BREAKING THE "IRON RICE BOWL" AND PRECAUTIONARY SAVINGS: EVIDENCE FROM CHINESE STATE-OWNED ENTERPRISES REFORM

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ABSTRACT. We use China's large-scale reform of state-owned enterprises (SOEs) in the late 1990s as a natural experiment to identify and quantify the importance of precautionary saving for wealth accumulation. Before the reform, SOE workers enjoyed similar job security as government employees. Following the reform, over 35 million SOE workers were laid off, although government employees kept their "iron rice bowl." The change in unemployment risk for SOE workers relative to that of government employees before and after the reform provides a clean identification of income uncertainty that helps us estimate the importance of precautionary saving. In our estimation, we correct a self-selection bias in occupational choice and disentangle the effects of uncertainty from pessimistic outlook. We obtain evidence that precautionary savings account for about 30 percent of the wealth accumulation for SOE workers between 1995 and 2002.

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I. Introduction

Precautionary savings are potentially important for wealth accumulation, especially for a country like China that has experienced large structural changes associated with policy reforms, which may have led to increases in economic uncertainty.

However, to identify and quantify the importance of precautionary saving can be challenging. In particular, it is difficult to clearly identify observable and exogenous sources of income risks that vary significantly across households (Lusardi, 1998; Carroll and Kimball, 2006). Many studies use the cross-sectional variance of income as a proxy for income uncertainty (Carroll and Samwick, 1998). But this proxy may be subject to measurement errors and potential endogeneity bias (Kennickell and Lusardi, 2005).

To quantify the importance of precautionary saving also requires correcting a self-selection bias. Theory implies that precautionary wealth accumulation depends not just on risk, but also on risk preferences (Caballero, 1990, 1991). An individual with high risk aversion has an incentive to choose a job with low income risk. Similarly, a worker with low risk aversion may want to choose a job with high income risk (with potentially high expected income). Failing to control for self-selection in occupational choices may lead to a significant downward bias in estimating the importance of precautionary saving (Fuchs-Schündeln and Schündeln, 2005).

Partly reflecting the difficulties in measuring income uncertainty and correcting self-selection biases, the existing literature obtains mixed evidence of precautionary saving. For example, some studies report weak or no evidence of precautionary saving (Dynan, 1993; Guiso, Jappelli, and Terlizzese, 1992), while some other studies attribute a large fraction (50% or more) of household wealth accumulation to precautionary savings (Carroll and Samwick, 1998; Gourinchas and Parker, 2002).

In this paper, we present a new approach to identifying and quantifying the contribution of precautionary saving to wealth accumulation. We use the large-scale reform of state-owned enterprises (SOEs) in China in the late 1990s as a natural experiment to achieve identification. Before the reform, workers in SOEs and in the government sector (GOV) had similar job security, including guaranteed pensions and near-free health care and housing. In this sense, workers in both sectors held an “iron rice bowl” before the reform. After the reform, however, over 35 million workers in the SOEs were laid off between 1995 and 2002. Those workers lost not just their jobs, but also the associated benefits. In contrast, workers in the government sector—where few layoffs occurred—were little affected by the reform; they were able to hold on to their iron rice bowl. The massive layoffs in the SOEs significantly changed the perceived income risk for the remaining SOE workers. The reform was exogenous and largely unexpected to individual workers, and it created significant variations of unemployment risk for workers across the SOE and GOV sectors. Thus, the reform provides us with a clean identification of relative income risk and perceived job uncertainty.

To estimate the importance of precautionary saving, we use data from the Chinese Household Income Project (CHIP) survey. We focus on the years 1995 and 2002. The large-scale SOE reform started to have significant impact on SOE employment in 1997, with the effects gradually phasing out by 2002. Our sample thus covers both the pre- and post-reform periods. To identify and quantify the contribution of precautionary saving to wealth accumulation, we exploit the differences in saving behavior both across sectors (SOE vs. GOV) and across time (before and after the SOE reform)—a difference-in-differences (DID) approach. The time variations (between 1995 and 2002) of the relative saving behavior of workers in the two sectors capture the magnitude of precautionary saving stemming from the breaking of the iron rice bowl.

To control for self-selection biases that may arise from correlations between occupational choices and individual workers' risk attitude, we use an important feature of the CHIP survey. The survey contains a question about how a worker obtained her current job. Some workers find jobs through a search and matching process; but in our sample, a majority of workers (over 70 percent) have jobs assigned by the government. For assigned jobs, the government has the final power to determine the worker's occupation and compensation. Thus, for those workers whose jobs are assigned by the government, the occupational choice is likely unrelated to worker preferences. We restrict our sample to include only those households whose jobs were assigned by the government. This restriction helps us control for the effects of self-selection in occupational choice.

The SOE reform might affect not only the perceptions of future income uncertainty, but also the expected future income growth rate. For example, after witnessing the impact of the reform on the relative job security, an SOE worker might expect not only an increase in future income risks but also a decline in future income levels. Declines in expected income would also raise current saving. Such saving behavior, however, is driven by the worker's desire for intertemporal consumption smoothing (i.e., the PIH effect), not by precautionary motives.

To isolate the effects of precautionary motives on saving from the PIH effects, we use a unique question in the 2002 CHIP survey that asks households about their expectations of income paths in the next five years. We restrict our sample to include only those workers who do not expect income to decline. This approach enables us to mitigate the effects of the PIH channel that could cause an upward bias in the estimation of precautionary saving.

By identifying job uncertainty caused by the SOE reform, correcting the self-selection bias in occupational choices, and controlling for PIH effects, we obtain estimates of precautionary savings that are significant both statistically and economically. We estimate that precautionary savings accounted for about 30 percent of total financial wealth accumulations for SOE workers during the period from 1995 to 2002.

II. Related Literature

The main contribution of our paper is that we identify and quantify the importance of precautionary saving using the large-scale SOE reform in China as a natural experiment. We further exploit the microeconomic details of the CHIP survey data to correct the self-selection bias in occupational choices and to disentangle the effects of uncertainty from pessimistic outlooks.

Our approach to correcting self-selection biases is complementary to that used by Fuchs-Schündeln and Schündeln (2005), who use the event of the German reunification as a natural experiment to identify the presence of self selection. Before the reunification, civil servants in East Germany had government-assigned jobs, while civil servants in West Germany were free to choose their occupations (and they chose to work in the government). Therefore, the occupational choices for civil servants in East Germany were not subject to self-selection, but those in West Germany were. The reunification significantly increased income risks for all workers in the former East Germany but civil servants, most of whom were able to keep their jobs in the newly unified Germany. By comparing wealth accumulations of civil servants with those of individuals in other occupations between East and West Germany, they identify the effects of self-selection on precautionary saving.

Our approach to controlling for self-selection biases is different. We focus our sample on workers whose jobs were assigned by the government. In our sample, about 83% of jobs were assigned by the government in 1995 and 72% in 2002. For assigned jobs, the government has the final power to determine the worker's occupation and compensation. Thus, for those workers whose jobs are assigned by the government, the occupational choice is likely unrelated to the risk attitudes of individual workers. Despite the differences in approach and the data set used, our results are similar to those obtained by Fuchs-Schündeln and Schündeln (2005), both indicating that self-selection can cause a significant downward bias for estimating the quantitative importance of precautionary saving.

Our work also contributes to a growing literature that attempts to quantify the importance of precautionary saving for explaining China's rising saving rate. Earlier studies obtain mixed evidence that supports the precautionary-saving view (Kraay, 2000; Carroll, Dynan, and Krane, 2003; Meng, 2003).

A more recent study by Chamon and Prasad (2010) uses the annual Chinese Urban Household Survey from 1995 to 2005 to disentangle different motives behind the rising urban household saving rate. Chamon and Prasad (2010) find that the increases in private burden of education, health and housing expenditures seem among the strongest candidates for explaining the increases in saving rates. Chamon, Liu, and Prasad (2013) document a sharp increase in income uncertainty associated with increases in transitory idiosyncratic shocks among Chinese urban households. They argue that rising income uncertainty induces younger households to raise their saving rate significantly.

