents of a son strongly prefer for their son to be married[2]. As the competition for a dwindling pool of potential wives intensifies, the parents of a son (or a son himself) will try harder to improve the son's prospects for marriage. Accumulating more wealth may be regarded as a way to do it, and this idea is not unique to China. For example, pop star Madonna has declared in the first two verses of the hit song, "Material Girl," that: "Some boys kiss me, some boys hug me, I think they're o.k./ *If they don't give proper credit, I just walk away/*They can beg and they can plead/But they can't see the light, that's right/*'Cause the boy with the cold hard cash/Is always mister right*." So she sees a connection between a man's ability to generate wealth and his success in dating. In Japan, ex-Prime Minister Taro Aso said, in a meeting with college students just before the 2009 election, "if a youngster doesn't have money, he should not get married. It seems to me that a man without means would have a hard time gaining respect from a potential marriage partner."[3] To be clear, our hypothesis does not require women to be material girls. Indeed, other factors can be important for success in finding a spouse. But other things being equal, as long as more wealth improves a man's likelihood for marriage, parents with a son (or the son himself) will be encouraged to find ways to generate wealth.

For the sex ratio to affect household labor supply or entrepreneurial propensity, people don't have to know local sex ratio statistics. There is an invisible hand at work. Consider two otherwise identical households with a son, one in a region with a high sex ratio, and the other in a region with a low sex ratio. Parents in the first region would observe or be told by relatives or colleagues with a son that it would cost at least a certain amount for their sons to find a girlfriend and to marry. Which level of wealth is perceived to be necessary for marriage success would differ between the two regions. In other words, even without the knowledge of local sex ratio statistics, parents with a son may make decisions on labor supply and occupational choices that reflect the local sex ratio.

Existing Literature

Several theoretical papers have pointed out a connection between concerns for status

[2] In Darwin's sexual selection theory, reproduction is the utmost objective for species after survival. Many individual and societal behavior, in his view, could be explained by a desire to reproduce.
[3] "Aso advises poor young people not to marry," Japan Today, August 24, 2009.

(one's relative position in a society), the savings rate, and the economic growth rate (Cole, Mailath and Postlewaite, 1992; Cornero and Jeanne, 1999; and Hopkins and Kornienko, 2009). When wealth defines one's status in the marriage market, a greater concern for status may lead to an increase in the growth rate. All three papers use competition in the marriage market to illustrate the idea. However, the sex ratio is always assumed to be balanced in the above papers. In other words, no explicit comparative statistics are derived in terms of a rise in sex ratio imbalance. Sex ratio imbalance could potentially introduce non-trivial complications. In particular, while a rise in the ratio of boys to girls is an unfavorable shock to men or families with a son, it is a favorable shock to women or families with a daughter. In abstract, the behavioral response by these two groups of people could offset each other, making the effects on aggregate savings or other aggregate variables of interest less clear cut.

Du and Wei (2009a) develop a model that explicitly considers the aggregate effects of a rise in the sex ratio imbalance – a widening gap between the number of men and the number of women in the marriage cohort. They examine an overlapping generation model in which every one lives two periods. An exogenous parameter, φ, represents the ratio of boys and girls at birth, and $\varphi \geq 1$. It is assumed that everyone wants to be married with a member of the opposite sex in the same generation. Because there are more men than women, all women get married in equilibrium, and φ-1 fraction of the men do not get married. Life-time bachelors solve a standard two-period intertemporal optimization problem. Parents maximize their life-time utility which includes a concern for the utility of their offspring, and make a financial transfer to their child which is taken into account in the parents' first-period savings decision.

Du and Wei (2009a) show that families with a son both supply more labor and save more than families with a daughter, which in turn supply more labor and save more than life-time bachelors. As the sex ratio increases, son-families raise their labor supply and savings, while life-time bachelors do not change their labor supply and savings rate. The effect of a rise in the sex ratio on daughter-families is ambiguous. However, if the emotional utility of having a daughter and that of having a son are the same for parents, then the daughter-families also raise their savings and labor supply. In general, if the emotional utility of having a daughter is not too far from that of having a son, the daughter-families would raise their labor supply and savings in response to a sex ratio imbalance. If the sex ratio is not too high (below 1.5 – no country in the data has a sex ratio at birth that exceeds 1.4), the economy-wide labor supply and savings rate both increase in response to a rise in the sex ratio.

In terms of the empirical literature, we have found no paper that links the sex ratio imbalance with economic growth. At the same time, it is common in the empirical growth papers to consider other demographic variables, in particular, the age structure of the population. We will include such variables as controls in our analysis. There is an extensive literature in demography that documents the phenomenon of unbalanced sex ratios in Asia (for example, Gu and Roy, 1995; Guilmoto, 2007; and Li, 2007). Several papers have examined the determinants of sex ratio imbalance (including Das Gupta, 2005; Edlund, 2001; and Ebenstein, 2008). In an influential paper, Oster (2005) proposes that the prevalence of Hepatitis B is a significant cause of the sex ratio imbalance in Asia. But this conclusion is later shown to be incorrect, including by Lin and Luoh (2008) and Oster, Chen, Yu and Lin (2008). In a paper with a clever instrumental variable approach, Qian (2008) shows that an improvement in the economic status of women tends to reduce the sex ratio imbalance. Her instrument for the economic status of women is the world price of tea, whose production is apparently particularly suitable for women laborers. Wei and Zhang (2009) document that a higher sex ratio induces more savings, which is consistent with our story. However, the paper does not examine how labor supply and entrepreneurship may respond to a change in the sex ratio, which is the central focus of the current paper. The introduction of this paper has reviewed other papers that study the consequences of a rise in the sex ratio, which we do not repeat here.

This discussion has clear implications for the empirical work. First, it is useful to check whether and how households with a son and those with a daughter respond differently to a rise in the sex ratio. Second, it is useful to check the general equilibrium effect – whether the economy wide work effort and entrepreneurial activities increase, as reflected in a higher overall growth rate, in response to a rise in the sex ratio. Third, it is important to take into account both the population growth and the age structure of the population in our analysis.

3. Statistical Evidence

We start by providing some basic facts about the Chinese growth, which are summarized by two 70% rules. We then use data from the two most recent censuses of manufacturing firms (in 1995 and 2004) to investigate whether local sex ratio imbalance is a predictor of the extent of local entrepreneurial activities. To zoom in on possibly distinct responses by families with a son versus those with a daughter, we turn to household-level evidence. Finally, to capture the general equilibrium effect of a rise in the sex ratio, we conduct a panel growth regressions across Chinese provinces over 1980-2005.

Background information: the two 70% rules about the Chinese growth

Since our first piece of evidence has to do with regional variations in entrepreneurial activities, we work with the two most recent censuses of firms in 1995 and 2004, respectively, so we can compute the growth in the number of firms by region. During this period, the country's industrial value added (at the current price) grew by 266%.

The growth of the private sector is a major part of the overall growth story. The private sector is not just restricted to firms that were legally registered private firms. In fact, very few firms were registered as private firms in the 1980s. According to Huang (2009), many private entrepreneurs at least in the 1980s and 1990s found it necessary to set up firms as nominally owned by local governments (in the form of "township-and-village enterprises," or "collectively owned firms"). The goal was presumably to buy "protection" from the local government and to minimize the risk of state expropriation. Such a practice was widespread and was called "private entrepreneurs wearing a red hat." Most entrepreneurs later engineered or attempted to engineer a change in the firm ownership through which they could become a majority shareholder without injecting much additional personal capital. Wu (2007) provides fascinating accounts of many entrepreneurs both when they first "wore a red hat," setting up a nominally collectively owned firm, and when they tried to take off the hat, with varying degrees of success. Because of the recognition that most newly established "collectively owned firms" in the 1990s and early 2000s were private firms in disguise, we adopt a broad definition of domestic private firms to include all such firms.

