Journal of Monetary Economics 000 (2017) 1–20

Contents lists available at ScienceDirect

Journal of Monetary Economics

journal homepage: www.elsevier.com/locate/jmoneco

Breaking the "iron rice bowl:" Evidence of precautionary savings from the chinese state-owned enterprises reform☆

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ARTICLE INFO

Article history:
Received 24 March 2016
Revised 18 December 2017
Accepted 19 December 2017
Available online xxx

JEL classification:
E21
P31
D81

Keywords:
Precautionary savings
China's SOE reform
Natural experiment
Self-selection bias
Difference-in-differences methods

ABSTRACT

China's large-scale reform of state-owned enterprises (SOE) in the late 1990s provides a natural experiment for estimating precautionary savings. Before the reform, SOE workers enjoyed similar job security as government employees. The reform caused massive SOE layoffs, but government employees kept their "iron rice bowl." The changes in the relative unemployment risks for SOE workers provide a clean identification of income uncertainty. With self-selection biases mitigated by focusing on government assigned jobs, precautionary savings account for about 40 percent of SOE households' wealth accumulation. Moreover, demographic groups more vulnerable to the reform also accumulated more precautionary wealth.

1. Introduction

Precautionary savings are potentially important for wealth accumulation, especially for an emerging market economy like China that has experienced large structural changes associated with policy reforms, which may have led to substantial increases in economic uncertainty. However, estimating the importance of precautionary savings has been a challenge in the empirical literature. One difficulty is to identify large and exogenous variations in income uncertainty (Carroll and Kimball,

☆ We are grateful to an anonymous referee, the Associate Editor Fatih Guvenen, and the Editor Urban Jermann for helpful comments and suggestions. We also thank David Card, Chris Carroll, Marcos Chamon, Russell Cooper, Nicola Fuchs-Schündeln, Bart Hobijn, Oscar Jorda, Dirk Krueger, Adam Shapiro, David Slichter, Dan Wilson, Shang-Jin Wei, Motohiro Yogo, Xiaobo Zhang, Kai Zhao, and Xiaodong Zhu, and seminar participants at the 2014 NBER Chinese Economy Meeting, the 2014 NBER Summer Institute EFACR Program Meeting, the 2014 SED Meeting, University of Pennsylvania, University of Rochester, the Federal Reserve Bank of San Francisco, and a number of other institutions. We thank Hanya Li for research assistance and Anita Todd for editorial assistance. Hui He acknowledges research support by Shanghai Pujiang Program, the Program for Professor of Special Appointment (Eastern Scholar) at Shanghai Institutions of Higher Learning. Hui He, Feng Huang and Dongming Zhu acknowledge research support by Key Laboratory of Mathematical Economics (SUFE), Ministry of Education. The project is partly supported by the National Science Foundation of China (Project #71633003). The views expressed in this paper are those of the authors and do not necessarily reflect the views of the IMF, the Federal Reserve Bank of San Francisco, or the Federal Reserve System.

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https://doi.org/10.1016/j.jmoneco.2017.12.002
0304-3932/

2008; Lusardi, 1998). The literature typically uses the cross-sectional variances of income as a proxy for income uncertainty (Carroll and Samwick, 1998), and it is well known that such proxies suffer from measurement errors and potential endogeneity biases for estimating precautionary savings (Kennickell and Lusardi, 2004).

A second difficulty stems from a self-selection bias related to job choices. Precautionary savings depend not just on risk, but also on risk preferences (Caballero, 1990; 1991). Risk preferences affect not just saving behaviors, but also job choices. A more risk averse individual would save more for given income risks, but she is also likely to choose a job with lower income risks. The correlations between risk preferences and job choices imply a self-selection bias, and failing to control for this self-selection can lead to a significant downward bias in estimating precautionary savings (Fuchs-Schündeln and Schündeln, 2005).

Partly reflecting the difficulties in measuring income uncertainty and correcting self-selection biases, the existing literature has obtained mixed evidence of precautionary savings. Some studies report weak or no evidence of precautionary savings (Dynan, 1993; Guiso et al., 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).[1]

This paper presents new empirical evidence for precautionary savings using Chinese data. We argue that the Chinese large-scale reform of state-owned enterprises (SOEs) in the late 1990s provides a natural experiment for identifying exogenous variations in income uncertainty. Prior to the reform, workers in the SOEs and the government sector (GOV) enjoyed similar job security, with near-free health care, education, housing, and retirement benefits. In this sense, workers in both sectors held an "iron rice bowl" before the reform. Following the reform, over 27 million SOE workers—equivalent to 27% of SOE employment in 1997—were laid off between 1997 and 2002. Those workers lost not just their jobs, but also the associated benefits. In contrast, few workers in the government sector were affected by the reform; they were able to hold on to their iron rice bowl. The massive layoffs in the SOE sector significantly changed the perceived job security for the remaining SOE workers. The reform was largely unexpected to an individual worker and it created significant variations of unemployment risks for workers across the SOE and GOV sectors. Thus, the reform provides a clean identification of variations in perceived income uncertainty across time and across sectors.