An alternative explanation for the rising Chinese saving rate is provided by Wei and Zhang (2011), who present evidence that sex imbalances caused by China's one-child policy have induced a "competitive savings motive": with a shortage of girls, parents with a son save more to increase the relative attractiveness of their son in a tighter marriage market. They show that this competitive savings motive is empirically important.

Complementary to this literature, we provide evidence that increases in economic uncertainty associated with large structural changes in China have contributed to substantial precautionary wealth accumulation.

III. Some Background of the SOE Reform

From 1949 to 1978, China's economy was under a central-planning regime. The government maintained tight controls over production and factor allocations. Most jobs were assigned by the government. To support the goal of industrialization, workers were paid subsistence wages and, in exchange, they were guaranteed life-time employment along with near-free housing, education, health care, and pension (Cai, Park, and Zhao, 2008). This cradle-to-grave regime is known as the "iron rice bowl," which has long been advocated as one advantage of China's socialist system.

The "open door" economic policy and nationwide reform initiated by Deng Xiaoping in the late 1970s initiated China's transition to a free-market economy. In the mid-1980s, the government introduced a labor contract system, under which workers were permitted to search for jobs and employers gained some flexibility in hiring (Meng, 2000). These reform policies led to a large-scale urban migration and increased competition facing SOEs. Following Deng Xiaoping's tour of the south in 1992, more liberalization policies were adopted by the government, leading to a boom in urban economies, which further intensified competition for SOEs. At that time, with soft budget constraints and the requirement to implement the government's goal of full-employment, the SOE sector had substantial redundant labor and many SOE firms were making losses. In 1995 and 1996, around 50% of the SOEs (mostly small or medium sized) reported losses (Meng, 2003). The Asian financial crisis in 1997 exacerbated the situation.

The Chinese government was forced to take actions to improve efficiency of the SOEs and to stem losses. Specific actions were laid out at the Fifteenth Communist Party Congress held in September 1997. A central spirit of the restructuring policy was to "grasp the large and let go of the small." Large (and usually more profitable) SOEs in strategic sectors such as electricity, oil, raw materials, and telecommunications were corporatized and maintained under state controls, while smaller (and often loss-making) SOEs were either privatized or let go bankrupt (see Hsieh and Song (2013)). These policy changes led to a massive layoff (*xia gang* in Chinese) of SOE workers starting in 1997, the scale of which was unprecedented. In 1997, a cumulative of about 6.92 million SOE workers were laid off. The wave of layoffs reached a peak in 1999 and 2000, each year with over 6.5 million SOEs workers losing their jobs. The massive layoffs started to subside by 2002. During the 5-year period from 1997 to 2002, a remarkable total of 35.52 million SOE workers had been laid off (Cai, Park, and Zhao, 2008).

There is evidence that the SOE layoffs were concentrated in small and loss-making firms and in some demographic groups. For example, female, less educated, less skilled, less healthy workers, and non-members of the communist party were more likely to be laid off than others. Workers in SOEs owned by local governments were also more likely to be laid off than those in SOEs owned by the central government (Appleton, Knight, Song, and Xia, 2002). However, the scale and the breadth of the layoffs were largely unexpected by individual workers (see Appendix A for a case study of the SOE layoff experience). Thus, for the SOEs workers who were fortunate to keep their jobs, the reform that broke the iron rice bowl had led to significant changes in their perceptions about future job security and substantially increased their perceived income uncertainty.

IV. Data and Empirical Strategy

IV.1. **Data.** The data that we use are taken from Chinese Household Income Project (CHIP) surveys. Those surveys were conducted by the Chinese Academy of Social Science (CASS) and National Bureau of Statistics (NBS) through a series of questionnaire-based interviews done in rural and urban areas in China in 1988, 1995, 2002 and 2007. The households in each survey are randomly selected following a strict sampling process so that they are nationally representative. The surveys cover a sample of about 15,000 to 20,000 households in about 10 provinces in China. The surveys contain detailed data on rural and urban households' economic status, employment, levels of education, sources of income, household compositions, household expenditures and wealth. The CHIP data have been frequently used in the empirical literature (e.g., Wei and Zhang 2011).

In this paper, we focus on the sample of urban households in the CHIP surveys of 1995 and 2002. These two years span the period of China's large-scale SOE reforms that had led to massive layoffs in the SOEs. As we described in Section III, before the reform, workers in SOEs had similar life-long employment status as those in GOV sectors; in both sectors, workers faced little unemployment risks and income uncertainty. However, since the reform started in 1997, a large number of SOE workers were laid off while GOV workers were able to keep their iron rice bowl. The reform thus injected substantial unemployment risks to SOE workers relative to GOV workers. The different impact of the reform on workers across the two sectors provides an ideal "natural experiment" for us to identify precautionary saving due to a sudden and substantial increase in unemployment risks.

To estimate the quantitative importance of precautionary saving, we exploit changes in saving behaviors associated with the SOE reform between SOE workers and GOV workers. Thus, we restrict our sample to include only those households whose heads work in either the SOE sector or the GOV sector. The SOE sector includes firms that are directly owned by the government (including central, provincial, and local governments), those in which the government holds a controlling share of stocks, and those under collective ownership. The GOV sector includes all levels of government and public institutions.[1] We further restrict our sample to include prime-age workers, whose ages are between 25 and 55 years. This choice is partly driven by concerns of measurement errors in wealth and permanent income for younger workers. It is also driven by concerns that the saving behaviors of workers close to retirement ages change dramatically for reasons more closely related to life-cycle factor than to income uncertainty (Carroll and Samwick, 1998; Gourinchas and Parker, 2002).[2]

With these sample restrictions, we end up with 4390 household-level observations in 1995, consisting of 2977 SOE workers and 1413 GOV employees; and in 2002, we have 3027 observations consisting of 1702 SOE workers and 1325 GOV employees.

IV.2. **Construction of Variables.** In this subsection, we discuss the construction of all variables used in our empirical studies. Table 1 provides a brief description of these variables;

[1]According to the *China Labor Statistics Year Book*, the SOE and the GOV sectors together employed about 94.1% of total urban workers in 1995. This share declined to 75.5% in 2002. During this period, however, the large-scale SOE reform has led to a substantial decline in the relative share of employment in the SOE sector from 70.5% to 42.4%.

[2]The normal retirement age for female workers in China is between 50 and 55; for male workers, it is between 55 and 60.

Table 2 reports summary statistics of the full sample; and Table 3 compares some key characteristics between GOV and SOE workers.

To estimate the quantitative importance of precautionary saving, we focus on the empirical relation between a measure of savings and a measure of income uncertainty, while controlling for a few demographic characteristics.

We measure household saving behavior by the ratio of financial wealth to permanent income (i.e., the W/P ratio; see Table 1 for a description of these variables).[3] We use the stock of financial wealth reported in the surveys instead of the flow of saving or the saving rate to measure an individual's saving behavior for two reasons. First, unlike the flow of saving, financial wealth is not influenced by high-frequency fluctuations in income and expenditures. Thus, it is better able to capture long-run (or average) saving behavior in which we are interested. Second, financial wealth is a direct measure of cumulative savings and is thus less subject to measurement errors than the flow of saving or the saving rate, which are indirectly calculated based on income and consumption expenditures. Moreover, our measure of financial wealth includes mostly liquid assets, which are relevant for studying precautionary saving (Carroll and Samwick, 1998).

We normalize financial wealth by permanent income to obtain a measure of average savings. We thus need to construct a measure of permanent income. In the CHIP datasets, survey participants report incomes earned by household heads during the current year and the recent past. In the 1995 survey, we observe household head earnings from 1990 to 1995; in the 2002 survey, we observe earnings from 1998 to 2002.[4] We use this information to construct a measure of permanent income following a similar approach used by Fuchs-Schündeln and Schündeln (2005). This is done in three steps. First, we calculate a household head's earnings relative to the average earnings of all households in each year with reported earnings. Second, we take the time-series average of the household relative earnings. Third, we multiply the household head's income in each of the survey years (1995 or 2002) by the average relative earnings to obtain an annual permanent income for the household in that year.[5]

[3]We deflate all nominal variables in the sample by the urban household consumer price index (CPI), with 2002 as the base year.

