Son Preference and the Missing Market for Social Insurance: Evidence from China's Rural Pension Program

Avraham Ebenstein and Steven Leung*

Harvard University

August, 2008

Abstract

China's high sex ratio at birth and the dramatic rise in recent years has alarmed policymakers worldwide. Many argue that the persistence of son preference is driven by greater anticipated old age support from sons relative to daughters and the absence of formal financial mechanisms for families to save for retirement (Das Gupta et al. 2003). We exploit the introduction of voluntary old-age insurance in rural China in the 1990's to examine whether (a) parents with sons are less likely to participate in pensions plans and (b) whether providing access to pension plans affects parental sex selection decisions. Consistent with the first hypothesis, we find that parents' with sons are less like to participate in the pension program and have less financial savings for retirement, after controlling for observable characteristics of the parents. Consistent with the second hypothesis, we find that the 1991 implementation of rural old age pension programs mitigated the increase in the sex ratio observed in counties that had at least partial adoption of the program, controlling for observable characteristics of the counties.

* We are grateful to Jonathan Gruber, Nancy Hearst, Jonathan Pohl, Jonathan Schlesinger, and Deng Quheng.
Corresponding author is Ebenstein: aebenste@rwj.harvard.edu.

1 Introduction

In 1979, Chinese instituted a One Child Policy in order to slow the growth of its immense population. Communist party officials feared that the rapid population growth was inhibiting China's economic growth and modernization, and so initiated coercive and restrictive measures (White 2006). While the policy restricted fertility for both rural and urban residents, its primary aim was to slow fertility among China's rural peasant population, who represent over two-third's of China's population. Though fertility in China declined dramatically in the subsequent period from 2.75 to 1.5 births per mother, the male fraction of births began to rise. Sex-selective abortion became common in the countryside (Zeng et al. 1993, Chu 2001) and the sharpest increase occurring in areas with the strictest policy enforcement (Ebenstein 2008). These patterns indicate that parents in modern rural China are reluctant to complete fertility without having at least one son.

The persistence of son preference in spite of China's modernization is at first surprising, but many point to economic realities in rural China that make sons valuable to parents. While son preference is in part a religious or cultural preference that is unexplained by incentives, a large component of son preference is the economic value they provide to parents, and this occurs through three primary channels. First, sons are potentially more valuable as farmers, and so they provide more labor income to parents[1] (Qian 2005). Second, since all land in China is formally owned by the State, a son provides parents the opportunity to reap financial benefits from their land through an important bequest mechanism, as daughters generally join their husband's

[1] Qian (2005) exploits variations in regional and sex-specific incomes in China and finds that higher relative income for females mitigates the distortion in the sex ratio of births.

family upon marriage.[2] Third, sons are traditionally expected to care for their parents in old age, and so a motivation for having a son is the desire to secure a viable source of support when parents can no longer care for themselves, and this is an important motivation in many developing countries (Cameron and Cobb-Clark 2001). This third motivation is the focus of this paper. We exploit quasi-random variation in China's introduction of voluntary old-age insurance in hundreds of villages throughout rural China to examine whether sons and formal savings mechanisms appear to be substitutes to parents, and by extension, the degree to which the female deficit is related to a missing market for social insurance.

Using the China 2002 Rural Household Income Survey (CHIS), a recent survey which contains detailed information on village programs and demographic characteristics, we find that parents without sons are more likely to participate in these programs, and also have a higher amount of savings specifically devoted for the purpose of old age security than those who have a son. Both facts suggest that son preference is in part driven by fears concerning old age support, and may consequently partially explain the dramatic increase in the sex ratio in rural areas in the 1990's following the introduction of ultra-sound. As such, in the second phase of the analysis we attempt to answer a potentially more important policy question: has the experiment with rural old age insurance impacted the share of parents willing to complete fertility without a son? Using a matched sample of China's county level 2000 Census and the social program information in the CHIS, we find that the increase in the sex ratio between the program's inception in 1991 and 2000 was more severe in counties with fewer villages participating in the program, with program adoption lowering the sex ratio at birth (SRB) by 8%, even after controlling for county-level variation in average education, share employed in agriculture, and other factors affecting the sex

[2] Because most households in rural China see themselves not as landowners but as merely having the rights contracted to them by the village collective, (Kung 2000), the legitimacy of an intergenerational transfer of these contracted rights may be predicated on the presence of a male heir who will continue to farm a specific plot of land.

ratio at birth. The results suggest that further expansion of the program may mitigate the sex ratio distortion by providing an additional mechanism for parents to prepare for old age.

The remainder of the paper is laid out as follows. Section 2 provides a review of the existing literature on the role of sons as a source of old age support in China, and institutional details regarding the market-based pension schemes in rural China introduced in 1991. Section 3 describes the two primary data sources for the analysis: CHIS survey data used to investigate take-up of these programs, and the China 2000 census data used to examine the effect of program availability on the sex ratio at birth. Section 4 presents the empirical results of the CHIS analysis, documenting higher program participation among those without a son. Section 5 examines the relationship between the availability of these programs and the male fraction of births. We conclude in section 6 with a brief discussion of the policy implications of the findings.

2 Background: Son's as Old Age Support and Efforts to Introduce Market Pension Schemes in Rural China

2.1 Sons as providers of Old Age Support

Son preference is deeply rooted in traditional Chinese culture, and the need to have at least one son is especially strong for both religious and practical reasons. Religious motivations for having sons are rooted in the Confucian custom that only a son can perform tasks such as ancestor worship, filial support of elders[3], and continuation of the family line as patriarch of the family[4] (Milwertz 1997). The logic of patrilinearity is rigid to the point that a daughter can never

[3] A notable passage from Confucius's analects regarding filial piety reports Confucius as stating, "If there is work to be done, young people shoulder the burden, and when wine and food are served, elders are given precedence (Confucius, Analects 2.8).

[4] An alternative to continuing the familial lineage by way of son is the uxorilocal marriage. A uxorilocal marriage, as opposed a virilocal marriage in which a couple resides with the husband's parents after marriage, involves residence with the wife's parents after marriage. Burton Pasternak (1985) suggests that uxorilocal marriages are not as rare as one would expect. Uxorilocality is said to be exercised only by the relatively small number of couples who are unable to bear or adopt a son, and is considered a less than optimal way to ensure lineage continuity. Uxorilocal

be a substitute for a son in a perfectly patrilineal framework (Das Gupta et al. 2003, Greenhalgh and Winckler 2005).

Survey evidence also indicates that the preference for a son is partly related to the more economically quantifiable need for old age support. Indeed, over half of respondents (51%) to a fertility survey in Hubei province professed that the primary motivation for a son is the desire for old age support, and place continuation of the family line (20%) as the distant second most common response (Davin 1985).

The fact that parents in rural China rely on sons rather than daughters for elderly care is perhaps due to the traditionally patrilocal culture[5] (Bray 1997). In patrilocal cultures, daughters join their new family upon marriage and are expected to provide care to their in-laws. As such, a son will generally provide parents with two caregivers (himself and his bride), while a daughter marries off and provides parents with no caregivers. Those who fail to have a son are in a precarious situation in terms of old age care, as parents in rural China are reluctant to rely on public support for care in old age – a sentiment that is further reinforced by historically common government warnings that families in rural areas should not anticipate government old age support (Lin 1994). In combination with the traditional role of the son as caregiver, it is perhaps unsurprising that the forced reduction in fertility is associated with an increase in the sex ratio in rural areas.

While in most parts of rural China sons provide support to parents in old age, this custom is not universal. Li et al. (2004) analyze parental support in rural Songzi county, and find that

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marriages have high moral costs – the groom is considered to enter the brides family as a "parasite," and he commits an unforgivable act by turning his back on his own ancestral line.

[5] Bray (1997) explains traditional kinship in China as being determined according to rules of patrilineal descent. The closest bonds are those trace along the lines of male descent, and relatives on the mother's or wife's side were considered more distant relatives. In some areas, it was even customary to enforce restrictions on visits between a wife and her natal family.

sons and daughters provide similar levels of care to parents[6]. Interestingly, this county has maintained a stable sex ratio of births in spite of the rigid fertility restrictions imposed on its inhabitants. The authors suggest that perhaps the village's diversified marriage forms and the apparent weak son preference are a result of a perception that daughters and sons are equivalent in their provision of old-age support.

The Songzi case-study suggests that the rise in the sex ratio may in part be driven to the reliance on sons, and that changing patterns in old age support may reduce the willingness of parents to engage in sex selection in the face of fertility restrictions. In most of rural China, the role of the son as caregiver is a deeply entrenched custom and will not change in the immediate future, but the recent attempts to introduce market-based reforms in rural reforms may indeed mitigate this son preference. China's efforts to establish rural pension programs are outlined in the next section.

2.2 China's Market Based Pension Schemes in Rural Villages

In 1986, China began to explore alternative means to the traditional family-based old age support systems in rural areas (Wang 2006).[7] The Rural Old Age Insurance Program[8] was initiated under the supervision of the Ministry of Civil Affairs, and aimed to develop an

[6] Specifically, they find the counter intuitive result that sons in uxorilocal marriages and daughters in virilocal marriages are more likely to provide financial support for older parents; that is, daughters support their natal families financially even in virolocal marriages, and sons who 'marry out' uxorilocally support their natal families as well.

[7] Whereas consistently growing urban old age security systems have been in place since 1951, as a part of the nation building and socialist rhetoric of the Mao era, old age programs have been largely inaccessible to the rural areas. Access to any old age security program is further limited by the extremely rigid urban/rural hukou registration that prevents rural residents from reregistering as urban residents.