To implement the idea that the SOE reform can be used as a natural experiment for estimating precautionary savings, we use the Chinese Household Income Project (CHIP) survey data and design a difference-in-differences (DID) approach, focusing on urban households in two sectors (SOE and GOV) and two CHIP surveys (1995 and 2002). The large-scale SOE reform started to have significant impacts on SOE employment in 1997, with the effects tapering gradually through 2002. Thus, our sample covers both the pre- and post-reform periods. This data structure allows us to estimate the differences in household savings both across sectors (SOE vs GOV) and across time (before and after the reform). The time variations (between 1995 and 2002) of the relative saving behavior of workers across the two sectors capture the magnitude of precautionary savings caused by the SOE reform.

To mitigate the self-selection bias in estimating precautionary savings, we restrict our sample to those households whose jobs were assigned by the government, following the approach by Fuchs-Schündeln and Schündeln (2005). Similar to the case of the former German Democratic Republic (GDR) studied by Fuchs-Schündeln and Schündeln (2005), job assignments by the Chinese government were often restricted by political considerations and job outcomes were often unrelated to individual preferences. Since the final job outcome was determined by the local governments rather than individual workers, self-selection was unlikely. In practice, however, job assignments by the government were not completely independent of worker preferences because workers could signal their preferred job positions to the government before actual assignments took place. Therefore, focusing on the subsample with government assigned jobs mitigates, but does not eliminate the effects of self-selection. In the subsample with government-assigned jobs and thus with less prevalent self-selection, the estimated magnitude of precautionary savings is significantly greater than that obtained from the full sample without correcting for self-selection. This finding using Chinese data confirms that obtained by Fuchs-Schündeln and Schündeln (2005) from German data.

With changes in income uncertainty for SOE workers identified by the SOE reform and with self-selection mitigated by focusing on government assigned jobs, the estimated precautionary savings are significant both statistically and economically. In particular, precautionary savings account for about 40% of the total financial wealth accumulation for urban SOE households during the period from 1995 to 2002. Moreover, our evidence suggests that self-selection results in a downward bias of the estimated precautionary savings of at least 30%. Thus, both precautionary wealth and self-selection biases are quantitatively important for Chinese households.

Our identification and estimation rely on institutional features in China during a period with large structural transformations. In this sense, this approach is novel and contributes to the literature. The magnitudes of precautionary savings and self-selection biases obtained from the Chinese data turn out to be very similar to what Fuchs-Schündeln and Schündeln (2005) found from the German data. Thus, our study lends further empirical support to the importance of precautionary savings and self-selection biases.

Our study also reveals substantial heterogeneity of precautionary savings across different demographic groups. First, consistent with the life-cycle consumption theory, precautionary savings for younger households (25–44 years) are much stronger than for older households in the CHIP sample, confirming the finding of Gourinchas and Parker (2002) obtained

[1] See Carroll and Kimball (2008) for a survey.

Please cite this article as: H. He et al., Breaking the "iron rice bowl:" Evidence of precautionary savings from the chinese state-owned enterprises reform, Journal of Monetary Economics (2017), https://doi.org/10.1016/j.jmoneco.2017.12.002

from U.S. data. Second, workers in local SOEs have stronger precautionary savings motives than workers in SOEs owned by the central government or provincial governments, consistent with the fact that layoffs were more likely observed in small and local SOEs (Hsieh and Song, 2015). Third, the demographic groups more exposed to unemployment risks following the reforms, including female, low-skilled, or less educated SOE workers (see Appleton et al., 2002) accumulated more precautionary wealth in response to the SOE reform. A consistent message emerges from these exercises: the more vulnerable groups to the SOE reform tend to have a stronger precautionary savings motive.

This study further contributes to the literature by explicitly examining the extent to which changes in income expectations could affect the estimation of precautionary savings. Theory suggests that an increase in future unemployment risks not only raises savings through precautionary motives, but also through a permanent income hypothesis (PIH) channel since it reduces expected future income. We control for PIH effects by using information on both short-term income expectations and pension participation reported in the CHIP surveys. The evidence suggests that short-term income expectations do not seem to affect precautionary savings significantly, but pension participation is relatively more important.

The evidence of precautionary savings is robust when potential sample-selection biases are taken into account. It is also robust to a few alternative model specifications and alternative variable measurements.

Our study also adds to the literature on Chinese saving rate, although it does not intend to directly address the specific issue of what drives the rising Chinese saving rate. The recent studies by Chamon and Prasad (2010) and Chamon et al. (2013) show that the increased private burdens of expenditures on housing, education, and health care, combined with the lack of social safety net in China, help explain the rising Chinese saving rate. Some other studies examine the importance of life-cycle and other demographic factors for explaining China's high and rising saving rate (Horioka and Wan, 2007; Kraay, 2000; Modigliani and Cao, 2004). Imrohoroğlu and Zhao (2017) argue that an aging population and declines in within-family insurance under China's one-child policy have led to increased long-term care risks, and thus increases in China's saving rate. Curtis et al. (2015) present an overlapping generations model calibrated to Chinese data and show that demographic changes in China (such as changes in the dependency ratio caused by the one-child policy and population aging) account for a substantial fraction of the observed rise in China's saving rate. Wei and Zhang (2011) provide evidence that sex-ratio imbalances due to the one-child policy have led to 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. Our focus is instead on the general issue of identifying and quantifying precautionary savings. Our study provides empirical evidence that increases in income uncertainty associated with large structural changes in China have contributed to substantial precautionary wealth accumulation for urban Chinese households.