[4]For a single-earner family, the household head is the bread winner. For a multiple-earner family, the head is the person with the highest income.

[5]We use box plot to detect possible outliers in the data of wealth measures and permanent income. We first determine the first and third quartiles (denoted by Q_1 and Q_3, respectively) for the data set. Define the interquartile range $IQR = Q_3 - Q_1$, which is a measure of noise or scale for the data set. Observations that are more than three IQR's are treated as potential outliers and excluded from the sample.

To mitigate potential measurement errors introduced in the process of constructing permanent income, we follow Fuchs-Schündeln and Schündeln (2005) by instrumenting permanent income using education dummies and interactions of education with age and age-squared as instruments in all the regressions.

We consider two types of income uncertainty. The first relates to the cross-sectional variance of the log income used in the literature (Carroll and Samwick, 1998). Since the average household income in our sample has grown from 1995 to 2002 and different demographic groups might have experienced different growth, directly using the cross-sectional variance of income would be inappropriate, especially for making cross-group comparisons. We thus use a unit-free measure, which is the coefficient of variation (CV) of log income, defined as the ratio of the standard deviation to the mean of log household head's income over the past six (or five) years as reported in the 1995 (or 2002) CHIP surveys.

The second type of income uncertainty that we consider captures the uncertainty stemming from unemployment risks specific to SOE workers. In particular, we include in the regression an SOE dummy variable, which takes a value of one if the household head works in the SOE sector and zero if the household head works in the GOV sector. If our hypothesis is correct, then we should expect SOE workers to increase their savings relative to GOV workers after the reform took place. We use this source of income uncertainty associated with the large-scale SOE reform as the key to identifying precautionary saving.

In our estimation, we control for the effects of a number of demographic characteristics of households, including the household head's age, age-squared, gender, marital status, education, occupation, the household size, status of children (the ages of children, the number of boys, and the number of children at school), health care (public health care, public health insurance, or own payments), home ownership status, and others. Table 2 shows some details of these demographic variables.

We categorize the education level of a household head into four groups: elementary school and below, junior middle school, senior middle school, and post-secondary (college). We take the first group as our reference group and construct four education dummies.

We also divide the occupations of the household heads into five groups: professional, director or manager, skilled or office workers, unskilled or service workers, and others. The group of "others" is our reference group in the regressions.

The health care reform enacted in 1998 significantly changed the share of household expenditures on health care. We categorize the types of health care that the households receive into three groups: public health care (almost free), public health insurance, and self-financing of health care. As shown in Table 2, in 1995, 71.3% of households in our sample had access to free public health care. This share was halved to about 35.0% in 2002, reflecting the impact of the health care reform on household health expenditures.

To control for the effects of rising education expenditure on households' saving rate, we include in the regressions the mean age of children and the number of children at school. To control for effects of potential competitive savings motive emphasized in Wei and Zhang (2011), we add the number of boys among children as an independent variable.

Purchasing a house is argued to be one of the major motives of saving for Chinese households (Wei and Zhang, 2011). The housing reform that started in 1998 has led to extensively privatized housing market. As shown in Table 2, the homeownership rate in our sample doubled over the seven year period, from 42.0% in 1995 to 80.4% in 2002. We control for the potential effects of saving for home purchases by including a housing dummy that takes a value of one if the household is a home owner and zero otherwise. We also include in our regressions an interaction term between the SOE dummy and non-homeownership to control for the effects of potential savings by SOE workers for home purchases rather than for precaution against future unemployment risks.

Since the SOE reform and the massive layoffs hit some industries and geographic areas more heavily than others, we include in our regression two dummy variables that indicate the industries and provinces where the household head worked.

As revealed by Table 3, the reform has impacted GOV workers and SOE workers differently. In 1995, before the reform took place, GOV employees had on average modestly more financial wealth and higher permanent income than SOE workers. They were also more likely to own houses than the SOE workers. Nearly 90% of the GOV jobs were assigned by the government, while 80% of the SOE jobs were assigned by the government. In 2002, after the reform, the wealth and income gaps between the two sectors widened. The homeownership rate rose sharply for both groups (from 45% to 83% for GOV workers and from 40% to 78% for SOE workers). The reform also led to different perceptions of future income across the two groups. In the 2002 survey, about 24% of the SOE workers expected to have declines in income in the next five years, while just a bit over 11% of GOV employees expected income to decline. As we discuss below, pessimistic income outlooks can also raise saving, but such saving behavior represents a desire for intertemporal consumption smoothing (or PIH effects) rather than a motive for precautionary saving. To obtain a clean estimation of precautionary saving, we need to disentangle the PIH effects from the precautionary motive.

IV.3. **Empirical Strategies.** Following Lusardi (1998) and Carroll, Dynan, and Krane (2003), we estimate the importance of precautionary saving using the ratio of financial wealth to permanent income as the dependent variable. The estimation equation for each household

i is given by

$$W_i/P_i = \beta_0 + \beta_1 SOE_i + \beta_2 RISK_i + \beta_3 \log(P_i) + \beta_4' Z_i + v_i, \tag{1}$$

where W_i/P_i is the ratio of financial wealth (W_i) to permanent income (P_i), SOE_i is a dummy variable, which takes a value of one if the household head works for an SOE and zero if the household head works for a government or public institution (GOV), $RISK_i$ denotes individual income risks measured by the coefficient of variations (CV) of log annual income of the household head for the past 5 or 6 years, Z_i is a vector of demographic characteristics, and v_i is the error term.

We choose financial wealth/permanent income (W/P) ratio as the dependent variable for three reasons. First, precautionary saving model predicts that W/P ratio should be a function of age and other household characteristics (Lusardi, 1998; Carroll and Samwick, 1998), and therefore using W/P ratio as the dependent variable is consistent with the theoretical framework. Second, as a normalized measure of wealth, W/P ratio makes wealth of households with different income levels comparable. Third, W/P ratio captures a household's cumulative savings, which helps to establish a natural link between wealth accumulation and saving behavior.

As we argue above, the SOE reform in the late 1990s mainly affected the job security for SOE workers, but not for GOV workers. Thus, the reform provides a natural experiment that helps identify changes in precautionary savings for SOE workers relative to GOV workers. This aspect of the data allows us to use a difference-in-difference (DiD) approach to identify changes in the relative precautionary savings across the two sectors before and after the reform.

To implement this idea, we estimate the regression equation (1) using the CHIP survey data for each of the two years in our sample, one before the reform (1995) and the other after (2002). The estimated coefficient (β_1) of the SOE dummy variable from each regression captures the excess savings by SOE workers relative to GOV workers. All else equal, changes in the estimated value of β_1 from 1995 to 2002 captures changes in the relative magnitude of precautionary savings of the SOE workers caused by increases in their unemployment risks following the breaking of the iron rice bowl.[6]

[6]Our approach is slightly different from the standard DiD approach, which pools data in all sample years and thus puts an implicit restriction that the coefficients on all variables but the SOE dummy should be identical across time. With our approach, we estimate a separate regression for each of the two sample years and thus we do not impose such restrictions. Since China has gone through large structural changes between 1995 and 2002, many demographic aspects of our sample are likely to have changed during that period. Thus, taking a more flexible DiD approach as we do here is appropriate.

To estimate the quantitative importance of precautionary saving, it is necessary to correct a self-selection bias in occupational choices. An individual with high risk aversion has an incentive to choose a job with low income risk. Similarly, a worker with low risk aversion may want to choose a job with high income risk (with potentially high expected income). Failing to control for self-selection in occupational choice may lead to significant downward bias in estimating the importance of precautionary savings (Fuchs-Schündeln and Schündeln, 2005). We address the self-selection issue by restricting our sample to include only workers whose jobs were assigned by the government. For assigned jobs, the government has the final power to determine the worker's occupation and compensation; and thus, the worker's occupational choice is likely unrelated to her preferences.