[8] A direct translation of the program would be 'insurance'. It may have been a deliberate gesture on the part of the Ministry of Civil Affairs to name the program 农村养老保险制度. The title is directly translatable to Rural(Farm) Old Age Insurance, and it avoids 退休金, the word for 'retirement pension.' Most western scholars have subsequently referred the program as a 'pension,' which is more in line with the actual function of the program. Interestingly, some villages intentionally propagandized the programs as 'insurance' to tempt pensioners into believe that they there was a larger element of government protection involved (Leisering et al. 2002).

institutional framework for administering a pension scheme based on voluntary contribution, defined-contribution, and fully funded individual accounts (Shi 2006). In 1991, the Ministry launched successful preliminary programs in Weihai and Yantai counties in the Shandong Province, and subsequently expanded the program to include most counties of Jiangsu province, before outlining plans for expansion throughout China (Leisering et al. 2002). The program did not provide a complete social safety net, but allowed individuals to contribute to the pension scheme with tax exempt earnings, and supplemented these contributions with subsidies from employers and local communities. Despite the Ministry of Civil Affair's announcement of a nationwide initiative to establish these programs, participation was voluntary at the individual level, and was voluntarily and locally administered at the city- and county-level by social insurance agencies.

In spite of the initial success of preliminary rural pension schemes and a successful expansion during the early 1990s, the growth of these programs stagnated in the late 1990s as participants began complaining about poor management and alleged that fund managers were embezzling pension funds. The Asian financial crisis of 1997 further impeded the growth of the program, as the large decline in asset prices made village leaders wary of placing resident assets in modern financial vehicles (Wang 2006). Additionally, the village administrators in many localities were also dis-incentivized from participating, since the administrative costs of the pensions schemes were large, and were not subsidized by the provincial or national government. In addition to the start-up costs, the programs also carried an annual operating expense of 3% of total local pension contributions (Leisering et al. 2002). These administrative difficulties made it difficult to convince the rural population to participate, and were further compounded by the

prevailing belief among rural peasants that land, family, and market provision should remain the primary providers of old age security (Shi 2006).

The program, admittedly, has encountered difficulty, but it has still attracted a sizable population of insurants even though participation is fully voluntary. The size of the Rural Old Age Insurance program peaked in 1997 with 82.8 farmer insurants but declined to approximately 53.8 million in 2004 (Shi 2006). The percentage of the rural farmers insured under these schemes declined as well, from 15.4 percent in 1997 to around 11 percent in 2004 (Wang 2006). In a recent survey, Wang (2006) reports that roughly 7% of rural elders age 60 and older receive either social old age insurance or pension benefits, implying that the program has maintained a presence in spite of the difficulties in the implementation. Today, there is an ongoing dialogue in China regarding how to improve these programs, or potentially to switch to a unified and centrally backed rural old age pension system.

While participation rates of these programs are reported by aggregate data, very little is known regarding the patterns of take-up in these programs. In the next section, we describe the data we use to analyze this take-up decision, which combine micro-level information on program availability, participation, and the number and gender of living children among those who choose to contribute to the rural pension scheme. We also describe the demographic information available at the county-level in the census, which we use to explore whether these programs have mitigated the female deficit in areas where the program was available.

3 Data

3.1 Rural China Household Income Survey

The 2002 China Household Income Survey (CHIS) provides rich demographic information for individual survey respondents in rural areas, and is complemented by a village-level survey that collected information regarding features of villages and their participation in social insurance programs. The CHIS sample is comprised of 39,969 individuals across 996 villages in various parts of China. Table 1 summarizes the sample means of these responding villages, and the map in Figure 1 reflects the wide coverage of the survey. The survey comprises villages that were selected in a multi-staged random sampling, and represent six geographic regions and a total of 95 counties in China (Li et al. 2008).

The CHIS individual survey contains detailed information regarding individual participation in the old age pension program, individual level savings specifically for the purpose of old age security, and whether the individual lived in a village where the pension program was available.[9] Approximately 6.7% of responding villages report that they offer the program to their residents and 5.4% of the respondents report being participants in the old age pension program. Importantly, the survey also contains detailed information on the number and gender of children in each family surveyed.[10] In section 4, we examine the individual participation decision exploiting the rich demographic and policy information for each survey respondent.

[9] Availability of the program is reported by the party branch secretary, the head of village committee, or the village accountant.

[10] In order to examine family structure and fertility outcomes, the Integrated Public Use Microdata Series mother matching algorithm was used to match mothers with their children, limiting the dataset to 15,948 observations of mothers and their matched children.

Because the CHIS is limited in size,[11] the institutional information in the survey is combined with county-level data from China's 2000 census, which provides a more comprehensive assessment of fertility patterns.

3.2 China 2000 Census County Level Data

In order to assess the impact of the availability of the program on China's sex ratio, we exploit a county-level dataset for China that contains tabulations of fertility rates and sex ratios for China's 2,870 counties.[12] The CHIS information is provided at the village-level and is aggregated and matched to the census data at the county level. Table 2 reports the sample means in the census sample, as well as the means for the 95 counties which had at least one village surveyed in the CHIS.

In Figure 2, we briefly summarize the county-level patterns in the sex ratio at birth in a color-coded map of China. The map reflects dramatic regional heterogeneity across counties in the missing women problem among children of age zero to nine.[13] The map indicates that the sex ratios are particularly high along the southern coast of China and along the central eastern coast, areas where son preference is known to be intense and that are under stricter enforcement of China's fertility restrictions (Gu et al. 2007). In contrast, the map reflects that Western provinces such as Tibet, Xinjiang, and Inner Mongolia – all with large populations of policy-exempt ethnic minorities – have less skewed sex ratios. The spatial heterogeneity in the sex ratio distortions and

[11] Though the CHIS is a smaller sample, the patterns in fertility are similar to what is observed in the census. In Appendix Figure 3.1 and 3.2, I report the conditional sex ratio of births among respondents, which indicates that among parents with a daughter, over 58% of second births are male. Given parents with 2 daughters, almost 70% of third births are males. These results are similar to those found in China's 2000 Census, reported in Appendix Figure 4.1 and 4.2. However, the small share of villages participating make it difficult to rely on the CHIS for the analysis of the program's effect on the sex ratio of births.

[12] The analysis is based on data provided by the Harvard Geospatial Library by merging the China 2000 Census Historical data and the China 2000 9.5% Long Form data using county *guobiao* indicators.

[13] Though more precise measures for the SRB are available, I choose to depict the sex ratio of this birth cohort as a proxy for SRB because of its significance in later analyses of the pension programs announced in 1991.

the missing women phenomenon suggests that a county-level analysis of China's fertility trends may be informative regarding the factors underlying the increase in the sex ratio. These sex ratio distortions are explored in more detail in section 5.

4 Old Age Preparation and Fertility Outcomes

In the following section, we explore the relationship between the number and gender of a family's children and financial decisions to prepare for old age. Survey respondents in the CHIS who were not currently enrolled in the Rural Old Age Insurance program were asked whether they would be interested in participating in this program in the future. Over 50% of respondents who were not already participants expressed a desire to participate in the future, suggesting that the old age insurance is still underdeveloped. Interestingly, Figure 3 reflects the strong link between fertility outcomes and a respondent's interest in participating in these pension schemes. Among parents with one or two children, those with only sons or only daughters appear similar in terms of their desire to participate. However, parents with 3 or 4 daughters are markedly more likely to express interest in future participation than those with 3 or 4 sons. Among those with 3 or 4 daughters, 89% and 91% respectively express interest in participating in the future, whereas only 76% and 50% of those with 3 or 4 sons wish to participate in the future. This is suggestive that parents anticipate more elder care from sons, a hypothesis we explore more rigorously in the next section.[14]

[14] An alternative hypothesis is that the birth of multiple sons in China is often deemed expensive because of bequest expectations, and only wealthier families could afford more sons. A family with many sons has a greater the denominator of inheritors, and requires a larger accumulation of wealth. As such, mothers with four sons may be wealthy enough to not demand separate old age support. Arthur Wolf (1985) suggests that historically, the fertility of families with greater accumulated wealth (he specifically mentions farm size) has been higher than that of less wealthy families. This alternate hypothesis partially informs the decision to control for household income, total financial assets, and household agricultural output in the following analyses.

4.1 Program Participation and Fertility Outcomes

To examine the relationship between program participation and a couple's fertility outcome, we estimate a set of Linear Probability Model regressions (ordinary least squares). In the first specification, we estimate the effect of a binary variable for failure to have a son on a binary variable for participation in the rural old age pension program.

$$participation_i = \alpha + \beta_1(failed_i) + \varepsilon_i \qquad (1)$$

The estimates of this specification are reported in column 1 of Table 3. We find that parents who fail to have a son are 1.1 percentage points more likely to participate in the program, which implies that they are roughly 28% more likely to participate (the sample average for participation is 3.9%). The coefficient is significant at the 1% confidence level.

In a second specification, we add explanatory for a set of observable family characteristics such as household income measured in RMB, annual hours spent in agriculture as a proxy for unreported household agricultural output, financial assets, the mother's age and age squared, and the mother's educational attainment measured in years of education.

$$\begin{aligned} participation_i = {} & \alpha + \beta_1(failed_i) + \gamma_1(age_i) + \gamma_2(age^2{}_i) + \gamma_3(yrseduc_i) \\ & + \gamma_4(income_i) + \gamma_5(finassets_i) + \gamma_6(planthrs_i) + \varepsilon_i \end{aligned} \qquad (2)$$

The results are reported in column 2 of Table 3, with the coefficients on the controls suppressed. Again, the results indicate that failing to produce a son positively increases the likelihood of participating by 1.38 percentage points. The coefficient is significant at the 1% level.