2. Some background of China's labor market and SOE reforms

Since this study exploits some institutional features of China's labor market to help identify changes in income uncertainty and self-selection biases for estimating precautionary savings, it is useful to provide some background information, with a brief description of the history of reforms in China's labor market and the SOE sector.

2.1. Labor market reforms

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. Job assignments were made typically through educational institutions (high schools or colleges) or local communal offices where potential workers registered their residency. The Ministry of Labour and Personnel assigned employment quotas to local governments, which then allocated the quotas to each school and local communal offices. Jobs were allocated to individuals who "need jobs," and individuals were not allowed to search for a job on their own. State-sector firms and government divisions were not allowed to recruit workers either. Instead, each working unit was assigned an annual employment quota. Final decisions of quota assignments were made by local Bureaus of Labor and Personnel. Once assigned to a job, a worker could not quit or switch jobs and a firm could not dismiss workers unless a crime was convicted (Meng, 2000). For those workers who obtained jobs through government assignment, they could not choose their jobs freely, and thus self-selection was unlikely.

To support the goal of industrialization, workers under the central-planning regime were paid subsistence wages and, in exchange, they were guaranteed life-time employment along with near-free housing, education, health care, and retirement benefits (Cai et al., 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.

In the late 1970s, the Chinese government under Deng Xiaoping's leadership initiated an "open door" economic policy and systematic economic reforms, setting off China's transition to a free-market economy. In the early 1980s, some experimental labor market reforms to the state sector started, in order to relax the rigid life-time employment rules under the central planning regime. In 1986, a systematic labor contract system was introduced to the state sector.[2] Under the rules of the labor contract system, state-sector employers were allowed to use examinations and conduct interviews in recruiting

[2] In particular, China's State Council announced the "Interim Provisions for State-Sector Recruiting" on July 12, 1986, which introduced the basic framework of a new labor contract system for hiring new workers in the state sector. The labor contract system became effective on October 1, 1986.

new workers. Labor contracts could be terminated if a worker was deemed incompetent during probation, violated work rules, or committed crimes while employed. A similar set of rules were applied to government jobs.

The new labor contract system was implemented gradually, and it was applied only for new hires. The share of workers in the state sector covered by the labor contact system was 3.7% in 1985, which gradually grew to 13% in 1990 (Meng, 2000). The labor contract reform fundamentally changed the mechanism for labor allocations in China. As labor contracts were more widely implemented over time, the share of government assigned jobs gradually declined.

In our CHIP sample, however, the majority of jobs were still assigned by the government. For example, in the 1995 CHIP survey, government assigned jobs accounted for 89% of jobs in GOV and 80% in SOE. Those workers who obtained jobs through other channels, including taking over from parents or relatives, through own job searching, or through employment agencies or other means, had different demographic characteristics than those with government assigned jobs. In the CHIP sample, they are typically younger, less educated, and less skilled, with lower income. In addition, GOV jobs had different characteristics than SOE jobs: GOV jobs were more concentrated in administrative services, whereas SOE jobs were mainly in production and other services.

2.2. SOE reforms

The labor market reforms implemented after 1986 also relaxed the tight controls over rural-to-urban migration flows. The influx of rural workers fueled expansions of private firms in urban areas. Furthermore, a wide-range of liberalization policies were adopted following Deng Xiaoping's "Tour of the South" in 1992. The boom in the private sector in urban areas intensified competition faced by SOE firms. 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. Although the labor contract rules gave SOE managers more flexibility in hiring new workers, they could not dismiss a worker on the ground of over-staffing (Meng, 2000). Indeed, very few SOE managers chose to fire workers unless their firms face serious financial stress or under the threat of closure. As competition from private firms intensified over time, 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, 2015).

These policy changes led to massive layoffs (*xia gang* in Chinese) of SOE workers starting in 1997, the scale of which was unprecedented. By the end of 1997, a cumulative of about 6.92 million SOE workers were laid off. The wave of layoffs reached a peak in 1999, and about 6.2 million SOE workers lost their jobs in that year. The layoff waves started to subside by 2002. According to the 2003 China Labor Statistical Yearbook, a remarkable total of over 27 million SOE workers had been laid off during the 5-year period from 1997 to 2002, which is equivalent to about 27% of total SOE employment in 1997. There were also large variations of the extent of SOE layoffs across regions and industries (Appleton et al., 2002). For example, in Fushun, a medium-sized city and a heavy industry base, layoffs accounted for about 42% of SOE employment in 2000 (see our case study in Appendix A).

However, government employees were little affected by the reform. According to the CHIP survey, the dataset used here for estimating precautionary savings, 58% of the individuals who had layoff experience prior to 2002 worked in SOEs, whereas only 2.3% of those individuals worked for the government.[3]

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, and low skilled workers 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 et al., 2002).

