Deadly Discrimination: Implications of "Missing Girls" for Workplace Safety

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Abstract

In the absence of gender discrimination, millions of more girls would have been born and grown up in China, India, Vietnam, South Korea and other countries. The "missing girls" phenomenon itself is a consequence of gender discrimination, and has been well studied by Sen and others. In this paper, we look into an indirect but potentially deadly consequence of "missing girls" phenomenon. Due to unnaturally low ratio of girls to boys at birth, the subsequent shortage of brides causes many parents with sons of marriageable age to be more tolerant of unsafe workplace practices than they otherwise would. In response, employers may underinvest in workplace safety, which in turn could increase work-related mortality. Using four datasets and taking advantage of large regional and temporal variations in sex ratios in China, we demonstrate that in areas with a more severe shortage of marriage-age women, the parent cohort suffers a higher incidence of accidental injuries and death.

***Keywords*: sex ratio; accidental death; competitive pressure; risk taking**

JEL Classification: I1; J1; J2

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1. Introduction

In the absence of gender discrimination, millions of more girls would have been born and grown up in China, India, Vietnam, Korea and other countries. The "missing girls" phenomenon itself is a deadly consequence of gender discrimination, and has been well studied by Sen (1992) and others. In this paper, we look into an indirect but also potentially deadly consequence of "missing girls". Our hypothesis can be stated as follows: due to unnaturally low girls-to-boys ratio at birth, the subsequent shortage of brides causes many young parents with sons of marriageable age to be more tolerant of unsafe workplace practices than they otherwise would. In response, employers may underinvest in workplace safety, which in turn will cause more injuries and death at work places in general. This effect works through a general equilibrium channel. As far as we know, this is the first paper that investigates this effect.

Accidental death is one of the 10 leading reasons for mortality, claiming around 1.1 million lives across the world every year. The problem is particularly serious in China. For instance, Chinese coal mines are among the most deadly in the world. Although China contributes about 35 percent of world coal output, it accounts for 70 percent of deadly accidents in coal mines. In 2000, for every million tons of coal produced, 5.9 workers died in China. The mortality rate of coal production was as high as 195 times the level in the United States and nearly 13 times that in India (see Table 1.1).

The literature on accidental death focuses on either environmental causes (Howland and Hingson 1987a, 1987b; Blake, Morgan, and Bendall 1988; Myers et al. 1991; O'Loughlin et al. 1992; Fuller 2000) or personal factors (Agnew and Suruda 1993; Fuller 2000; Lord and Dayhew 2001; Dong, Wang, and Daw 2012). Few studies have examined the particular impact of competitive pressure on accidental death, although there is a debate in the medical literature as to whether stress, often as a result of competitive pressure, affects people's health outcomes. In general, the findings are mixed and none of the studies traces the causes of accidental deaths to sex ratio imbalance.

Aro and Hasan (1987) find that stressors at work are associated with mental-stress symptoms among blue-collar workers, and Stansfeld, Fuhrer, and Shipley (1998) show that psychological morbidity can be attributable to work-related stressors among civil servants as well. In both studies, the health outcome is self-reported. However, there are likely some confounding factors at play. For instance, some people may have certain unobserved personality traits, such as negative affectivity or neuroticism, making them more likely to report both work-related stress and adverse health outcome variables (Watson 1984; Burke and Brief 1990; Brett, Brief, and George 1993). When those individuals are excluded from the sample, House et al. (1979) reveal that the links between stress and dermatological symptoms diminish.

When measuring the health outcome not by self-reporting but based on physicians' observations or laboratory tests, or both, however, House et al. (1979) fail to detect a positive relationship between work-related stress and health outcomes. Without a clear measure of stressful environment independent of idiosyncratic factors, it would be difficult to tease out the causal impact of work-related stress on health outcomes.

Another flaw of such studies is that only the health outcomes of living people are observed. If an excess mortality occurs among those overstressed, then the aforementioned studies on the living population may be subject to a selection bias. In this paper, we do not debate whether work-related stress leads to health outcomes of those who are alive. Instead, we investigate whether exposure to a competitive environment - induced by the "missing girls" phenomenon - contributes to excess mortality, an extreme form of adverse health outcome.

One methodological challenge is to measure the exogenous competitive pressure. China's rising gender imbalance, which varies greatly over time and across regions, provides us with a natural experiment with which to tackle this issue.[1] With looming marriage likelihood owing to rising sex ratio imbalances (greater numbers of excess men in the marriage market), families with sons must accumulate more wealth in order to attract potential brides for their sons (Wei and Zhang, 2011b). However, the desire to accumulate wealth may lead people to work hard and undertake risky activities. Moreover, if more potential employees are willing to tolerate dangerous work conditions, employers may respond by underinvesting in workplace safety. This

[1] China's rising sex ratio imbalance is caused by a combination of one-child policy introduced in the late 1970s, the spread of ultrasound gender-identification technology and son preference culture (Li and Zheng, 2009; Bulte, Heerink, and Zhang, 2011). Until recently, China adopted a very stringent one-child policy. Couples were only allowed to have one child except that both parents were only children; parents were from officially recognized ethnic minorities, or the first child was a girl or disabled in rural areas. The sex ratio at birth increased from a normal level in the late 1980s to about 120 (boys per 100 girls) in 2010. In 2013, the Chinese government loosened the one-child policy by allowing married couples in which just one parent is an only child to have two children.

is a form of negative spillover from the behavior of parents with an unmarried son to the safety environment of all people in the same workplace. Hence, increasing competitive pressures may result in higher rates of accidental death in the local population. In this paper, we provide a theoretical model to illustrate above ideas and empirically test them. Since our focus is mainly on the empirical parts, we put the theoretical model into the appendix.

We also contribute to the literature on the impacts of sex ratio imbalance. Sex ratio studies cover a wide range of topics, such as sexually transmitted infections (Tucker et al. 2005; South and Trent 2010), marriage and labor markets (Becker 1973; Grossbard-Shechtman 1984; Chiappori, Fortin, and Lacroix 2001; Angrist 2002), crime (Edlund et al. 2007), savings (Wei and Zhang 2011a), current account imbalance (Du and Wei 2010), entrepreneurship (Wei and Zhang 2011b), and housing prices (Wei, Zhang, and Liu 2012), to name a few. To our knowledge, the negative impact on mortality has not been mentioned in the literature.

Our empirical tests are carried out in two steps. First, using disability data from the China National Survey on Disabled People (CNSDP), mortality data from China's National Disease Surveillance Points System dataset (NDSPS), and sex ratio variables inferred from China's population census, we establish that the disability rate due to work injury and the death rate owing to fires and flames and accidental falls are significantly higher in regions with a more skewed premarital sex ratio (a greater relative shortage of young women). Moreover, the effect is stronger for males than for females and it mainly operates through the cohort of 25 to 44 years old, or the parents' cohort.

Second, we explore the mechanisms underlying how sex ratio affects accidental death rates from both the employee and employer perspective. In terms of employees, using the 2010 China Family Panel Studies (CFPS), a national survey about economic and social situations at the individual and family level, we find that parents with unmarried sons in regions with more men than women in the premarital cohort are more likely to work outside their hometowns, put in longer hours, and take higher-risk jobs. Families with unmarried daughters do not exhibit a similar tendency. Since working extended hours and participating in dangerous jobs are risk factors of accidental death, we interpret these patterns as evidence that the "missing girls" phenomenon contributes to a higher accidental death rate.

On the employer side, we investigate whether employers respond to the weak bargaining power of potential employees who are more willing to endure high-risk jobs by compromising workplace safety. A testable implication is that this should occur more in regions with a higher ratio of young men to young women. We first show that the proportion of trade unions[2] setting up labor protection, supervision, and inspection committees (LPSIC) in charge of workplace safety is significantly lower in regions with higher male-to-female sex ratios. Next, we examine the impacts of imbalanced sex ratios on work-related injury insurance coverage at the firm and county level, a critical facet of employers' investment in workplace safety, using a survey by the All China Federation of Trade Unions. The empirical results demonstrate that the coverage ratio of work-related injury insurance is significantly lower in regions with a more skewed sex ratio (more young men than young women).

When we divide the sample into two groups using the median of sex ratio as the cutoff, we find that the difference in the sex ratio between the two groups explains about 78.1 percent and 67.2 percent of the between-group difference in accidental death incidence due to fires and accidental falls, respectively. An increase in one standard deviation of the sex ratio would elevate the accidental death rate by 0.2 standard deviation. Therefore, the sex ratio imbalance plays an economically important role in explaining accidental deaths.

The rest of the paper is organized as follows. In Section 2, we review the literature and connect it to our paper. In Sections 3 and 4, we introduce our data and methodology. Section 5 presents the baseline empirical results and robustness checks. Underlying mechanisms are explored in Section 6. Section 7 concludes.

[2] As of 2011, 90 percent of firms in China had established trade unions. The data about trade union committees and their functions are more accessible than those about firms.

2. The Connection between the Hypothesis and the Existing Literature

Our research is connected to several strands of literature. Each is too vast to be reviewed comprehensively here. Instead, we select a few to highlight the most relevant topics for our empirical investigation.

There is an extensive body of literature discussing the impact of sex ratio on marriage and labor markets. If men outnumber women, the prospects of finding a mate are more promising for a woman than for a man (Angrist 2002); consequently women gain greater bargaining power (Chiappori, Fortin, and Lacroix 2001) in the marriage market. Becker (1973) develops a model illustrating how sex ratios shape marriage rates and family income. Grossbard-Shechtman (1984) empirically shows that increasing the sex ratio boosts demand for female labor and increases women's shadow wages of home production, thereby lowering their labor participation rate. Angrist (2002) also finds that decreases their labor participation rate thanks to their increased bargaining power.

The economic impact of the sex ratio goes beyond the labor market. An imbalanced sex ratio towards more males than females induces people to save more (Wei and Zhang 2011a), stimulates entrepreneurship (Wei and Zhang 2011b), generates a current account imbalance (Du and Wei 2010), and boosts housing prices (Wei, Zhang, and Liu 2012), to name a few. A rising gender imbalance may also contribute to many social ills. For instance, Edlund et al. (2007) reveal that increasing sex ratios may account for up to one-seventh of the overall rise in crime in China. By comparison, studies on the health impact of sex ratios are limited. A few existing studies focus on the effect of imbalanced sex ratios on sexually transmitted infections (Tucker et al. 2005; South and Trent 2010), a relatively direct link through the channel of sexual behavior. In this paper, we investigate the negative effect of marriage market squeeze, measured by local sex ratios in the premarital cohort, on accidental deaths.

The literature on accidental deaths primarily focuses on social-demographic factors and personal characteristics. For instance, Howland and Hingson (1987a, 1987b) list alcohol as a risk factor for injury or accidental death due to falls, fires, and burns. Most studies on accidental falls concentrate on the elderly, while a small proportion focuses on the working-age population. Blake, Morgan, and Bendall (1988) explore the effects of medical and anthropometric variables on falls. Myers et al. (1991) examine the role of social-demographic factors, functional status, and medications in falls. O'Loughlin et al. (1992) expand the list by including physical activity, alcohol consumption, acute and chronic health problems, dizziness, mobility, and medications on accidental falls. Aging, medication use, cognitive impairment, and sensory deficits are mentioned as the major risk factors of accidental falls for the elderly (Fuller 2000). Lord and Dayhew (2001) list visual impairment as another major risk factor.

Some studies investigated work-related falls. For instance, Agnew and Suruda (1993) find that fatality rates from falls begin to rise for workers aged 45 to 54, earlier than for other work-related death causes. Dong, Wang, and Daw (2012) show that in the US construction industry, older workers (older than 55 years) have a higher fatality rate owing to accidental falls than young workers (16 to 54 years). None of the studies on accidental deaths mentions competitive pressure in general or rising sex ratios in particular as a risk factor.

3. Data

We chose China for our empirical analysis for the following reasons. First, nowhere else has a more skewed sex ratio than China today, where a strict family planning policy has restricted the number of children most families can have to one or two, thereby greatly reinforcing the incentives for sex-selective abortions. Second, the family planning policy is unevenly enforced in different regions of China, resulting in regional variations in the sex ratio. The premarital (5–19 age cohort) sex ratio ranges from 100 boys per 100 girls in some places to 126 boys per 100 girls in other regions in our sample. Moreover, a within-country study has advantages over cross-country studies as the institutions are more similar across regions within a country than across countries. Furthermore, due to government restrictions on household registrations, internal mobility for the purpose of marriage is low in China. Nearly 90 percent of marriages take place between men and women from the same county (Wei and Zhang 2011a). This makes the marriage market highly localized. Therefore, the premarital sex ratio at the county level can serve as a good proxy for the measurement of local marriage market squeeze.

Sex Ratios and Family Planning Policy Variables

We use the 1990 census data to infer predetermined sex ratios. Using the 2000 population census data may lead to the problem of reverse causality because the difference in gender-specific mortality rates across age cohorts in previous years (1991–2000) would affect sex ratios observed in 2000.

Since our data structure is cross-sectional, we first carry out the analysis for 1991 and then use the pooled sample from 1991 to 2000. In the pooled data, since the population census is available only at the cohort level (with a five-year interval), it is impossible to control the precise premarital cohort in each year. Our approximation is that in the 1991–1995 period, the sex ratio is calculated as the ratio of males to females aged 5 to 19 in the 1990 population census, while for the period 1996–2000, the sex ratio is obtained from the 0-to-14 age cohort in the 1990 census.[3] Although we use a dummy to differentiate the sample between the period 1996–2000 and the period 1991–1995, measurement errors may occur during this process. The use of the two-stage least squares (2SLS) method is a partial solution to such a problem and other omitted-variable problems. Our choice of instrumental variables (IVs), motivated by the family planning policy in China, is as follows.

A key determinant of the sex ratio imbalance is a strict family planning policy introduced at the beginning of the 1980s. Whereas the goals of family planning are national, the enforcement is local. We employ three determinants of local sex ratio that are unlikely to be correlated with the omitted factors such as personal lifestyle and environmental factors, which may contribute to accidental death, as instrument variables (IVs). First, Ebenstein (2011) 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 third or fourth child in a one-child zone, or the fourth or fifth child in a two-child zone. Such a nonlinear financial penalty scheme was introduced by different local governments in different years (if at all), thereby varying across regions and over time. We call this variable "premium to higher-order birth". Second, we use the amount of the one-child bonus, the amount of reward going to a family if the family abides by the family planning policy, as an instrument. Such a bonus also varies across counties and over time. Third, most of the ethnic minority groups in China do not face a strict family planning policy (the government allowed this exemption possibly to avoid being criticized for marginalizing minority groups with the use of a one-child policy). Since the distribution of non-Han Chinese is not uniform across regions, we can employ such variations (share of minority population) as another possible IV. Our IV data start from 1979,[4] and a person with the age of 5 to 19 or 0 to 14 in 1990 was born during 1971–1985 or 1976–1990. To fully capture the effects of the family planning policy, our IVs take the average value of 1979 to 1985 in wave 1 and the average value of 1979 to 1990 in wave 2.