In Table 2, we report a simple exercise that decomposes the contributions to the growth by firm ownership (domestic private firms, majority state-owned firms, and foreign-invested firms). Let X(total, t) be the industrial value added for the country as a whole in year t. Define X(private, t), X(FDI, t), and X(SOE, t) to be the industrial value added in year t by the domestic private sector, foreign invested firms, and state-owned firms, respectively. X(total, t) = X(private, 04) + X(FDI, 04) + X(SOE, 04). Let s(private, 95), s(FDI, 95), and s(SOE, 95) be the share of the domestic private sector, foreign firms, and state-owned firms, respectively, in the natural industrial output in 1995. We can decompose the overall growth rate into a weighted average of the growth rates from the three types of firms:

(1) G(total) = X(total, 04)/X(total, 95) – 1
= g(private)*s(private, 95) + g(FDI) s(FDI, 95) + g(SOE) s(SOE, 95)

From this equation, we can compute the contribution of the domestic private sector to the overall growth as: Private sector's share of the contribution = g(private)s(private,95)/g(total) = 6.22*30.7%/2.66 = 71.9%. Similarly, foreign invested firms account for 30.8% of the overall growth. The state sector accounts for -2.7% as many state-owned firms were either closed or taken over by foreign or domestic private firms. (Note that the decomposition of the real or nominal growth rates gives the same result.)

We can next decompose the private sector growth into the extensive margin (the growth in the number of firms) and the intensive margin (the growth of average output per firm)

(2) Ln[X(private, 04)/X(private, 95)] = Ln[N(private, 04)/N(private, 95)] +
Ln{[X(private,04)/N(private, 04)]/ [X(private, 95)/N(private, 95)]}

The first term on the right hand side denotes the extensive margin, while the second term denotes the intensive margin growth. In Table 3, we report the result of decomposing the real growth of the private sector. The contribution of the extensive margin = the first term on RHS/LHS = 0.499/0.713 =70.1%.

To summarize, a little over 70% of the Chinese growth is attributable to the rise of the private sector. In addition, a little over 70% of the private sector growth is attributable to the birth and the growth of new private firms. Therefore, the birth and the growth of new private firms are a significant part of the Chinese growth story.

<u>Where are domestic private firms most likely to emerge?</u>

We now examine whether there is any connection between the sex ratio and the extensive margin of the private sector growth. In Figure 1, we plot the growth in the number of private firms in a province from 1995 to 2004 against the sex ratio imbalance for the age cohort of 10-24 at the beginning of the period. In Figure 2, we plot the growth of the firm count against the average sex ratio in 1995 and 2004. In both cases, a strong positive association between the two is clearly visible.

Many factors could affect the birth and the growth of new firms. The age structure of the local population, the growth rate of the population, local income and education levels, local industrial structure, and initial scale of the private sector could all matter. After controlling for these factors, we are interested in investigating whether local sex ratio also plays a role. We do so by looking at variations in the growth rate of the count of private firms and local sex ratios (measured in three different ways) across 1788 counties, conditional on other factors. The specification is as follows:

$$(3) \quad Growth\_in\_firm\text{-}count_{k,\,95\text{-}04,} = \beta\ Sex\_ratio_{k,\,95,} \ +X_k\,\Gamma \ +e_k$$

The result is reported in Column 1 of Table 4. The coefficient on initial sex ratio is 1.95 and statistically significant. We also employ two alternative measures of the sex ratio. The first alternative is the average sex ratio for the age cohort 10-24 in 1995 and 2004 (inferred from the 1990 and the 2000 censuses, respectively). The second is the growth rate in the sex ratio (or the increase in the log sex ratio, ln(sex ratio in 2004)- ln(sex ratio in 1995)). The corresponding regressions are reported in Columns 2 and 3 of Table 4, respectively. In both cases, the coefficients are positive and statistically significant. In other words, more domestic private firms were established in regions with a higher average sex ratio during the period, or a higher growth rate in the sex ratio. In Columns 4-6, we add province fixed effects (a province has several counties). While the coefficients on the sex ratio variable become smaller, they remain positive and statistically significant.

Possible problems with the OLS estimation and solutions

The OLS estimation may produce biased estimates. First, there could be errors in measuring the sex ratio for the pre-marital age cohort at the county level. For example, with migration in and out of a county (in spite of the policy restrictions), the sex ratio recorded in the population census may not exactly correspond to the sex ratio in the local marriage market. The measurement errors tend to lead to underestimated coefficients. Second, the sex ratio might be endogenous. In particular, the positive association between the local sex ratio and the rate of growth of private firms may reflect a reverse causality. For example, if private entrepreneurs have a stronger urge to have a male heir to take over their business when they retire, then regions that happen to see a lot of private firms may also exhibit a strong son preference and a high sex ratio imbalance. Third, the sex ratio may be endogenous if it is correlated with some missing regressors. For example, in spite of our best effort to control for determinants of the growth of private firms, there may be other variables that are good predictors of future profitability in a region that are not captured by our list of control variables. If these variables happen to be correlated with the local sex ratio, we may find a positive association between the local sex ratio and the growth of local private firms even when there is no direct economic association between the two variables.

To address these problems associated with the OLS estimation, we adopt three approaches. First, we implement a two-stage least squared (2SLS) procedure in which the local sex ratio is instrumented by variables that affect regional variations in the sex ratio but are otherwise unlikely to affect directly the growth of local private firms. Second, we adopt a placebo test on the growth of other firms. If the local sex ratio is simply a proxy for missing regressors that help to forecast local growth potential, then the sex ratio should also forecast the extensive margin growth of foreign-invested firms. On the other hand, if the connection between the local sex ratio and the extensive margin growth of domestic private firms reflects primarily our theory, then the local sex ratio would not forecast the extensive margin growth of foreign firms. Third, we go to household-level data where we can check possible interactions between local sex ratios and son-families in ways that will also help us to rule out the endogeneity story. In particular, our theory suggests that son families and daughter families may respond differently to a rise in the local sex ratio. We will discuss these approaches in turn.

Instrumental variable approach

A strategy to address both the measurement error problem and the endogeneity problem is to employ an instrumental variable approach. A key determinant of the sex ratio imbalance is a strict family planning policy introduced at the beginning of the 1980s[4]. We explore three variables determinants of local sex ratios that are unlikely to be affected by the growth of local private firms, and for which we can get data. First, while the goals of family planning are national, the enforcement is local. Ebenstein (2008) proposes to use regional variations in the monetary penalties for violating the birth quotas, originally collected by Scharping (2003), as instruments for the local sex ratio. The idea is that, in regions with stiff penalties, parents may engage in more sex-selective abortions, rather than paying a penalty and having more children. The monetary penalty is often on the order of between one to five times the local average annual household income. In addition, Eberntein (2008) coded a dummy for the existence of extra fines for violations at higher-order births. For example, an additional penalty may kick in on a family for having the 3rd or 4th child in a one child zone, or the 4th or 5th child in a two-children zone. Such a non-linear financial penalty scheme was introduced by different local governments in different years (if at all), generating variations across regions and over time. These two monetary penalty variables constitute the first two candidates for our instrumental variables. [5]