Since 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), for the SOE 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. Furthermore, our estimation below suggests that demographic groups more exposed to unemployment risks also had more precautionary savings in response to the SOE reform.

3. A simple model of precautionary saving

To illustrate how changes in income risks and job uncertainty could affect an individual's wealth accumulation, consider a simple two-period endowment economy with a continuum of households.

An individual household has the expected utility function

$$U = u(c_1) + \beta \mathbb{E} u(c_2), \tag{1}$$

[3] The remaining 39.7% worked in the private sector.

where $u(\cdot)$ is the period-utility function, $c_1 \geq 0$ and $c_2 \geq 0$ denote consumption in the two periods, $\beta \in (0, 1)$ is a subjective discount factor, and $\mathbb{E}$ is an expectation operator. Assume that the period utility function is strictly increasing, strictly concave, and continuously differentiable, with a positive third derivative.

The household chooses consumption plans c_1 and c_2 and savings s to maximize the utility function in Eq. (1), subject to the budget constraints

$$c_1 + s = w_1, \tag{2}$$

$$c_2 = (1+r)s + w_2, \tag{3}$$

where $r > 0$ denotes the net real interest rate.[4] The period-1 endowment $w_1 = \bar{w}$ is a constant, while the period-2 endowment w_2 is a random variable, the realization of which depends on the individual's employment status. An individual faces a probability p of unemployment, in which state she receives zero income. All else equal, an increase in p reduces expected future income, which may increase household savings for consumption smoothing. However, that increase in savings might reflect the effects of both a reduction in permanent income and an increase in future income uncertainty associated with an increase in p. In order to isolate the precautionary savings effects from the permanent income effects on saving, we restrict the period-2 income process such that the unconditional mean of w_2 is kept at $\bar{w}$, the same as that in period 1.

Labor income conditional on staying employed is itself a random variable, capturing idiosyncratic income risks that an employed individual might face. Specifically, the period-2 endowment follows the process:

$$w_2 = \begin{cases} \tilde{w}_h, & \text{with probability} \quad 1-p, \\ 0, & \text{with probability} \quad p, \end{cases}$$

where the income conditional on being employed $\tilde{w}_h$ is itself a random variable given by

$$\tilde{w}_h = \begin{cases} \frac{\bar{w}}{1-p} + \sigma, & \text{with probability} \quad \frac{1}{2}, \\ \frac{\bar{w}}{1-p} - \sigma, & \text{with probability} \quad \frac{1}{2}, \end{cases}$$

where σ is the standard deviation of the employment income in the second period.

The interior optimizing decisions for consumption and saving imply the intertemporal Euler equation

$$u'(\bar{w} - s) = \beta(1+r)\mathbb{E}u'((1+r)s + w_2), \tag{4}$$

where c_1 and c_2 are substituted out using the budget equalities.

Given the exogenous endowments and the interest rate, Eq. (4) determines the equilibrium savings s. Define the functions $f(s)$ and $g(s)$

$$f(s) = u'(\bar{w} - s), \tag{5}$$

$$g(s) = \beta(1+r)\mathbb{E}u'((1+r)s + w_2). \tag{6}$$

The concavity of $u(\cdot)$ implies that $f(s)$ increases with s whereas $g(s)$ decreases with s. The equilibrium savings s^* satisfies $f(s^*) = g(s^*)$.

To understand the effects of the two different types of risks (unemployment risks and idiosyncratic income risks conditional on employment) on precautionary savings decisions, we use the case with no risks in the second period (i.e., with $p = \sigma = 0$ so that $w_2 = \bar{w}$) as a baseline. The baseline equilibrium saving is determined by the intersection of the $f(s)$ curve and the $g(s)$ curve, with the latter evaluated at $w_2 = \bar{w}$. This equilibrium corresponds to point E_0 in Fig. 1, with the equilibrium savings given by s_0^*.

First, consider the effects of idiosyncratic income risks condition on being employed, which is captured by an increase in σ, while holding p constant.

Since $u'' < 0$ and $u''' > 0$, the marginal utility of period-2 consumption is a decreasing and convex function of the level of income w_2. It follows from the Jensen's inequality that a mean-preserving spread in w_2 (i.e., raising σ while keeping the mean of w_2 at $\bar{w}$) raises the expected marginal utility $\mathbb{E}u'((1+r)s + w_2)$ and shifts the $g(s)$ function upward, leading to an increase in equilibrium savings to s_1^* from s_0^* shown in the top panel of Fig. 1.

This example illustrates that precautionary savings increase with the variance of income shocks, provided that the third derivative of the utility function is positive. This is consistent with the textbook model of precautionary savings (Carroll and Kimball, 2008; Kimball, 1990). The variance of income σ^2 in this model corresponds to the idiosyncratic income risks conditional on employment in our empirical model below.

Next, consider the effects of an increase in the probability of unemployment on precautionary savings, while holding σ constant.

[4] To ensure non-negative consumption in the second period, we impose a borrowing constraint $s \geq -b$ for some non-negative borrowing limit b. The non-negativity of consumption implies that $b \leq w_2/(1+r)$ for all realizations of w_2, which corresponds to the natural borrowing constraint of Aiyagari (1994).