The CHIP survey contains a question that asks an interviewee whether his current job was assigned by the government or found through a searching and matching process such as passing an exam, responding to a vacancy post, or referred by friends. We use answers to this question to identify the subsample of workers with government-assigned jobs. As shown in Table 2, 82.6% of workers had jobs assigned by the government in 1995 and this fraction declined somewhat to 71.7% in 2002. With this restriction imposed, we obtain a modestly smaller sample of 3627 in 1995 (with 2369 SOE workers and 1258 GOV employees) and 2170 in 2002 (with 1171 SOE workers and 999 GOV employees).

It is important to recognize that, while the SOE dummy (SOE_i) in the regression equation (1) captures income uncertainty specific to SOE workers, the $RISK_i$ variable reflects idiosyncratic income risks for all workers. These two variables are indeed uncorrelated in our sample, with a correlation coefficient of about -0.04 in each of the two sample years, consistent with our view that they capture different aspects of the risks for individual households.

In estimating the model, we also need to address the issue that arises with observations of zero wealth. In our sample, 11.3% of households have zero wealth in 1995 and this share declined to 4.5% in 2002. We treat this issue as a "censored data" problem. We address the issue by estimating an instrumental variable Tobit regression (IV-Tobit). In a robustness check, we also estimate the baseline model in equation (1) by eliminating the zero-wealth observations from our sample and then applying the standard two-stage least squares (2SLS) method (see Section VI.4).

V. Empirical Results

We now discuss the main empirical results and provide evidence of precautionary saving. We first discuss the estimation results with self-selection corrected in Section V.1. We then examine the quantitative importance of the self-selection bias in Section V.2. Finally, we discuss our approach to disentangling the permanent income effects from precautionary saving in Section V.3.

V.1. **Evidence of precautionary saving.** We now present evidence of precautionary saving when we correct the self-selection bias by focusing on the subsample with government assigned jobs. The estimation results for 1995 and 2002 are shown in Table 4 (columns (i) and (iii)).

The parameter of interest is the coefficient of the SOE dummy, β_1, which captures the difference in wealth accumulation between SOE and GOV workers when we control for the effects of all the demographic characteristics in the empirical model described by equation (1). The estimated value of β_1 is statistically insignificant in 1995 (column (i)), indicating that wealth accumulations of SOE workers and GOV workers were similar in 1995. By 2002, however, SOE workers had accumulated significantly more financial wealth than GOV employees (reflected by a much large estimate of β_1, see column (iii)). This evidence suggests that the relative saving behaviors of SOE workers has changed during that period. In particular, the difference between the two estimated values of β_1 is large ($0.723 - 0.09 = 0.633$) and statistically significant, with a p-value of 0.048. The substantial increase in β_1 reflects the effects of the large-scale SOE reform on workers' unemployment risks and thus captures the importance of precautionary saving.

We now discuss the interpretations of estimated coefficients for the control variables. In addition to the demographic controls such as the age, gender and occupation of the household head, we highlight here a few important control variables. These controls include an indicator of idiosyncratic income risks (CV), the permanent income (P) that captures non-homothetic preferences, and additional income or expenditure risks introduced by reforms between 1995 and 2002, such as health care reforms, education reforms, and housing reforms.

We continue to focus on the case with self-selection bias controlled for (columns (i) and (iii) in Table 4). The estimated coefficient β_2 of idiosyncratic income risks (CV) is positive and significant at the 1% level for both years. The estimated coefficient β_3 of $\log(P)$ is positive, but it is significant only in 2002, implying a significant income effect for that year.

To control for the impact of health care spending on households' saving behavior, we include in the regression a dummy variable indicating public health care (mostly free) and another dummy indicating public health insurance. The coefficients of both dummy variables are small and insignificant in 1995 but become significantly negative in 2002. This result is intuitive. In 1995, most workers were covered under a near-free public health care system, so that the health care status did not impose any significant impact on households' saving behavior. However, after the health care reform that started in 1998, a significant fraction of health care spending was shifted to private households. Thus, households not covered by public health care or public health insurance had a strong incentive to save. This finding is consistent with that obtained by Chamon and Prasad (2010), who report that declining public provisions of health care in the late 1990s in China created strong motives for precautionary saving against potential health expenditure shocks.

To control for the effects of education reforms on households' saving behavior and potential competitive saving motive in the marriage market emphasized by Wei and Zhang (2011), we include in our regression three additional variables: the mean age of children, the number of children enrolled in schools, and the number of boys in each household. Our estimation shows that the mean age of children does not explain wealth accumulation. The number of children enrolled in schools tends to reduce wealth accumulation in both years, although the effects were significant only in 2002. Having more children at school requires more expenditure on education after the education reforms in the late 1990s, which led to lower disposable income and reduced wealth accumulation. The number of boys contributes positively to savings in 1995, consistent with the findings in Wei and Zhang (2011), although the estimated coefficient is insignificant for that year. In 2002, however, having more boys in the household actually reduced savings and the effect is significant at the 10% confidence level. A possible explanation lies in the reforms of social security and the pension system, which substantially weakened the public safety net for retirees. In the Chinese culture, sons are supposed to take responsibility of taking care their elderly parents. Therefore, facing an uncertain future of safety net, having more boys means having better insurance for their parents. Parents thus do not need to save that much for their old-age consumption. In our 2002 sample, this self-insurance effect of having more boys dominates the potential competitive savings motive highlighted by Wei and Zhang (2011).

Finally, to control for the effects of housing reform on saving, we include in the regression a home ownership dummy and an interaction term between a non-homeowner dummy and the SOE dummy. The coefficients for these two variables are not significant for both years. A possible explanation is that, in 2002, the housing market was not fully developed and home purchases were still heavily subsidized. This result indicates that the saving motive for home purchases was weak in both 1995 and 2002.

V.2. **The self-selection bias.** The literature shows that self-selection in occupational choices can lead to a substantial downward bias in the estimated magnitude of precautionary saving (Fuchs-Schündeln and Schündeln, 2005). An individual with high risk aversion has an incentive to choose a job with low income risk. Similarly, a worker with low risk aversion may want to choose a job with high income risk (with potentially high expected income). To correct the downward bias caused by self-selection, we restrict our sample to workers whose jobs were assigned by the government. To the extent that the government's job assignments are not systematically correlated with individual risk attitude, our sample restriction should mitigate the bias caused by self-selection in occupational choices.

Our estimation shows that the downward bias caused by self-selection was not statistically significant in 1995, but it became significant in 2002. This can be seen by comparing the estimated value of β_1 from the subsample with government-assigned jobs to the estimate obtained from the full sample (i.e., the difference between β_1 in columns (i) and (ii) for 1995 and in columns (iii) and (iv) for 2002). In 1995, self-selection did not cause a significant downward bias in the estimated value of β_1 (the magnitude of the bias is $0.039 - 0.09 = -0.051$). In 2002, self-selection bias became statistically significant, with a magnitude of about $0.327 - 0.723 = -0.396$, which is equivalent to a little under 5 months of permanent income.[7] Thus, without correcting the self-selection bias, we would have substantially underestimated the importance of precautionary saving, especially for the post-reform period in 2002.

V.3. **Disentangling PIH Effects from Precautionary Saving.** The large-scale SOE reform not only led to significant changes in the relative job security between GOV and SOE workers, they might also produce potentially large differences in future income expectations between the two groups. All else equal, a worker who expects declines in future income would like to increase saving, but such increases in saving reflects a desire for intertemporal consumption smoothing (i.e., a permanent income effect) rather than a motive of precautionary saving. To the extent that the difference in perceived job security and income expectations between the two groups of workers were both caused by the SOE reform, disentangling the PIH effect from precautionary saving is particularly important for the post-reform period in 2002.

To isolate the effects of precautionary motives on saving from the PIH effects, we use a unique question in the 2002 CHIP survey that asks households about their expectations of income paths for the next five years. As Table 3 shows, a significant fraction of SOE workers surveyed in 2002 expected future income declines (23.8%), although a much smaller fraction of GOV workers expected income declines (11.4%). Thus, the reform has caused different income expectations in addition to different unemployment risks across the two groups of workers.