The third and fourth specifications examine the impact of the total number of sons and daughters on the probability of participation. Specification 3 estimates the effect of additional number of sons and daughters that parents have on pension participation.

$$participation_i = \alpha + \beta_2(sons_i) + \beta_3(daughters_i) + \varepsilon_i \qquad (3)$$

The results of the specification, reported in column 3 of Table 3, indicate that additional daughters and sons both induce a reduced participation rate, but each additional son reduces the probability of participation by 1.8 percentage points, whereas an additional daughter reduces the probability of participation by 1.5 percentage points. The coefficients on additional sons and daughters are both significant to the 1% level.

A fourth specification includes the controls that were added in specification 2, and estimates the coefficients on sons and daughters with the addition of observable characteristics.

$$\begin{aligned} participation_i = {} & \alpha + \beta_2(sons_i) + \beta_3(daughters_i) + \gamma_1(age_i) + \gamma_2(age^2{}_i) \qquad (4) \\ & + \gamma_3(yrseduc_i) + \gamma_4(income_i) + \gamma_5(finassets_i) + \gamma_6(planthrs_i) + \varepsilon_i \end{aligned}$$

With the added family characteristic descriptors as controls, the coefficients on additional sons and daughters remain negative, but their magnitudes are larger. They remain statistically significant to the 1% level. We find that an additional son decreases the probability of participation by 2.4 percentage points, whereas an additional daughter decreases the probability of participation by 2.2 percentage points.

Specification 5 examines the effects of binaries for families that include one son, two sons, one daughter, or two daughters, with the set of *j* additional control explanatory variables listed above, on pension program take-up.[15]

[15] For convenience and to avoid repetition, I abbreviate the set of 6 control regressors and their coefficients using summation notation.

$$participation_i = \alpha + \beta_4(oneson_i) + \beta_5(twosons_i) + \beta_4(onedaught_i) \quad (5)$$
$$+ \beta_5(twodaught_i) + \sum_{j=1}^{6} \gamma_j(control_{j,i}) + \varepsilon_i$$

The results of this specification are reported in column 5 of Table 3, and suggest that binaries for two daughters reduce participation by lower amounts than binaries for two sons. The coefficients on binaries for one daughter and one son are similar, but diverge at two son and two daughter fertility outcomes, suggesting that pension take-up decisions are more influenced by two daughter fertility outcomes than by one daughter fertility outcomes.

These findings are also graphically depicted in Figure 4, which indicates that families with 1 child are more likely to participate and that among families with two children, those with two daughters are more likely to participate. Families who have an only daughter are more likely to participate than those who have an only son. Notably, the histogram depicts participation in a stepwise pattern where mothers with a fertility outcome of two sons are less likely to participate than those with a mixed gender two child fertility outcome, who are in turn less likely to participate than those with two daughters.

4.2 Savings for Old Age and Fertility Outcomes

To further analyze the influence of fertility outcomes on decisions to prepare for old age, We examine whether savings for old age in general is responsive to the number and gender of a couple's children. We estimate the effect of failing to have a son on log savings in the specification, with and without the set of added controls.

$$oldsave_i = \alpha + \beta_2(failed_i) + \varepsilon_i \quad (1)$$

$$oldsave_i = \alpha + \beta_2(failed_i) + \sum_{j=1}^{6} \gamma_j(control_{j,i}) + \varepsilon_i \quad (2)$$

The results of these regressions are listed in columns 1 and 2 of table 4, and suggest that failure to produce a son, leads to a 22% increase in savings. Controlling for observable characteristics, failure to produce a son leads to a 13% increase in savings for old age. The results are not statistically significant, but their directions and magnitude are important.

Specifications 3 and 4 examine the impact of number of daughters and number of sons on log savings for old age, with and without additional age, income, and financial assets regressors.

$$oldsave_i = \alpha + \beta_2(sons_i) + \beta_3(daughters_i) + \varepsilon_i \qquad (3)$$

$$oldsave_i = \alpha + \beta_2(sons_i) + \beta_3(daughters_i) + \sum_{j=1}^{6} \gamma_j(control_{j,i}) + \varepsilon_i \qquad (4)$$

Like the results on participation, the results in Table 4 suggest that more children, regardless of gender, induce decreases in savings devoted for old age. Parents with more children are markedly less likely to save in financial vehicles for old age, as measured by the total log savings for old age of the household head. Importantly, the magnitude of the reduction in savings induced by additional sons is significantly larger than the magnitude of the daughter-induced reduction. Unfortunately, the results on the regressions are not statistically significant, but the discrepancy in coefficients between daughters and sons evidences a differential perception of sons and daughters as providers for old age.

We then repeat the specification involving binaries for number of sons and daughters (Section 4.1, specification 5), replacing participation with log savings for old age as the dependent variable.

$$oldsave_i = \alpha + \beta_4(oneson_i) + \beta_5(twosons_i) + \beta_4(onesdaught_i) \qquad (5)$$
$$+ \beta_5(twodaught_i) + \sum_{j=1}^{6} \gamma_j(control_{j,i}) + \varepsilon_i$$

The results are reported in column 5 of Table 4, and suggest that a two daughter family is associated with a large increase in savings, whereas a two son family is associated with a slight decrease.[16]

4.3 Interpretation and Limitations

The results indicate that fertility outcomes are very much related to the decisions to prepare for old age, both in the form of participation in rural old age pension programs, as well as decisions to save more specifically for the purpose of old age preparations. Fertility outcomes that either include daughters, or lack sons, induce increased old age preparation, whereas the presence of sons in a fertility outcome tends to crowd out other forms of old age security.

Though the evidence suggests strong correlations, they unfortunately cannot be interpreted as the causal impact of fertility outcomes on the probability of program participation. First, program participation is partly influenced by access to the program, and it may be that villages with stricter fertility regulations (and consequently lower fertility rates) were more likely to adopt the program. This channel affecting program participation is interesting in its own right, since it is consistent with the hypothesis that fertility outcomes affect a couple's perceived need for financial resources for old age. However, the empirical evidence regarding general savings for old age would not be affected by such endogeneity, and are consistent with the results regarding program take-up.

Second, even if parents can participate in old age preparation regardless of program availability, it could be the case that parents who desire larger families are more traditional and have more skepticism for financial savings in general. An unobserved correlation between

[16] Table 1 of the appendix displays results from estimations of the Section 4.2 specifications using non-log savings.

preference for fertility and a general distaste for formal savings would also generate the negative correlation between savings and family size. However, the China Household Income Survey allows for a rich set of controls and we are able to control for a household's total financial assets. As such, the results we present indicate that fertility outcomes affect program participation and savings for old age, even when information on the full financial portfolio of the household is included as a regressor. Still, the exact channels through which fertility outcomes affect financial preparation for old age are not observable given the data limitations.

A third potential weakness is reverse causality: the families with sons may have engaged in sex selection, and their decisions to engage in sex selection were affected by program availability. In order to address this weakness in the identification strategy, and as an additional robustness check, we include an examination of whether couples who completed their first fertility decision before the introduction of the program exhibit similar patterns. We restrict the regression sample to those mothers whose first child was at least 11 years old at the time of the survey, and must have been born before the 1991 pilot date of rural pension programs. The results of the restricted sample regressions are reported in Table 5. The restricted sample produces similar coefficients, though the results are not significant.

5 Rural Old Age Insurance and the Sex Ratio at Birth

The China 2000 Census reflects a remarkable increase in the sex ratio at birth during the 1990's, potentially due to rigid enforcement of the One Child Policy among parents who wished to have at least one son. The introduction of the Rural Old Age Insurance Program in 1991 represents a natural experiment that provides variation within China for the desire to have at least one son. In the following section, we examine whether areas with the program had a more muted

increase in the number of "missing girls". By comparing the births within each county that occurred prior to the introduction of the program (the birth cohort of 1981 to 1990) with the births following the introduction of the program (the birth cohort of 1991 to 2000), we examine the impact of the program. We first describe overall patterns in the increase in the sex ratio that suggest that county-level incentives may play a role in this phenomenon, and then present the main empirical results which indicate that program availability mitigated the increase in the sex ratio.

5.1 Changes in the Sex Ratio Pre and Post the Rural Old Age Insurance Program

We first examine the overall patterns in the increase in missing women by county. Using the census tabulations on children born between 1991 and 2000 as a post-program sex ratio, and tabulations for children born between 1981 and 1990 as a pre-program sex ratio, we examine the increase in the sex ratio by county.[17] Figure 5 reports the percent change in sex ratio of pre- and post- program birth cohorts, as coded by color intensity. The results indicate that the increase in the sex ratio is most pronounced in the southern coastal region of China. Again, this is potentially due to intense son preference and rigid policy enforcement in these areas. The figure also reflects a high degree of spatial variation in the sex ratio changes in the last two decades, suggesting that county-level incentives may help explain the overall increase or decrease in the sex ratio.