[3] When analyzing the CFPS 2010 survey, we use the sex ratio of 10-to-19-year-olds inferred from China's 2000 population census. Because the cohort data at the county level for the 2010 census is not available, we have to compute the sex ratio variable using the 2000 census. The youngest cohort in 2000 reached the age of 10 by 2010. So we cannot use the same sex ratio of 5-to-19-year-olds as in other analyses. All the major conclusions hold if we use the sex ratio of 10-to-19-year-olds instead of 5-to-19-year-olds in analyses based on other data sources.

[4] Some provinces put forward family planning as early as 1979, although the policy was formulated as the basic state policy in 1982.

Work-Related Disability

The individual-level disability data used in this paper come from the China National Survey on Disabled People (CNSDP), conducted jointly in 2006 by the National Bureau of Statistics, the Ministry of Civil Affairs, and the China Disabled Persons' Federation. The survey uses probabilistic proportional sampling and covers all the 31 provinces, autonomous regions, and super-municipalities in China.

The survey includes individuals' basic information and their economic, social, and disability status. Basic information refers to an individual's age, gender, family, and address. Economic and social status is reflected by school attainment, personal income, and marital status. Disability information is more detailed, consisting of disability status, pathogenic causes and degree of severity. All of the disabilities are classified into five categories: visual, hearing and speech, cognitive, mental, and physical. The key variable of interest for this research is physical disability. There are roughly 21 pathogenic causes, including work-related injuries, cerebral palsy, development deformity, dwarfism, other congenital abnormalities, and so on. We primarily examine physical disability caused by work-related injuries.

Mortality

The death data come from China's National Disease Surveillance Points System (NDSPS), which includes exhaustive information about roughly 500,000 deaths due to different causes in 145 sites of 30 provinces in China during the 1991–2000 period. The 145 sites are chosen to be nationally representative. The NDSPS records all the deaths and population counts in a site. Several studies (for example, Ebenstein 2012; Chen et al. 2013) use this dataset. Because an urban sample may include migrants whose places of origin are not known, we focus on the rural sites of the NDSPS.

The measurement error of a location's mortality owing to migration is not a big challenge to our results. First, the NDSPS assigns a death to an individual's *hukou* (household registration system in Chinese) or place of birth. If a death occurs away from the hometown, a family is instructed to report the death as wherever it happens and a site official is expected to accurately record the death at the site of the deceased's original registration. Furthermore, the *hukou* system requires that the death certificate report the *hukou* or place of birth, and that information is checked by a local official who provides an official stamp of verification to the death certificate. Second, according to Meng (2012), migration did not become common until the end of the 1990s and the direction of migration is from inland to coastal areas and from rural to urban areas. To address the concern of outmigration, we also restrict our sample to the early period of the 1990s and exclude the largest destination provinces of migration as robustness checks.

The NDSPS dataset does not have direct measurement for work-related deaths. We mainly use deaths caused by fires, flames and accidental falls to proxy work-related deaths. In order to check if these types of death to a large extent represent work-related deaths, we gather the records of work-related deaths from the website of the State Administration of Work Safety[5] aggregate them to the provincial level and compare them with the number of deaths due to fires, flames, and accidental falls at the provincial level reported in NDSPS. The correlation coefficient between the two series in 2000[6] is 0.52, significant at the 5 percent level. In comparison, the correlation coefficient of deaths caused by accidental poisoning with the aforementioned work-related deaths is as low as 0.09 and not statistically significant. So we use death owing to accidental poisoning as a placebo in our analyses.

Edlund et. al. (2013) find that China's sex ratio imbalances are an important driver of crimes. So there is a possibility that fire-related deaths are due to arsons induced by high sex ratios rather than people's hard work and rick-taking behaviors. The official crime statistics does not list arson in a separate category because it rarely occurs. So we cannot directly test the relationship between arsons and fire-related deaths. According to various issues (1982-2005) of *China Law Yearbooks*, the share of crimes in "other category" to which arsons belong ranged between nine percent and thirteen percent in the period of 1981-2004. It is likely that the incidence of arsons is too small to make a difference to the total number of fire-related deaths.[7]

[5] http://media.chinasafety.gov.cn:8090/iSystem/shigumain.jsp. For instance, a record from the website looks like this "At around twelve o'clock on June 8th, the scaffold of Longshan Coke Company in the city of Xuzhou (Jiangsu province) collapsed and three men died."

[6] It is the first year that the record from State Administration of Work Safety is available.

[7] As a robustness check, we also control for the crime rate on the right hand side and redo the baseline analysis. Crime rate refers to violent and property crime rate (arrests/10,000 population) by province and year.

Working outside the Hometown and Risky Jobs

The data on working outside the hometown (probability and time) and participation in risky jobs are from the CFPS 2010, a large-scale survey of 14,798 families and 33,600 adults in 25 Chinese provinces in 2010. It was funded by the 985 Program of Peking University – a special government research grant to the university - and carried out by the Institute of Social Science Survey of Peking University. It contains specific information not only about family structure and the demographics of individuals, but also about working outside the hometown and career choices. In the survey, people were asked, “Did any family member work outside hometown last year?”, “What is the total number of months of working outside hometown last year for all family members?”, “What is your job?”, and “What does your working place look like?”. We generate dependent variables according to those questions. Since working outside the hometown and taking risky jobs is more common for rural people, we concentrate on the rural sample of the CFPS 2010.

Employers' Investment in Workplace Safety

Work-related injury insurance is an important aspect of employers' investment in workplace safety. The Regulations on Work-Related Injury Insurance in China, implemented in 2004, stipulate that all the enterprises in China pay for the premium of work-related injury insurance (Article 2) for their employees. However, in practice, many firms do not purchase such insurance for all employees. Using a survey implemented by the All China Federation of Trade Unions in 2009, which covers firms in Beijing, Shanghai, and Tianjin and the provinces of Liaoning, Jiangsu, Zhejiang, Guangdong, Hebei, Henan, Hubei, Sichuan, and Shaanxi and includes detailed information about work-related injury insurance coverage and situations of unions, we investigate the impacts of imbalanced sex ratios on employers' provisions of work-related injury insurance. The survey asks firms about the number of workers covered by work-related injury insurance, the share of union members, female union members, rural union members, female employees, rural employees in total employees, and the share of grass-roots employees in the board of directors and supervisors. To obtain firm-level financial indicators, such as sales, operating profits, fixed assets, short-term debt, and three expenses (operating expense, administrative expense and financial expense), we match the firms to the Industrial Firms Census Dataset compiled by the National Bureau of Statistics.[8]

[8] The survey by All China Federation of Trade Unions does not include detailed firm operational information. We match the firms in the survey with the Industrial Firms Census Dataset to obtain such additional information.

4. Methodology

In the analysis of sex ratio imbalance and accidental death or disability caused by work-related injury at the individual level, our baseline specification is

$$y_i = \beta_1 sex\,ratio_i + \beta_2 I_i + \beta_3 C_i + \varepsilon_i . \quad (1)$$

In equation (1), our dependent variable y_i is a dummy, indicating whether the person has died from fire or fall accidents or has been disabled by work-related injuries. The sex ratio of the premarital cohort is the variable of interest. I_i are individual-level characteristics. The NDSPS and disability survey datasets contain individual information about gender, age, ethnicity, marital status, years of schooling, and occupation, allowing us to control for the effects of demographic factors on death and injury. C_i are other county-level controls, such as total population of relevant cohorts, share of prime-age population, and dummy variables indicating the eastern and middle region with the western region as the base group. The share of prime-age population (15–64) in total population from the 1990 population census is calculated for the accidental death analysis and for the disability analysis from the 2000 population census. For our baseline regressions, we use a linear probability model and employ ordinary least squares (OLS) to estimate equation (1).[9]

The NDSPS has a drawback in that it lacks comprehensive information about the deceased's living style (such as drinking and smoking behavior, exercise habits, and so on), medication usage, and family health history, and so forth. So omitted variables are a potential problem. We exploit three methods to deal with the problem.

First, we use placebo tests. Since our transmission channel is that parents work hard and take risks to help their sons win a competitive edge in the marriage market, the effects should mainly operate through the parents' cohort with unmarried children. We test whether the impacts are significant only in the parents' cohort but not in older cohorts. In addition, we use accidents and diseases that are unrelated with hard work, such as accidental poisoning and death due to congenital anomalies, as placebo tests.

[9] We also use probit and logit models as robustness checks, and the results are similar for sex ratio. Due to space limitations, we do not report these results separately.

Second, we construct a pseudo cohort-region-wave panel by calculating the death incidence of accidents in each age cohort during the 1991–2000 period. We use the corresponding total cohort population (each cohort includes five years) in a county to compute the death rates owing to a certain cause during 1991–1995 (wave 1) and 1996–2000 (wave 2). To be concrete, the specification is shown in equation (2). y_{ciw} represents the death rate of cohort *c* in surveillance spot *i* of wave *w*, and *sex ratio* is corresponding to the surveillance spot *i*. Therefore, we can control for cohort, region, and wave dummies (fixed effects) α_c , μ_r , and η_w in the pseudo-panel analysis.

$$y_{ciw} = \beta sexratio_i + \alpha_c + \mu_r + \eta_w + \varepsilon_{ciw}. \tag{2}$$

Third, we employ the IV approach in the pseudo-panel regressions. The three instruments include a dummy for the existence of extra fines for violations at higher-order births, the amount of one-child bonus that a family receives if the family abides by the family planning policy, and the share of minority population who are less subject to the strict family planning policy.

Another weakness of the NDSPS dataset is that we are unable to obtain information about whether the dead had a son or a daughter, so it is impossible to contrast between families with unmarried sons and families with unmarried daughters. To remedy this concern, we use the 2010 CFPS as a complement to delve into the underlying transmission mechanisms. Our explanatory variable is the sex ratio of the premarital age cohort. We can infer the sex ratio of the 10-to-19-year-old age cohort in 2010 from the 0-to-9-year-old age cohort in China's 2000 population census.

In the analysis of likelihood of working outside one's hometown, we use a probit model. The dependent variable is a dummy indicating whether the family has any members working outside the hometown. The control variables include log of family income per person, family size, household head's age, years of schooling, square of years of schooling, gender, ethnicity, poor health dummy, number of children, a dummy for whether the household has a child between the ages of 10 and 14, whether the household has a child between the ages of 15 and 19, and a dummy for the gender of the first child. To explore the difference between families with at least one son and families having only daughters, we examine these two groups separately.

In exploring the time spent working outside of the hometown, the dependent variable is the total number of months spent working outside the hometown in the last year for all family members. Since the lower limit of the dependent variable is censored at 0, we use a Tobit model. Control variables are the same as those in the analysis of the probability of working outside the hometown.

We exploit probit models to analyze the impact of sex ratios on participation in risky jobs. Our dependent variable equals 1 if any family member has a dangerous job or the workplace is in the open air and equals 0 otherwise. Dangerous jobs encompass firefighting, public security, mineral mining, well drilling, oil and gas exploration, construction, electricity supply, installation and debugging of electrical equipment, and operation of mechanical equipment. Control variables are the same as those in the analysis of working outside the hometown. Since urban people take part in risky jobs much less frequently than rural residents, we mainly focus on the rural sample.

In the analysis of employers' investment in workplace safety, we first investigate the association between sex ratio and the share of trade unions that have set up LPSICs, whose data are available at the province level. The key responsibility of such committees is to participate in the formulation and implementation of workplace safety measures. They convey workers' concerns and suggestions about safety and health in the workplace to upper management with the purpose of improving working conditions. We include only the provinces covered by the CFPS 2010 for ease of comparison.[10] The sex ratio of the 10-to-19-year-old cohort at the province level is calculated from the 2010 population census. All data about trade unions and committees are from the *Chinese Trade Unions Statistics Yearbook 2011.*

We then examine the impacts of sex ratio on work-related injury insurance coverage, an essential part of employers' investment in workplace safety. Our dependent variable is the share of employees who are covered by work-related injury insurance, and the analysis is carried out at the firm level and county level (by averaging insurance coverage to the county level). We use OLS as our baselines, and because the insurance coverage ratio is censored between 0 and 1, we employ Tobit models as well. We control for the share of total union members, female union members, and rural union members in total employees, to separate the impacts of strength and composition of unions on insurance coverage. The shares of female

[10] The results based on all provinces are similar.

employees, rural employees, and grass-roots employees in the board of directors and supervisors are also controlled for with consideration of corporate governance factors. We control for the log of sales and other firm-level financial indicators, which are normalized with sales to isolate the influence of profits, liquidity, leverage, and expenses on insurance coverage.

5. Empirical Results on Work-Related Disability and Mortality

Sex Ratios and Work-Related Disability

Table 5.1 presents the estimation results on sex ratio and work-related disability rates using the 2006 CNSDP. The left panel covers the full sample, which includes both healthy and disabled people, and the right panel focuses only on physically disabled people. In the individual-level analysis, we have controlled for the respondent's age, gender, log of family income per person, education level, marital status, occupation, and province dummies.[11] We have restricted the sample to respondents aged 25 to 44 and from rural areas. The dependent variable equals 1 if the respondent is physically disabled due to work injury and equals 0 otherwise. In the county-level analysis based on the full sample, the dependent variable is the disability rate due to work injury (number of workers physically disabled due to work injury divided by average number of workers, unit: 1/100,000). In the county-level analysis based on the physically disabled sample, the dependent variable is the ratio of people who are physically disabled due to work injury to the total number of physically disabled people (unit: 1/100). Considering many counties have no people who are physically disabled due to work injury, we also report Tobit results with the lower limits set at 0 in the county-level analysis. The variable "sex ratio (5-19 cohort)" refers to the male to female ratio among the 0-to-14-year-old cohort, inferred from the 2000 population census. Since we will investigate the transmission channels with the use of the 2010 CFPS dataset, we also report the results based on the provinces in the CFPS.

The coefficient for the sex ratio variable is significant and positive in all the regressions. The results are robust regardless of whether individual-level or county-level analysis is conducted and regardless of which sample is used. An increasing (male-to-female) sex ratio imbalance is associated with a higher incidence of work-related disability.

[11] If additional controls such as log of population, share of prime-age (15–64) population to total population at the county level, and western and middle region dummies are included, the results are similar. Due to space limitations, they are not reported.

Having examined the impact of sex ratio imbalance on work-related disability, we next investigate its impact on accidental deaths. Before coming to the regression analysis, we first scrutinize relevant figures and summary statistics of our key variables to obtain a general idea about the associations between sex ratio and accidental death.