We disentangle the PIH effects from the precautionary motive on saving, we separate the sample of SOE workers in the 2002 survey into two groups based on their reported expectations of future income. One group expected income declines in the next five years, and the other group expected non-declines. We run IV-Tobit 2SLS regressions based on the

[7]The dependent variable (W/P) is the ratio of financial wealth to annual permanent income. A value of $W/P = 0.4$, for example, is equivalent to $0.4 * 12 \approx 5$ months of permanent income.

empirical model in equation (1) for each group of the SOE workers in 2002, using all GOV workers in that year as the control group.

The estimation results are reported in Table 5. The first column shows the estimation results for the group of SOE workers who expected their income to decline. The second column shows the results for the group that did not expect their income to decline. In both cases, we restrict our sample to those workers whose jobs were assigned by the government to control for the self-selection bias.

Our parameter of interest is again β_1, the coefficient of the SOE dummy. For the group of SOE workers who expected their income to decline, the estimated value of β_1 (Table 5, column (i)) significantly exceeds the benchmark estimate reported in Table 4, column (iii) (1.257 vs. 0.723). This finding is consistent with the PIH theory because this group of households increased their savings not just for precautionary reasons, but also for consumption smoothing. In contrast, the estimate of β_1 for those households who did not expect future income to decline (Table 5, column (ii)) is lower than the benchmark estimate (0.603 vs 0.723). The difference between the two estimates (0.723-0.603=0.12) is statistically significant at the 5% level. Since the PIH theory predicts that, all else equal, a household who does not expect future income to decline should save less and consume more, our estimate of $\beta_1 = 0.603$ provides a lower bound of the precautionary motive for saving. We use this estimated value of β_1 to provide a lower-bound estimate of the quantitative contribution of precautionary savings to wealth accumulation, as we discuss in the next section.[8]

V.4. **Importance of Precautionary Saving Motive.** Using the SOE reform as a natural experiment, we are able to identify the presence of precautionary saving. But an important question remains to be answered: To what extent does precautionary saving account for the observed increases in financial wealth for SOE workers between 1995 and 2002? To answer this question, we follow the literature (Carroll and Samwick, 1998; Fuchs-Schündeln and Schündeln, 2005) to quantify the contributions of precautionary saving to wealth accumulation. The idea is to compute the difference between (1) the model's predicted change in financial wealth held by SOE workers from 1995 to 2002 and (2) the counterfactual change in financial wealth had SOE workers enjoyed the same job security as GOV workers before and after the reform.

[8]Using the group of SOE workers that did not expect their income to decline in the 2002 survey might cause a downward bias in estimating precautionary saving, for two reasons. First, we do not exclude workers who expected their future income to rise; whereas for this group, the PIH channel should induce them to save less. Second, workers who expected their future income to fall might be the group who also faced higher probability of being laid-off and thus higher future income uncertainty; these workers might have stronger motives for precautionary saving than the group who expected their income not to decline.

To implement this idea, we go through the following steps. First, we calculate the model's predicted wealth held by SOE workers in 1995 (denote this by $\hat{W}^{soe}_{1995}$) using the benchmark estimation results reported in column (i) of Table 4, where we have corrected the self-selection bias related to occupational choice. Second, we calculate the predicted wealth held by SOE workers in 2002 (denote this by $\hat{W}^{soe}_{2002}$) using the estimation results reported in column (ii) of Table 5, where we have controlled for both the self-selection bias and the PIH effects. Third, we compute the counterfactual wealth holdings by SOE workers in each year of the surveys by assuming that those workers had the same job security as GOV employees, while keeping all the other characteristics unchanged. In particular, we use the same estimated coefficients as in the first two steps, except that we replace the estimated value of β_1 by zero (and thus assuming the SOE dummy did not affect the wealth accumulation at all). Denote by $\tilde{W}^{soe}_t$ the counterfactual wealth holdings of SOE workers in year $t \in \{1995, 2002\}$. In the fourth (and final) step, we compute the magnitude of wealth accumulation for precautionary reasons according to the relation

$$W^{ps} = (\hat{W}^{soe}_{2002} - \hat{W}^{soe}_{1995}) - (\tilde{W}^{soe}_{2002} - \tilde{W}^{soe}_{1995}), \tag{2}$$

where W^{ps} denotes the wealth accumulation from precautionary savings. The ratio $\frac{W^{ps}}{\hat{W}^{soe}_{2002} - \hat{W}^{soe}_{1995}}$ then measures the fraction of the (model-predicted) changes in financial wealth held by the SOE workers that can be accounted for by precautionary saving.

Our estimation implies that precautionary saving accounts for 30.3% of financial wealth accumulation for SOE workers between 1995 and 2002, which is statistically significant with a standard error of 0.166. This result suggests that the SOE reform in the late 1990s have led to quantitatively important precautionary savings.

VI. Robustness

We have presented evidence that increases in unemployment risks for SOE workers relative to GOV workers following the large-scale SOE reform in China have led to substantial increases in precautionary saving for the SOE workers. Our estimation suggests that such saving behavior accounts for about 30% of the cumulative increase in financial wealth held by SOE workers from 1995 to 2002. In our estimation, we have exploited the micro-level details of our dataset to control for self-selection biases and to disentangle PHI effects from precautionary saving.

We now examine the robustness of our results by running a few more experiments. In each experiment, we focus on the sample with all workers' jobs assigned by the government so that we can control for self-selection biases. We also remove those SOE workers in the 2002 sample who expected their incomes to decline in the next five years so that we can control for PIH effects.

VI.1. **Pension effects.** Pension benefits can also affect saving behaviors through a channel similar to that of income expectations: they both reflect PIH effects. To get a clean estimate of precautionary saving, we would also need to take into account of pension effects. Unfortunately, our dataset does not provide direct pension information at the individual household level. However, we do observe average pension benefits for SOE workers and GOV employees in both 1995 and 2002. We use these aggregate observations to obtain a rough estimate of the extent to which differences in pension benefits across the two sectors and the changes of those benefits over time would affect our estimates of precautionary saving.

According to the 2003 China Labor Statistical Yearbook, the average pension replacement ratio—defined as the ratio of pension income to annual salaries—was about 99.5% for GOV employees in 1995, and it declined slightly to 94.4% in 2002. The average pension replacement ratio for SOE workers was much lower at 74.2% in 1995, and it further declined to 64.3% in 2002. Thus, the pension replacement ratio for SOE workers was about 25.4% lower than GOV employees in 1995 ($(0.995 - 0.742)/0.995 = 0.254$), and this gap widened to 31.9% in 2002 ($(0.944 - 0.643)/0.944 = 0.319$).

The PIH hypothesis implies that, all else equal, a worker with lower pension benefits should save more, although such saving behavior reflects a desire for intertemporal consumption smoothing (i.e., a wealth effect), which is different from precautionary saving. To adjust for the pension effects in our estimation of precautionary saving, we assume that the positive estimated value of $\beta_1 = 0.090$ in 1995 reflects mainly the lower pension benefits for SOE workers than for GOV workers. This assumption seems reasonable because, in our regression, we have controlled for all other demographic characteristics for both groups of workers. By 2002, the pension gap has widened by a factor of 1.26 ($0.319/0.254 = 1.26$). Thus, to a first-order approximation, the pension effects for the 2002 sample should be 1.26 times that for the 1995 sample (i.e., $1.26\beta_1^{1995} = 0.1134$). With these pension effects taken into account, the estimated precautionary saving—which corresponds to the pension-adjusted differences between the estimated values of β_1 in 2002 and 1995 becomes smaller. In particular, the pension-adjusted estimate of precautionary saving should be $\beta_1^{2002} - 1.26 \times \beta_1^{1995} = 0.603 - 1.26 \times 0.09 = 0.49$, which is slightly smaller than our benchmark estimate of 0.513 ($0.603 - 0.09 = 0.513$).