In Figure 6, we demonstrate a widening in the gap between places with highly distorted sex ratios and those with more normal sex ratios. Figure 6 depicts a line plot of the means by

[17] Differential mortality rates by age and by gender may potentially bias the sex ratio of the earlier birth cohort. Hill and Banister (2004) explore mortality rates in China, and find that mortality rates for males and females from ages 5-19 tend to be similar, suggesting that mortality did not have a large meaningful impact on the sex ratios 1981 to 1990 birth cohort.

quartile of county level sex ratios (weighted by village population) of the two birth cohorts. The population-weighted mean county-level sex ratio has increased in the past two decades, from 107 in the earlier cohort to 117 in the later cohort. Notably, the difference in first and fourth quartile means of the zero to nine birth cohort is much greater than that of the ten to nineteen birth cohort. The line plots suggest that not only has the sex ratio increased on average over the two cohorts, but it seems that the variance in county level sex ratios across China is increasing. The divergence in sex ratios by county provides further motivation for a county-level exploration of the missing women phenomenon.

5.2 Village Level Program Availability and Fertility Outcomes

To examine the effects of pension program availability on sex ratio changes, we estimate a 'first-difference' regression of sorts on counties from which villages have been sampled in CHIS. We use as a measure of county level program availability the aggregated average of village level availability, weighted by village population. This coverage score is the percentage of villages within a county that have adopted the old age pension program, and represents the treatment for the analysis. As a dependent variable, we compute the percent change in pre- and post- program announcement sex ratios. We first estimate the specification

$$\Delta sexratio_i = \alpha + \beta_1(oldavail_i) + \varepsilon_i \qquad (1)$$

where the change in sex ratio is equivalent to the differences in sex ratio between the two birth cohorts.

$$\Delta sexratio_i = \frac{sexratio_{post} - sexratio_{pre}}{sexratio_{post}} = \frac{sexratio_{0to9} - sexratio_{10to19}}{sexratio_{0to9}} \qquad (1.1)$$

The results of this specification are reported in column 1 of Table 6, and initially indicate that an increase in county level program availability to 100% of village is associated with a positive 2.5% change in the sex ratio.

Additional specifications include regressors for average number of villages that offer rural health insurance programs, and average number of villages that offer poverty relief programs

$$\Delta sexratio_i = \alpha + \beta_1(oldavail_i) + \gamma_1(healthavail_i) + \gamma_2(povavail_i) + \varepsilon_i \qquad (2)$$

and regressors for proportion of total laborers that are employed in the construction and manufacturing sectors at the county level as proxies for urban development, as well as explanatory variables for total fertility rate and log of the fine rate[18] for an additional child.

$$\Delta sexratio_i = \alpha + \beta_1(oldavail_i) + \gamma_1(sharemanu_i) + \gamma_2(shareconstr_i) + \gamma_3(tfr_i) + \gamma_4(\log fine_i) + \varepsilon \qquad (3)$$

Finally, we estimate a specification that includes the above listed controls, as well as additional controls for educational attainment, village income, and share of village households that are communist cadre households (as a measure of voluntary adherence to fertility restrictions).

$$\Delta sexratio_i = \alpha + \beta_1(oldavail_i) + \sum_{j=1}^{11} \gamma_j(control_{j,i}) + \varepsilon_i \qquad (4)$$

Table 3 reports the results of specifications 2 through 4 in their respective columns. Including additional controls, particularly controls for fertility (TFR) and policy restrictions (log fine rate), results in a coefficient that is significant and positive. Controlling for a rich set of attributes, an increase in county level pension program availability from zero percent of villages 100 percent

[18] The fine rates are taken from Scharping (2003) with an imputation as performed by Ebenstein (2008). The code for the fine imputation is available on Ebenstein's website, as well as a detailed explanation of the calculation of the fine rate.

of villages is associated with an 8% decline in the sex ratio from pre-program to post-program. The coefficient is significant, but only at the 10% level.

5.3 Interpretations and Limitations

The results presented suggest that pension program availability is indeed associated with a negative change in the sex ratio (fewer missing women). Additionally, the directions and effects of other regressors are informative. Policy intensity, represented by the fine rate, is positively associated with changes in the sex ratio, whereas fertility is negatively associated with the sex ratio. These results imply that restricted fertility may lead to sex ratio deteriorations. The share of construction labor, where men would earn higher sex specific incomes, is associated with more missing women.

However, there are several caveats to interpreting the regression results above as the causal impact of the availability of the pension scheme on the sex ratio at birth. One limitation is that the adoption of the pension program was entrusted to village-level administrators who may have responded to the needs of the village inhabitants. For example, village decisions to offer the program could be related to the dissatisfaction of parents with their fertility outcomes. However, the availability of the program is almost certainly driven by the phased introduction of the program, and must be at least partially exogenous to specific decisions of village-level administrators. It is also unlikely that village elders would acquiesce to demands of village inhabitants if it is economically unfeasible or impossible for them to implement such programs in their village. As such, program availability in 2002 is most likely not influenced by village patterns in fertility.[19]

[19] Also note that if parents are demanding the program in response to undesirable fertility outcomes, this reinforces the paper's central claim that the sex ratio may be partly driven by fears regarding care in old age.

A more likely explanation is that villages vary mostly in their ability to successfully implement a structured and self-funded pension program, as is evidenced in the differences in means of Table 7, and the correlations between different social insurance programs presented in Table 8. Villages that participated in the program have richer and better educated inhabitants. Also, the participating villages are more likely to have established infrastructure; evidence of this is found in the greater availability of running water, health insurance, and poverty relief programs in villages that offer the rural pension program. This idea is plausible given the administrative complexity of running a pension scheme. As such, there could be certain factors, not properly accounted for in the regression, that are driving both the availability of pensions and the lower sex ratios. Note however that the CHIS and the county level census data allow for a rich set of county level and aggregated village level attributes to be included as controls in the regressions. For the estimates to be biased, the correlation between the sex ratio and the availability of the program must be operating through channels not already captured.

To further assess the claim that these village characteristics may confound the causality of the analysis, we present an additional set of propensity score weighted first-difference regressions as a robustness test. Using counties in which at least one village adopted the old age pension program, we construct a binary for county level availability. We then estimate the effects of the binary for availability on sex ratio in the following specifications, with the same set of control variables as in specification 4 above.

$$\Delta sexratio_i = \alpha + \beta_1(oldbinary_i) + \varepsilon_i \tag{5}$$

$$\Delta sexratio_i = \alpha + \beta_1(oldavail_i) + \sum_{j=1}^{11} \gamma_j(control_{j,i}) + \varepsilon_i \tag{6}$$

We then examine pension program availability as a treatment effect that has a propensity-of-adoption score based on county-level and aggregated village-level observables. The first-stage logistic estimate of the propensity score includes attributes that may impact both program adoption and sex ratio: average educational attainment, share of labor that is manufacturing and construction, availability of health insurance, availability of poverty relief programs, and share of households that are communist cadres. We then employ an inverse probability weighted differences regressions strategy (weighted by propensity score), as suggested by Rosenbaum and Rubin (1983), and re-estimate the above two specifications.

The results of this robustness test are listed in Table 9. The coefficients are smaller than those in Table 6 and most lack statistical significance, but the signs on the coefficients remain largely the same. Though the results cannot be directly compared to those of specifications 1 through 4 above because of the nature of constructing the binary variable for program availability, they still serve to reinforce the claim of a causal impact of the program on the sex ratio.

6 Conclusion

Although rapid industrialization and large changes in fertility have reshaped China in the last 40 years, son preference has survived the transition. The 2005 sex ratio reached 118 boys for every 100 girls, possibly related to stricter enforcement of China's One Child Policy in recent years (Ebenstein 2008).[20] Recently, the Chinese government has also both re-instated the One Child limit and declared that correcting the sex ratio at birth by 2016 is a national priority (Li 2007). Such goals may be in conflict if economic realities making sons valuable are not

[20] Report issued by Chinese State Council and Central Committee (January 2007).

addressed. This paper presents several pieces of evidence that one aspect driving son preference is the son's role as provider of support to his elderly parents. We find that parents who fail to produce a son are more likely to participate in old age pension programs and that the number of children is negatively related to pension program participation. We also find evidence that the implementation of the rural old age pension programs have had a mitigating effect on the sex ratios in the affected areas.

Policymakers should perhaps be more realistic about the current role of the son as the primary form of elderly care, and either increase the number of children allowed under the one child policy or provide parents with alternative means to support themselves in old age. Future policy should attempt to implement and financially support larger-scale pension schemes that would allow parents in smaller villages to participate. With more alternatives to sons as vehicles of old age support, perhaps parents will feel less compelled to have a son, and the sex ratios in China may return to more normal levels. Failure to make institutional and governmental attempts at change could lead to a further deterioration in the sex ratio at birth.