The third instrumental variable explores the legal exemptions in the family planning policy. While the policy imposes a strict birth quota on the Han ethnic group (the main ethnic group in the country), the rest of the population (i.e., some 50 ethnic minority groups) do not face or face much less stringent quotas. (The government allowed the exemption, possibly to avoid criticisms for using the family planning policy to marginalize the minority groups). As a result, the share of non-Han Chinese in the total population has risen from 6.7% in 1982 to 8.5% in 2000 (Bulte, Heerink, and Zhang, 2009). Non-Han Chinese are not uniformly distributed across space. In regions with relatively more ethnic minorities, marriages between Han and non-Han peoples are not uncommon, reducing the competitive pressure for men in the marriage market. Therefore, the share of non-Han Chinese in the local population offers another possible instrument. (This instrument is also used in Bulte, Heerink and Zhang, 2009).[6] Since the same

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[4] China's family planning policy, commonly known as the "one-child policy," has many nuances. Since 1979, the central government has stipulated that Han families in the urban areas should normally have only child (with some exceptions). Ethnic Han families in rural areas can have a second child if the first one is a daughter (this is referred to as the "1.5 children policy" by Eberstein, 2008). Ethnic minority (i.e., non-Han) groups are generally exempted from birth quotas. Non-Han groups account for a relatively significant share of local populations in Xinjiang, Yunnan, Ganshu, Guizhou, Inner Mongolia, and Tibet.

[5] Edlund et al (2007) conduct some diagnostic checks and conclude that the level of financial penalties is uncorrelated with a region's current economic status. We will perform and report a formal test on whether the proposed instruments and the error term in the second stage regressions are correlated.

[6] In principle, variations in the cost of sex screening technology especially the use of an Ultrasound B machine, and the economic status of women (such as that documented in Qian 2008) could also be candidates for instrumental variables. Unfortunately, we

monetary penalties may generate more sex selective abortions by Han families in regions with a less competitive marriage market, we also include interaction terms between the two monetary penalty variables and the share of the non-Han Chinese in the local population.

The first stage regressions are reported in Table 5. The dependent variables for the three regressions are, respectively, the initial sex ratio for the age cohort 10-24, the average sex ratio of the same age cohort over 1995 and 2004, and the increase in the log sex ratio from 1995 to 2004. The coefficients on the share of the local population not subject to birth quotas are negative in all three regressions, and statistically significant in the first two regressions. This is consistent with the notion that sex selective abortions are less prevalent when more people are not subject to birth quotas.

The financial penalties for violating birth quotas generate a positive and statistically significant coefficient in all three regressions. The dummy for the existence of extra penalties for violations at higher-order births also produces a positive coefficient in all three regressions (which are significant in two cases). These results imply that a more severe penalty for violating legal birth quotas tends to induce parents to more aggressively abort girls, resulting in a higher sex ratio imbalance. In other words, when the penalties are light, many couples with daughters may opt to keep the daughter, pay the penalties, and have another child, rather than abort the female fetus. The interaction terms between the share of non-Han Chinese in the local population and the two financial penalty measures produce statistically significant coefficients in five out of six cases.

The adjusted R^2's are in the range between 0.09-0.24. The F statistics (for the null that all slope parameters are jointly zeros) are 53, 50, and 20, respectively. The Kleibergen-Paap statistics (for weak instruments) for the first two regressions are 48.1 and 33.4, respectively, which are greater than the Stock-Yogo 10% critical value of 19.9. The Kleibergen-Paap statistic for the third regression is 14.5, which is greater than the Stock-Yogo 15% critical value of 11.6. Across the three first-stage regressions, it appears that the instruments perform better in the first two cases.

The second stage regressions are reported in Table 6. While the first three regressions do not include province fixed effects, the last three regressions do. All three measures of the local

do not have the relevant data. Note, however, for the validity of the instrumental variable regressions, we do not need a complete list of the determinants of the local sex ratio in the first stage.

sex ratio in both specifications (six coefficients in total) are positive and statistically significant. The Durbin-Wu-Hausman test easily rejects the null that the 2SLS and OLS estimates are the same in all six regressions, implying that the sex ratio variable is likely to be endogenous. The Hansen's J statistics in the first three regressions (without the province fixed effects) indicates that the instruments may be correlated with the error term. On the other hand, the same test in the last three columns (with the province fixed effects) does not reject the null that the instruments and the error term are uncorrelated[7]. This leads us to prefer the estimation results in the last three regressions. The point estimates in Table 6 are generally larger than their OLS counterparts in Table 4. This suggests that the bias in Table 4 generated either by missing regressors or by measurement errors is substantial.

We can compute the economic significance of the estimates. Using the most conservative estimate in Column 5, an increase in the sex ratio by 3 basis points (e.g., from 1.08 to 1.11), which is equal to the increase in the average sex ratio from 1995 to 2004 (see Table 1), generates an increase in the natural log number of private firms by 0.249 (=8.3x0.03). Since the actual increase in log number of firms in this period is log 254988- log 807821 = 0.499 (see Table 2), the rise in the sex ratio can potentially explain 49.9% (=0.249/0.499) of the actual increase in the number of private firms in China during this period. In other words, the economic impact of the rise in sex ratio in promoting entrepreneurial activities is potentially very big.

Placebo tests

We now turn to placebo tests. The basic idea is to examine the birth of new foreign invested firms and that of state-owned firms, and to check if they are related to the local sex ratio imbalance. If the positive association between the local sex ratio and the growth in the number of domestic private firms is purely an artificial outcome of missing regressors that predict relative profitability across regions and happen to be correlated with the local sex ratio, we would expect to also find a similarly positive association between the growth in the number of foreign firms and the local sex ratio. On the other hand, if our theory is right that a higher sex ratio imbalance drives more domestic entrepreneurial activities, then the local sex ratio won't necessarily affect how foreign-invested firms choose to locate their productions in China, and won't necessarily

[7] In household-level regressions to be reported in Tables 10 and 11, we check if the ethnic minorities have a different labor supply pattern from Han Chinese, holding constant local sex ratio imbalance and other determinants of labor supply, and cannot reject the null that there is no difference.

affect how bureaucrats decide where to set up new state-owned firms (or shut down existing ones).

The placebo tests are reported in Table 7. In the first three regressions, the dependent variable is the growth in the number of foreign-invested firms from 1995 to 2004. The right-hand-side regressors are identical to those in Table 4. In none of the cases can we reject the null that the coefficient on the sex ratio variable is zero. In other words, statistically speaking, the location of new foreign-invested firms is uncorrelated with the local sex ratio imbalance. In Columns 4-6 of Table 7, we perform similar placebo tests with the growth in the number of state-owned firms as the dependent variable. In two out of the three cases, the coefficient on the sex ratio is not statistically different from zero. Taken together, the placebo tests make it unlikely that the local sex ratio is a proxy for missing regressors that predict future profitability in a region.

In Table 8, we apply the same instrumental variable approach to the placebo tests. Again, the local sex ratio is uncorrelated with the growth in the number of foreign invested firms during 1985-2004. While the sex ratio is significant in two cases when we look at the growth in the number of state-owned firms, the coefficients have a negative sign. In other words, a strong sex ratio imbalance may be associated with a faster contraction of the state sector in the local economy (though this result is not robust to alternative ways to measure the sex ratio). Overall, the placebo tests further bolster our confidence in the interpretation that a higher sex ratio imbalance stimulates more entrepreneurship.