Figure 5.1 depicts the evolution of sex ratio and the work-related death rate over the period 1976–2003.[12] The sex ratio is defined as the ratio at birth 20 years earlier and the work-related death rate is computed as the number of deaths caused by work-related accidents divided by the average number of workers in a certain period (unit: 1/1,000,000), which is available from various issues of *China's Work Safety Yearbook* and the *China Statistical Yearbook*. To exclude the effects of time trend, we rescale the two variables by subtracting the mean and dividing by the standard deviation. Figure 5.1 reveals that these two variables follow similar patterns for most of the time over the period except for the late 1990s. Higher male to female sex ratios are usually associated with higher rates of work-related deaths.

[Figure 5.1 about here]

Figure 5.2 shows a scatter plot of the proportion of each province's work-related deaths in nationwide total work-related deaths over sex ratio. We use the Frisch-Waugh theorem to exclude the impacts of the log of gross domestic product (GDP) per capita. It is clear from Figure 5.2 that the share of work-related deaths is higher in provinces with more skewed sex ratios.

[Figure 5.2 about here]

Table 5.2 provides the summary statistics of our key variables. We calculate relevant county-level death rates (number of deaths due to a certain cause ÷ population × 100,000) from the NDSPS dataset. As shown by the standard error in the table, the sex ratio varies widely across counties. In addition, there exist significant differences in the mean death rate of accidents caused by fires and accidental falls between "the low-sex-ratio group" and "high-sex-ratio group." The high-sex-ratio group exhibits a greater death rate caused by fires and accidental falls. Since our story

[12] The choice of this time span is due to data availability and the consistency of definition of work-related death.

is about the health consequences of hard work and risk taking when people respond to marriage market competition, sex ratios should not be correlated with deaths that are unrelated with hard work and risk taking. Thus we use accidental poisoning due to the intake of a poisonous substance as a placebo test. In accordance with expectations, it is not significant at all. Furthermore, the aforementioned relationship is relatively robust to the choice of different time windows. Regardless of whether wave 1, wave 2, or the whole sample is used, the basic pattern all holds. However, the results in Table 5.2 are just suggestive because we do not control for any other factors.

[Table 5.2 about here]

We next come to the results of the regression analysis with more control variables. The estimates of the linear probability models at the individual level are presented in Table 5.3. In the first four columns, we report the results based on the whole sample in year 1991 and the pooled sample from 1991 to 2000. In the last four columns, the regression results in the parents' cohort are exhibited, both in the 1991 sample and in the pooled sample.

[Table 5.3 about here]

Table 5.3 demonstrates that regardless of deaths due to fires and flames or accidental falls, in a county with a higher sex ratio, the probability of deaths attributable to these causes is higher. Furthermore, such a relationship remains robust in the parents' cohort (with the age of 25 to 44) with unmarried children. Since we have controlled for age in the regressions and the incidence of accidental falls does not increase until after 45 years old (Agnew and Suruda 1993), it is unlikely that age factors, as pointed out by Dong et al. (2012), set the aforementioned results in motion.

As for other control variables, in the parents' cohort, the coefficient for the male dummy is positive in all cases, implying that men have a larger likelihood of dying from accidents. This is probably because men are more likely to be involved in dangerous activities.

Considering that crime rates are potentially associated with local sex ratios as uncovered by Edlund et al. (2013), the effect of sex ratios on deaths because of fires and flames may actually work through the channel of arsons instead of hard work and risk-taking behaviors. However, due to lack of arson data at the local level, we cannot directly test this out. As an alternative, we add the violent and property crime rate as a regressor in the regressions of Table 5.3. The sex ratio variable remains significant and positive, whereas the crime rate is insignificant. This suggests that the effect of sex ratios on deaths is not through the channel of crimes.

We next analyze the associations between sex ratio and accidental death by gender (presented in Appendix Table A.1). It is apparent from Table A.1 that in 1991, for fires and flames, the sex ratio's impacts are significant in the male sample but not in the female sample. For accidental falls, the magnitude of the impact is larger and also more significant in the male sample than the female sample, which is probably due to men's higher appetite for risk. In the pooled sample of 1991–2000, for both fires and accidental falls, the sex ratio coefficient is larger for the male sample as well. Clearly, males respond more than females to higher sex ratios, complying with the view of Becker (1981) and Chiappori, Fortin, and Lacroix (2001). Imbalanced sex ratios enhance the bargaining power of women in the marriage market and men are induced to become more "efficient." Hard work and risk taking are two critical ways to enhance earnings but may come at a health cost. For the parents' cohort, similar patterns hold. Moreover, the coefficients on sex ratio for the parents' sample are larger than those based on the whole sample, implying that the channels through the parents' cohort are more dominant. Although both fathers and mothers are subject to their sons' marriage market squeeze, fathers are more likely to assume the burden of working harder and undertaking high-risk jobs.

As a robustness check, we combine accidents due to fires and flames and accidental falls. We then rerun previous regressions controlling for gender and list the results in Appendix Table A.2. The impact of sex ratio remains significantly positive, and is especially stronger among the parents' cohort and males.

Furthermore, considering the potential measurement errors due to migration issues, we also conduct a robustness check by restricting our sample to the interior regions of China and before 1997. According to Meng (2012, page 76),

> "In the mid to late 1990s, economic growth in the cities began to accelerate and the demand for unskilled labor rose substantially. In particular, after China became a member of the World Trade Organization in November 2001, China's labor-intensive, export-led growth generated major demand for unskilled labor. Migration restrictions then relaxed considerably. Between 1990 and 1997, rural migrants working in cities increased slightly from 25 million to 37 million, but by 2009 the number of rural migrants to China's cities almost quadrupled to reach 145 million."

So large-scale migration did not take place until 1997 and the direction of migration is from inland to coastal areas. The results are presented in Appendix Table A.3. In sample 1, we drop the observations in Guangdong, Zhejiang, and Jiangsu, the three largest destinations of migration. In sample 2, we drop the observations in Shandong, Jiangsu, Shanghai, Zhejiang, Fujian, and Guangdong, the rich and coastal areas of China. We rerun the regressions in the above subsamples in 1991, the pooled sample of 1991–2000, and the pooled sample of 1991–1996 in turn. The basic patterns concerning sex ratio continue to hold in all these scenarios.

Since our story is mainly about the consequences of hard work and risk taking in response to skewed sex ratios, sex ratios should play no role in deaths caused by factors unrelated with hard work and risk taking. Deaths caused by accidental poisoning or owing to congenital anomalies are not linked with hard work and risk taking. Therefore, we use them as placebo tests, the results of which are displayed in Table 5.4. Unlike in the previous table, none of the sex ratio variables is significantly positive regardless of whether we use the 1991 cross-sectional sample or the pooled sample of 1991–2000.

[Table 5.4 about here]

Furthermore, we carry out a falsification test by replacing the sex ratio of the cohort aged 5 to 19 with the cohort aged 25 to 29 and rerunning the regressions in Table 5.3. Since the majority of people in the rural area get married before age 25, our logic of marriage market competition implies that the accident incidence of the parents cohort should not respond to the sex ratio of the 25-to-29-year-old cohort. In compliance with our expectations, the coefficient for the sex ratio of the 25-to-29-year-old cohort is not significant in all cases. Due to space limitations, these results are not reported in tables.

The preceding analyses are based on individual-level deaths. The areas with a higher mortality rate may be overrepresented in the sample. In our sample, accidental deaths are counted as one and other deaths as zero. The number of deaths by other causes may bias the inference on the accidental death incidence. To address this concern, we compute the death rates at the county level by different causes using the death data from the NDSPS and the local population from China's 1990 population census. Average years of schooling, share of ethnic Han, and the share of employment in industry and agriculture are controlled for in the county-level regressions. Table 5.5 presents such results.

[Table 5.5 about here]

From Table 5.5, in accordance with previous individual-level results, sex ratio exhibits a positive correlation with death from accidental falls in the regression based on the 1991 sample. When we combine accidents related to fires and accidental falls, the sex ratio variable remains significantly positive. In the pooled sample of 1991–2000, regardless of whether we use the separate or the combined accidental death incidence, the coefficient for the sex ratio variable remains significant. The results at the county level and the individual level are consistent with each other, demonstrating the robustness of our results.

In terms of the magnitude of the effect, we first divide our whole 1991–2000 sample into lower- and higher-sex-ratio groups using the median value of sex ratio as a threshold. We then compute the difference in averaged sex ratios between the two groups and the difference in death incidence by fires and accidental falls predicted by the model. Dividing this by the actual difference in death incidence between the two groups, we can attain the fraction of accidental deaths explained by sex ratio. We find that the difference in sex ratio accounts for 78.1% $[(110.10 - 104.34) \times 0.08 \div (1.53 - 0.94)]$ and 67.2% $[(110.10 - 104.34) \times 0.21 \div (4.39 - 2.59)]$ of the difference in accidental death incidence due to fires and falls, respectively.

Since complete data about social and economic variables are not systematically available at the county level until 1999, we try to add more control variables and rerun the regressions as before using data obtained from the *China County Statistical Yearbook* in 2000 and 2001 to check whether our results are sensitive to the additional control variables. We include GDP growth rate, log of GDP per capita, share of secondary industry in total GDP, fiscal capacity (ratio of fiscal income to fiscal expenditure), log of number of beds in local hospitals, dummy for hilly county, and dummy for coastal county. In our baseline regressions, we lag these control variables by one period (these variables take their values in 1999) to partly reduce the potential reverse causality problem. The results are robust to these additional control variables. Owing to space limitations, we do not report these results separately.

Although the county-level regressions have included population, the overall death rates may mask the age and gender differences in mortality. For instance, in a region with an aging population, the death rate owing to accidental falls tends to be higher than in a location with a young population. To remedy this concern, we construct a cohort-county-wave pseudo-panel based on the individual-level death information. We use the corresponding total cohort population (each cohort includes five years) in a county to compute the death rates owing to a certain cause during 1991–1995 (wave 1) and 1996–2000 (wave 2) by cohort and gender. Table 5.6 presents the results of OLS regressions based on the pseudo-panel by controlling for cohort, region, and wave fixed effects.[13] Since the channel we emphasize is mainly about hard work and risk taking of parents with unmarried children, rather than old people, we look at the parents' cohort (aged 25 to 44) and the older-than-59 cohort separately. Similar as before, we use deaths due to accidental poisoning as a placebo test. Table 5.6 reveals that for the parents' cohort, sex ratio is still significantly positive in the regressions for both accidents due to fires and accidental falls, but not for accidental poisoning (placebo). In sharp contrast, such impacts are no longer significant in the regressions on the older-than-59 cohort, who are too old to be the parents of the young cohort (aged 5 to 19), complying with our logic that the sex ratio mainly operates through the parents' cohort.[14] That is, parents respond to their sons' marriage market squeeze by working hard and taking greater risks, so as to accumulate wealth and become "efficient".

[Table 5.6 about here]

Even though we have controlled for cohort, wave, region, and some county-specific factors, we may still omit some important variables. In addition, the sex ratio variable may not precisely capture the competitive marriage market pressure facing families with sons. To address the problem of omitted variables and measurement error, we implement 2SLS regressions in the pseudo-panel as well, with the control of cohort, region, and wave fixed effects. Table 5.7 presents the first-stage results, and Table 5.8 reports the results of 2SLS estimations.

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[13] We also try Tobit regressions (censored at 0) with the control of cohort, regional, and wave fixed effects since many cohort-county pairs' death incidence is 0, although such models are subject to the incidental parameter problem. According to Greene (2004), the estimation of the fixed effects Tobit model's slope is consistent, but not for the variance. In case of outliers, we also exclude the observations of the upper 5 percentiles as a robustness check. In these exercises, the results for sex ratio are similar.

[14] A spillover to the older cohort might occur when grandparents also endeavor to share some of the burdens of their children by devoting more time and effort to look after their grandchildren. For example, they may take their grandchildren to school or pick them up after school. During that process, some accidents may occur. With the use of CHARLS, a survey carried out by the National School of Development of Peking University that provides comprehensive information about grandparents' care for grandchildren, we do not find evidence that sex ratio is significantly linked with grandchildren care. Therefore, the spillover effects of sex ratio to the older cohort are not straightforward and may be of second-order concern, complying with our findings that sex ratio's impacts on the older cohort's accidental deaths are not significant.

In Table 5.7, we add possible IVs stepwise. In accordance with expectations, the coefficients on the amount of one-child bonus and the premium for higher-order birth are significantly negative and positive, respectively, at the 1 percent significance level. That is, looser implementations of family planning policy in the context of son preferences exert negative impacts on abortions and on sex ratios. Although the coefficient on the share of minority population is not significant, the three IVs are jointly significant. Since the implementation of the one-child policy is much looser in the minority population, a negative coefficient on the minority population share makes sense.

[Table 5.7 about here]

The endogeneity tests listed in Table 5.8 reject the null hypothesis that sex ratios are exogenous in the regressions of accidental falls and overall accidents in the parents' cohort, and in the regressions of accidents caused by fire in the older-than-59 cohort (p values are 0.06, 0.06, and 0.03, respectively). In these circumstances, the IV method is a better alternative. Furthermore, overidentification tests displayed in Table 5.8 do not reject the null hypothesis of no overidentification in all cases.

Table 5.8 implies that the same pattern remains even after taking the omitted-variable problem and measurement errors into consideration. Reassuringly, the sex ratio variable is significant in the parents' cohort but plays no significant role in the older-than-59 cohort. In the placebo tests of accidental poisoning, the coefficient on the sex ratio variable is not significant in any cases.

[Table 5.8 about here]

As complementary evidence, we check the associations between sex ratio and deaths due to mental disorder and disease of the nervous system by age cohorts.[15] The logic is that parents' hard work and risk taking in response to their sons' dismal prospects in the marriage market are likely to result in severe stress and ultimately impose a deleterious effect on the nervous system. In addition, such an effect should be relevant in the parents' cohort, who react to the challenge of the imbalanced sex ratio of their sons' cohort, rather than in the older-than-59 cohort.

Since the effects of stress are gradual and longstanding, stress does not lead to the immediate consequence of death due to mental disorder and disease of the nervous

[15] A comprehensive analysis of the links between sex ratio and nervous system disease is not the focus of this paper, so we present only preliminary results here as complementary evidence.

system. Taking account of such a time lag, we examine whether the death incidences related to the above causes in 1995 and 2000 are significantly linked with sex ratio. In other words, severe stress at a young age could reduce an individual's life span. A person who died in 1995 aged 30 to 49 or in 2000 aged 35 to 54 is actually in the parents' cohort of our interest in 1991. Sex ratio calculated from the 5-to-19-year-old cohort in the 1990 population census is predetermined in this setup.

Table 5.9 demonstrates that the preceding logic is corroborated by data. No matter whether the death sample from year 1995 or that from year 2000 is used, sex ratio is significant in the parents' cohort, rather than the older-than-59 cohort, which lends complementary support to our baseline results.

[Table 5.9 about here]

6. Underlying Mechanisms

Previous results have established the association between sex ratio and accidental death rates. In this section, we delve into the underlying transmission channels by exploring the connection between sex ratio and hard work, between sex ratio and involvement in risky jobs on the employee side, and between sex ratio and investment in workplace safety on the employer side.