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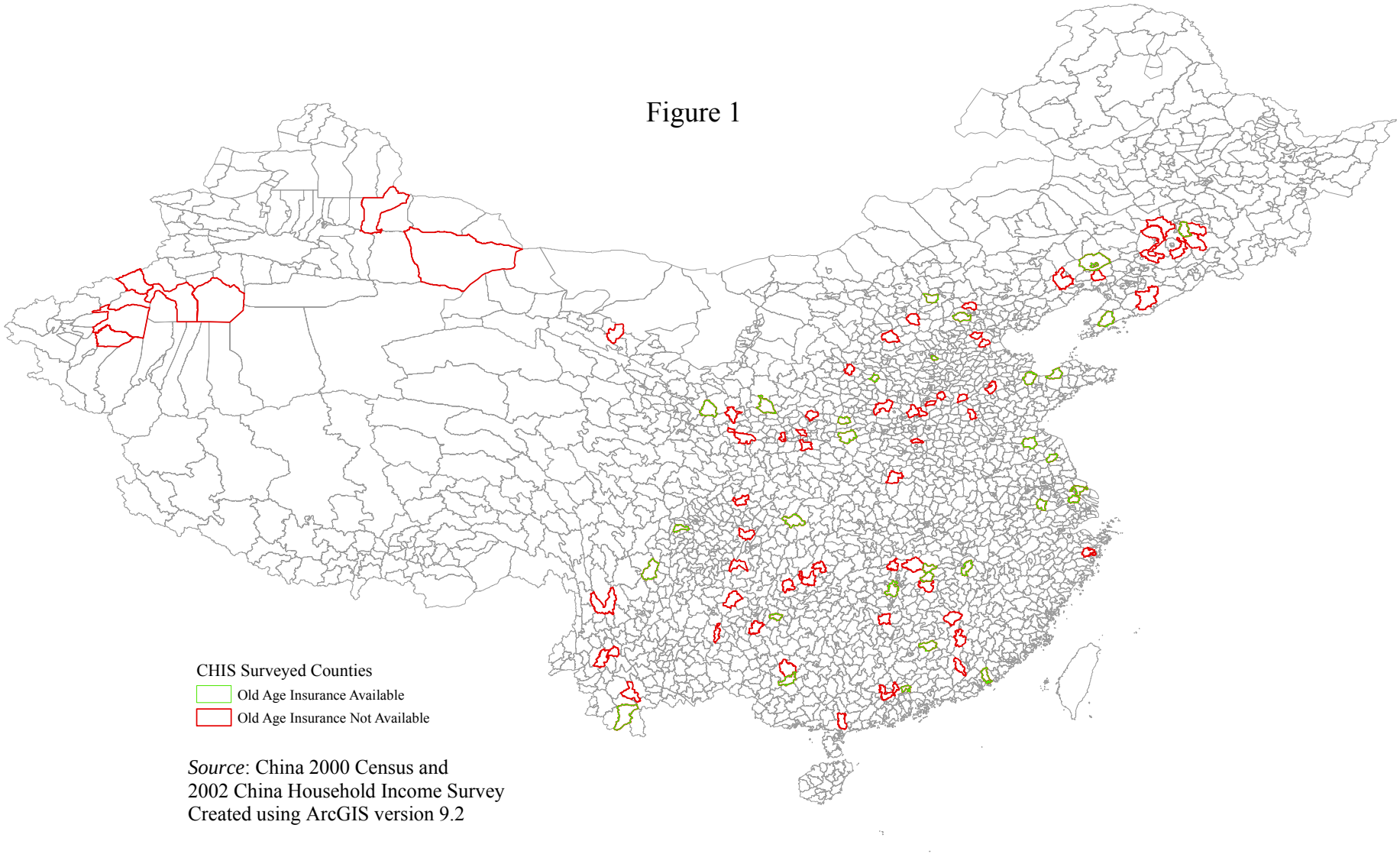
8 Appendix

The Missing Women Problem as a Rural-Driven Phenomena

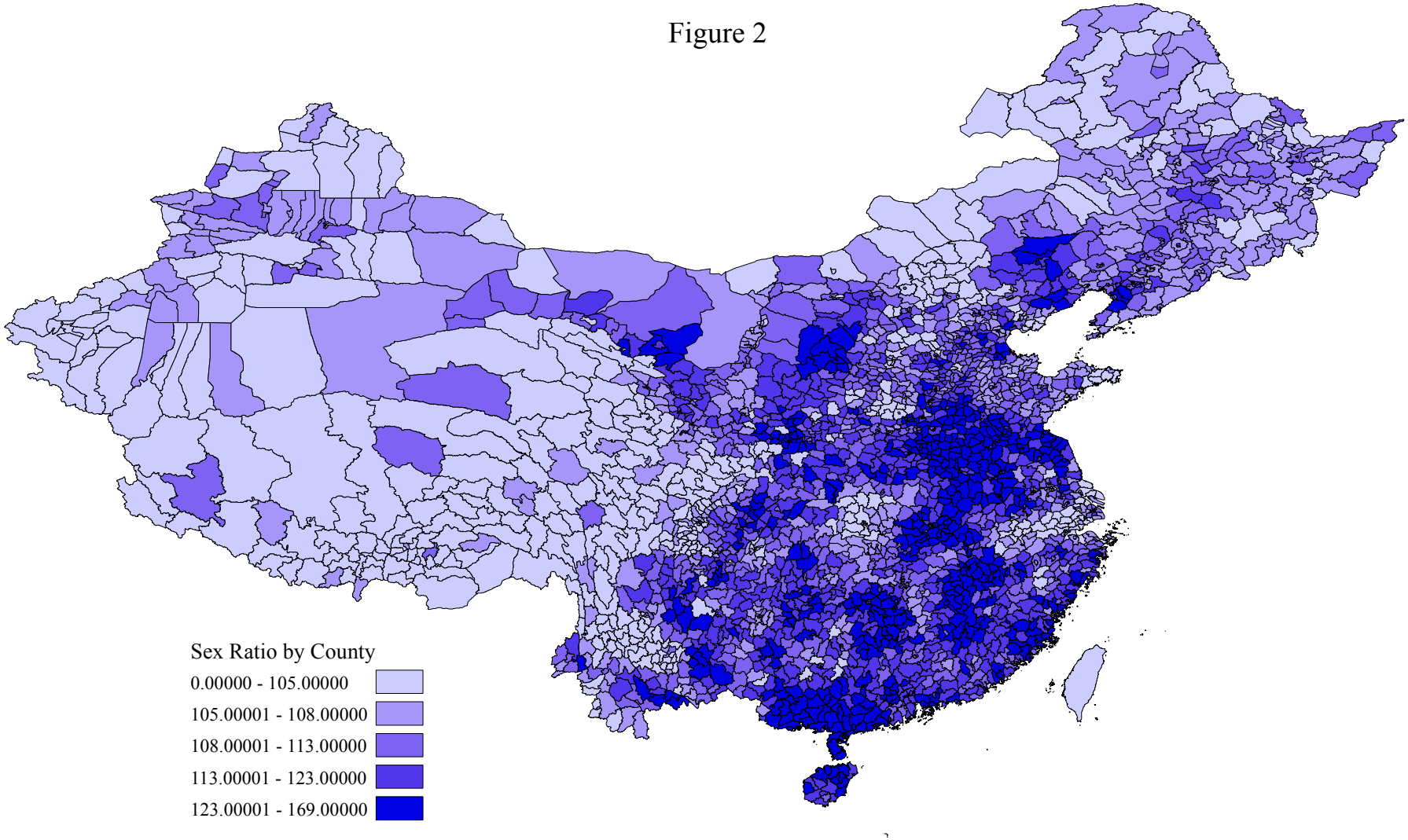
This paper claims that the high sex ratio in China is driven by patterns observed in rural China. A quick examination of the 2000 Census .10% micro level Data give insight into this phenomenon. Appendix Figure 1.1 shows the sex ratio by birth order and by the rural/urban divide. Appendix Figure 1.2 shows the proportion of families in either rural/urban areas that are producing higher order births. Though sex ratios by parity appear to be slightly worse in urban areas, appendix Figure 1.2 indicates that it is in the rural areas that most higher order births are occurring. The proportion of rural to urban families sampled was approximately .54 rural to .46 urban, the proportion of rural families that consist of more than one child was .78, as opposed to .41 of urban families. Thus, even if sex ratios are slightly higher among all birth orders in urban areas, the effect is overwhelmed by the small share, relative to rural areas, of urban families with multiple children.

Figure 1



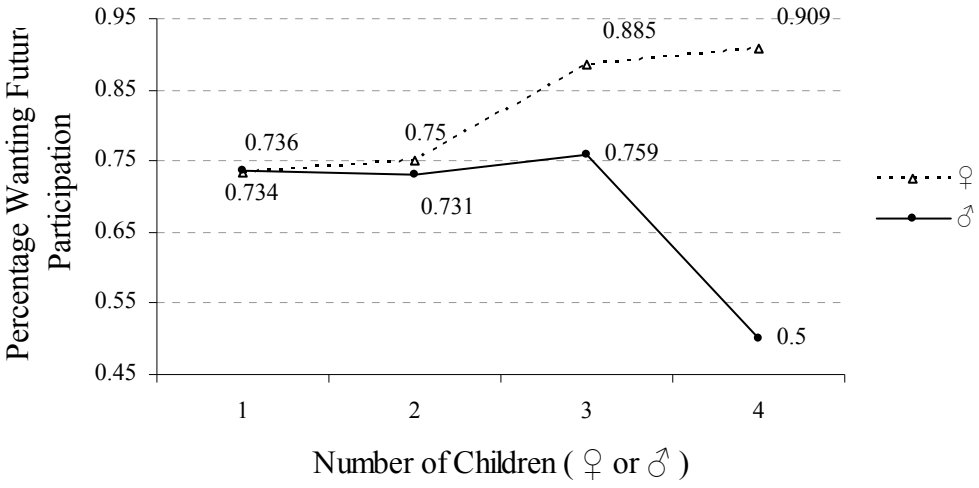
Source: China 2000 Census and
2002 China Household Income Survey
Created using ArcGIS version 9.2

Figure 2



Source: China 2000 Census
Created using ArcGIS version 9.2

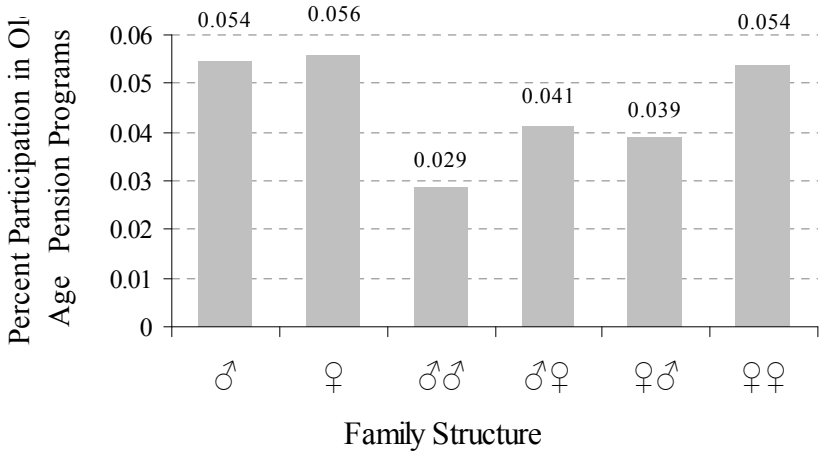
Figure 3



Source: China 2002 Household Income Survey

Notes: Sample limited to mothers matched with children, who have completed their fertility outcomes.