VI.2. **Size effects.** The SOE reform in the late 1990s had very different impact on large SOE firms than on medium and small firms. The spirit of the reform was to "Grasp the Large and Let Go of the Small." Accordingly, large and profitable SOEs in strategically important sectors (such as oil, electricity, and telecommunications) were corporatized or consolidated into large state-owned conglomerates, while smaller and loss-making SOE firms were shut down or privatized (Hsieh and Song, 2013). Those large SOEs that survived the reorganization typically gained stronger government protections of their monopoly power, leading to higher profits than their privatized counterparts (Li, Liu, and Wang, 2012). Thus, our higher estimate of β_1 in 2002 than in 1995 might reflect partly the fact that workers in large and surviving SOEs were richer, which is not the same as precautionary saving.

To address this concern, we divide SOEs in our sample into two groups based on their sizes: central or provincial SOEs (CSOE) vs. local SOEs (LSOE).[9] CSOEs are typically larger than LSOEs, and workers in CSOEs typically face a lower unemployment risks than those in LSOEs (Appleton, Knight, Song, and Xia, 2002).[10] Therefore, we should expect to see higher precautionary saving motives for LSOE workers than for CSOE workers.

To examine the importance of this size effect, we modify the benchmark model in equation (1) by replacing the SOE dummy variable with the two dummy variables, indicating whether the household head worked in a CSOE or an LSOE. The regression model is now

$$W_i/P_i = \beta_0 + \beta_1^{CSOE} CSOE_i + \beta_1^{LSOE} LSOE_i + \beta_2 RISK_i + \beta_3 \log(P_i) + \beta_4' Z_i + v_i \qquad (3)$$

where $CSOE_i$ and $LSOE_i$ are the two dummy variables indicating the type of the SOE firm in which the household head i worked.

Table 6 reports the regression results. From 1995 to 2002, β_1^{CSOE} increased from 0.0001 to 0.088, but it is not significant in both years. In contrast, β_1^{LSOE} was estimated to be 0.160 and in significant in 1995, but it rose sharply to 1.082 in 2002 and became significant, with the p-value of 0.046 for the Chow test. This finding is consistent with the view that workers in LSOEs had stronger precautionary saving motives than those in CSOEs because they faced higher unemployment risks.

VI.3. **Survival Bias.** To obtain a clean identification of precautionary saving caused by the SOE reform, we need to control the characteristics of SOE workers before and after the reform. In particular, in the 2002 sample, we should include workers who share the same characteristics as those in the 1995 sample except that they face higher unemployment risks. There is evidence that workers with lower educational attainment or lower skills had a higher chance of being laid off (Appleton, Knight, Song, and Xia, 2002). Therefore, it is plausible that the SOE workers in the 2002 sample who survived the layoffs had on average

[9]LSOE also includes urban collective enterprises.

[10]In the 2002 sample, only 3.4% of workers in CSOEs had experience of prior layoffs, while 7.4% of workers in local SOE and 16.4% workers in urban collective enterprises experienced layoffs.

higher skills and thus higher income, which may affect their saving behavior and thus create a survival bias in the estimation of precautionary saving.

To correct the survival bias, we use a propensity score approach to adjust the 1995 sample to include only those workers who are likely to survive the massive layoffs. We estimate the probability of being laid off for a SOE worker in 1995 based on the 2002 sample, extended to include also those workers who had worked in the SOE sector but had experienced layoffs between 1995 and 2002. We use the extended 2002 sample to estimate the Probit model

$$\Pr(\text{laid-off}_i = 1 \mid Z_i) = \Phi(Z_i\delta) \tag{4}$$

where Z_i is the individual i's characteristics, such as age, gender, education levels, occupation, and industry and province dummies. The dependent variable in the Probit model is a dummy variable that takes a value of 1 if an individual had prior or current layoff experience, and equals zero otherwise.

Based on the estimated probability of being laid off as a function of individual characteristics using the extended 2002 sample, we go back to use the 1995 sample to predict each SOE worker's probability of being laid-off conditional on their characteristics.

According to Giles, Park, and Zhang (2005), urban household unemployment rate reached 11.1% in 2002. This implies that for SOE workers in 1995, at least 10% of them would not survive until 2002. Thus, we drop the SOE workers in the 1995 sample who, according to the estimated layoff probability, are the top 10% of the sample that are most likely to be laid off. In other words, we keep the top 90% of SOE workers in 1995 sample who are most likely to survive the massive layoffs. We argue that the subsample of potential survivors of the layoffs in 1995 share similar characteristics with the 2002 sample (who are ex post survivors of the layoffs), except that they face different levels of unemployment risks.

Table 7 shows the estimated results of equation (1) when the survival bias is corrected. Column (1) keeps all workers in 1995 sample and therefore simply replicates the results in column (i) of Table 4. Column (2) shows that if we drop those SOE workers who had the top 10% probability of being laid-off in future years, the coefficient β_1 of the SOE dummy increases from 0.090 to 0.122, although remains insignificant. To further examine the importance of survival bias, we drop the SOE workers with the top 20% and top 30% of layoff probabilities and reestimate the benchmark model. The results are reported in columns (3) and (4) in Table 7, respectively. The estimated value of β_1 increases to 0.192 and 0.195 respectively, but remains insignificant. The difference in β_1 between 2002 and 1995 becomes somewhat smaller than that obtained in the benchmark model (0.408 v.s. 0.513), but remains significant both statistically and economically.

Thus, correcting the survival bias modestly reduces the quantitative magnitude of precautionary savings, but precautionary saving caused by the large-scale SOE reform remains evident.

VI.4. **Excluding Zero Wealth Observations.** Recall that all the empirical results above are based on the sample including zero wealth observations and the estimations from the (IV) Tobit model. To test whether these results are driven by zero wealth observations, we exclude zero wealth observations from the sample and run the commonly used IV (2SLS) regression respectively for 1995 and 2002 data. Our sample size thus reduces to 3221 and 1807 observations for 1995 and 2002, respectively. The results are summarized in panel A of Table 8.

The estimated value of β_1 is 0.100 (not significant) in 1995 and 0.467 (significant at 10% level) in 2002. The difference is $0.467-0.100 = 0.367$, which is modestly smaller than that our benchmark estimate of $0.603 - 0.090 = 0.513$ when we included zero-wealth observations. Thus, excluding zero-wealth observations from the sample tends to reduce the estimated magnitude of precautionary saving. Nonetheless, the estimated difference between β_1 in 1995 and 2002 remains significant, suggesting evidence of precautionary saving stemming from increased unemployment risks for SOE workers following the reform.

VI.5. **Conventional Risk Measure.** So far we keep using CV (ratio of the standard deviation to the mean) of the logarithm of a household head's labor income over the past five or six years to measure idiosyncratic income risk. An alternative measure of risk, which is also widely used in the literature, is the variance of the logarithm of permanent income (Carroll and Samwick, 1998; Fuchs-Schündeln and Schündeln, 2005). In this subsection, we check the sensitivity of our results to this conventional risk measure. More specifically, we follow Carroll and Samwick (1998) to divide our data sample into 20 subsamples corresponding to the five occupation categories and four education groups (see Table 1) in both years. For each household, we calculate the log variance of log of annual income with respect to the mean income within the group that it belongs to. We use this within-group variance of income to proxy risk. The results are shown in panel B of Table 8.

As we can see from the table, the results are similar to those in Tables 4 and 5. The estimated value of β_1 increases from 0.083 (not significant) in 1995 to 0.713 (5% significant) in 2002. This finding indicates that using the conventional risk measure does not significantly change our results.

VI.6. **Alternative Wealth Measures.** Some alternative measures of wealth such as very liquid assets and non-housing non-business wealth are also commonly used in the literature

(Carroll and Samwick, 1998). We shall check the sensitivity of our empirical results to these alternative measures of wealth (see Table 1 for the construction of these variables in CHIP).[11]

Panel C of Table 8 presents the results using very liquid assets as wealth measure to construct the dependent variable in equation (1). The estimated value of β_1 increases from 0.062 (not significant) in 1995 to 0.439 (significant at 10% level) in 2002. Panel D of Table 8 shows the results employing non-housing non-business wealth to construct the dependent variable in (1). The estimated value of β_1 is 0.210 and insignificant in 1995 and it increases substantially to 0.632 (significant at 10% level) in 2002, which are similar to the benchmark estimates.

In summary, we find that using alternative wealth measures does not significantly affect our main result.