Household-level evidence

We now examine evidence from the Chinese Household Income Project (CHIP) of 2002, which covers 9,200 households in 122 rural counties in 22 provinces. Household survey data offers a new angle to check our theory as it allows us to examine families with a son and families with a daughter separately. As the survey does not contain information on business ownership and probably does not cover many business owners anyway, we are not able to examine which households set up firms. However, we can examine a household's supply of labor and willingness to accept a relatively dangerous job (in exchange for a relatively good pay), and their connections with the local sex ratio.

To make the households as comparable as possible, we construct a sub-sample of households with two living parents and a child.[8] This sub-sample consists of 480 families with a son and 262 families with a daughter in 122 rural counties. Since most unmarried young people live with their parents, the survey does not contain many observations of an unmarried young man or woman as the household head. Therefore, we are not able to analyze single-person households directly.

We focus on two key aspects of a household's labor supply that are captured in the survey. The first is a household's willingness to accept a relatively dangerous job. A dangerous job is defined in the survey as one in the mining or construction sector, or one with exposure to extreme heat, extreme cold, or hazardous materials. While the survey does not contain occupation-specific wage information, we may expect that, in equilibrium, the wage rate is higher for a dangerous (or unpleasant) job than other jobs, holding constant skill requirement and other determinants of the wage. In other words, people presumably accept a more dangerous (or a less pleasant) job in exchange for a higher pay. The second variable that we look at is the total number of days in a year that members of a household worked off the farm (mostly as a migrant worker). Off farm work usually pays better, but one has to endure all the difficulties and inconvenience associated with working away from the hometown. Given the policy restrictions on internal migration in China, most migrant workers treat out-of-town jobs as temporary, do not expect to settle in the cities where they work, and likely return to their hometowns eventually.

The summary statistics on these two variables across the rural counties are reported in Table 9. The last panel indicates that, on average, 27.6% of all three-person households in a county have at least one family member working in a relatively dangerous job. The average fraction is only moderately higher for households with a son (27.7%) than households with a daughter (27.4%). However, the standard deviation across the counties (around 45%) is big. As for the total number of days members of a household worked off the farm, the unconditional average is 35.6 days per household. The son families work more days off farm (41.4 days) than the daughter families (24.9 days). From the summary statistics, we cannot rule out the possibility that the differences across the two types of households simply reflect a greater ability for a man to work away from home than for a woman. Our theory, however, implies a particular regional

[8] We place an age limit of 40 for heads of households with an aim to exclude households who may have older children that we do not observe because they have left home. Such households may not be comparable to those with only a stay-at-home child.

variation in the labor supply: son families are more willing to take a relatively dangerous job or work more days if they are located in a region with a more unbalanced sex ratio. Therefore, to test our theory, we have to explore interactions between a household's labor supply and the local sex ratio, while controlling for other determinants of labor supply.

In Table 10, we perform three Probit regressions on household propensity to accept a relatively dangerous job in 2002. All the regressions control for family income, children's ages, and characteristics of the head of the household (age, education, and ethnic background). It also controls for health shocks to the family by a dummy denoting "poor health" if the family has a disabled or severely ill member. In Column 1, we focus on households with a son. The local sex ratio has a positive and significant coefficient, implying that a son-family is more willing to take a relatively dangerous job if it lives in a region with a higher sex ratio imbalance. An increase in the sex ratio by 3.5 basis points (which is equal to one standard deviation across the rural counties in the sample) is associated with an increase in the probability for a son-family to accept a dangerous job by 17.5 percentage points (e.g., an increase from 20% to 37.5%). Since the unconditional mean in the sample is 27.7% (as reported in the first column in Table 9), this effect is economically large.

The regression in Column 2 of Table 10 looks at households with a daughter. The coefficient on the local sex ratio is not statistically significant. In other words, the willingness to accept a dangerous job for a daughter-family is unrelated to the local sex ratio imbalance. In the last column of Table 10, we combine the two sets of households and add a dummy for households with a son and an interaction term between the dummy and the local sex ratio. The local sex ratio is insignificant while the dummy for son-families has a negative coefficient. Most interestingly, the interaction term between the local sex ratio and the dummy for son families is positive and statistically significant. Our interpretation is that it is not having a son per se that motivates families to be more willing to accept a dangerous job. Rather, it takes a combination of having a son and living in a region with a high sex ratio imbalance to induce families to be more eager to accept a relatively dangerous job.

Table 11 performs Tobit estimations on the total number of days in a year that household members worked off farm. In the first column, we look at households with a son. The coefficient on the local sex ratio is positive and statistically significant. An increase in the local sex ratio by 3.5 basis points is associated with an increase in the supply of off farm labor by 1.0 day/year

(=27.8x0.035). Since the unconditional mean in the sample is 35.6 days per year per household (Column 2 of Table 9), this represents a non-trivial although not a huge effect. Across all rural counties in the sample, the difference between the maximum and minimum sex ratio is 13 basis points. This would translate into a difference in the supply of off-farm labor by 3.6 days per year per household.

In the second column of Table 11, we look at households with a daughter. The coefficient on the local sex ratio is not statistically significant. This implies that the supply of off farm labor by daughter families is uncorrelated with the local sex ratio. In the third column of Table 11, we combine the two sets of households, and add a dummy for son families and an interaction term between the dummy and the local sex ratio. Similar to Table 10, only the interaction term is positive and statistically significant. In other words, a combination of having a son and living in a region with a high sex ratio imbalance motivates these rural households to be more willing to work away from home.

Over all, the patterns in Tables 10 and 11 are consistent with each other, and consistent with our theory. Of course, accepting a relatively dangerous job and working more days are not mutually exclusive. Taken together, the estimation results suggest that, as the sex ratio imbalance increases, son families respond by increasing moderately the number of days in off-farm work, but increasing significantly the willingness to accept a relatively dangerous job, presumably in pursuit of a higher pay.

<u>Sex ratios and per capita GDP growth</u>

So far, we have discussed evidence on how a higher sex ratio stimulates the extensive margin of economic growth in the form of the birth of new private firms, and have also presented some evidence on how it increases the intensive margin of economic growth in the form of a greater supply of work effort and a greater tolerance of hardship and hazardous work environment. To capture the general equilibrium effect, we now examine the direct connections between sex ratios and local income growth by using a panel data on provincial GDP per capita from 1980 to 2005. We organize the data into five 5-year periods, 1980-85, 1985-90, 1990-95, 1995-2000, and 2000-05.

Let y(k, t) be the log GDP per capita for province k in period t. We run the following regression:

$$[y(k, t+5)-y(k, t)]/5 = \beta\ sr(k,t) + X(k,t)\Gamma + \text{province fixed effects} + \text{period fixed effects} + e(k,t)$$

where the dependent variable is the average annual growth rate in a 5-year period, sr(k,t) is the sex ratio for the age cohort 15-24 in province k and period t (inferred from the 1990 Population Census), and X(k,t) is a vector of control variables which includes the beginning-of-period log income, y(k,t), the share of working age population in local population, the ratio of local investment to local GDP, the ratio of local foreign trade to local GDP, and birth rate. β is a scalar parameter and Γ is a vector of parameters to be estimated, and e(k,t) is an error term that is assumed to be independent and identically normally distributed. Each period is a five-year interval. The choice of the control variables is based on the set of robust predictors of growth from the empirical growth literature (Sala-i-Martin, xxx). One key missing regressor is human capital, of which we do not have a good measure that is both across provinces and over time. We will implement a 2SLS estimation that aims to address this (and other) problems.