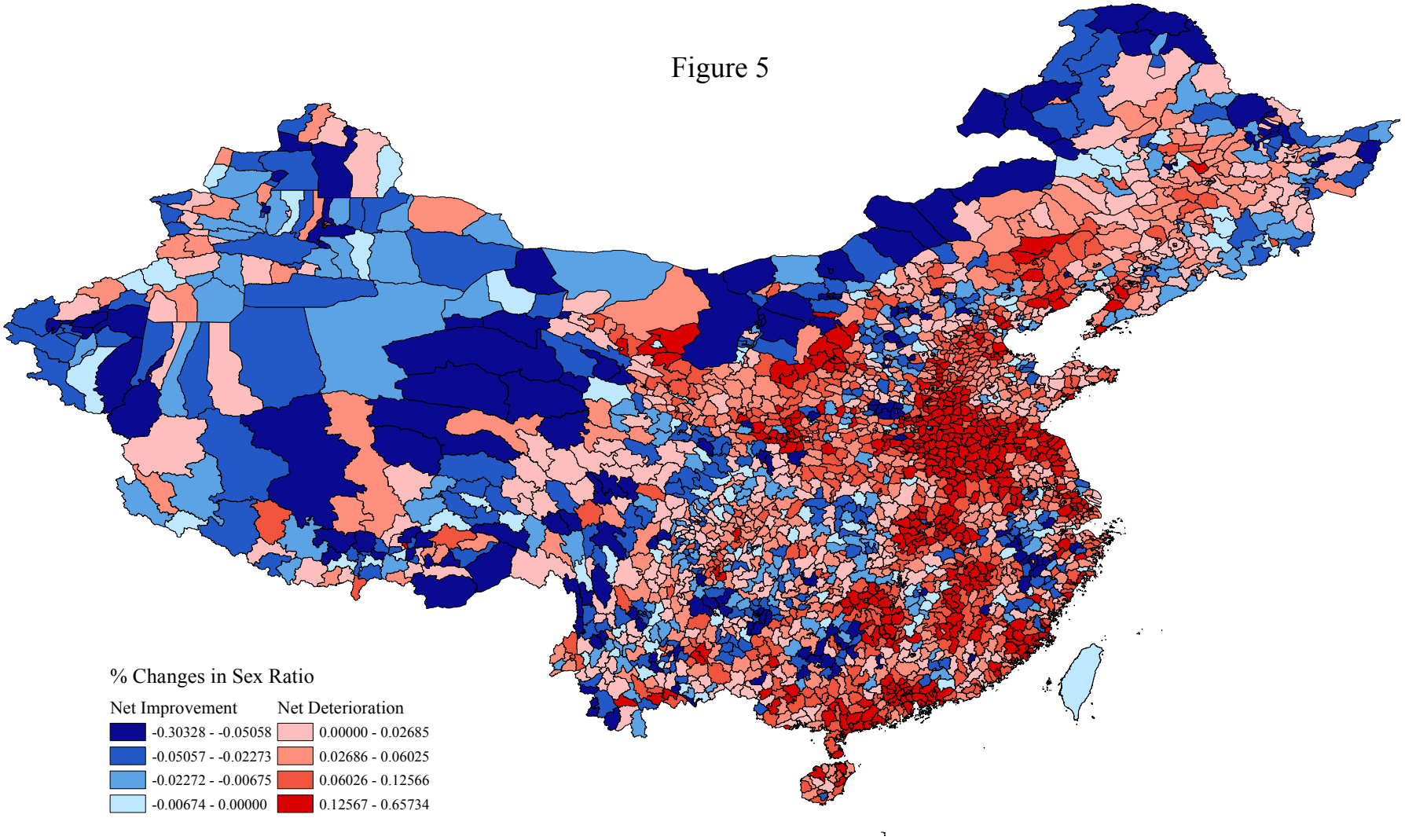
Figure 4



Source: China 2002 Household Income Survey

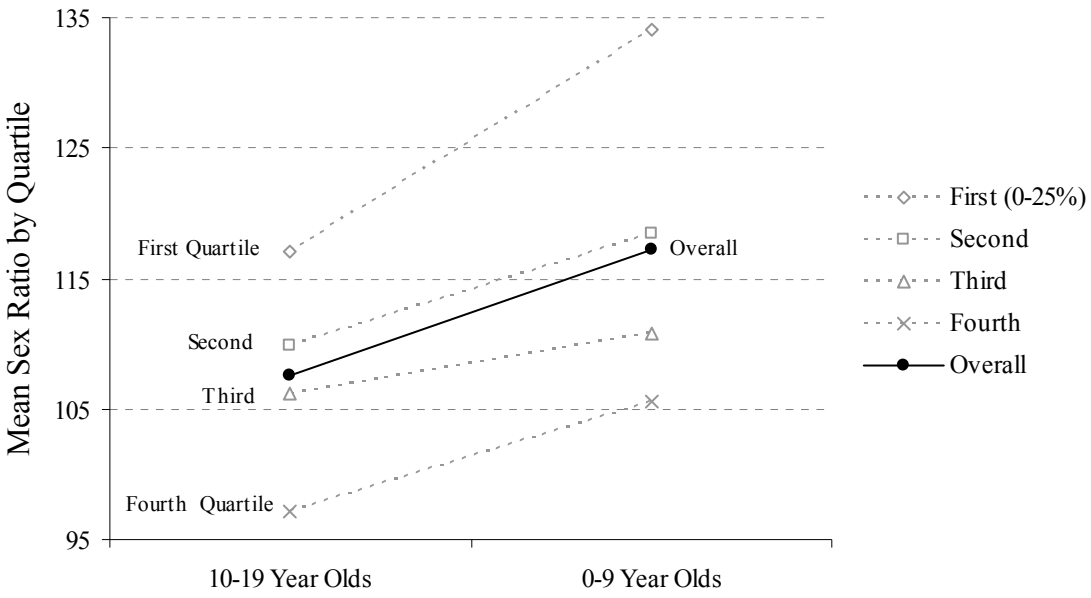
Notes: Sample limited to mothers matched with children, who have completed their fertility outcomes.

Figure 5



Source: China 2000 Census
Created using ArcGIS version 9.2

Figure 6



Source: China 2000 Historical Census Data and China 2000 Census

Note: Analytic weights of county population are assumed when calculating mean sex ratios.

Table 1. Individual level Descriptive Statistics of CHIS 2002 Data

Variables	Means
Age	33.041
Years Education	5.509
Income	1,016.80
Financial Assets	7,189.69
Savings for Old Age	2,530.66
Days Per Year Spent Planting	53.301
Old Age Program Participation	0.068
Observations	37,969

Source: China Household Income Survey (2002).

Notes: Income measured in RMB. Financial assets are provided as a measure of total savings. Savings for Old Age are those savings that are self reported as specifically for the purpose of old age security. Days planting are provided as a proxy for unreported agricultural income.

Table 2. Descriptive Statistics of China 2000 Census County Level Data and Aggregated CHIS Attributes

	Means		
Variables	All Counties	CHIS Counties without Pension	CHIS Counties with Pension
Population	425,800	555,817	738,406
Total Fertility Rate	1.310	1.464	1.292
Sex Ratio	114.1	115.3	114.0
Years Education	6.865	6.675	6.883
Share Urban	0.389	0.238	0.300
Share Manufacturing	0.115	0.064	0.106
Share Construction	0.027	0.020	0.030
CHIS Attributes			
Income		2193	2771
Health Insurance		0.094	0.168
Poverty Relief		0.234	0.253
Share of Cadre Households		0.029	0.022
Observations	2870	61	34

Source: China 2000 Census County Level Data, China 2002 Household Income Survey

Notes: CHIS counties are those counties from which at least one village was sampled by the CHIS. CHIS statistics reported are village level statistics aggregated to the county-level (weighted by village population). Sex Ratio is calculated for children aged 0-9. Share urban and share manufacturing refer to fraction of total employees that are employed in their respective sectors. Income is measured in RMB and is individual income from CHIS, collapsed to the county level. Health insurance and poverty represent the fraction of villages, in each county, that provide public health insurance and poverty relief programs. Share of cadre households is included as a proxy measure of the extent of communist party influence.