We first turn to the relationship between sex ratio and probability of working outside the hometown, which we present in Table 6.1. We restrict our sample to those families having at least a child aged 5 to 19 and household heads and mothers (men usually marry younger women in China) aged between 25 and 44. We contrast the behavior of families having at least a son and families having only daughters. Columns (1) through (3) and columns (4) through (6) present the results based on household head's age and mother's age, respectively. Table 6.1 indicates that for families with at least a son, the probability of working outside the hometown is significantly greater in a county with higher sex ratios. In sharp contrast, such a relationship is not significant at all and is negative for families with only daughters. The results of the interaction term between the son dummy and sex ratio when pooling all families together, which are reported in columns (3) and (6), further corroborate this pattern: in a region with a more skewed sex ratio, it is more likely for families with sons to work outside the hometown. These results imply that families with unmarried sons take action in response to higher sex ratios so as to win a competitive advantage in the marriage market. Since working outside the hometown is a critical method to earn higher income and increase bargaining power in the marriage market, parents with sons are more likely to choose to do so when facing skewed local sex ratios.

[Table 6.1 about here]

Next, we turn to the links between sex ratio and time spent working outside of one's hometown. Because many families have no member working outside their hometown, we use a Tobit model and the lower limit of the model is set to zero. The results based on the Tobit model are shown in Table 6.2. From the table, it is still apparent that the significantly positive effects of a skewed sex ratio mainly apply to families with at least a son, rather than to families with only daughters. Hence, no matter whether we look at the probability of or the time of working outside the hometown, sex ratio plays a significant role in families with at least a son, rather than in families with only daughters.

[Table 6.2 about here]

Aside from working outside one's hometown, taking part in risky jobs is another alternative approach to accumulate wealth, but doing so may result in a higher risk of accidental death. Risky jobs often offer higher income. After excluding the respondent's experience, age, gender, years of education, father's and mother's years of education, and county fixed effects, the mean difference of the parent cohort's income between a risky job and nonrisky job in the rural area is 3,307 yuan, significant at the 1 percent level.

We next investigate the connections between sex ratio and participation in risky jobs. The regression results for sex ratio and probability of participating in risky jobs are presented in Table 6.3. In the first three columns, we focus on families whose household head's age is between 25 and 44, and we restrict our sample based on the mother's age in the last three columns as a robustness check since women are usually a few years younger than their husbands in China. As before, we examine the families with at least a son and families with only daughters separately. Although the coefficient on sex ratio in the son sample is not significant, in the daughter sample, it is significantly negative. That is, in a region with a higher sex ratio, families with daughters are less likely to work at risky jobs. Furthermore, we pool the son sample and daughter sample together and add a son dummy and an interaction term between sex ratio and the dummy. Consistent with our story, the interaction term is significantly positive in both cases. In other words, families with a son are more prone to participate in risky jobs in a region with more unbalanced sex ratios.

[Table 6.3 about here]

However, a potential challenge to the preceding results exists: if sex ratio imbalances prompt people to take risks in order to earn more money, with a greater number of people swarming into the risky sector, the premium of risky jobs might disappear and as a result it would not be remunerative to take risks. This concern is not corroborated by the data. As shown in Figure 6.1, the scatter plot of risk premium averaged to the county level over sex ratio is basically flat (p = 0.30).

[Figure 6.1 about here]

Imbalanced sex ratios affect not only employees but also employers. The eagerness to accumulate wealth triggered by skewed premarital sex ratios could decrease the bargaining power of parents in the labor market. Thus they may accept jobs with adverse working conditions. Accordingly, employers are inclined to underinvest in workplace safety. We next scrutinize the impact of sex ratio on the employers' side. We first examine the relationship between sex ratio and the share of trade unions that have set up labor protection, supervision, and inspection committees (LPSICs), which are mainly responsible for workplace safety issues, as introduced in the fourth section of the paper.

Figure 6.2 presents a scatter plot of those two variables conditional on log of GDP per capita. Clearly, a negative relationship exists. That is, the higher the sex ratio, the worse the situation of workplace safety investment.

[Figure 6.2 about here]

A challenge of this analysis is whether the same functions of LPSICs are performed by other committees in a trade union. Besides LPSICs, trade unions may also establish "warmth project" committees (WPCs), which are primarily responsible for helping employees cope with economic difficulties in life; labor legal supervision committees (LLSCs), which supervise the implementation of the labor law in the workplace; technical cooperation committees (TCCs), which encourage technical cooperation among employees to improve productivity and carry out technical contests to boost technical innovation and invention; labor dispute mediation committees (LDMCs), which deal with the renewal, dissolution, termination, and continuation of labor contracts, the dismissal or resignation of staff and workers, labor rewards, insurance, working hours, rest and vacation, labor safety and sanitation, occupation training, special protection for underage workers, and female workers. In terms of prevalence, LPSICs are similar to LDMCs, and are more common than other types of committees.[16] Although LDMCs cope with labor safety and sanitation problems as well, their role is mainly reflected in the field of renewal, dissolution, termination, and continuation of labor contracts, rather than safety-related issues.[17] In other words, LPSICs' functions of labor protection are not performed or easily

[16] In 2011, the average share of trade unions that had set up LPSICs, WPCs, LLSCs, TCCs, and LDMCs across provinces was respectively 24.17 percent, 6.33 percent, 18.80 percent, 1.79 percent, and 23.72 percent.

[17] In 2011, the average proportion of labor contract dispute cases in total cases handled by LDMCs across provinces was 57.4 percent, whereas the figure for labor safety and sanitation disputes was 2.17 percent. LDMCs coped with 5,866 safety cases nationally, whereas the figure for LPSICs was 59,207.

replaced by other committees in a trade union, and they are remarkable forces to protect workers' rights in safety.

Since we can obtain the above information only at the province level, the sample size is too small to carry out a robust regression analysis. As an alternative, we implement a placebo test to examine whether sex ratio is significantly connected with the share of LDMCs. Since LDMCs do not primarily deal with safety-related issues, we should not observe a significantly negative pattern. Figure 6.3 confirms that this is indeed the case. Applying the same conditional plot for LMDCs in Figure 6.3, the fitted line is fairly flat and the p value for sex ratio is 0.94, much larger than 0.1.

[Figure 6.3 about here]

Because work-related injury insurance is also an important facet of workers' rights and a critical investment that employers should make in workplace safety, we analyze how imbalanced sex ratios affect it. Table 6.4 shows the regression results on sex ratio and work-related injury insurance coverage. We report OLS results as baselines, and since insurance coverage ratios are censored between 0 and 1, we employ Tobit models as well. At the firm level, we add control variables stepwise.

Table 6.4 demonstrates that regardless of controlling for firm-level characteristics, industry and owner type dummies, and province dummies, the coefficient on the sex ratio variable is negative and highly significant at 1 percent. Even if we average insurance coverage to the county level, the variable remains significantly negative. Moreover, the major results still hold when we employ the Tobit models. That is, a higher sex ratio is linked with worse situations of work-related injury insurance coverage.

[Table 6.4 about here]

7. Conclusions

We explore the negative health consequences of ”missing women” of the premarital age cohort. More specifically, we investigate if and how higher male-to-female ratios of the premarital age cohort lead to a higher incidence of accidental death among the parent cohort. Empirical evidence based on four large datasets demonstrates that skewed male-to-female ratios have a positive impact on the work-related injury and death incidence of accidents due to fires and accidental falls, especially for the parents' cohort. Parents react to the imbalanced sex ratio by working harder and taking more risks, which are important risk factors for death owing to accidents, in order to accumulate wealth and improve their sons' attractiveness in the marriage market. Imbalanced sex ratios boost the probability of and time working outside the hometown and participation in risky jobs for families with sons, but not for families having only daughters. Employers also underinvest in workplace safety because more potential employees put up with dangerous working conditions.

Considering the empirical challenges of omitted-variable bias and endogeneity issues, we not only implement placebo tests and carry out IV regressions, but also construct a pseudo-panel to control the cohort, region, and wave specific fixed effects. Regardless of our focus, whether at the individual level, the county level, or the cohort-region-wave level, the connection between imbalanced sex ratio and accidental death due to fires and falls remains robust. The detailed and separate investigations of the underlying mechanisms on the employees' side (that is, in families with sons and families with only daughters) and the employers' side lend further support to our baseline findings.

The impact of sex ratio is quantitatively large, explaining about 78.1 percent of the variations in the mortality rate of accidents caused by fires and 67.2 percent of the death incidence due to accidental falls. Competitive pressure as a cause of accidental death has not been systematically explored in the literature. Our findings shed some new light on the causes of the abnormally high work-related death rates in China.

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Table 1.1 Coal mine death rates across countries (Unit: Deaths Per Million Tons)

Year	China	United States of	India
1988	6.78	0.06	0.93
1989	6.67	0.07	0.84
1990	6.1	0.07	0.54
1991	5.21	0.07	0.63
1992	4.65	0.06	0.69
1993	4.78	0.06	1.18
1994	5.15	0.05	0.75
1995	5.03	0.05	0.54
1996	4.67	0.04	0.48
1997	5.1	0.03	0.54
1998	5.02	0.03	0.47
1999	5.3	0.03	0.45
2000	5.86	0.03	0.46
2001	4.67	0.05	0.70
2002	5.46	0.03	0.47
2003	4.41	0.04	0.52
2004	3.42	0.03	0.41
2005	3.06	0.03	0.48
2006	2.25	0.05	0.54
2007	1.68	0.04	0.28
2008	1.38	0.03	0.24
2009	1.06	0.02	0.20
2010	0.90	0.06	0.29
2011	0.67	0.03	0.18

Sources: China State Administration of Coal Mine Safety Bulletin (1988-2011); United States Mine Safety and Health Administration Bulletin (1988-2011); and BP Statistical Review of World Energy (1988-2011).

Table 5.1—Sex ratio and disability due to work injury

	Full sample				Physically disabled sample			
	Individual level		County level		Individual level		County level	
	All provinces	Provinces in CFPS	OLS results	Tobit results	All provinces	Provinces in CFPS	OLS results	Tobit results
Sex ratio	1.194**	1.134**	1.393***	3.153***	0.071*	0.066*	0.085**	0.192**
(5–19 cohort)	(0.511)	(0.553)	(0.537)	(1.218)	(0.043)	(0.035)	(0.042)	(0.091)
Observations	835,029	732,652	733	733	11,507	9,951	733	733

Source: China National Survey on Disabled People (2006).

Notes: CFPS = China Family Panel Studies; The full sample encompasses both healthy and disabled people in the survey, while the physically disabled sample focuses only on people who are physically disabled. In the individual-level analysis, we have controlled for the respondent's age, gender, log of family income per person, education level, marital status, occupation, and province dummies. If we control for the preceding individual-level characteristics, log of population, and share of prime-age (15–64) population to total population at the county level, and western and middle region dummies, the results are similar. Due to space limitations, they are not reported. We have restricted the sample to respondents aged 25 to 44 and from the rural area. The dependent variable equals 1 if the respondent is physically disabled due to work injury, and equals 0 otherwise. In the county-level analysis based on the full sample, the dependent variable is the disability rate due to work injury (number of workers physically disabled due to work injury divided by average number of workers, unit: 1/100,000). In the county-level analysis based on the physically disabled sample, the dependent variable is the share of people who are physically disabled due to work injury (unit: 1/100). Considering many counties have no people who are physically disabled due to work injury, we also report Tobit results with the lower limits setting at 0 in the county-level analysis. Sex ratio of the 5–19 cohort refers to the male-to-female ratios among the 0–14 cohort, inferred from the 2000 population census. For ease of exposition, in the individual-level analysis, we have rescaled our independent variables by dividing 100,000 in the full sample and by dividing 100 in the physically disabled sample. Since we will investigate the transmission channels with the use of the CFPS dataset, we also report the results based on provinces in the CFPS.

Table 5.2—Summary statistics of key variables

	Whole sample		Lower-than-median	Greater-than-median	
	Mean	Standard error	Mean	Mean	Mean difference
			1991–2000 (Whole sample)		
Sex ratio of 5–19 age cohort	107.20	4.50	104.34	110.10	-5.77***
Death rate of fires and flames	1.23	5.77	0.94	1.53	-0.59*
Death rate of accidental falls	3.49	5.02	2.59	4.39	-1.79***
Death rate of accidental poisoning	3.68	5.15	3.79	3.56	0.23
			1991–1995 (Wave 1)		
Sex ratio of 5–19 age cohort	106.45	4.32	103.87	109.10	-5.23***
Death rate of fires and flames	1.64	7.74	0.96	2.33	-1.37*
Death rate of accidental falls	3.56	5.27	2.70	4.45	-1.75***
Death rate of accidental poisoning	3.81	5.61	3.99	3.62	0.38
			1996–2000 (Wave 2)		
Sex ratio of 5–19 age cohort	107.97	4.55	104.95	111.00	-6.04***
Death rate of fires and flames	0.82	2.42	0.95	0.70	0.25
Death rate of accidental falls	3.42	4.77	2.65	4.19	-1.54***
Death rate of accidental poisoning	3.54	4.63	3.39	3.69	-0.30

Source: National Disease Surveillance Points System (1991-2000).

Notes: The county-level death rates are defined as the number of deaths due to a certain cause per 100,000 people based on the National Disease Surveillance Points System dataset. We use the median of sex ratio in the full sample to divide the full sample into two subsamples. Lower/Greater-than-median indicates the subsample that the sex ratio is lower/greater than the median of sex ratio in the whole sample. The sex ratio variable during 1991–1995 is the sex ratio among the 5–19 cohort in the 1990 population census, while it is defined as the male-to-female ratio among the 0–14 cohort in the 1990 population census for the late period of 1996–2000. The symbols ***, **, and * indicate that we can reject the null hypothesis that mean difference is 0 at the significance level of 1%, 5%, and 10%, respectively.