Some summary statistics for the panel are reported in Table 12. During 1980-2005, the average annual growth rate of per capital GDP across the provinces was 8.6% with a standard deviation of 2.8%. The average sex ratio for the age cohort 7-21 in 1980 was 107 boys/100 girls (only slightly higher than the normal ratio), but there were already variations across the provinces with the standard deviation being 3.5 and the maximum ratio being 114 boys/100 girls. As indicated earlier, the sex ratio deteriorates over time.

The panel growth regressions with both province fixed effects and period fixed effects are reported in Table 13. In Columns 1 and 4, we report regressions without the sex ratio measure (for comparison). In Columns 2-3 and 5-6, two different measures of the local sex ratio (the average of a 5-year interval and the value at the beginning of each interval, respectively) are used. In all four regressions that include the sex ratio, the coefficients on the sex ratio are positive and statistically significant. This is consistent with the idea that a higher local sex ratio is associated with a higher income growth rate.

To address concerns with missing regressors (in particular, the absence of human capital), and also possible endogeneity and measurement errors associated with the sex ratio measure, we implement a 2SLS approach, where the instruments for the sex ratio are the same as in Table 5. The (second-stage) estimation results are reported in Table 14. The coefficients on the sex ratio measure in all regressions are positive and statistically significant. The Hansen's J statistics indicates that one cannot reject the null that the instruments are uncorrelated with the error term, which is what we want from the instruments.

We can assess the economic significance of the estimates. Take the first column as an illustration, an increase in the sex ratio by 2 basis points (which is about the average increase in a five-year interval), holding other variables constant, would raise the growth rate by 1.38 percentage points per annum (= 0.69 x 0.02 x 100). This accounts for 15.8% (=1.98/8.75) of the actual mean increase in the annual income growth during this period. In other words, the effect of a rise in the sex ratio on the growth of per capita GDP is not only statistically significant but also economically significant.

4. Concluding Remarks

Robert M. Solow, the Nobel Prize winner for his pioneering work on the theory of economic growth, is quoted as having said, "Everything reminds Milton (Friedman) of the money supply. Well, everything reminds me of sex, but I keep it out of the paper."[9] Well, Solow might have missed something economically significant by not linking sex with economic growth. This paper proposes that an unbalanced sex ratio (in the pre-marital age cohort) may be one of the significant drivers for economic growth.

A strong sex ratio imbalance is present in China, Vietnam, Korea, India, Taiwan, Singapore and several other economies due to a combination of a parental preference for sons, easy availability of technology to screen the sex of a fetus, and a limit on the number of children that a couple either desires to have or is allowed to have. As men face a diminishing prospect of finding a wife, parents of a son or the son himself are more eager to do something to improve the son's standing in the marriage market relative to other men in the same age cohort. Since family wealth is a significant measure of one's relative standing, parents with a son and men respond to a rise in the sex ratio by engaging in more entrepreneurial activities, supplying more labor, and becoming more willing to take unpleasant or relatively dangerous jobs, all in pursuit of a higher expected pay. The economic significance of this mechanism has thus far not been empirically examined.

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[9] "The concise encyclopedia of economics: Robert Merton Solow (1924-)." www.ecolib.org/library/Enc/bios/Solow.html.

We find strong supportive evidence across regions and households in China. Using data from two censuses of industrial firms in 1995 and 2004, we find that the local sex ratio is a statistically significant predictor of which regions are more likely to have new domestic private firms (beyond other determinants of the birth and growth of new firms) since entrepreneurship brings a higher expected income. The economic impact is also significant: an increase in the sex ratio by one standard deviation can potentially explain 50% of the difference in the rates of growth of new private firms across regions. Across rural households, we find that families with a son respond to a higher sex ratio by moderately increasing the number of days that they work off farm (mostly as migrant workers) in a year but significantly increasing their willingness to take a relatively dangerous job, presumably in exchange for a higher pay. Households with a daughter do not respond to a higher sex ratio in the same way. These patterns are consistent with our story.

The sex ratio imbalance may have a mean reverting feature. That is, it is unlikely that the sex ratio at birth can deteriorate beyond 140 boys per 100 girls. Parents would re-evaluate their preference for sons before the society reaches that point, and the increasing difficulty for a man to find a wife is a key reason for the adjustment in the preference[10]. Korea is an example of mean reversion: it used to have a worse sex ratio imbalance a decade ago that it does today, though it still has too many boys at birth relative to girls. At the same time, both the Korean example and cross regional data in China suggest that the reversion of the sex ratio is a slow process[11]. In any case, we know with a high degree of confidence that the sex ratio for the pre-marital age cohort in China will be getting worse in the next twenty years than it is today, since the sex ratio at birth today is significantly worse than the ratio for today's 10-year-olds, which in turn is worse than the ratio for today's 20-year-olds. This means that both the extensive margin (more births of new private firms) and the intensive margin of economic growth (greater tolerance of hard work and dangerous jobs, and greater supply of work effort) due to the sex ratio imbalance will continue to be a force to reckon with in the foreseeable future. This will partially offset the natural force of a declining growth rate that one may expect from the Solow growth model.

Accumulating more wealth is not the only way for men or households with a son to compete in the marriage market. Parents may also invest more in the education of their sons, and

[10] In a theoretical model, Edlund (2009) suggests that as the sex ratio rises, poor families may start to switch their preference over gender of their children in favor of having girls.

[11] A regression of the change in sex ratio at birth on lagged sex ratio across Chinese provinces over 1980-2005 suggests a slow reversion. The half life for the convergence is estimated to be 25 years.

push them to work harder in schools. There may also be a spillover from a boy's education to a girl's education. Such mechanisms have not been empirically investigated. In addition, as noted early, many other economies such as Vietnam, Korea, India, Singapore, and Taiwan also have a strong sex ratio imbalance. Some of them are also known to have a high rate of economic growth. It will be interesting to examine rigorously whether a sex ratio imbalance has played a significant role both in these economies' development and in cross country variations in the growth rates generally. We leave these topics for future research.

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Table 1: Sex Ratios for Age Cohort 10-24 at the County Level in 1995 and 2004

Census Year	Mean	Median	Standard deviation
1990	1.08	1.07	0.05
2000	1.11	1.10	0.08

Note: Authors' calculation based on China Population Censuses in 1990 and 2000.

Table 2: Contributions to the growth of Chinese industrial output by ownership

		Private	SOEs	Foreign	total
1	Value added in 1995 (billion yuans)	2956	5470	1210	9636
2	Value added in 2004 (billion yuans)	21350	4790	9080	35220
3	growth rate =(2)/(1)-1	6.22	**-0.12**	6.50	2.66
4	Initial share in total VA (%)	30.7	56.8	12.6	100.0
5	**Share in growth rate (%)**				
	= [(3)*(4)]/Growth of total VA	71.9	-2.7	30.8	100.0

Table 3: The extensive versus intensive margins in the growth of the Chinese private sector

Decompose private sector real growth

		Number (a)	Avg Output (million yuans/firm) (b)=(c)/(a)	Total VA (billion 1995 yuans) (c)
1	1995	807821	3.66	2956
2	2004	2549888	5.98	15250
3	Growth rate =log(row 2)-log(row 1)	0.499	0.213	0.713
4	Share in growth (%)	70.1	29.9	100.0

assuming P2004/P1995=1.4

Table 4: Sex Ratios and the Growth in the Number of Private Firms at the County Level from 1995 to 2004