Table 3. Linear Probability Estimates of Fertility Outcomes and Pension Takeup

	Participation in Pension Program				
Variable	(1)	(2)	(3)	(4)	(5)
Failed to Have a Son	0.0114**	0.0138**			
	(0.0042)	(0.0043)			
Number of Sons			-0.0186**	-0.0247**	
			(0.0026)	(0.0029)	
Number of Daughters			-0.0152**	-0.0221**	
			(0.0023)	(0.0026)	
One Boy					-0.0120*
					(0.0053)
Two Boys					-0.0216**
					(0.0073)
One Girl					-0.0143*
					(0.0068)
Two Girls					-0.0144*
					(0.0068)
Number of Kids					-0.0195**
					(0.0035)
Constant	0.0358**	-0.0103	0.0655**	0.0145	0.0634**
	(0.0021)	(0.0097)	(0.0041)	(0.0098)	(0.0135)
Controls	No	Yes	No	Yes	Yes
Mean Participation	0.0390	0.0390	0.0390	0.0390	0.0390
Observations	11453	11057	11453	11057	11057
R^2	0.0007	0.0312	0.0042	0.0371	0.0393

* significant at the 5% level. ** significant at the 1% level

Source: China 2002 Household Income Survey matched mothers sample, restricted to mothers whose oldest child is younger than the age of 20 to prevent generational overlapping of mother matching.

Notes: Robust standard errors are listed in parentheses under coefficients. Failed to have a son is a binary with a score of 1 if fertility was completed and no son was born. Suppressed controls included are age, age squared, years of education, total household income, total financial assets, and total hours spent planting (as

Table 4. OLS Estimates of Fertility Outcomes and Savings for Old Age

	Log Savings for Old Age				
Variable	(1)	(2)	(3)	(4)	(5)
Failed to Have a Son	0.2266	0.1352			
	(0.2414)	(0.2557)			
Number of Sons			-0.4114*	-0.3772	
			(0.1974)	(0.2147)	
Number of Daughters			-0.2083	-0.2061	
			(0.1888)	(0.2113)	
One Boy					0.1441
					(0.3805)
Two Boys					-0.0524
					(0.5991)
One Girl					0.0892
					(0.3477)
Two Girls					0.3411
					(0.5916)
Number of Kids					-0.4092
					(0.2601)
Constant	7.5202**	6.8720**	8.0659**	7.1901**	8.0130**
	(0.1369)	(0.5701)	(0.2632)	(0.5559)	(0.7590)
Controls	No	Yes	No	Yes	Yes
Observations	541	510	541	510	510
R^2	0.0016	0.0592	0.0088	0.0654	0.0747

* significant at the 5% level. ** significant at the 1% level

Source: China 2002 Household Income Survey matched mothers sample, restricted to mothers whose oldest child is younger than the age of 20 to prevent generational overlapping of mother matching.

Notes: Robust standard errors are listed in parentheses under coefficients. Failed to have a son is a binary with a score of 1 if fertility was completed and no son was born. Controls included are age, age squared, years of education, total household income, total financial assets, and hours spent planting daily.

Table 5. Restricted- and Non Restricted- Sample Estimates for Robustness

	Participation in Pension Program			
Variable	(1)	(2)	(3)	(4)
Failed to Have a Son	0.0114**	0.0138**	0.0082	0.0100
	(0.0042)	(0.0043)	(0.0051)	(0.0052)
Age		-0.0015		0.0005
		(0.0009)		(0.0012)
Age^2		0.0000**		0.0000
		(0.0000)		(0.0000)
Years of Education		0.0041**		0.0052**
		(0.0008)		(0.0010)
Total Household Income		0.0000**		0.0000**
		(0.0000)		(0.0000)
Hours Spent Planting		-0.0000*		-0.0000
		(0.0000)		(0.0000)
Constant	0.0358**	-0.0025	0.0371**	-0.0491**
	(0.0021)	(0.0095)	(0.0024)	(0.0171)
Restricted Sample	N	N	Y	Y
Observations	11057	11057	7991	7991
R^2	0.0007	0.0116	0.0007	0.0312

* significant at the 5% level. ** significant at the 1% level

Source: China 2002 Household Income Survey matched mothers sample, restricted to mothers whose oldest child is younger than the age of 20 to prevent generational overlapping of mother matching.

Notes: Robust standard errors are listed in parentheses under coefficients. Failed to have a son is a binary with a score of 1 if fertility was completed and no son was born. Columns (3) and (4) estimate coefficients based on a further restricted sample of mothers whose first child was born before the 1991 program announcement date.

Table 6. First-Difference Regressions:
Percent Changes in Sex Ratio and Program Availability

	Percent Change in Sex Ratio			
Variable	(1)	(2)	(3)	(4)
Pension Availability	0.0250	0.0175	-0.0902**	-0.0810†
	(0.0889)	(0.0723)	(0.0394)	(0.0436)
Health Insurance Availability		0.0200		-0.0244
		(0.0539)		(0.0388)
Poverty Relief Availability		-0.0729**		-0.0399
		(0.0303)		(0.0345)
TFR			-0.0265	-0.0258
			(0.0204)	(0.0236)
Log Fine Rate			0.0199	0.0224
			(0.0134)	(0.0140)
Share Construction			0.4298**	0.4514**
			(0.1054)	(0.1712)
Share Manufacturing			-1.2306*	-1.1827**
			(0.5615)	(0.5670)
Constant	0.0648**	0.0781**	0.0822*	0.1034**
	(0.0097)	(0.0132)	(0.0324)	(0.0413)
Additional Controls	No	No	No	Yes
Observations	95	95	95	95
R^2	0.0027	0.0595	0.2123	0.2289

† significant at the 10% level.** significant at the 5% level. * significant at the 1% level.

Source: China 2000 Census County Level Data, China 2002 Household Income Survey

Notes: Robust standard errors listed in parentheses under coefficients. CHIS village-aggregated attributes include health insurance availability and poverty relief availability. Health insurance and poverty represent the fraction of villages in each county that provide public health insurance and poverty relief programs. Additional CHIS attributes included, but not reported, are village income and share of households that are cadre households. An additional county level control that is suppressed is years of education. Change in sex ratio calculated as percent change of sex ratio from 10 to 19 year olds to 0 to 9 year olds. Share urban and share manufacturing refer to fraction of total employees that are employed in the respective sectors.

Table 7. Differences in Means for Village Characteristics, by Rural Old Age Insurance Availability

Characteristic	*Insurance Available*	*Insurance Unavailable*	*Differences in Means*
	(1)	(2)	(1) - (2)
Yearly Income (RMB)	1767.006	968.486	798.52
	(1988.574)	(1152.82)	(.000)**
Age	35.586	33.066	2.52
	(4.790)	(4.234)	(.000)**
Proportion Male	0.522	0.532	-0.01
	(.049)	(.053)	(.169)
Years of Education	6.823	6.614	0.209
	(3.623)	(3.069)	(.145)
Electric Lighting (1=yes)	.999	.997	0.002
	(.036)	(.009)	(.738)
Running Water (1=yes)	.613	.554	.059
	(.451)	(.441)	(.009)**
Observations	66	895	

* significant at the 5% level. ** significant at the 1% level

Source: China 2002 Household Income Survey

Notes: Standard deviations are listed in parentheses under the means in columns 1 and 2. P-values of a difference-in-means test are listed under the differences in the (1) - (2) column. Income calculated in RMB.

Table 8. Correlations Between Availability of Different Social Insurance Programs

Program	*Old Age Pension Program*
Collective Health Insurance	0.0891 (.000)**
Poverty Relief Loans	0.0211 (.000)**

* significant at the 5% level. ** significant at the 1% level

Source: China 2002 Household Income Survey

Notes: Correlations given refer to OLS estimates of old age pension program availability at the village level on the listed programs.

Table 9. Unweighted and Weighted First-Difference Regressions: Pension Availability and Sex Ratio Changes

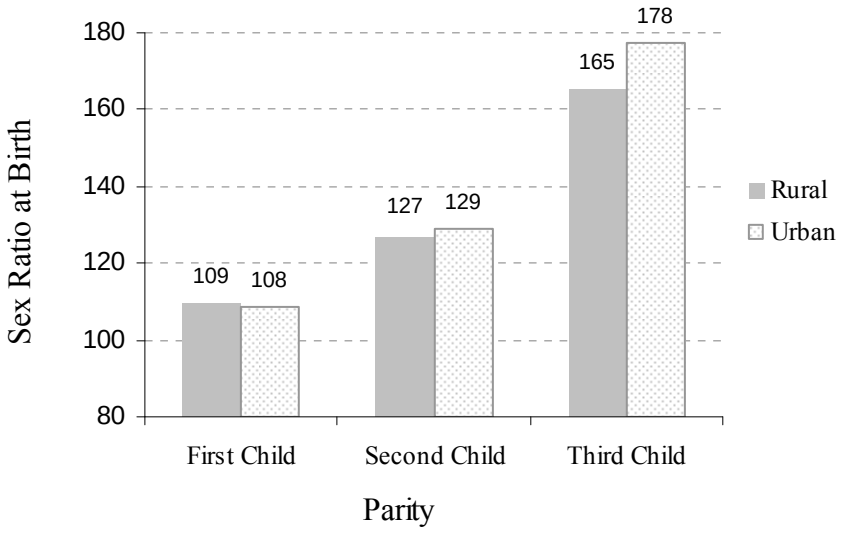
	Percent Change in Sex Ratio			
Variable	(1)	(2)	(3)	(4)
Pension Availability Binary	0.0098	-0.0136	0.0057	-0.0169
	(0.0214)	(0.0201)	(0.0215)	(0.0167)
Health Insurance Availability		-0.0376		-0.0430
		(0.0397)		(0.0335)
Poverty Relief Availability		-0.0484		-0.0532
		(0.0355)		(0.0340)
TFR		-0.0342		-0.0289
		(0.0304)		(0.0266)
Log Fine Rate		0.0228		0.0268
		(0.0144)		(0.0137)
Share Construction		-0.9554		-1.1591*
		(0.5515)		(0.4408)
Share Manufacturing		0.3869*		0.3874**
		(0.1639)		(0.1375)
Constant				
Propensity Score Weighted	No	No	Yes	Yes
Observations	95	95	95	95
R^2	0.0037	0.2178	0.0013	0.3277