Table 5.3—Sex ratio and accidental death: OLS regressions at the individual level in the whole sample and parents' cohort

	Whole sample				Parents' cohort			
	Year 1991		Pooled 1991–2000		Year 1991		Pooled 1991–2000	
	Accident 1	Accident 2	Accident 1	Accident 2	Accident 1	Accident 2	Accident 1	Accident 2
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sex ratio for 5–19 cohort	0.02**	0.06***	0.01**	0.03***	0.09***	0.16***	0.05*	0.05*
	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.03)	(0.03)	(0.03)
Male	0.08	-0.15	0.04	0.26***	0.21	0.72	0.26	0.02***
	(0.07)	(0.14)	(0.03)	(0.05)	(0.26)	(0.46)	(0.16)	(0.26)
Age	-0.01*	-0.01	-0.01**	-0.03***	-0.01	0.00	-0.02	-0.02
	(0.01)	(0.01)	(0.00)	(0.00)	(0.02)	(0.03)	(0.02)	(0.02)
Years of schooling	0.00	0.03	-0.00	-0.00	-0.00	0.04	0.02	-0.00
	(0.01)	(0.02)	(0.00)	(0.01)	(0.04)	(0.08)	(0.02)	(0.03)
Labor participation rate	0.00	-0.06*	-0.01	-0.02	0.03	-0.07	-0.02	-0.01
	(0.01)	(0.03)	(0.01)	(0.02)	(0.05)	(0.06)	(0.03)	(0.06)
Log of population	-0.07	-0.17	-0.01	-0.20***	0.46**	0.39	0.48	0.23
	(0.05)	(0.13)	(0.02)	(0.06)	(0.19)	(0.40)	(0.37)	(0.24)
East	-0.17	0.39	-0.03	0.08	-0.04	0.78	-0.55	0.33
	(0.12)	(0.35)	(0.05)	(0.20)	(0.35)	(0.53)	(0.37)	(0.51)
Central	-0.20*	-0.34	0.09	-0.04***	-0.10	0.23	0.01	-0.70*
	(0.11)	(0.25)	(0.11)	(0.13)	(0.31)	(0.46)	(0.44)	(0.34)
Observations	26,436	26,436	262,117	262,117	2,603	2,603	19,286	19,286
Adjusted R^2	0.05	0.01	0.01	0.01	0.06	0.09	0.07	0.13

Source: National Disease Surveillance Points System (1991-2000).

Notes: The parents' cohort age is between 25 and 44. The dependent variable equals 1 if a person died of accident 1 or accident 2 and 0 otherwise. Accident 1 and accident 2 indicate death caused by fire and flames and accidental falls, respectively. We have rescaled all the independent variables by dividing 100. For the regression on year 1991, we define the sex ratio as the male to female ratio among age cohort 5-19, calculated from the 1990 population census. In the pooled sample, the sex ratios for the period 1991–1995 and for the period 1996–2000 refer to the male-to-female ratios among the cohorts 5–19 and 0–14 inferred from the 1990 population census, respectively. In the regressions, we also control for wave dummy, which equals 1 in the period 1991–1995 and 0 in the period 1996–2000, and other individual-level characteristics, such as ethnicity (dummy variables for Han, Hui, Mongolian, Tibetan, Uigur, Zhuang, others), occupation (dummy variables for production, farmer, sheepherder, fisherman, medical staff, food preparation and serving, business and social service, manager, retired, housework, jobless, children at kindergarten, children not at kindergarten, student of primary school, student of middle and high school, undergraduate or more, others) and marital status (dummy variables for married, single, widowed, divorced, others). We do not report the coefficients for these variables for brevity. Labor participation rate is defined as the share of prime-age (15–64) working population in total population. The regional dummy variables *East* and *Central* indicate whether the county is in the eastern or central region, with the western region as the base group. Robust standard errors clustered at the county level are in parentheses. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%.

Table 5.4—Sex ratio and poisoning death (placebo tests): Individual-level OLS

	Whole sample				Parents' cohort	
	Year 1991		Pooled 1991–2000		Year 1991	Pooled 1991–2000
	a1	d1	a1	d1	a1	a1
	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio for 5–19 cohort	-0.01	0.004	0.01	-0.02**	-0.09*	-0.01
	(0.02)	(0.03)	(0.01)	(0.01)	(0.05)	(0.03)
Observations	26,436	26,436	262,117	262,117	2,603	19,286
Adjusted R^2	0.01	0.03	0.01	0.01	0.03	0.01

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square; a1 = poisoning death; d1 = death of newborn baby due to congenital anomalies. Parents' cohort stands for those who died between 25 and 44 years old. Other regressors and notes are the same as those in Table 5.3.

Table 5.5—Sex ratio and accidental death: County-level OLS regressions

	Year 1991			Pooled 1991–2000		
	Accident 1	Accident 2	Accident	Accident 1	Accident 2	Accident
	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio for 5–19 cohort	0.02	0.35**	0.37**	0.08***	0.21***	0.29***
	(0.08)	(0.15)	(0.16)	(0.03)	(0.04)	(0.04)
Urbanization	0.79	-1.78	-0.98	-0.13	-0.55**	-0.68**
	(0.73)	(1.27)	(1.55)	(0.19)	(0.23)	(0.30)
Years of schooling	-0.07	0.25	0.18	0.10	-0.185*	-0.09
	(0.29)	(0.78)	(0.73)	(0.10)	(0.10)	(0.15)
Share of Han	-0.06	1.92	1.86	-0.14	0.48	0.34
	(0.90)	(2.78)	(3.18)	(0.40)	(0.68)	(0.80)
Share of industrial employment	1.18	-15.60	-14.42	-15.91	-5.80	-21.71
	(4.94)	(14.00)	(14.60)	(13.20)	(3.73)	(13.50)
Share of agri. employment	4.51**	-5.25	-0.74	-3.50	-0.39	-3.89
	(2.13)	(5.33)	(6.04)	(4.23)	(1.09)	(4.27)
Observations	67	67	67	640	640	640
Adjusted R^2	0.12	0.13	0.09	0.03	0.06	0.06

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square. The regressions are at the county level. The dependent variable is the accidental mortality rate. *Accident 1* means deaths caused by fire and flames. *Accident 2* indicates death as a result of accidental falls. *Accident* encompasses *Accident 1* and *Accident 2*. For the regressions on year 1991, we define the sex ratio as the male-to-female ratio among the 5–19 age cohort, calculated from the 1990 population census. In the pooled sample, the sex ratios for the period 1991–1995 and for the period 1996–2000 refer to the male-to-female ratios among the 5–19 and 0–14 cohorts inferred from the 1990 population census, respectively. We also control for wave dummy, which equals 1 in the period 1991–1995 and 0 in the period 1996–2000. *Urbanization* takes 1 to 5, indicating the degree of urbanization. *Years of schooling* is defined as the average years of schooling among those older than 20. Robust standard errors are in parentheses. ***, **, and * indicate significance level at 1%, 5%, and 10%.

Table 5.6—Sex ratio and accidental death: OLS cohort-county-wave pseudo-panel regressions

	Parents' cohort				Cohort older than 59			
	Accident 1	Accident 2	Accident	Poisoning	Accident 1	Accident 2	Accident	Poisoning
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sex ratio for 5–19	0.107	0.649*	0.756*	0.623	1.491	3.596	5.087	1.572
	(0.08)	(0.38)	(0.44)	(0.49)	(1.09)	(4.92)	(5.58)	(1.45)
Observations	528	528	528	528	1,039	1,039	1,039	1,039
Adjusted R^2	0.03	0.18	0.17	0.18	0.02	0.03	0.04	0.01

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square;. We use the total population corresponding to each cohort (five years) at the county level to scale the total number of deaths owing to a certain cause during 1991–1995 (wave 1) and 1996–2000 (wave 2) and construct the cohort-county-wave level death incidence by cause. The dependent variable is death incidence. *Accident 1* indicates accidental deaths caused by fire and flames, while a*ccident 2* deaths result from accidental falls. *Accident* encompasses the preceding two types of deaths. As a placebo test, we look at poisoning death. The sex ratios in the period 1991–1995 and the period 1996–2000 refer to the male-to-female ratios among the 5–19 and 0–14 cohorts inferred from the 1990 population census, respectively. Cohort, regional, and wave fixed effects are controlled in the regressions. We also control log of cohort population in case of the inaccuracy of death incidence due to small cohort size. The results for sex ratio are similar. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%.

Table 5.7—Sex ratio and accidental death: First stage of 2SLS cohort-county-wave pseudo-panel regressions

Dependent variable: sex ratio for 5–19 cohort						
Amount of one-child bonus	-0.008***	-0.007***	-0.006**	-0.008***	-0.007***	-0.006**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Premium to higher-order birth		2.635**	2.603**		2.635**	2.603**
		(1.140)	(1.163)		(1.140)	(1.161)
Share of minority population			-0.007			-0.007
			(0.018)			(0.018)
Observations	528	528	528	1,039	1,039	1,039
Adjusted R^2	0.05	0.09	0.09	0.05	0.09	0.09
Kleibergen-Paap Wald F statistic	10.00	5.86	4.19	10.00	5.86	4.19

Source: National Disease Surveillance Points System (1991-2000).

Notes: 2SLS = two-stage least squares. We use the total population corresponding to each cohort (five years) at the county level to scale the total number of deaths owing to a certain cause during 1991–1995 (wave 1) and 1996–2000 (wave 2) and construct the cohort-county-wave level death incidence by cause. The dependent variable is death incidence. *Accident 1* indicates accidental deaths caused by fire and flames, while a*ccident 2* deaths result from accidental falls. *Accident* encompasses the preceding two types of deaths. As a placebo test, we look at poisoning death. The sex ratios in the period 1991–1995 and the period 1996–2000 refer to the male-to-female ratios among the 5–19 and 0–14 cohorts inferred from the 1990 population census, respectively. Cohort, regional, and wave fixed effects are controlled in the regressions. We also control log of cohort population in case of the inaccuracy of death incidence due to small cohort size. The results for sex ratio are similar. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%.

Table 5.8—Sex ratio and accidental death: 2SLS cohort-county-wave pseudo-panel regressions

	Parents' cohort				Cohort older than 59			
	Accident 1	Accident 2	Accident	Poisoning	Accident 1	Accident 2	Accident	Poisoning
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Amount of one-child bonus as IV							
Sex ratio for 5–19 cohort	0.601*	1.473**	2.074**	0.205	0.296	-1.785	-1.489	2.428
	(0.33)	(0.65)	(0.91)	(0.55)	(3.04)	(10.70)	(11.90)	(4.26)
Observations	528	528	528	528	1039	1039	1039	1039
Adjusted R^2	0.09	0.02	0.09	0.13	0.02	0.03	0.03	0.01
Durbin-Wu-Hausman (*p* value)	0.14	0.16	0.12	0.52	0.69	0.63	0.59	0.85
	Amount of one-child bonus and premium to higher-order birth as IVs							
Sex ratio for 5–19 cohort	0.614*	1.564**	2.179**	0.33	2.001	4.948	6.95	-0.522
	(0.32)	(0.66)	(0.90)	(0.43)	(3.01)	(10.40)	(11.50)	(1.99)
Observations	528	528	528	528	1039	1039	1039	1039
Adjusted R^2	0.11	0.04	0.11	0.16	0.02	0.03	0.03	0.01
Durbin-Wu-Hausman (*p* value)	0.13	0.13	0.10	0.62	0.87	0.87	0.85	0.39
Hansen J-statistics (*p* value)	0.82	0.60	0.61	0.47	0.11	0.32	0.22	0.20
	Amount of one-child bonus, premium to higher-order birth, and share of minority population as IVs							
Sex ratio for 5–19 cohort	0.610*	1.439**	2.049**	0.335	1.682	4.925	6.607	-0.739
	(0.31)	(0.64)	(0.87)	(0.43)	(2.95)	(10.70)	(11.70)	(2.03)
Observations	528	528	528	528	1039	1039	1039	1039
Adjusted R^2	0.10	0.03	0.10	0.16	0.02	0.03	0.04	0.01
Durbin-Wu-Hausman (*p* value)	0.14	0.06	0.06	0.73	0.03	0.49	0.22	0.17
Hansen J-statistics (*p* value)	0.68	0.19	0.13	0.62	0.95	0.88	0.88	0.33

Source: National Disease Surveillance Points System (1991-2000).

Notes: 2SLS = two-stage least squares. Cohort, regional, and wave fixed effects are included in 2SLS regressions. We also control log of cohort population in case of the inaccuracy of death incidence due to small cohort size. Due to space limitations, the estimations for cohort, regional, and wave dummies are not reported. Robust standard errors clustered on the county level are in parentheses. Symbols ***, **, and * indicate significance level at 1%, 5%, and 10%. Parents' cohort includes those died at the age of 25 to 44.

Table 5.9—Sex ratio and stress-related deaths: Individual-level cross-sectional OLS regressions

Dependent variable: death due to mental disorder or disease of nervous system				
	Death in year 1995		Death in year 2000	
	Parents' cohort	Older-than-59 cohort	Parents' cohort	Older-than-59 cohort
	(1)	(2)	(3)	(4)
Sex ratio for 5–19 cohort	0.11*	-0.02	0.06*	-0.001
	(0.05)	(0.03)	(0.03)	(0.02)
Male	0.76*	-0.34**	-0.49	-0.34**
	(0.43)	(0.15)	(0.43)	(0.13)
Age	-0.07	0.06**	-0.13***	0.01
	(0.05)	(0.03)	(0.04)	(0.02)
Years of schooling	-0.32**	0.02	-0.20*	-0.06**
	(0.12)	(0.03)	(0.10)	(0.02)
Urbanization	0.54	0.98	0.20	0.16
	(0.47)	(0.66)	(0.41)	(0.15)
Labor participation rate	-0.11	-0.11	0.17	0.05
	(0.09)	(0.07)	(0.10)	(0.05)
Log of population	-1.35**	0.33*	-0.31	0.10
	(0.60)	(0.17)	(0.43)	(0.13)
East	0.55	-149**	-0.13	0.03
	(0.82)	(0.70)	(0.89)	(0.51)
Central	-0.47	-0.36	-0.58	-0.10
	(0.63)	(0.50)	(0.71)	(0.43)
Observations	2,915	16,223	3,487	14,072
Adjusted R^2	0.02	0.01	0.01	0.01

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square; The regressions are at the individual level. The dependent variable equals 1 if a person died of mental disorder or nervous system disease and 0 otherwise. The parent cohort aged 30 to 49 in 1995 and 35 to 54 in 2000. Sex ratio is the male-to-female ratio among the cohort aged 5 to 19, calculated from the 1990 population census. Other notes are the same as in Table 5.3.