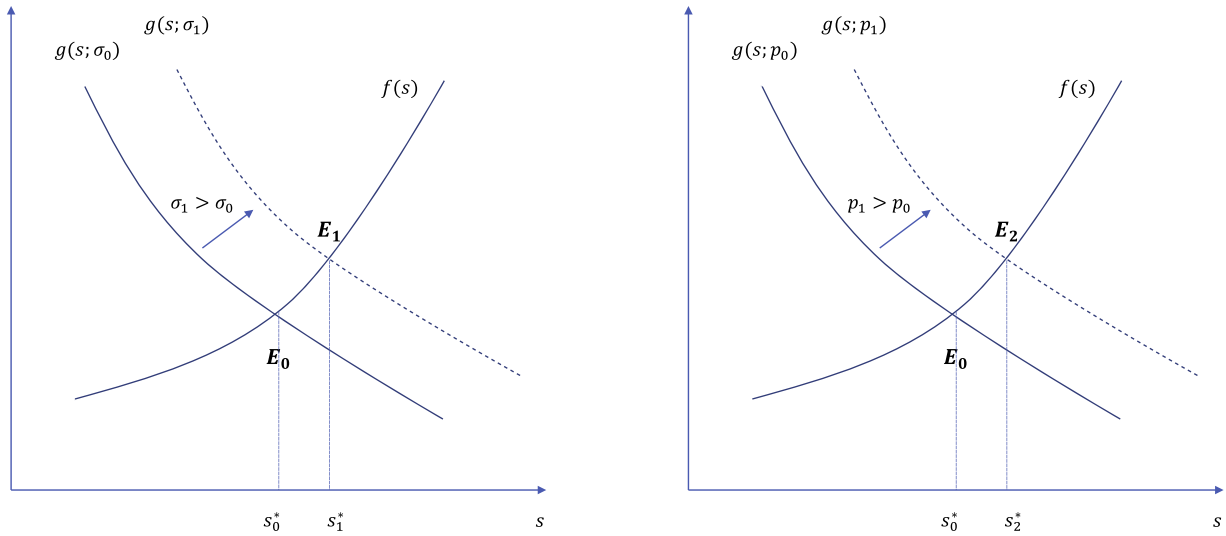

Fig. 1. Precautionary savings illustrated: Effects of an increase in income variance (left panel) and an increase in the probability of unemployment (right panel).

Eq. (6) implies that

$$g(s) \propto (1-p)\left[\frac{1}{2}u'\left((1+r)s+\frac{\bar{w}}{1-p}+\sigma\right)+\frac{1}{2}u'\left((1+r)s+\frac{\bar{w}}{1-p}-\sigma\right)\right]+pu'((1+r)s). \tag{7}$$

Differentiating $g(s)$ with respect to p leads to

$$\frac{\partial g}{\partial p} \propto \frac{1}{2}\left[u'(c_{2l})-u'(c_{2h+})+u''(c_{2h+})\frac{\bar{w}}{1-p}\right]+\frac{1}{2}\left[u'(c_{2l})-u'(c_{2h-})+u''(c_{2h-})\frac{\bar{w}}{1-p}\right], \tag{8}$$

where $c_{2h+}=(1+r)s+\frac{\bar{w}}{1-p}+\sigma$ and $c_{2h-}=(1+r)s+\frac{\bar{w}}{1-p}-\sigma$ denote period-2 consumption in the good and bad states conditional on employment, and $c_{2l}=(1+r)s$ denotes consumption in the unemployment state.

Since $u'(\cdot)$ is continuously differentiable, it follows from the Lagrangian Mean Value Theorem that there exist some $\bar{c}_+ \in (c_{2l}, c_{2h+})$ and $\bar{c}_- \in (c_{2l}, c_{2h-})$, such that

$$u''(\bar{c}_+)=\frac{u'(c_{2h+})-u'(c_{2l})}{c_{2h+}-c_{2l}}=\frac{u'(c_{2h+})-u'(c_{2l})}{\frac{\bar{w}}{1-p}+\sigma}, \tag{9}$$

and

$$u''(\bar{c}_-)=\frac{u'(c_{2h-})-u'(c_{2l})}{c_{2h-}-c_{2l}}=\frac{u'(c_{2h-})-u'(c_{2l})}{\frac{\bar{w}}{1-p}-\sigma}. \tag{10}$$

Then, Eq. (8) implies that

$$\frac{\partial g}{\partial p} \propto \frac{\bar{w}}{1-p}\left\{\frac{1}{2}\left[u''(c_{2h+})-u''(\bar{c}_+)\right]+\frac{1}{2}\left[u''(c_{2h-})-u''(\bar{c}_-)\right]\right\}+\frac{1}{2}\sigma\left[u''(\bar{c}_-)-u''(\bar{c}_+)\right]. \tag{11}$$

Since $u''(\cdot)>0$, the first two terms are both positive. The last term is negative if $\bar{c}_- < \bar{c}_+$. For a sufficiently small value of σ, however, the difference between $u''(\bar{c}_-)$ and $u''(\bar{c}_+)$ would be small and the last negative term would be dominated by the first two positive terms, implying a positive $\frac{\partial g}{\partial p}$. For example, in the extreme case with $\sigma=0$, $\frac{\partial g}{\partial p}$ is unambiguously positive.