VII. Conclusion

Using China's large-scale reform of the state-owned enterprises (SOEs) in the late 1990s as a natural experiment, we identify and quantify the importance of precautionary saving in a rapidly growing transition economy. With self-selection in occupational choices corrected and with expected income effects controlled for, we obtain significant evidence of precautionary saving stemming from sudden increases in unemployment risk for SOE workers relative to that for government employees. Our estimation suggests that precautionary saving can account for about 30 percent of the actual increase in wealth accumulation by urban SOE households in China for the period from 1995 to 2002. Thus, precautionary saving associated with large structural changes in the Chinese economy is quantitatively important.

[11]Another widely used measure of wealth is total net worth, which is NHNBW plus estimated market value of owner-occupied housing and fixed assets of farms and business. Market value of owner-occupied housing accounts for a significant portion of total net worth. However, before 1998, housing in China was mostly assigned by the government. There was no housing market. The privatization of housing market began in 1998 and moved slowly until the mid 2000s. Therefore it is very hard to accurately estimate market value of owner-occupied housing back to the time period we consider in this paper. In addition, since housing market has not been fully established until recently, it is even harder to sell a house once one purchase it. This makes housing in China is extremely illiquid. In other words, precautionary saving is not the main reason why Chinese purchase house. Therefore, a caveat should be noticed for using total net worth as an appropriate measure for identifying precautionary wealth in Chinese economy.

Appendix A. A Case Study: Massive Lay-off in Fushun, Liaoning

Smyth, Zhai, and Wang (2001) demonstrate a case study of massive lay-off happened in Fushun, Liaoning. Fushun is a medium sized city located 45 kilometers northeast of Shenyang, the capital city of Liaoning. It was well known as a state-owned heavy industrial base in the "rust belt" of China. In 2000, nearly 91% of workers in Fushun were employed by SOEs. And SOEs produced 88.5% of gross industrial output.

The wave of layoffs (*xia gang*) hit Fushun very severely. In 2000, laid-off workers from SOEs accounted for about 42% of total workers in SOEs in Fushun, which was the highest in Liaoning. The industries saw largest number of laid-off workers were coal, textiles, light industry, electronics, machinery and chemicals. For example, of the 71000 workers in SOEs in the coal sector in Fushun, 35000 or 49.7% of workers were classified as *xia gang*.

What differentiates *xia gang* from official unemployment (known as "registered unemployment") is that *xia gang* workers still retain their ties with SOEs they used to work. In practice, there were different ways to lay off SOE workers. 1) *fang jia*, firms make workers on temporary leave; 2) *xia gang*, defined as those on long-term leave; 3) *tui yang*, which refers to workers who have taken voluntary early retirement. 4) *mai duan,* which refers to firms pay a lump-sum amount (usually not exceeding three year salary) to buy out or terminate the labor contract with workers.

Allowances were paid to *xia gang* workers by their former employer, the local government, and the central government, each was supposed to contribute one-third. However, many SOE firms had financial difficulties in making the payments to the laid-off workers. For example, of the 35,000 laid-off workers from state-owned coal mines in Fushun, 33,000 did not receive basic living allowances from their former employers.

In Fushun, the main avenue for laid-off workers to find new jobs was through re-employment centers sponsored by the local government. The re-employment centers offered various training classes. However, there are several problems hinder the effectiveness of government-sponsored re-employment structure. The majority of laid-off workers were middle-aged and female accounted for a high proportion. It is very hard for them to find a job given the discrimination against age and gender in Chinese labor market. And they were reluctant to take jobs in non-state-owned sector worrying that it is going to cut their ties with their original enterprises. Among the laid-off workers who have registered at re-employment centers in Fushun, 50% are middle-aged. Among them, only 50% of these middle-aged workers found jobs.[12]

[12]This is consistent with the official number of national reemployment rate, see Lee (2000). However, a survey of 54,000 workers carried out by the Chinese Federation of Labor Unions in 1997 reports that only about 18% of the laid off have found new employment. See Lee (2000) for details.

Table 1. Definition of variables

Variable	**Description**
Financial wealth (W)	Balances in checking accounts, saving accounts, CDs, bonds, stock holdings, etc.
Very liquid assets (VLA)	Financial wealth minus business investment, housing fund, etc.
Nonhousing, nonbusiness wealth (NHNBW)	Financial wealth plus estimated market value of durable cons. goods and other assets, minus total debt
Annual income	Annual income of household head and revenues from business, farming, fishing, gardening, livestock, non-retirement wages, retirement income, subsidies, and other income
Income risk	Coefficient of variation (CV) of log annual income of past 5 or 6 years
SOE	Dummy variable for employers of HH, 1 for State Owned Enterprises (SOE), 0 for Government & Institutions
Permanent income (P)	See text
W/P	Wealth / permanent income ratio
Age	Age of HH
Male	Dummy variable for the gender of HH, 1 for male, 0 otherwise
Married	Dummy variable for the marital status of HH, 1 for married, 0 otherwise
Education	Four dummy variables for college, senior middle school, junior middle school, and elementary school or below (see text)
Occupation	Five dummy variables for professional, director or manager, skilled or office workers, unskilled or service workers, and the others (see text)
Health care	Three dummy variables for public health care, public health insurance and own payment (see text)
No house owned	Dummy variable for housing ownership, 1 for no house owned, 0 otherwise
Age of children (mean)	Mean age of children in household
Num. of boys	Number of boys in household
Num. of children at school	Number of children at school in household

Table 2. Summary statistics of the full sample

Variable	**1995**			**2002**		
	Obs.	**Mean/%**	**SD**	**Obs.**	**Mean/%**	**SD**
Financial wealth (W)	4390	10042	10165	3027	32826	32140
Annual income	4390	7034	3349	3027	12985	6658
CV×100	4390	2.61	2.07	3027	2.9	7.67
Age	4390	40.91	7.37	3027	42.61	6.88
Age of children (mean)	4390	11.65	6.94	3027	12.5	7.58
Num. of boys	4390	0.57	0.58	3027	0.47	0.53
Num. of students	4390	0.65	0.48	3027	0.69	0.54
Household size	4390	3.18	0.68	3027	3.03	0.61
Male	4390	63.4%		3027	68.8%	
Married	4390	97.6%		3027	96.7%	
Education						
College	4390	24.6%		3027	37.2%	
Senior middle school	4390	39.5%		3027	38.8%	
Junior middle school	4390	30.8%		3027	21.5%	
≤Elemen. School	4390	5.1%		3027	2.4%	
Occupation						
Professional	4390	24.3%		3027	24.7%	
Director or manager	4390	14.3%		3027	15.3%	
Skilled worker	4390	44.7%		3027	44%	
Unskilled worker	4390	13.6%		3027	15%	
Other occupation	4390	3.1%		3027	0.9%	
Health Care						
Own payment	4390	9.9%		3027	23.1%	
Public health care	4390	71.3%		3027	35%	
Public health insurance	4390	8.8%		3027	41.9%	
Own house	4390	42%		3027	80.4%	
SOE	4390	67.8%		3027	56.2%	
Job assigned by Gov.	4375	82.9%		3018	71.9%	

Notes: Monetary values are in constant RMB Yuan, base year = 2002.

TABLE 3. Comparison between employees in GOV vs. SOEs

		1995			2002		
	Variable	**Obs.**	**Mean**	**SD**	**Obs.**	**Mean**	**SD**
GOV	Financial wealth (W)	1413	10457	10209	1325	34677	32351
	Annual permanent income	1413	7545	3215	1325	14752	6698
	W/P	1413	1.376	1.386	1325	2.559	2.36
	Non homeowners	1413	0.546	0.498	1325	0.165	0.372
	Job assigned by Gov.	1408	0.893	0.309	1319	0.757	0.429
	Expected income to decline	N.A	N.A	N.A	1321	0.114	0.318
SOE	Financial wealth (W)	2977	9845	10140	1702	31386	31910
	Annual permanent income	2977	6791	3385	1702	11610	6294
	W/P	2977	1.382	1.448	1702	2.703	2.906
	Non homeowners	2977	0.597	.491	1702	0.220	.414
	Job assigned by Gov.	2967	0.798	0.401	1699	0.689	0.463
	Expected income to decline	N.A	N.A	N.A	1699	0.238	0.426

Notes: Data are taken from CHIP surveys. Monetary values of financial wealth and permanent income are in constant Chinese Yuan units, with 2002 as the base year.