Table 6.1—Sex ratio and working outside the hometown: Probit regressions based on CFPS rural sample

Dependent variable: whether any family member worked outside the hometown last year						
	Household head's age 25 to 44			Mother's age 25 to 44		
	Son	Daughter	Whole	Son	Daughter	Whole
	(1)	(2)	(3)	(4)	(5)	(6)
Have a son * Sex ratio			0.022**			0.024**
(10–19 age cohort)			(0.011)			(0.010)
Have a son			-2.294*			-2.411**
(dummy)			(1.235)			(1.147)
Sex ratio	0.011**	-0.010	-0.011	0.013**	-0.010	-0.011
(10–19 age cohort)	(0.006)	(0.010)	(0.010)	(0.005)	(0.009)	(0.009)
Per capita income (log)	-0.073	-0.135*	-0.096*	-0.031	-0.107	-0.060
	(0.067)	(0.081)	(0.054)	(0.061)	(0.079)	(0.052)
Family size	0.138**	0.128	0.137***	0.147**	0.089	0.127**
	(0.060)	(0.095)	(0.051)	(0.063)	(0.096)	(0.056)
Household head's age	0.025	-0.029	0.007	0.035***	-0.007	0.022**
	(0.018)	(0.025)	(0.014)	(0.013)	(0.017)	(0.009)
Household head's year of	0.059	0.170***	0.088**	0.031	0.189***	0.072*
schooling	(0.041)	(0.054)	(0.035)	(0.049)	(0.056)	(0.040)
Square of above variable	-0.008**	-0.018***	-0.011***	-0.005	-0.020***	-0.009***
	(0.004)	(0.005)	(0.003)	(0.004)	(0.005)	(0.003)
Male family head	-0.485***	-0.300	-0.408***	-0.533***	-0.382*	-0.468***
(dummy)	(0.157)	(0.204)	(0.134)	(0.159)	(0.199)	(0.134)
Han family head	0.383*	0.361	0.381**	0.313	0.351	0.321**
(dummy)	(0.225)	(0.233)	(0.173)	(0.199)	(0.233)	(0.154)
Poor health	0.067	0.177	0.074	-0.192	0.307	-0.014
(dummy)	(0.185)	(0.304)	(0.157)	(0.204)	(0.322)	(0.186)
Have a child 10 to 14	-0.117	0.064	-0.058	-0.106	-0.142	-0.121
(dummy)	(0.128)	(0.152)	(0.091)	(0.129)	(0.137)	(0.094)
Have a child 15 to 19	0.030	0.624***	0.235*	0.019	0.517***	0.177
(dummy)	(0.159)	(0.215)	(0.133)	(0.153)	(0.192)	(0.114)
Number of children	0.060	0.022	0.051	0.116	0.185	0.150
	(0.153)	(0.161)	(0.116)	(0.146)	(0.156)	(0.113)
Observations	621	366	987	610	350	960

Source: China Family Panel Studies (2010).

Notes: The dependent variable equals 1 if any family member worked outside the hometown and 0 otherwise. We restrict our sample to those families with at least a child aged 5 to 19 and household head or mother aged 25 to 44. The columns marked with "son" mean that a family only has sons and those with "daughter" indicate that a family only has daughters. Sex ratio for the 10–19 cohort in 2010 is inferred from the male-to-female ratio among the cohort aged 0 to 9 in 2000 based on China's 2000 population census. Robust standard errors clustered at the county level are in parentheses. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%.

Table 6.2—Sex ratio and hard work: Tobit regression results based on CFPS rural sample

Dependent variable: total number of months working outside the hometown last year for all family members						
	Household head's age 25 to 44			Mother's age 25 to 44		
	Son	Daughter	Whole	Son	Daughter	Whole
	(1)	(2)	(3)	(4)	(5)	(6)
Have a son * Sex ratio			0.285**			0.279**
(10–19 age cohort)			(0.138)			(0.130)
Have a son			-30.012*			-28.934**
(dummy)			(15.829)			(14.697)
Sex ratio	0.150**	-0.134	-0.138	0.143**	-0.133	-0.138
(10–19 age cohort)	(0.059)	(0.133)	(0.137)	(0.057)	(0.119)	(0.119)
Per capita income (log)	-0.509	-1.152	-0.782	-0.085	-0.817	-0.399
	(0.821)	(0.748)	(0.609)	(0.743)	(0.783)	(0.590)
Family size	2.175***	2.477*	2.255***	1.613**	1.378	1.474**
	(0.704)	(1.412)	(0.655)	(0.687)	(1.150)	(0.595)
Household head's age	0.490**	-0.327	0.220	0.394***	-0.064	0.250**
	(0.215)	(0.290)	(0.174)	(0.141)	(0.209)	(0.105)
Household head's year of	1.067**	2.311***	1.365***	0.579	2.280***	0.996**
schooling	(0.490)	(0.645)	(0.430)	(0.534)	(0.664)	(0.466)
Square of above variable	-0.121***	-0.248***	-0.151***	-0.071*	-0.245***	-0.112***
	(0.043)	(0.059)	(0.038)	(0.043)	(0.064)	(0.039)
Male family head	-5.931***	-3.647	-5.015***	-6.239***	-4.412*	-5.454***
(dummy)	(1.882)	(2.353)	(1.661)	(1.871)	(2.302)	(1.628)
Han family head	4.565*	3.778	4.384**	3.572	3.475	3.528**
(dummy)	(2.356)	(2.680)	(1.805)	(2.176)	(2.623)	(1.710)
Poor health	0.739	2.251	0.947	-2.134	3.261	-0.300
(dummy)	(2.231)	(3.313)	(1.786)	(2.275)	(3.422)	(2.036)
Have a child 10 to 14	-1.859	0.455	-1.113	-0.961	-2.198	-1.376
(dummy)	(1.433)	(1.678)	(1.088)	(1.344)	(1.557)	(1.108)
Have a child 15 to 19	-0.405	6.867***	2.092	0.536	5.427**	2.062*
(dummy)	(1.826)	(2.558)	(1.545)	(1.602)	(2.340)	(1.248)
Number of children	0.401	-0.515	0.208	0.794	2.126	1.430
	(1.770)	(2.412)	(1.546)	(1.631)	(2.059)	(1.382)
Observations	621	366	987	610	350	960

Source: China Family Panel Studies (2010).

Notes: . We use Tobit models, and the left limit is set to 0. We restrict our sample to those families with at least a child aged 5 to 19 and household head or mother aged 25 to 44. The columns marked with "son" mean that a family only has sons and those with "daughter" indicate that a family only has daughters. Sex ratio for the 10–19 cohort in 2010 is inferred from the male-to-female ratio among the cohort aged 0 to 9 in 2000 based on China's 2000 population census. Robust standard errors clustered at the county level are in parentheses. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%.

Table 6.3—Sex ratio and participation in risky jobs: Probit regression results based on CFPS rural sample

Dependent variable: whether any family member has a dangerous job or workplace is in the open air or a workshop						
	Household head's age 25 to 44			Mother's age 25 to 44		
	Son	Daughter	Whole	Son	Daughter	Whole
	(1)	(2)	(3)	(4)	(5)	(6)
Have a son * Sex ratio			0.038***			0.034***
(10–19 age cohort)			(0.012)			(0.012)
Have a son			-4.168***			-3.707***
(dummy)			(1.394)			(1.405)
Sex ratio	0.003	-0.040***	-0.035***	0.003	-0.038***	-0.031***
(10–19 age cohort)	(0.007)	(0.012)	(0.012)	(0.007)	(0.012)	(0.012)
Per capita income (log)	0.339***	0.489***	0.371***	0.360***	0.461***	0.374***
	(0.081)	(0.112)	(0.063)	(0.078)	(0.108)	(0.062)
Family size	0.006	0.096	0.034	-0.045	0.085	-0.007
	(0.074)	(0.086)	(0.064)	(0.071)	(0.097)	(0.060)
Household head's age	-0.012	0.007	-0.006	-0.024	0.011	-0.011
	(0.019)	(0.027)	(0.016)	(0.016)	(0.024)	(0.015)
HH's years of schooling	0.092**	0.070	0.085**	0.065	0.028	0.051
	(0.047)	(0.051)	(0.035)	(0.045)	(0.051)	(0.033)
Square of above variable	-0.006	-0.007*	-0.006**	-0.003	-0.004	-0.003
	(0.004)	(0.004)	(0.003)	(0.004)	(0.004)	(0.003)
Male family head	-0.030	0.206	0.050	0.001	0.204	0.068
(dummy)	(0.135)	(0.191)	(0.105)	(0.138)	(0.197)	(0.109)
Han family head	0.337	0.778***	0.462***	0.352*	0.903***	0.522***
(dummy)	(0.211)	(0.254)	(0.158)	(0.214)	(0.237)	(0.157)
Poor health	-0.327	-0.799**	-0.474**	-0.263	-0.754**	-0.419**
(dummy)	(0.228)	(0.344)	(0.193)	(0.252)	(0.366)	(0.207)
Have a child 10 to 14	-0.094	0.064	-0.022	0.040	0.075	0.054
(dummy)	(0.130)	(0.191)	(0.116)	(0.128)	(0.201)	(0.108)
Have a child 15 to 19	-0.039	0.128	0.035	0.110	-0.006	0.068
(dummy)	(0.162)	(0.222)	(0.142)	(0.153)	(0.215)	(0.135)
Number of children	-0.116	-0.187	-0.125	-0.052	-0.146	-0.074
	(0.161)	(0.198)	(0.124)	(0.161)	(0.190)	(0.127)
Observations	621	366	987	610	350	960

Source: China Family Panel Studies (2010).

Notes: The dependent variable equals 1 if any family member has a dangerous job or the workplace is in the open air or a workshop and 0 otherwise. Dangerous jobs are jobs related to firefighting, public security, mineral mining, well drilling, oil and gas exploration, construction, electricity supply, installation and debugging of electrical equipment, and operation of mechanical equipment. We restrict our sample to those families with at least a child aged 5 to 19. The columns marked with "son" mean that a family only has sons and those with "daughter" indicate that a family only has daughters. Sex ratio for the 10–19 cohort in 2010 is inferred from the male-to-female ratio among the cohort aged 0 to 9 in 2000 based on China's 2000 population census. Robust standard errors clustered at the county level are in parentheses. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%.

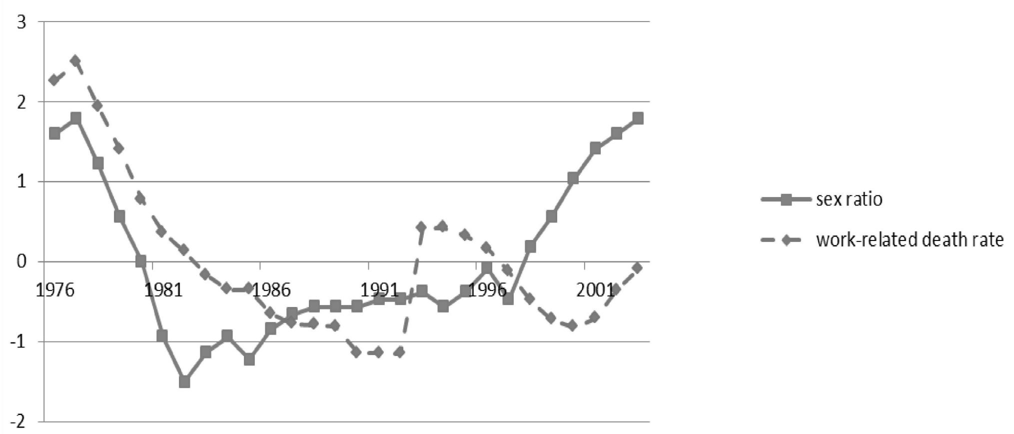
Table 6.4—Sex ratio and work-related injury insurance coverage

Dependent variable: work-related injury insurance coverage ratio (number of staff having this insurance ÷ total staff)								
	Firm-level analysis				County-level mean			
	OLS		Tobit		OLS		Tobit	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sex ratio	-1.327***	-1.232***	-1.031***	-3.360***	-3.270***	-2.595***	-0.775***	-1.312***
(10–19 age cohort)	(0.27)	(0.27)	(0.27)	(0.84)	(0.83)	(0.78)	(0.20)	(0.38)
Union members to total		0.258***	0.216***		0.653***	0.613***		
staff (unionization)		(0.07)	(0.07)		(0.20)	(0.20)		
Female staff		-0.048	-0.096		-0.128	-0.282		
to total staff ratio		(0.13)	(0.13)		(0.35)	(0.32)		
Staff from rural area		0.118	-0.071		0.255	-0.075		
to total staff ratio		(0.15)	(0.13)		(0.41)	(0.36)		
Female union members		0.044	0.043		0.122	0.170		
to total staff ratio		(0.13)	(0.13)		(0.35)	(0.32)		
Union members from		-0.032	0.127		-0.043	0.219		
rural area to total staff		(0.14)	(0.12)		(0.37)	(0.33)		
Ratio of grass-roots staff		0.080	0.072		0.218	0.222		
in board of directors		(0.07)	(0.07)		(0.19)	(0.18)		
Ratio of grass-roots staff		0.305***	0.270***		0.821***	0.731***		
in board of supervisors		(0.05)	(0.05)		(0.17)	(0.17)		
Log sales		0.036***	0.031***		0.094***	0.079***		
		(0.01)	(0.01)		(0.02)	(0.02)		
Operating profits		0.093	0.110		0.202	0.248		
to sales ratio		(0.08)	(0.07)		(0.21)	(0.19)		
Three expenses		0.521***	0.455***		1.542***	1.386***		
to sales ratio		(0.12)	(0.11)		(0.36)	(0.34)		
Short-term debt		0.0186*	0.012		0.0515*	0.035		
to sales ratio		(0.01)	(0.01)		(0.03)	(0.02)		
Fixed assets		-0.0359**	-0.022		-0.129***	-0.0932**		
to sales ratio		(0.02)	(0.02)		(0.05)	(0.05)		
Industry and owner type	no	yes	yes	no	yes	yes	no	no
Province dummies	no	no	yes	no	no	yes	no	no
Observations	3,341	3,329	3,329	3,341	3,329	3,329	463	463

Source: All China Federation of Trade Unions survey (2009).

Notes: OLS = Ordinary least squares. The survey used covers Beijing, Shanghai, Tianjin, Liaoning, Jiangsu, Zhejiang, Guangdong, Hebei, Henan, Hubei, Sichuan, and Shaanxi, to get the firm-level work-related injury insurance information and union-related information. Firm-level financial indicators are matched from the Industrial Firms Census dataset compiled by the National Bureau of Statistics. Sex ratio of the 10–19 age cohort is inferred from the 0–9 age cohort in the population census in 2000. Robust standard errors clustered at the county level are in parentheses. ***, **, ** indicate significance level at 1%, 5%, and 10%, respectively. The upper limit and lower limit of the Tobit model are set at 1 and 0, respectively.