Therefore, as p increases while σ is held constant (at a sufficiently small value), the $g(s)$ curve shifts upward, as shown in the bottom panel of Fig. 1. As a consequence, equilibrium savings increase from s_0^* at point E_0 to s_2^* at point E_2. Since the expected income level is held constant at $\bar{w}$, the rise in savings following an increase in the probability of unemployment captures the precautionary savings effects rather than the response to changes in permanent income.

4. Empirical strategies

In light of the theoretical model in Section 3, we construct and estimate an empirical model to examine the effects on precautionary savings of both unemployment risks (i.e., p, captured by an SOE dummy) and idiosyncratic labor income risks conditional on staying employed (i.e., σ, captured by our RISK measure below).

4.1. The empirical model

Following Lusardi (1998) and Carroll et al. (2003), consider the empirical specification

$$\frac{W_i}{P_i} = \beta_0 + \beta_1 SOE_i + \beta_2 RISK_i + \beta_3 \log(P_i) + \beta_4' Z_i + v_i. \tag{12}$$

In this model, the dependent variable is the ratio of financial wealth W_i to permanent income P_i for household i, as in Lusardi (1998). This ratio measures the household's cumulative savings relative to her permanent income.

The explanatory variable SOE_i is a dummy variable that 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).[5] It captures the unemployment risks for SOE workers relative to GOV employees following the SOE reform in the late 1990s. The explanatory variable $RISK_i$ measures idiosyncratic income risks conditional on being employed. As discussed in the simple theoretical model in Section 3, the SOE dummy and the $RISK_i$ measure capture two different types of income risks. In our sample, the correlations between these two variables are very low (in absolute values), with a correlation of about −0.05 in 1995 and −0.17 in 2002.

The empirical model includes the log-level of permanent income P_i as an explanatory variable to control for the potential effects of non-homothetic preferences. It also includes a number of demographic control variables summarized by the vector Z_i. The term v_i denotes regression errors.

We use a difference-in-differences approach to estimating precautionary savings. The CHIP data do not have a panel dimension and thus do not keep track of individual households over time. Thus, the empirical model in Eq. (12) is estimated by running two separate cross-sectional regressions, one with the pre-treatment group in 1995 and the other with the post-treatment group in 2002.[6]

The key parameter of interest is β_1, the coefficient for the SOE dummy. The estimated β_1 from each regression (denoted by β_1^{1995} and β_1^{2002}, respectively) captures – all else equal – the average excess savings by SOE workers relative to GOV workers. The difference $\Delta\beta_1 = \beta_1^{2002} - \beta_1^{1995}$ then captures the magnitude of precautionary savings of the SOE workers caused by increases in their unemployment risks following the breaking of the iron rice bowl.

Following Fuchs-Schündeln and Schündeln (2005), the permanent income measure is instrumented using education dummies and interactions of education with age and age-squared as instrumental variables. We address the issue that arises with observations of zero wealth by treating it as a censored data problem and estimating an instrumental variable Tobit regression (IV-Tobit).[7] In a robustness check, the model in Eq. (12) is also estimated by eliminating the zero-wealth observations from our sample and then applying the standard two-stage least squares (2SLS) method (see Section 6.3).

4.2. The data

The data are taken from the Chinese Household Income Project (CHIP) surveys. The 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 four different years—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 15,000–20,000 households in 10 provinces in China. The surveys contain detailed data on households' employment status, education, income, expenditures, and other demographic information. The CHIP data have been frequently used in the empirical literature.[8]

We focus on the sample of urban households in the CHIP surveys of 1995 and 2002, which span the period of China's large-scale SOE reforms that had led to massive layoffs of SOE workers. More importantly, both surveys contain data on households' wealth and its compositions, allowing us to examine the quantitative importance of precautionary wealth accumulation caused by the SOE reform.[9]

Our sample is restricted to include those households whose heads work in the SOE sector or the GOV sector. The SOEs in the sample include 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. The sample is further restricted 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 savings behaviors of workers close to retirement ages change dramatically for reasons more closely related to life-cycle factors than to income uncertainty (Carroll and Samwick, 1998; Gourinchas and Parker, 2002).[10]

[5] 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.

[6] The lack of panel data implies that the treatment group (the SOE workers) may not be comparable across time. In particular, the post-treatment group includes only those who survived the SOE reform and those who chose not to quit from their SOE jobs. These issues may cause biases in our estimation. This sample selection issue is addressed by using the standard propensity score weighting approach (see Section 6).

[7] In our sample, 12.1% of households have zero wealth in 1995 and this share declined to 9.7% in 2002.

[8] The website http://www.icpsr.umich.edu/icpsrweb/ICPSR/series/243 lists some recent studies that use the CHIP survey data.

[9] The CHIP surveys in 1988 and 2007 do not report wealth information, and thus they are less useful for studying precautionary savings.