	R1	R2	R3	R4	R5	R6
Sex ratio for the age cohort 10-24 in 1995	1.95**			0.65*		
	(0.40)			(0.37)		
Sex ratio for the cohort 10-24 averaged over 1995 and 2004		1.76**			0.74**	
		(0.40)			(0.35)	
Increase in log sex ratio for the age cohort 10-24 from 1995 to 2004			0.92**			0.65*
			(0.40)			(0.36)
Log number of firms in 1995	-0.46**	-0.48**	-0.47**	-0.56**	-0.57**	-0.57**
	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)
Log GDP in 1995	0.34**	0.35**	0.36**	0.30**	0.30**	0.31**
	(0.05)	(0.04)	(0.05)	(0.05)	(0.05)	(0.05)
Average year of schooling based on 2000 census	-0.11**	-0.10**	-0.11**	0.09**	0.07*	0.05
	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)
Share of agricultural output in gross output values in 1995	-1.13**	-1.14**	-1.14**	-0.61**	-0.60**	-0.60**
	(0.13)	(0.12)	(0.13)	(0.11)	(0.11)	(0.11)
The ratio of local revenues to total government employees	0.10**	0.10**	0.10**	0.15**	0.14**	0.14**
	(0.04)	(0.04)	(0.04)	(0.05)	(0.04)	(0.04)
Population growth from 1990 to 2000	-0.01	0.00	0.00	0.00	-0.01	0.02
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
Share of labor force (aged 20-64) in total population in 1995	(0.69)			-2.20**		
	(0.52)			(0.63)		
Share of labor force (aged 20-64) in total population, averaged over 1995 and 2004		(0.20)			-1.34*	
		(0.61)			(0.73)	
Increase in log labor force share in local population from 1995 to 2004			(0.14)			0.76**
			(0.30)			(0.28)
Province fixed effects	no	no	no	yes	yes	yes
Adjusted R square	0.27	0.27	0.26	0.49	0.49	0.49
AIC	3901	3897	3921	3275	3282	3280
N	1788	1788	1788	1788	1788	1788

Notes: The growth in the number of private firms is measured by the increase in the log number of firms from 1995 to 2004. The sex ratio for the age cohort 10-24 in 1995 is inferred from the age cohort 5-19 in the 1990 population census; the sex ratio for the same age cohort in 2004 is inferred from the age cohort 6-20 in the 2000 population census. The definition of private firms includes township-and-village enterprises and other "collectively owned" firms. Provincial fixed effects are included in the last three regressions but not reported. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 5: Instrumenting for the Local Sex Ratio - First Stage Regressions

	R1	R2	R3
Share of minority population in 1990	-0.16**	-0.19**	-0.05
	(0.03)	(0.04)	(0.04)
Penalties for family planning violations at the provincial level	0.003**	0.003**	0.007**
	(0.00)	(0.00)	(0.00)
Dummy for extra penalty for higher order births at the provincial level	0.011**	0.012**	0.004
	(0.00)	(0.00)	(0.01)
Share of minority population in 1990 * penalties for family planning violations	0.127**	0.143**	-0.002
	(0.03)	(0.04)	(0.04)
Share of minority population in 1990 *Dummy for extra penalty for higher order births	0.057**	0.089**	0.054**
	(0.02)	(0.03)	(0.02)
Initial number of firms (log) in 1995	0.001	0.008**	0.007**
	(0.00)	(0.00)	(0.00)
Per capita GDP (log) in 1995	0.005**	0.004**	0.002
	(0.00)	(0.00)	(0.00)
Average year of schooling based on 2000 census	-0.004**	-0.005**	-0.013**
	(0.00)	(0.00)	(0.00)
Share of agricultural output in gross output values in 1995	-0.007	0.003	0.017**
	(0.01)	(0.01)	(0.01)
The ratio of local revenues to total government employees	-0.002	0.000	-0.006**
	(0.00)	(0.00)	(0.00)
Increase in log population	0.000	-0.008**	0.000
	(0.00)	(0.00)	(0.00)
Share of labor force (aged 20-64) in local population in 1995	-0.050		
	(0.06)		
Share of labor force (aged 20-64) in local population, averaged over 1995 and 2004		-0.508**	
		(0.05)	
Increase in log labor force share in local population from 1995 to 2004			0.116**
			(0.03)
Adjusted R square	0.12	0.24	0.09
F statistic	53.0	50.0	19.8
Kleibergen-Paap Wald statistic	48.1	33.4	14.5
AIC	-6128	-5594	-5318
N	1780	1780	1780

Notes: The share of minorities in local population is computed from the 1990 population census at the county level; the two financial penalty variables are only available at the province level. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively. For the Kleibergen-Paap Wald F test for weak instruments, the Stock-Yogo critical values are: 10% maximal IV size 19.93; 15% maximal size 11.59; 20% maximal size 8.75; 25% maximal IV size 7.25.

Table 6: 2SLS Estimation on Sex Ratios and Growth in the Number of Private Firms

	R1	R2	R3	R4	R5	R6
Sex ratio for the age cohort 10-24 in 1995	17.7**			19.2**		
	(2.17)			(7.23)		
Sex ratio for the cohort 10-24 averaged over 1995 and 2004		13.9**			8.3**	
		(1.66)			(2.78)	
Increase in log sex ratio for the age cohort 10-24 from 1995 to 2004			14.4**			11.2**
			(2.38)			(4.08)
Log number of firms in 1995	-0.57**	-0.67**	-0.64**	-0.61**	-0.62**	-0.61**
	(0.04)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)
Log GDP in 1995	0.27**	0.31**	0.33**	0.20**	0.24**	0.23**
	(0.04)	(0.04)	(0.05)	(0.06)	(0.05)	(0.05)
Average year of schooling of local population	-0.05	-0.04	0.06	0.17**	0.08**	0.13**
(inferred from the 2000 census)	(0.03)	(0.04)	(0.05)	(0.06)	(0.04)	(0.05)
Share of agriculture in gross output values in 1995	-1.01**	-1.15**	-1.35**	-0.68**	-0.69**	-0.80**
	(0.16)	(0.14)	(0.17)	(0.17)	(0.13)	(0.16)
The ratio of local government revenue to government employees	0.09*	0.06	0.14**	0.17**	0.16**	0.23**
	(0.05)	(0.05)	(0.05)	(0.05)	(0.04)	(0.06)
Increase in log population	0.00	0.10**	0.01	0.00	0.03	0.03
	(0.03)	(0.03)	(0.03)	(0.02)	(0.03)	(0.03)
Share of labor force (aged 20-64) in local population in 1995	0.11			-3.76*		
	(1.08)			(1.98)		
Share of labor force (aged 20-64) in local population, averaged over 1995 and 2004		6.23**			1.22	
		(1.14)			(1.18)	
Increase in log labor force share in local population from 1995 to 2004			-1.41**			(0.52)
			(0.55)			(0.63)
Province fixed effects	No	No	No	Yes	Yes	Yes
Durbin-Wu-Hausman test for endogeneity	0.00	0.00	0.00	0.00	0.00	0.00
Hansen's J statistic for over identification	0.00	0.00	0.00	0.48	0.27	0.22
N	1780	1780	1780	1780	1780	1780

Notes: Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 7: Placebo Tests - Sex Ratios and the Growth in the Number of Other Firms