Table 4. IV-Tobit regressions, 1995 and 2002 sample

Dep. variable: W/P	**1995 (i)**	**(ii)**	**2002 (iii)**	**(iv)**
SOE	0.090	0.039	0.723**	0.327*
	(0.117)	(0.114)	(0.298)	(0.221)
CV×100	0.111***	0.136***	0.124***	0.091***
	(0.040)	(0.038)	(0.045)	(0.028)
log(permanent income)	0.759	1.225	4.512***	3.533***
	(1.028)	(0.900)	(1.497)	(0.992)
Age	0.020	-0.020	0.028	0.240*
	(0.052)	(0.050)	(0.150)	(0.125)
Age squared(*100)	-0.030	0.019	-0.039	-0.274*
	(0.059)	(0.059)	(0.175)	(0.147)
Male	-0.362***	-0.463***	-1.180***	-1.176***
	(0.102)	(0.094)	(0.202)	(0.148)
Professional	0.102	0.031	4.776***	0.370
	(0.212)	(0.200)	(1.648)	(0.787)
Director	0.295	0.185	4.780***	0.183
	(0.214)	(0.208)	(1.636)	(0.800)
Skilled worker	0.042	0.004	4.993***	0.341
	(0.182)	(0.168)	(1.661)	(0.762)
Unskilled worker	-0.031	0.039	6.093***	0.981
	(0.201)	(0.179)	(1.770)	(0.767)
Public med service	0.047	0.036	-1.228**	-0.978***
	(0.192)	(0.166)	(0.501)	(0.362)
Public med insurance	0.031	0.102	-0.908**	-0.755**
	(0.166)	(0.150)	(0.434)	(0.318)
Married	0.520***	0.488***	0.637	0.406
	(0.192)	(0.161)	(0.429)	(0.363)
Age of children (mean)	0.008	0.005	0.004	-0.000
	(0.006)	(0.006)	(0.013)	(0.010)
Num. of boys	0.044	0.022	-0.253*	-0.198*
	(0.048)	(0.045)	(0.145)	(0.118)
Num. of children at school	-0.086	-0.035	-0.317*	-0.363***
	(0.066)	(0.063)	(0.176)	(0.140)
Household size	-0.037	-0.008	0.279	0.357***
	(0.051)	(0.048)	(0.171)	(0.136)
No house owned	0.080	0.138	-0.244	-0.221
	(0.101)	(0.097)	(0.264)	(0.228)
No house owned×SOE	-0.114	-0.106	0.356	0.300
	(0.109)	(0.104)	(0.376)	(0.300)
Industry & Province dummies	yes	yes	yes	yes
Log-Likelihood	-7167.03	-8875.88	-5803.38	-8240.22
p-value of Chow test for SOE			0.048	0.247
Number of observations	3627	4390	2170	3027

Notes: Results from instrumental variable Tobit regressions. Standard errors are in parentheses and are corrected for heteroskedasticity. Columns (i) and (iii) show the estimation results obtained from the subsample with government assigned jobs and thus correct for the self-selection bias in occupation choices. Columns (ii) and (iv) are obtained using the full sample and thus do not address the self-selection issue.

TABLE 5. Regressions with 2002 sample: Controlling for PIH effects

Dep. variable:	**expected future income**	
W/P	**decline**	**non-decline**
SOE	1.257**	0.603**
	(0.531)	(0.305)
CV×100	0.120**	0.123***
	(0.061)	(0.046)
log(permanent income)	5.339**	4.681***
	(2.194)	(1.665)
Age	0.103	0.068
	(0.172)	(0.167)
Age squared(*100)	-0.150	-0.088
	(0.200)	(0.195)
Male	-0.955***	-1.161***
	(0.224)	(0.224)
Professional	1.123	5.386***
	(1.649)	(1.872)
Director	1.219	5.439***
	(1.668)	(1.856)
Skilled worker	1.799	5.554***
	(1.602)	(1.891)
Unskilled worker	2.755*	6.797***
	(1.661)	(2.017)
Public med service	-1.212*	-1.378**
	(0.632)	(0.571)
Public med insurance e	-0.595	-1.055**
	(0.463)	(0.507)
Married	0.399	0.590
	(0.525)	(0.441)
Age of children (mean)	0.011	-0.001
	(0.015)	(0.013)
Num. of boys	0.058	-0.266*
	(0.191)	(0.160)
Num. of children at school	-0.251	-0.281
	(0.231)	(0.193)
Household size	0.164	0.317*
	(0.184)	(0.181)
No house owned	-0.373	-0.234
	(0.282)	(0.267)
No house owned×SOE	-0.170	0.548
	(0.620)	(0.415)
Industry & Province dummies	yes	yes
Log-Likelihood	-3182.68	-4925.80
p-value of Chow test for SOE	0.032	0.116
Number of observations	1284	1876

Notes: Results from 2SLS regressions. Standard errors are in parentheses and are corrected for heteroskedasticity.

TABLE 6. Precautionary saving by workers in central vs. local SOEs

Dep. variable: W/P	**1995**	**2002**
CSOE	0.0001	0.088
	(0.146)	(0.294)
LSOE	0.160	1.082**
	(0.180)	(0.425)
CV×100	0.116***	0.127***
	(0.045)	(0.047)
log(permanent income)	0.893	4.930***
	(1.184)	(1.744)
Industry & Province dummies	yes	yes
Log-Likelihood	-7094.60	-4901.44
p-value of Chow test for CSOE		0.790
p-value of Chow test for LSOE		0.046
Number of observations	3627	1876

Notes: Results from instrumental variable Tobit regressions. Standard errors are in parentheses and are corrected for heteroskedasticity. "CSOE" denotes SOEs owned by the central and provincial governments and "LSOE" denotes those owned by local governments.

Table 7. Controlling for the survival bias

Dep. variable: W/P	**(1)**	**(2)**	**(3)**	**(4)**
Keep 1995 Sample	**100%**	**90%**	**80%**	**70%**
SOE	0.090	0.122	0.192	0.195
	(0.117)	(0.122)	(0.131)	(0.133)
CV×100	0.111***	0.123**	0.165***	0.175***
	(0.040)	(0.050)	(0.053)	(0.048)
log(permanent inc)	0.759	1.117	2.130	2.266**
	(1.028)	(1.335)	(1.358)	(1.129)
Industry & Province dummies	yes	yes	yes	yes
Log-Likelihood	-7167.03	-6703.03	-6209.17	-5746.92
p-value of Chow test for SOE	0.116	0.143	0.215	0.220
Number of observations	3627	3415	3198	2971

Notes: Results from IV-Tobit regressions. Standard errors are in parentheses and are corrected for heteroskedasticity. In column 2, we eliminate the top 10 percent SOE workers in 1995 sample who are more likely being laid-off in future; the top 20 percent in column 3 and the top 30 percent in column 4.

Table 8. Robustness Checks

Case	**1995**	**2002**	**Chow Test**
A. Eliminating zero wealth			
	0.100	0.467*	0.150
	(0.104)	(0.268)	
	[N=3221]	[N=1807]	
B. Conventional risk measure			
	0.083	0.713**	0.085
	(0.117)	(0.346)	
	[N=3627]	[N=1876]	
C. Very liquid asset			
	0.062	0.439*	0.168
	(0.114)	(0.248)	
	[N=3627]	[N=1876]	
D. Non-housing Non-business wealth			
	0.210	0.632*	0.243
	(0.159)	(0.355)	
	[N=3627]	[N=1876]	

Notes: Reported is the coefficient on the SOE dummy from different wealth/income ratio regressions. Results from IV-Tobit regressions. Other controls are the same with Table 4. Standard errors are in parentheses and are corrected for heteroskedasticity. Numbers of observations are in squared brackets.

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