[10] In our sample periods, the normal retirement age for female workers in China is between 50 and 55; for male workers, it is between 55 and 60.

Please cite this article as: H. He et al., Breaking the "iron rice bowl:" Evidence of precautionary savings from the chinese state-owned enterprises reform, Journal of Monetary Economics (2017), https://doi.org/10.1016/j.jmoneco.2017.12.002

Table 1
Definition of variables.

Variable	Description
Financial wealth (*W*)	Balances in checking accounts, savings accounts, stocks, bonds, contributions to employer funds, and loans to others
Very liquid assets (VLA)	Financial wealth minus contributions to employer funds and loans to others
Non-housing non-business wealth (NHNBW)	Financial wealth plus estimated market value of durables and other assets, minus total debt
Income	Total annual income of the household head, including salaries and bonuses, subsidies, other labor income, property income, and transfer income
Income risk (*RISK*)	The log of the variance of log annual income over the past few years
SOE	Dummy variable that equals one if household head works for SOE and zero for Government
Permanent income (*P*)	Constructed based on earnings by household heads in the current year and the recent past
W/*P*	Ratio of wealth to permanent income
Age	Age of household head
Male	Dummy variable that equals one if household head is male and zero otherwise
Married	Dummy variable, equals one if household head is married and zero otherwise
Education	Dummy variable for household head's education: elementary school or below, junior middle school, senior middle school, or some college (or above)
Occupation	Dummy variable for 4 occupations: (1) professional, (2) director or manager, (3) skilled or office workers, and (4) unskilled, service workers or other
Health care	Dummy variable for health care status: public health care, public health insurance, or self pay (see text)
Non-homeowner	Dummy variable for housing ownership, equals 1 if not a house owner and 0 otherwise
Child age	Mean age of children in household
Num. of boys	Number of boys in household
Children at school	Number of children at school
Household size	Number of residents in a household
Job assigned by Gov.	Dummy variable that equals one if the household head obtained current job through government assignments and zero otherwise
Income decline	Dummy variable that equals one if the household head expects income to decline in the next 5 years and zero otherwise
No-pension	Dummy variable that equals one if the household head did not contribute to pension funds and zero otherwise

With these restrictions, the sample has 4390 household-level observations in 1995, consisting of 2977 SOE workers and 1413 GOV employees; and 3027 observations in 2002, consisting of 1702 SOE workers and 1325 GOV employees.

4.3. The measurement

The variables used in the regressions include wealth (*W*), permanent income (*P*), the SOE dummy, a measure of idiosyncratic risks (*RISK*), and a set of household characteristics. Table 1 shows the definitions of the variables used in our study. Tables 2 and 3 show the summary statistics of those variables, both in the full sample and for each sector (SOE or GOV).

To estimate the importance of precautionary savings, we focus on relatively liquid components of household wealth (Carroll and Samwick, 1998), measured in our sample by financial wealth (*W*), which is the sum of checking accounts, savings accounts, stocks, bonds, contributions to employer funds, and loans to others. Table 4 shows some summary statistics of the household portfolio compositions in our CHIP sample. Our measure of financial wealth corresponds to asset categories 1–6 in the table.

Our study uses the stock of financial wealth instead of the flow of savings (or the saving rate) for two reasons. First, unlike savings flows, financial wealth is not influenced by high-frequency fluctuations in income and expenditures. Thus, it is better able to capture long-run (or average) savings 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 savings or the saving rate, which is indirectly calculated based on income and consumption expenditures.

The measure of permanent income is constructed following the approach by Fuchs-Schündeln and Schündeln (2005). The CHIP surveys report earnings of the household heads in the current year and the recent past. In particular, the 1995 survey reports earnings in 1990 through 1995 and the 2002 survey reports earnings in 1998 through 2002. Permanent income is constructed in three steps. The first step calculates a household head's earnings relative to the average earnings of all

Table 2
Summary statistics of the full sample.

Variable	1995			2002		
	Obs.	Mean/%	SD	Obs.	SD	SD
Financial wealth (*W*)	4390	9556	9892	3027	25,669	27,443
Permanent income (*P*)	4390	7520	3131	3027	12,843	6018
Income risk (*RISK*)	4390	−3.41	1.28	3027	−3.76	1.92
Age	4390	40.91	7.37	3027	42.61	6.88
Child age	4390	11.65	6.94	3027	12.5	7.58
Num. of boys	4390	0.57	0.58	3027	0.47	0.53
Children at school	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	44.0%	
Unskilled/other worker	4390	16.7%		3027	15.9%	
Health Care						
Own payment	4390	19.9%		3027	23.1%	
Public health care	4390	71.3%		3027	35.0%	
Public health insurance	4390	8.8%		3027	41.9%	
Non-homeowner	4390	58.0%		3027	19.6%	
SOE	4390	67.8%		3027	56.2%	
Job assigned by Gov.	4375	82.9%		3018	71.9%	
Income decline	N.A	N.A		3020	18.4%	
No-pension	4390	63.0%		3027	32.2%	

Notes: Data are taken from CHIP surveys. Monetary values are in constant Chinese Yuan units, with 2002 as the base year.