	Foreign Invested Firms			State-owned Enterprises		
	R1	R2	R3	R4	R5	R6
Sex ratio for the age cohort 10-24 in 1995	0.00			0.55*		
	(0.53)			(0.33)		
Sex ratio for the cohort 10-24 averaged over 1995 and 2004		0.13			0.02	
		(0.47)			(0.32)	
Increase in log sex ratio for the age cohort 10-24 from 1995 to 2004			0.28			-0.47
			(0.48)			(0.32)
Log number of firms in 1995	-0.38**	-0.38**	-0.38**	-0.40**	-0.40**	-0.41**
	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)
Log GDP in 1995	0.21**	0.21**	0.21**	0.01	0.01	0.02
	(0.06)	(0.06)	(0.06)	(0.02)	(0.02)	(0.02)
Average year of schooling based on 2000 census	0.16**	0.16**	0.16**	0.02	0.04	0.02
	(0.07)	(0.07)	(0.07)	(0.03)	(0.03)	(0.03)
Share of agricultural output in gross output values in 1995	-0.56**	-0.56**	-0.57**	-0.10	-0.10	-0.08
	(0.17)	(0.17)	(0.17)	(0.11)	(0.11)	(0.11)
The ratio of local revenues to total government employees	0.30**	0.30**	0.30**	0.09**	0.10**	0.09**
	(0.07)	(0.07)	(0.07)	(0.04)	(0.04)	(0.04)
Increase in log population	0.00	0.00	0.00	0.02	0.02	0.01
	(0.02)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
Share of labor force (aged 20-64) in local population in 1995	0.04			0.27		
	(0.97)			(0.58)		
Share of labor force (aged 20-64) in local population, averaged over 1995 and 2004		-0.03			-0.56	
		(1.10)			(0.68)	
Increase in log labor force share in local population from 1995 to 2004			(0.10)			-0.54*
			(0.45)			(0.31)
Adjusted R square	0.29	0.29	0.29	0.28	0.28	0.28
AIC	2424	2423	2423	3816	3817	3812
N	1074	1074	1074	1799	1799	1799

Notes: Provincial fixed effects are included but not reported. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 8: 2SLS Estimation on the Placebo Tests - Sex Ratio and the Growth of Other Firms

	Growth in Number of Foreign Invested Firms			Growth in number of State-owned Enterprises		
	R1	R2	R3	R4	R5	R6
Sex ratio for the age cohort 10-24 in 1995	-3.96			-11.21**		
	(7.48)			(5.69)		
Sex ratio for the cohort 10-24 averaged over 1995 and 2004		-1.85			-3.98	
		(4.46)			(2.47)	
Increase in log sex ratio for the age cohort 10-24 from 1995 to 2004			-2.76			-5.53*
			(5.21)			(3.33)
Log number of firms in 1995	-0.40**	-0.39**	-0.39**	-0.42**	-0.41**	-0.41**
	(0.06)	(0.05)	(0.05)	(0.05)	(0.05)	(0.05)
Log GDP in 1995	0.24**	0.23**	0.24**	0.11*	0.06	0.07*
	(0.10)	(0.09)	(0.10)	(0.06)	(0.04)	(0.04)
Average year of schooling based on 2000 census	0.14*	0.16**	0.14*	0.01	0.05*	0
	(0.08)	(0.07)	(0.08)	(0.04)	(0.03)	(0.03)
Share of agricultural output in gross output values in 1995	-0.58**	-0.54**	-0.48**	-0.03	-0.05	0.02
	(0.19)	(0.17)	(0.21)	(0.14)	(0.11)	(0.13)
The ratio of local revenues to total government employees	0.30**	0.30**	0.28**	0.07*	0.09**	0.04
	(0.08)	(0.08)	(0.09)	(0.04)	(0.04)	(0.05)
Increase in log population	0.00	-0.02	-0.01	0.02	-0.01	0.01
	(0.02)	(0.04)	(0.03)	(0.02)	(0.03)	(0.02)
Share of labor force (aged 20-64) in local population in 1995	0.18			1.17		
	(1.20)			(1.36)		
Share of labor force (aged 20-64) in local population, averaged over 1995 and 2004		-1.02			-1.94*	
		(2.25)			(1.07)	
Increase in log labor force share in local population from 1995 to 2004			0.41			0.11
			(0.98)			(0.51)
Durbin-Wu-Hausman test for endogeneity	0.00	0.00	0.00	0.00	0.00	0.00
Hansen's J statistic for over identification	0.67	0.60	0.63	0.05	0.00	0.00
N	1067	1067	1067	1790	1790	1790

Notes: Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively. Province fixed effects are included but not reported.

Table 9: Some Summary Statistics for Three-person Households

	% of households with at least one member taking a dangerous job	Total number of days that a household worked off farms
Families with a son (480)		
Mean	27.7	41.4
Standard deviation	44.8	96.3
Families with a daughter (262)		
Mean	27.4	24.9
Standard deviation	44.7	69.8
All 3-person households in the sample (742)		
Mean	27.6	35.6
Standard deviation	44.7	88.1

Note: The sample consists of households with two living parents and a child. The child is at least 4 years old and the household head is younger than 40. A dangerous job is defined as one in a mining or construction sector, or with exposure to an extremely high or low temperature or hazardous material at work. The working day count refers to number of days members of the household working off-farms (as a migrant worker or in a factory).

Table 10: Probit Estimation of Household Propensity to Take a Dangerous Job in 2002 (Marginal Effect)

	One son	One daughter	**Total**
Local sex ratio for age cohort 12-26	0.05**	-0.02	-0.03
	(0.02)	(0.03)	(0.03)
Having a son			-8.0**
			(3.46)
Sex ratio*son			0.07**
			(0.03)
Log household income	0.06	-0.02	0.02
	(0.10)	(0.12)	(0.08)
Year of education	-0.02	-0.04	-0.03
	(0.03)	(0.04)	(0.02)
Household head as minority ethnic group	-0.22	-0.36	-0.3
	(0.33)	(0.38)	(0.25)
Poor health among at least one family member	-0.21	-0.74	-0.4
	(0.39)	(0.57)	(0.31)
Head younger than 35	-0.24*	-0.03	-0.16
	(0.15)	(0.19)	(0.12)
Age of a child 5-9	0.53**	0.21	0.40**
	(0.24)	(0.35)	(0.20)
Age of child 10 or older	0.31	0.35	0.32*
	(0.22)	(0.33)	(0.18)
N	480	262	742

Notes: Sex ratio for age cohort 12-26 is inferred from the age cohort 0-14 in the 1990 population census. Other data are derived from the rural sample of CHIP 2002. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 11: Tobit Estimation on the Number of Off-farm Working Days

	One son	One daughter	Total
Local sex ratio for age cohort 12-26	27.8**	7.5	5.3
	(6.7)	(8.5)	(8.4)
Having a son			-2434**
			(1144)
Sex ratio*son			23.1**
			(10.6)
Log household income	78.4**	-39.1	32.44
	(35.9)	(40.49)	(28.57)
Year of education	-4.76	13.03	0.46
	(10.25)	(15.77)	(8.82)
Household head as minority ethnic group	88.05	50.2	70.97
	(105.05)	(110.77)	(78.53)
Poor health among at least one family member	-56.21	-93.49	-72.6
	(136.43)	(147.22)	(104.54)
Household head younger than 35	-49.99	-54.20	-51.34
	(48.63)	(67.41)	(40.02)
Age of child 5-9	85.18	-56.56	42.01
	(82.97)	(121.29)	(69.62)
Age of child 10 or older	41.92	-66.54	10.36
	(76.07)	(111.70)	(63.47)
N	480	262	742

Notes: Sex ratio for age cohort 12-26 is inferred from the age cohort 0-14 in the 1990 population census. Other data are derived from the rural sample of CHIP 2002. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 12: Summary Statistics on Growth Rates of Per Capita GDP and Sex Ratios, 1980-2005

	Mean	Median	Standard deviation	Minimum	Maximum
Annualized growth rate of GDP/per capita (over a five-year interval)	8.75	8.71	2.77	0.41	16.84
Sex ratio for the 15-24 age cohort in 1980	107	107	3.5	103	114
Sex ratio for the 15-24 age cohort in 2005	109	110	2.x	102	116

Note: The per capita GDP growth rate is the LHS variable used in Table 13 and Table 14.