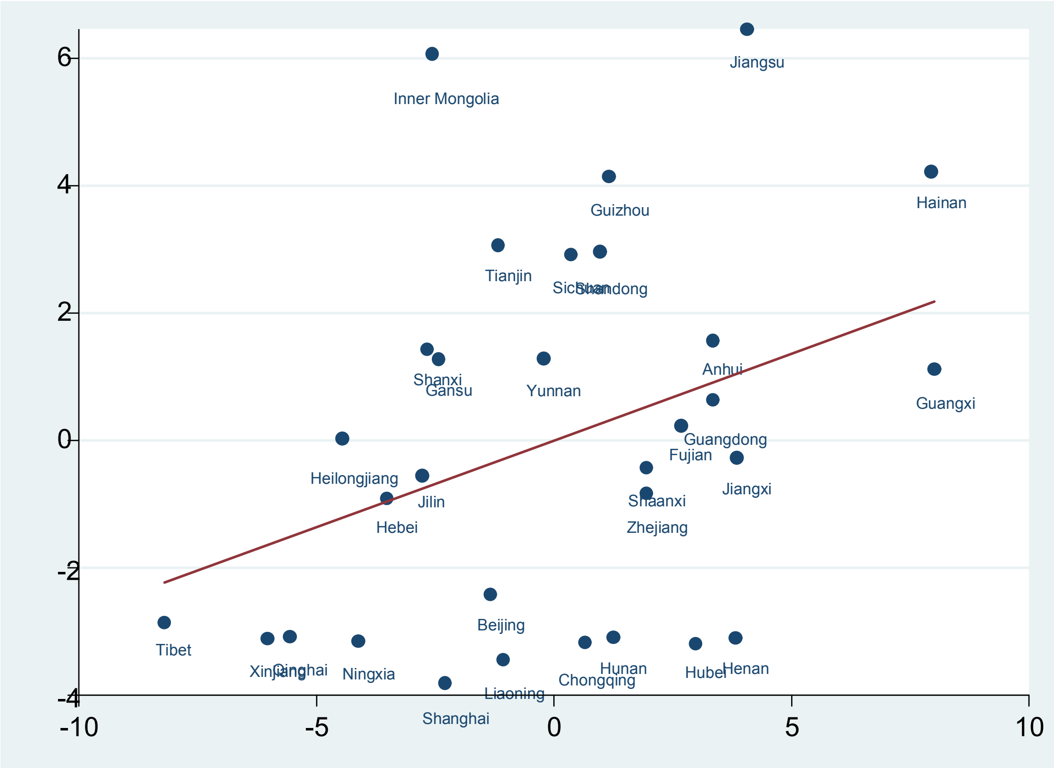
Figure 5.1—Sex ratio and work-related death rate at the national level



Source: China's Work Safety Yearbook (2000-2004) and the China Statistical Yearbook (1978-2004).

Notes: The sex ratio is defined as the ratio at birth 20 years earlier. The sex ratios at birth in 1982, 1990, and 2000 at the national level are published figures from China's population censuses. Since the disaggregate sex ratios at birth in 1980 and 1990 are not publicly available, we use sex ratios for the 20-year-old cohort and 10-year-old cohort, respectively, from the 2000 census to approximate them. Work-related death rate is defined as number of deaths caused by work-related accidents divided by average number of workers in a certain period (unit: 1/1,000,000). Both variables have been rescaled by subtracting the mean and dividing by the standard deviation.

Figure 5.2—Sex ratio and work-related deaths across provinces

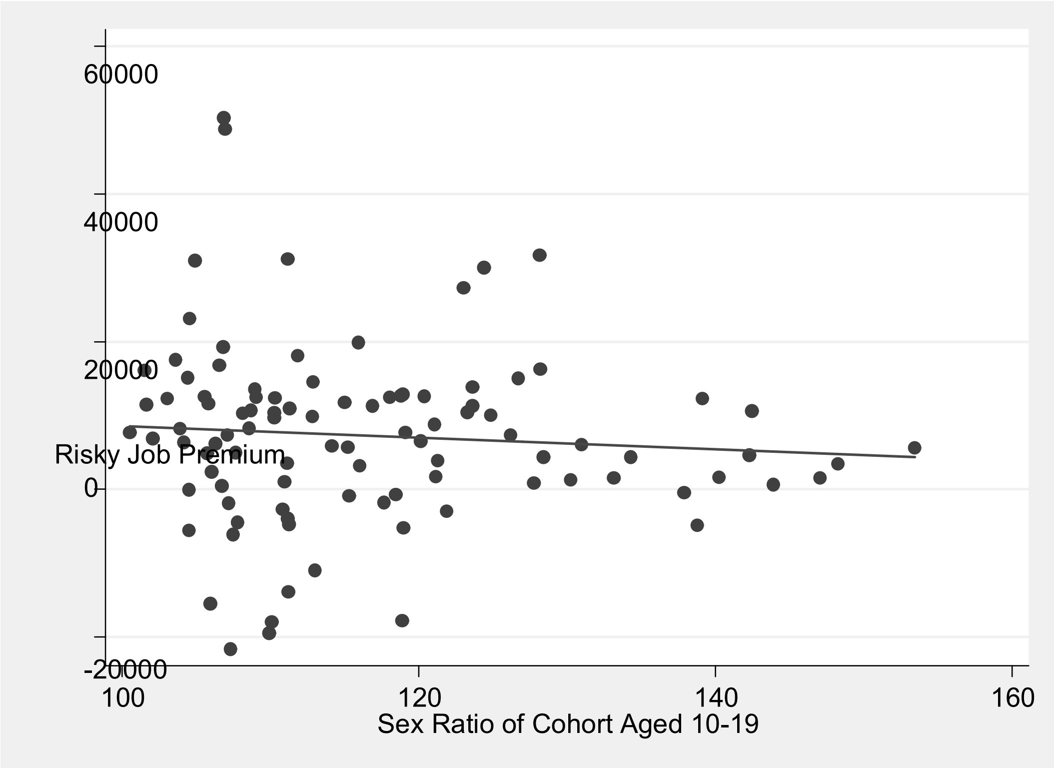


Source: website of the State Administration of Work Safety.

(http://media.chinasafety.gov.cn:8090/iSystem/shigumain.jsp)

Notes: We use the Frisch-Waugh theorem to exclude the impacts of log of gross domestic product (GDP) per capita. The horizontal axis is the residual of the regression of sex ratio on log of GDP per capita, and the vertical axis is the residual of the regression of the share of work-related deaths in each province to nationwide total work-related deaths on log of GDP per capita. Sex ratio of the 10–19 premarital cohort in 2009 is calculated as the male-to-female ratio for the 0–9 cohort in the 2000 population census. In the simple regression of share of work-related deaths on sex ratio (sex ratio and log of GDP per capita), sex ratio is significant at the 5% level.

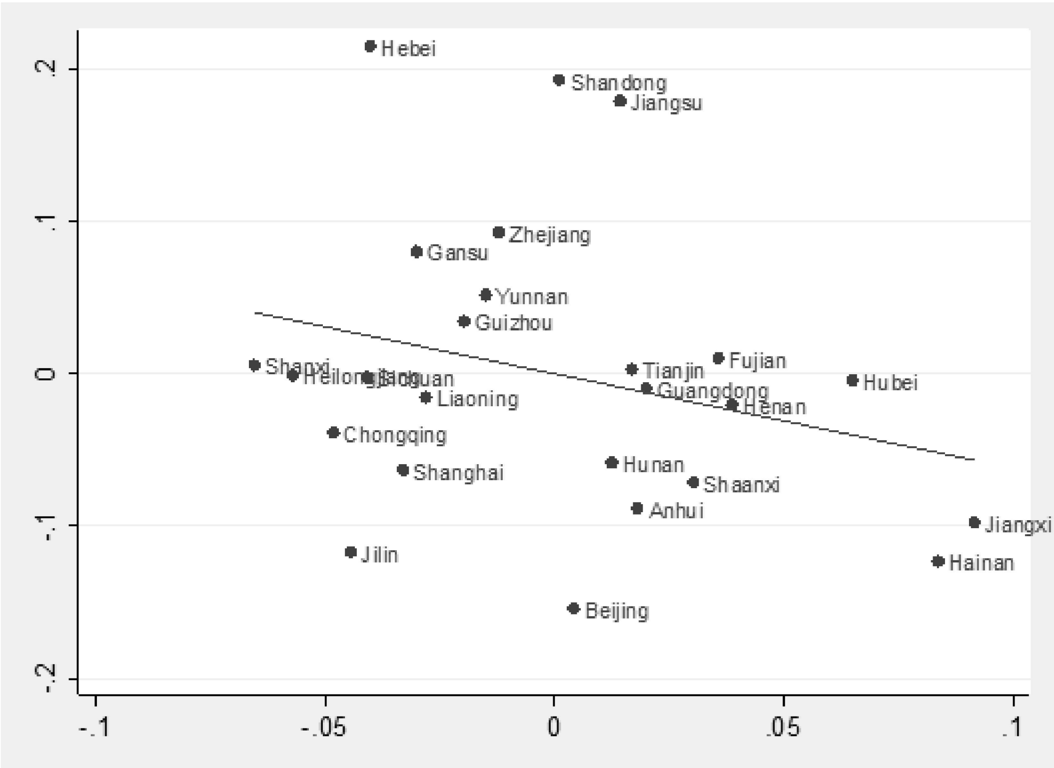
Figure 6.1—Sex ratio and risky job premium



Source: China Family Panel Studies (2010).

Notes: Risky jobs are jobs related to firefighting, public security, mineral mining, well drilling, oil and gas exploration, construction, electricity supply, installation and debugging of electrical equipment, and operation of mechanical equipment, or the working place of which is in the open air or a workshop. The vertical axis is the risky job income premium of the rural parent cohort averaged to the county level. The horizontal axis is the sex ratio for the 10–19 cohort in 2010, inferred from the male-to-female ratio among the cohort aged 0 to 9 in 2000 based on the 2000 China population census.

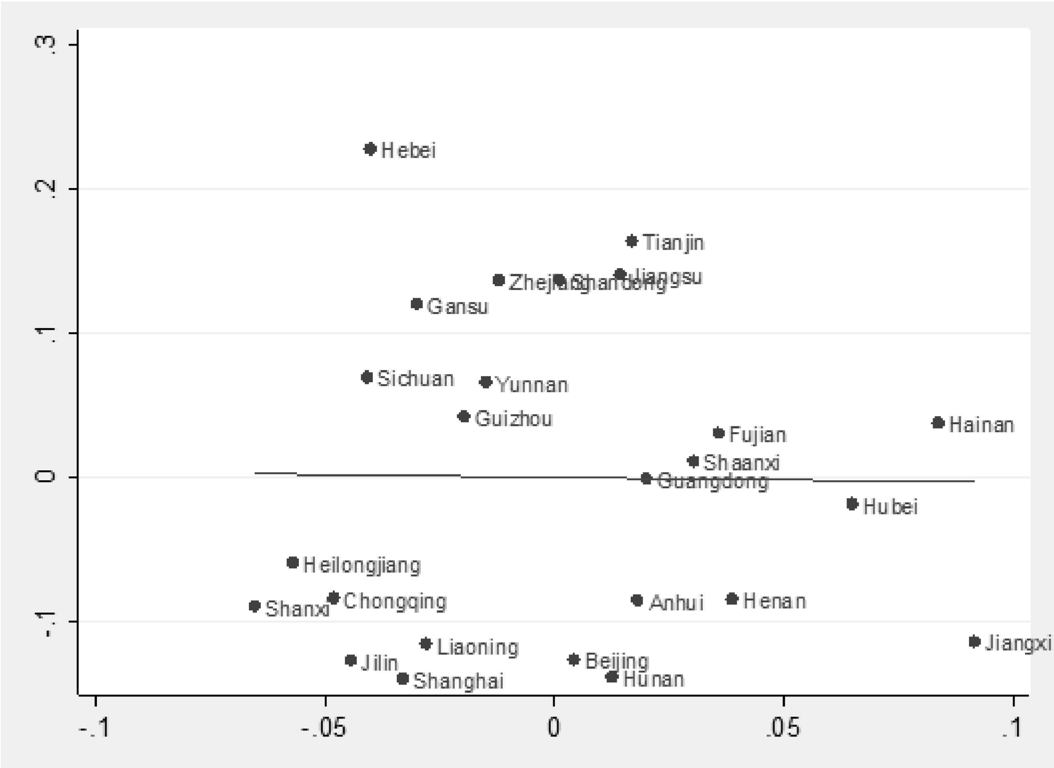
Figure 6.2—Sex ratio and labor safety protection



Source: Chinese Trade Unions Statistics Yearbook (2011).

Notes: We use the Frisch-Waugh theorem to exclude the impacts of log of gross domestic product (GDP) per capita. The horizontal axis is the residual of the regression of sex ratio on log of GDP per capita, and the vertical axis is the residual of the regression of the ratio of trade unions that have set up the labor protection, supervision, and inspection committees (LPSICs) on log of GDP per capita. Sex ratio is calculated as male-to-female ratio for the 10–19 cohort in the 2010 population census.

Figure 6.3—Sex ratio and labor dispute mediation: Placebo test



Source: Chinese Trade Unions Statistics Yearbook (2011).

Notes: We use the Frisch-Waugh theorem to exclude the impacts of log of gross domestic product (GDP) per capita. The horizontal axis is the residual of the regression of sex ratio on log of GDP per capita, and the vertical axis is the residual of the regression of the share of trade unions that have set up the labor dispute mediation committees (LDMCs) on log of GDP per capita. Sex ratio is calculated as male-to-female ratio for the 10–19 cohort in the 2010 population census.

Appendix A: Supplementary Tables

Table A.1 Sex ratio and accidental death by gender: OLS regressions

	Year 1991				Pooled 1991–2000			
	Accident 1		Accident 2		Accident 1		Accident 2	
	Male	Female	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Whole sample							
Sex ratio for 5–19 cohort	0.04***	0.01	0.07***	0.04*	0.02*	0.01***	0.03***	0.02*
	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)	(0.00)	(0.01)	(0.01)
Observations	14790	11645	14790	11645	148130	113978	148130	113978
Adjusted R^2	0.05	0.05	0.07	0.08	0.05	0.04	0.09	0.07
	Parents' cohort							
Sex ratio for 5–19 cohort	0.15***	0.01	0.25***	0.03	0.04	0.06*	0.08**	0.01
	(0.02)	(0.03)	(0.04)	(0.02)	(0.03)	(0.03)	(0.03)	(0.02)
Observations	1587	1016	1587	1016	12259	7027	12259	7027
Adjusted R^2	0.06	0.05	0.10	0.04	0.01	0.01	0.01	0.01

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square. The parents' cohort age is between 25 and 44. The dependent variable equals 1 if a person died of accident 1 or accident 2 and 0 otherwise. Accident 1 and accident 2 indicate death caused by fire and flames and accidental falls, respectively. We have rescaled all the independent variables by dividing 100. For the regression on year 1991, we define the sex ratio as the male to female ratio among age cohort 5-19, calculated from the 1990 population census. In the pooled sample, the sex ratios for the period 1991–1995 and for the period 1996–2000 refer to the male-to-female ratios among the cohorts 5–19 and 0–14 inferred from the 1990 population census, respectively. In the regressions, we also control for wave dummy, which equals 1 in the period 1991–1995 and 0 in the period 1996–2000, and other individual-level characteristics, such as ethnicity (dummy variables for Han, Hui, Mongolian, Tibetan, Uigur, Zhuang, others), occupation (dummy variables for production, farmer, sheepherder, fisherman, medical staff, food preparation and serving, business and social service, manager, retired, housework, jobless, children at kindergarten, children not at kindergarten, student of primary school, student of middle and high school, undergraduate or more, others) and marital status (dummy variables for married, single, widowed, divorced, others). We do not report the coefficients for these variables for brevity. Labor participation rate is defined as the share of prime-age (15–64) working population in total population. The regional dummy variables *East* and *Central* indicate whether the county is in the eastern or central region, with the western region as the base group. Robust standard errors clustered at the county level are in parentheses. The symbols ***, **, and * indicate significance level at 1%, 5%, and 10%. .