* significant at the 5% level. ** significant at the 1% level

Source: China Household Income Survey 2002 and China 2000 Census

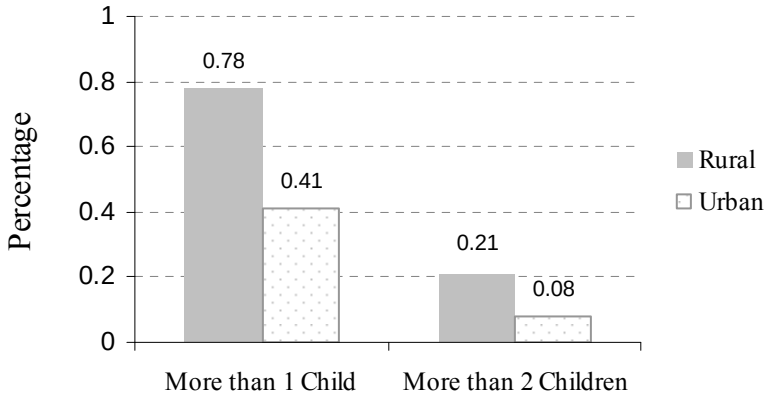
Notes: Robust standard errors are listed in parentheses under coefficients. First-stage logistic propensity score estimates include average educational attainment, share of labor that is manufacturing and construction, availability of health insurance, availability of poverty relief programs, and share of households that are communist cadres. Regressions (3) and (4) are inverse probability weighted to the propensity scores of each county,

Appendix Figure 1



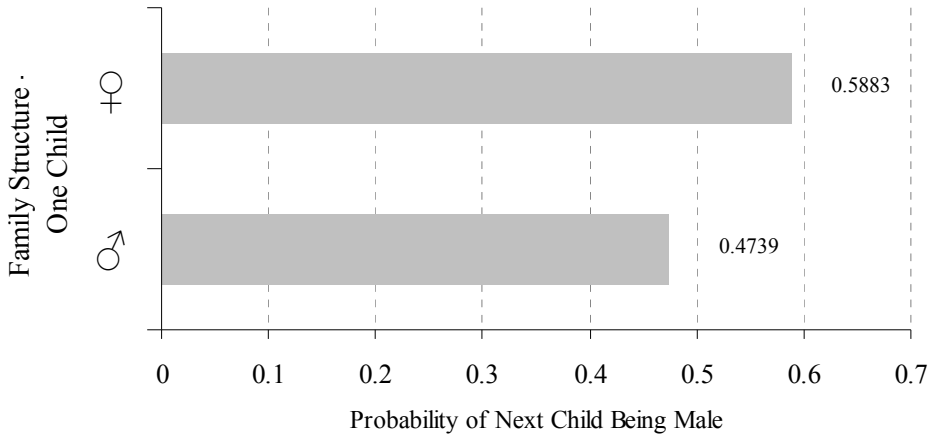
Source: China 2000 Census .1% Micro level data.

Appendix Figure 2

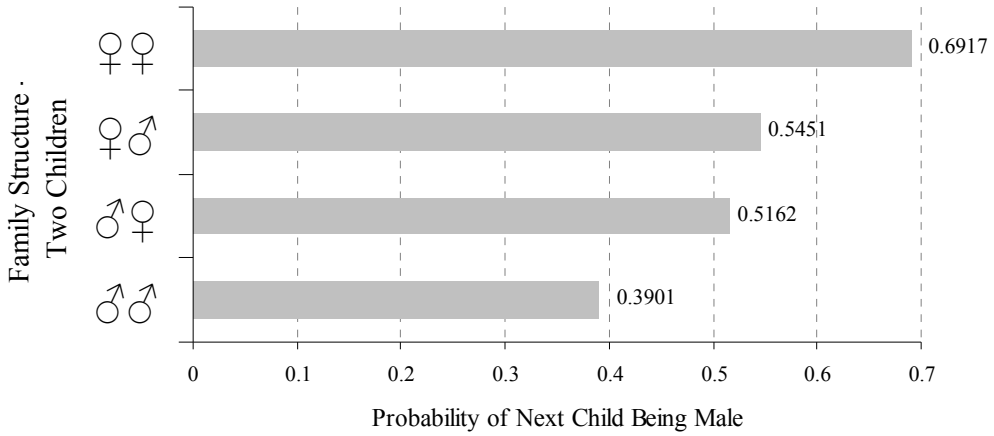


Source: China 2000 Census .1% Micro level data.

Appendix Figure 3.1

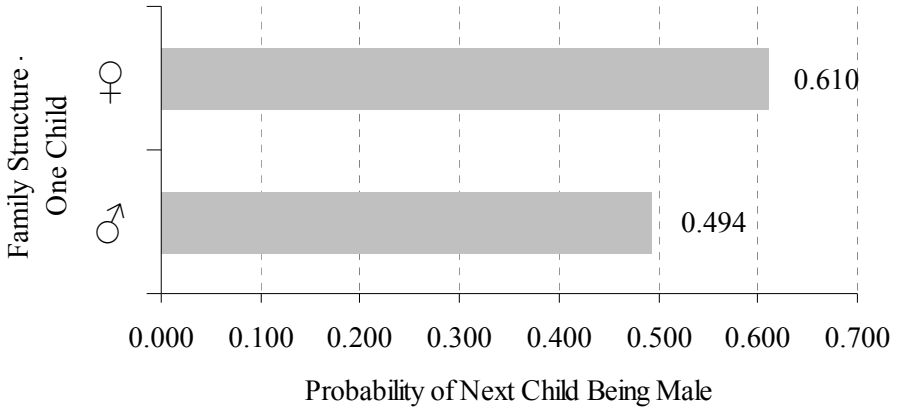


Appendix Figure 3.2

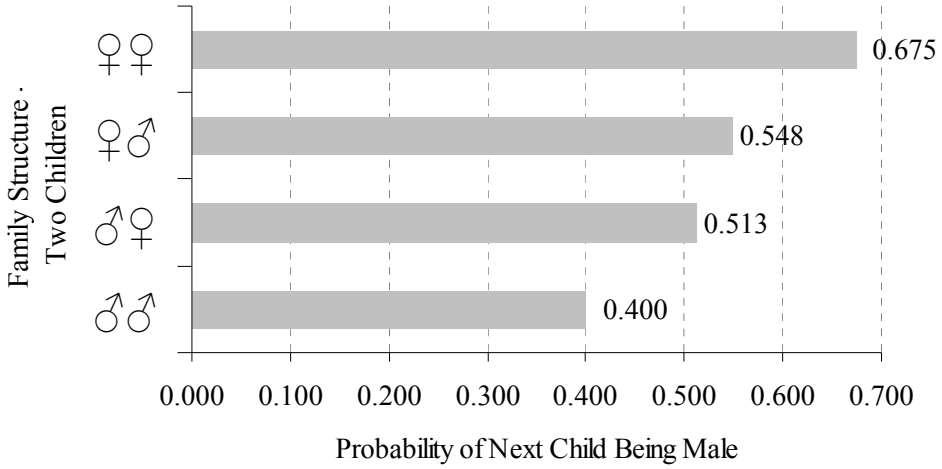


Source: China Household Income Survey 2002.

Appendix Figure 4.1



Appendix Figure 4.2



Source: China 2000 Census .1% Micro level data.

Appendix Table 1. OLS Estimates of Fertility Outcomes
and Non-log Savings for Old Age

	Savings for Old Age				
Variable	(1)	(2)	(3)	(4)	(5)
Failed to have a son	252.44	230.42			
	(267.94)	(276.70)			
Number of Sons			-29.13	73.03	
			(154.85)	(167.54)	
Number of Daughters			-44.25	-21.11	
			(167.03)	(192.18)	
One Boy					-76.61
					(300.26)
Two Boys					236.72
					(449.37)
One Girl					-306.68
					(303.99)
Two Girls					220.66
					(430.37)
Number of Kids					-1,067*
					(219.00)
Constant	1,366**	-762	3,849**	2,391**	2,652**
	(133.46)	(591.13)	(439.39)	(813.51)	(818.10)
Controls	No	Yes	No	Yes	Yes
Observations	3616	3507	3616	3507	3507
R^2	0.0003	0.0085	0.0121	0.0179	0.0187

* significant at the 5% level. ** significant at the 1% level

Source: China 2002 Household Income Survey matched mothers sample, restricted to mothers whose oldest child is younger than the age of 20 to prevent generational overlapping of mother matching.

Notes: Robust standard errors are listed in parentheses under coefficients. Failed to have a son is a binary with a score of 1 if fertility was completed and no son was born. Suppressed controls included are age, age squared, years of education, total household income, total financial assets, and total hours spent planting (as a proxy for agricultural income).