Table 13: Sex Ratios and Income Growth Rate from 1980 to 2005

	R1	R2	R3	R4	R5	R6
Average sex ratio for age cohort 7-21		0.23**	0.22**			
(average in five year interval)		(0.07)	(0.07)			
Initial sex ratio for age cohort 7-21					0.16**	0.17**
(initial in a 5-year interval)					(0.06)	(0.06)
Log initial per capital GDP	-5.59**	-8.28**	-8.18**	-6.63**	-7.35**	-8.43**
	(1.20)	(1.59)	(1.55)	(1.21)	(1.53)	(1.45)
Share of labor force (aged 20-64)	-0.15	-0.25**	-0.18	-0.31**	-0.23*	-0.30**
in total population	(0.14)	(0.11)	(0.13)	(0.15)	(0.12)	(0.13)
Investment/local GDP	0.08**	0.09**	0.09**	0.10**	0.09**	0.10**
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
Foreign trade/local GDP	0.11*	0.09	0.09	0.23**	0.10	0.22**
	(0.06)	(0.08)	(0.07)	(0.09)	(0.07)	(0.10)
Birth rate	-0.17		-0.15			
(average over a five-year interval)	(0.11)		(0.12)			
Birth rate in the initial year				-0.03		-0.06
				(0.12)		(0.11)
Province fixed effects	yes	yes	yes	yes	yes	Yes
Time fixed effects	yes	yes	yes	yes	yes	Yes
Adjusted R-square	0.63	0.66	0.66	0.64	0.64	0.66
AIC	510	499	498	509	505	501
N	140	140	140	140	140	140

Notes: The data is a panel of five 5-year periods, 1980-85, 1985-90, 1990-95, 1995-2000, and 2000-05. The dependent variable is the average annual growth rate of per capita GDP over a 5-year period. The sex ratio is inferred from the 1900 population census. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 14: 2SLS Estimation – Sex Ratios and Income Growth Rates over 1980 to 2005

	R1	R2	R3	R4
Average sex ratio for age cohort 7-21	0.69**	0.51*		
(average in five year interval)	(0.18)	(0.27)		
Initial sex ratio for age cohort 7-21			0.66**	0.61**
(initial in a 5-year interval)			(0.19)	(0.21)
Log Initial per capita GDP	-13.56**	-11.53**	-12.67**	-13.63**
	(2.30)	(3.09)	(2.30)	(2.26)
Share of labor force (aged 15-59) in total population in 1990	-0.29**	-0.02	-0.24*	-0.18
	(0.12)	(0.29)	(0.14)	(0.28)
Investment/local GDP	0.10**	0.09**	0.10*	0.09*
	(0.05)	(0.05)	(0.05)	(0.05)
Foreign trade/GDP	0.06	0.09	0.06	0.22*
	(0.09)	(0.07)	(0.10)	(0.13)
Birth rate		-0.56		
(average in five year interval)		(0.60)		
Birth rate in the initial year				-0.45
				(0.75)
Province fixed effects	yes	yes	yes	yes
Time interval fixed effects	yes	yes	yes	yes
Durbin-Wu-Hausman test for endogeneity	0.16	0.47	0.14	0.01
Hansen's J statistic for over identification	0.36	0.33	0.33	0.11
N	140	140	140	140

Notes: The data is a panel of five 5-year periods, 1980-85, 1985-90, 1990-95, 1995-2000, and 2000-05. The dependent variable is the average annual growth rate of per capita GDP over a 5-year period. The instrumental variables include the share of minority population, penalty for violating family planning policy (% of local yearly income), and a dummy for extra penalty for higher order births are used as instrumental variables. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

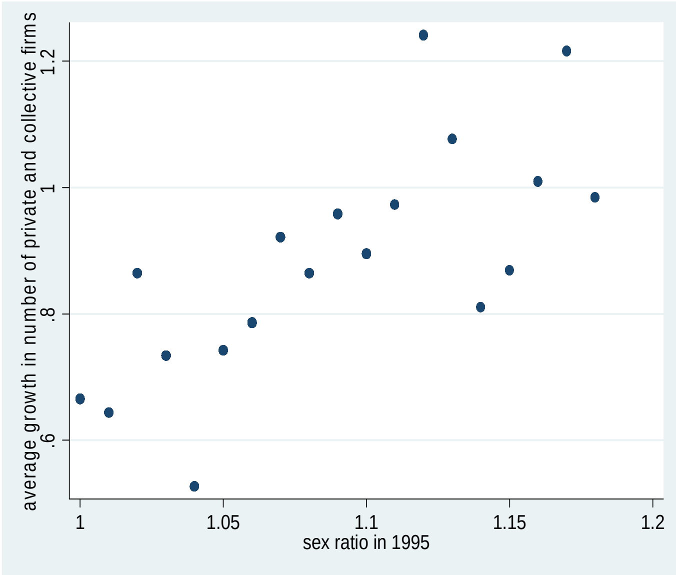

Figure 1: Initial Sex Ratios and Growth Rates of Private Firms during 1995-2004

On the horizontal axis is the sex ratio for the age cohort 10-24 in 1995 inferred from the 1990 Population Census. On the vertical axis is the growth rate in the number of private firms from 1995 to 2004, averaged over all counties that had the same value of sex ratio (up to a basis point).

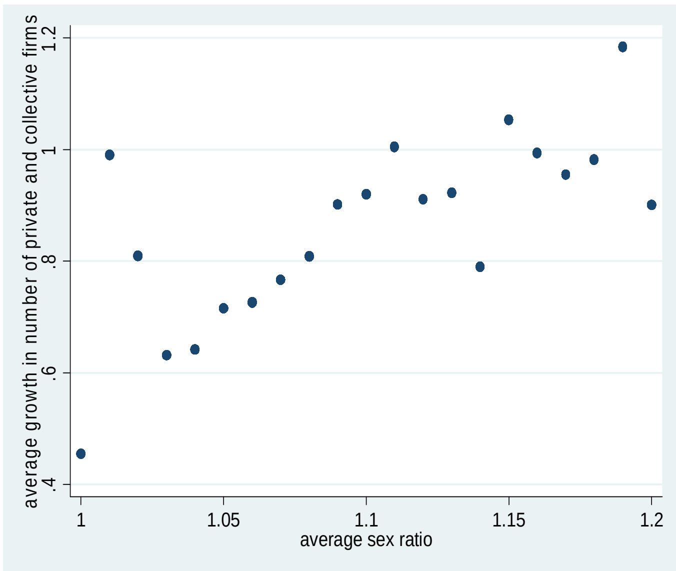

Figure 2: Average Sex Ratios and Growth Rates of Private Firms during 1995-2004

On the horizontal axis is the sex ratio for the age cohort 10-24, averaged over 1995 and 2004. On the vertical axis is the growth rate in the number of private firms from 1995 to 2004, averaged over all counties that had the same value of sex ratio (up to a basis point).