Appendix Table A.2—Sex ratio and accidental death (without distinction between the two types of accidents): OLS regressions

Dependent variable: accident						
	Year 1991			Pooled 1991–2000		
	Total	Male	Female	Total	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
	Whole sample					
Sex ratio for 5-19 cohort	0.08***	0.11***	0.05**	0.04***	0.05***	0.02***
	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)
Observations	26436	14791	11645	262117	148138	113979
Adjusted R^2	0.01	0.01	0.01	0.01	0.01	0.01
	Parents' cohort					
Sex ratio for 5-19 cohort	0.26***	0.39***	0.04	0.10**	0.12**	0.07*
	(0.03)	(0.05)	(0.03)	(0.04)	(0.05)	(0.04)
Observations	2603	1587	1016	19286	12259	7027
Adjusted R^2	0.06	0.09	0.07	0.01	0.01	0.01

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square. Accident equals 1 if a person died of fires and flames or accidental falls and 0 otherwise. Other regressors are the same as those in Table A.1.

Appendix Table A.3—Sex ratio and accidental death of parents' cohort (without distinction between the two types of accidents): OLS regressions in interior regions

Dependent variable: accident						
	Year 1991		Pooled 1991–2000		Pooled 1991–1996	
	Sample 1	Sample 2	Sample 1	Sample 2	Sample 1	Sample 2
	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio for 5–19 cohort	0.23***	0.25***	0.08**	0.07*	0.14**	0.14**
	(0.03)	(0.03)	(0.04)	(0.04)	(0.06)	(0.06)
Observations	2349	1874	21785	18097	13607	11043
Adjusted R^2	0.07	0.10	0.01	0.01	0.01	0.01

Source: National Disease Surveillance Points System (1991-2000).

Notes: OLS = ordinary least square. Accident equals 1 if a person died of accident due to fires and flames or accidental fall and 0 otherwise. Other regressors are the same as those in Table 5.3. In sample 1, we drop observations in Guangdong, Zhejiang, and Jiangsu, the three largest migrant destinations of China. In sample 2, we drop observations in Shandong, Jiangsu, Shanghai, Zhejiang, Fujian, and Guangdong, the rich and coastal areas of China.

Appendix B: Model

In this appendix, we set up a stylized model to illustrate the impact of sex ratio imbalance on the work-related deaths/injuries and its underlying mechanisms.

Consider an economy which is composed of families and firms. Unmarried children are not old enough to work. Parents work and earn income in order to help their unmarried children win a competitive edge in the marriage market. We first discuss the situations for parents with unmarried sons. Behaviors for parents with only girls will be explored in the following. Consider a representative parent worker and a firm (employer). The worker chooses how many efforts, denoted by e, that he would like to devote to the firm. The firm pays the worker wage, w, to compensate him and spends m to invest in workplace safety. The wage is a function of e and m. We assume that the harder the worker works, the higher the wage is but the marginal return is decreasing. That is, $w_1(e, m)>0$ and $w_{11}(e, m)<0$ (the subscript denotes the partial derivative with respect to the corresponding argument). Moreover, when the firm invests more in safety, it will offer lower wage for the same e. That is, $w_2(e, m)<0$. In addition, we assume that investment in workplace safety will increase the marginal benefit of efforts on wage and the absolute value of marginal effect on wage of investment in safety decreases with such investment. That is, $w_{12}(e,m)>0$ and $w_{22}(e,m)>0$ (note that $w_2(e,m)<0$). Furthermore, $w_{22}(e,m)$ increases with m, i.e., $w_{222}(e,m)>0$. Besides, we assume that the degree of complementarity of safety investment and hard work decreases with safety investment, and the decrease of marginal return to efforts gets slower when the firm invests more in workplace safety. That is, $w_{212}(e,m)<0$ and $w_{112}(e,m)>0$. For given e and m, with probability $f(e, m)$, the worker runs into an accident and suffers from a damage $D>0$. D can be interpreted as the cost to treat the injury due to accidents. Therefore, the expected income Y is $w(e, m)-f(e,m)D$. We assume that when the worker works harder, it is more easily to incur an accident and the marginal increase in risk gets larger with the increase of efforts. Besides, when the firm invests more in workplace safety, the probability of confronting with an accident is lower. That is, $f_1(e,m)>0$, $f_{11}(e,m)>0$, $f_2(e,m)<0$. Moreover, marginal effect on risks of hard work decreases with the investment in workplace safety and the absolute value of the marginal effect of investment in safety on probability of accidents is decreasing. That is, $f_{12}(e,m)<0$ and $f_{22}(e,m)>0$ (note that $f_2(e,m)<0$).

Following Frank (1985), Hopkins and Kornienko (2004), workers not only care about the absolute utility from income and damage to their health, but also about their relative position in the population. Let $S(Y)$ denotes the probability that the worker's expected income falls below or equals Y. The higher this probability is, the higher the utility is for the worker with income Y. We assume that harder work will bring in larger Y. That is, $w_1-f_1D>0$. Furthermore, the more the firm invests in workplace safety, the less the expected income is and the absolute value of marginal effect of investment in safety on expected income increases with such investment. That is, $w_2-f_2D<0$ and $w_{22}-f_{22}D<0$.

Therefore, the utility function takes the following form.

$$U = V[w(e,m) - f(e,m)D, f(e,m)D] + \alpha S[w(e,m) - f(e,m)D]$$

V is the usual utility function which satisfies $V_1>0$, $V_2<0$, $V_{1e}<0$, $V_{2e}<0$, $V_{1m}>0$, $V_{2m}>0$. $\alpha > 0$ is the weight that assigns to the relative status. We assume that the more imbalanced the pre-marital sex ratio of males to females (denoted by r, $r>1$) is, the larger α is. In other words, when their sons face more competition in the marriage market, parents care more about their relative income in the society. Sons in families with higher income are more attractive and will find a wife to get married. The sons in the poorest families are not able to get married. So when the sex ratio becomes larger, the relative income to others gets more important.

The firm chooses m to minimize its total cost.

Min $w(e,m)+m$

From the first order condition, we have $w_2(e,m)+1=0$.

The worker chooses e to maximize utility. We assume that the expected income, $w(e,m)-f(e,m)D$ follows a uniform distribution on [0,1], so the utility function can be simplified to:

$$U = V[w(e,m) - f(e,m)D, f(e,m)D] + \alpha[w(e,m) - f(e,m)D]$$

The worker maximizes his utility taking the firm's optimization into account. His optimization problem is as follows:

$$U = V[w(e,m) - f(e,m)D, f(e,m)D] + \alpha[w(e,m) - f(e,m)D]$$

s.t. $w_2(e,m)+1=0$

We can construct the Lagrangian function as follows:

$$L = V[w(e,m) - f(e,m)D, f(e,m)D] + \alpha[w(e,m) - f(e,m)D] + \lambda[w_2(e,m) + 1], \ \lambda < 0$$

The first order conditions with respect to e and m are:

$$E_1 \equiv V_1(w_1 - f_1D) + V_2 f_1 D + \alpha(w_1 - f_1D) + \lambda w_{21} = 0$$

$$E_2 \equiv V_1(w_2 - f_2D) + V_2 f_2 D + \alpha(w_2 - f_2D) + \lambda w_{22} = 0$$

From E_1 and E_2, we can substitute λ and obtain that:

$H_1 \equiv Aw_{22} - Bw_{21} = 0$,

in which $A \equiv V_1(w_1 - f_1D) + V_2f_1D + \alpha(w_1 - f_1D)$ and $B \equiv V_1(w_2 - f_2D) + V_2f_2D + \alpha(w_2 - f_2D)$.

The first order condition with respect to λ obtains $H_2 \equiv w_2 + 1 = 0$.

We have A>0 and B>0 from E_1 and E_2 and relevant partial derivatives can be calculated as follows.

$$\frac{\partial A}{\partial e} \equiv A_e = V_{1e}(w_1 - f_1D) + V_1(w_{11} - f_{11}D) + V_{2e}f_1D + V_2f_{11}D + \alpha(w_{11} - f_{11}D)$$

$$\frac{\partial A}{\partial m} \equiv A_m = V_{1e}(w_2 - f_2D) + V_1(w_{21} - f_{21}D) + V_{2e}f_2D + V_2f_{21}D + \alpha(w_{21} - f_{21}D)$$

$$\frac{\partial B}{\partial e} \equiv B_e = V_{1m}(w_1 - f_1D) + V_1(w_{12} - f_{12}D) + V_{2m}f_1D + V_2f_{12}D + \alpha(w_{12} - f_{12}D)$$

$$\frac{\partial B}{\partial m} \equiv B_m = V_{1m}(w_2 - f_2D) + V_1(w_{22} - f_{22}D) + V_{2m}f_2D + V_2f_{22}D + \alpha(w_{22} - f_{22}D)$$

$$\frac{\partial H_1}{\partial e} \equiv H_{1e} = A_e w_{22} + Aw_{221} - B_e w_{21} - Bw_{211}$$

$$\frac{\partial H_1}{\partial m} \equiv H_{1m} = A_m w_{22} + Aw_{222} - B_m w_{21} - Bw_{212}$$

$$\frac{\partial H_1}{\partial \alpha} \equiv H_{1\alpha} = (w_1 - f_1D)w_{22} - (w_2 - f_2D)w_{21}$$

$$\frac{\partial H_2}{\partial e} \equiv H_{2e} = w_{21}$$

$$\frac{\partial H_2}{\partial m} \equiv H_{2m} = w_{22}$$

$$\frac{\partial H_2}{\partial \alpha} \equiv H_{2\alpha} = 0$$

Since w_1-f_1D>0, V_{1e}<0, V_1>0, V_2<0, V_{2e}<0, f_1>0, w_{11}<0, f_{11}>0, D>0, we have A_e<0.

Since w_2-f_2D<0, V_{1e}<0, V_1>0, w_{21}>0, f_{21}<0, V_2<0, V_{2e}<0, f_2<0, D>0, we have A_m>0.

Since V_{1m}>0, w_1-f_1D>0, V_1>0, w_{12}>0, f_{12}<0, V_{2m}>0, f_1>0, V_2<0, D>0, we have B_e>0.

Since $V_{1m}>0$, $w_2-f_2D<0$, $V_1>0$, $w_{22}-f_{22}D<0$, $V_{2m}>0$, $f_2<0$, $V_2<0$, $f_{22}>0$, $D>0$, we have $B_m<0$.

Since $A_e<0$, $w_{22}>0$, $A>0$, $w_{221}=w_{212}<0$, $B_e>0$, $w_{21}>0$, $B>0$, $w_{211}=w_{112}>0$, we have $H_{1e}<0$.

Since $A_m>0$, $w_{22}>0$, $A>0$, $w_{222}>0$, $B_m<0$, $w_{21}>0$, $B>0$, $w_{212}<0$, we have $H_{1m}>0$.

Since $w_1-f_1D>0$, $w_{22}>0$, $w_2-f_2D<0$, $w_{21}>0$, we have $H_{1\alpha} > 0$.

Since $w_{21}>0$, $w_{22}>0$, we have $H_{2e}>0$, $H_{2m}>0$.

Taking total derivatives of H_1 and H_2 with respect to α, we have:

$$H_{1e}\frac{\partial e^*}{\partial \alpha}+H_{1m}\frac{\partial m^*}{\partial \alpha}+H_{1\alpha}=0$$

$$H_{2e}\frac{\partial e^*}{\partial \alpha}+H_{2m}\frac{\partial m^*}{\partial \alpha}+H_{2\alpha}=0$$

Solving these two equations, we can obtain (let * denotes the optimal value):

$$\frac{\partial e^*}{\partial \alpha}=\frac{H_{2m}H_{1\alpha}}{H_{2e}H_{1m}-H_{1e}H_{2m}}>0$$

$$\frac{\partial m^*}{\partial \alpha}=\frac{H_{2e}H_{1\alpha}}{H_{1e}H_{2m}-H_{2e}H_{1m}}<0$$

Therefore, $\frac{\partial e^*}{\partial r}=\frac{\partial e^*}{\partial \alpha}\frac{\partial \alpha}{\partial r}>0$

We then come to the behaviors for parents with only unmarried daughters. The set up is the same for them except that the more imbalanced the pre-marital sex ratio of males to females is, the smaller α is. In other words, when r gets larger, they have more bargaining power in the marriage market and care less about their relative income. Similar analysis as before shows that the larger the sex ratio is, the fewer efforts they will spend on the job.

Therefore, we have proposition 1.

Proposition 1: The larger the pre-marital sex ratio is, the harder parents with sons will work. The opposite applies to parents with only daughters.

Furthermore, for parents with sons, $\frac{\partial m^*}{\partial r} = \frac{\partial m^*}{\partial \alpha} \frac{\partial \alpha}{\partial r} < 0$.

Since f_1>0, and f_2<0, $\frac{\partial f(e^*, m^*)}{\partial r} = \frac{\partial f(e^*, m^*)}{\partial e^*} \frac{\partial e^*}{\partial r} + \frac{\partial f(e^*, m^*)}{\partial m^*} \frac{\partial m^*}{\partial r} > 0$. That is, the more imbalanced the pre-marital sex ratio is, the more frequently they will run into accidents. In other words, they will endure riskier working conditions.

For parents with only daughters, similar analysis as before demonstrates that the larger the pre-marital sex ratio is, the less frequently they will be involved in accidents. In other words, they will bear fewer risks of accidents.

Therefore, we have proposition 2.

Proposition 2: The larger the pre-marital sex ratio is, the more risks parents with sons will take. The opposite applies to parents with only daughters.

When we sum up the efforts in the whole economy, since there are more families with unmarried sons, the total efforts, e_T, will increase when the sex ratio of pre-marital cohort gets larger. That is, $\frac{\partial e_T^*}{\partial r} > 0$.

In the same vein, $\frac{\partial m_T^*}{\partial r} < 0$. We therefore have proposition 3.

Proposition 3: The larger the pre-marital sex ratio is, the less the firm will invest in workplace safety.

Furthermore, since $\frac{\partial e_T^*}{\partial r} > 0$, we have:

$$\frac{\partial f(e_T^*, m_T^*)}{\partial r} = \frac{\partial f(e_T^*, m_T^*)}{\partial e_T^*} \frac{\partial e_T^*}{\partial r} + \frac{\partial f(e_T^*, m_T^*)}{\partial m_T^*} \frac{\partial m_T^*}{\partial r} > 0$$

That is, the larger the pre-marital sex ratio is, the higher the probability of work-related accidents is. We summarize it in the following proposition.

Proposition 4: In regions with a more skewed pre-marital sex ratio, the probability of work-related deaths and injuries is higher.

To summarize, proposition 4 demonstrates the mortality cost of imbalanced sex ratio and propositions 1-3 reveal the underlying mechanisms. When the competition in the marriage market becomes more acute for families with sons, they will respond by working harder and taking more risks so as to gain a relatively higher position in the population. Taking this into account, firms will under-invest in workplace safety to minimize their cost. As a result, the probability of work-related deaths and injuries is higher in a region with a more skewed pre-marital sex ratio.