Table 3
Comparison of selected worker characteristics: GOV vs. SOE.

Variable	Gov			SOE		
	Obs.	Mean/%	SD	Obs.	Mean/%	SD
1995						
Financial wealth (*W*)	1413	10,004	9940	2977	9343	9864
Permanent income (*P*)	1413	7905	3063	2977	7337	3146
W/*P*	1413	1.306	1.296	2977	1.305	1.383
Non-homeowner	1413	54.6%		2977	59.7%	
Job assigned by Gov.	1408	89.3%		2967	79.8%	
Income decline	N.A	N.A		N.A	N.A	
No-pension	1413	82.9%		2977	53.5%	
2002						
Financial wealth (*W*)	1325	27,041	27,924	1702	24,600	27,023
Permanent income (*P*)	1325	13,979	5853	1702	11,958	5998
W/*P*	1325	1.981	2.117	1702	2.136	2.481
Non-homeowner	1325	16.5%		1702	22.0%	
Job assigned by Gov.	1319	75.7%		1699	68.9%	
Income decline	1321	11.4%		1699	23.8%	
No-pension	1325	51.0%		1702	17.6%	

Notes: Data are taken from CHIP surveys. Monetary values are in constant Chinese Yuan units, with 2002 as the base year.

households in each year with reported earnings. The second step takes the time-series average of the household relative earnings. The third step multiplies the household head's earnings 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. To mitigate potential measurement

Table 4
Wealth compositions.

Items	1995			2002		
	Mean	SD	% of W	Mean	SD	% of W
(1) checking accounts	6400	7844	67%	15,406	20,372	60%
(2) savings accounts	1244	2406	13%	4666	7674	18%
(3) stocks	343	1705	4%	3277	10,668	13%
(4) bonds	858	2484	9%	712	4427	3%
(5) contributions to employer funds	396	1904	4%	397	3097	2%
(6) loans to others	315	1456	3%	1211	5126	5%
Very liquid assets (VLA, items (1)–(4))	8845	9459		24,061	26,264	
Financial wealth (W, items (1)–(6))	9556	9892		25,669	27,443	
Non-housing, nonbusiness net worth (NHNBW)	19,429	15,876		39,111	40,337	
Sample size	4390			3027		

Notes: Data are taken from CHIP surveys. Monetary values are in constant Chinese Yuan units, with 2002 as the base year. Non-housing nonbusiness net worth (NHNBW) equals financial wealth plus estimated market value of durable goods and other assets, minus total debt.

errors, permanent income is instrumented by education dummies and interactions of education with age and age-squared as instruments (Fuchs-Schündeln and Schündeln, 2005).[11]

Idiosyncratic income risks conditional on being employed ($RISK_i$) are measured by the log of the variance of log annual household head income *across time* (in the current year and the recent past). This measure is different from the conventional measure of income risks based on *cross-sectional* variances of log income (Carroll and Samwick, 1998). For robustness, we also estimate our model using the conventional risk measure (see Section 6).

The regression controls for household demographic characteristics, including the household head's occupation (professional, director or manager, skilled or office worker, or unskilled or other workers), education level (elementary school or below, junior middle school, senior middle school, or some college (or above)), health care status (public health care, public health insurance, or own payments), home ownership status, age, age-squared, gender, marital status, the household size, the ages of children, the number of boys, the number of children at school, and the industry and the province where the household head worked.

The health care reform enacted in 1998 significantly changed the share of household expenditures on 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.

Purchasing a house is argued to be one of the major motives of savings 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 home-ownership 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 savings for home purchases by including a non-homeownership dummy that takes a value of one if the household is not a home owner and zero otherwise.

Since the SOE reform and the massive layoffs hit some industries and geographic areas more heavily than others, dummy variables that indicate the industries and provinces where the household head worked are included in the regressions as controls.

Table 3 compares some key characteristics between GOV and SOE workers. It shows that the reform 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. The wealth–income ratios (W/P), however, were similar (at around 1.3). In 2002, the gaps in both wealth and income widened substantially across the two sectors compared to the pre-reform year in 1995. More importantly, the wealth–income ratios diverged. In particularly, the W/P ratio for the SOE workers increased much more than that for the GOV workers, suggesting that SOE workers on average raised savings more than GOV workers did during the reforming years. Consistent with this suggestive evidence, our estimation below shows that SOE workers did increase their savings significantly relative to GOV workers in response to increased income uncertainty associated with the massive layoff waves.

[11] Potential outliers in the wealth measures and permanent income are detected and excluded by box plots. In particular, define the interquartile range as $IQR = Q_3 - Q_1$, where Q_1 and Q_3 denote the first and third quartile, respectively. Observations that are outside of the interval $(Q_1 - 3IQR, Q_3 + 3IQR)$ are treated as potential outliers and excluded from the sample.

Table 3 also shows the homeownership rate for the two types of workers. In 1995, the home ownership rate for GOV workers was slightly higher than for the SOE workers (45% vs. 40%). In 2002, the home ownership rate rose for both groups (to 83% for GOV workers and 78% for SOE workers), although the difference in the average home ownership rates across the two groups remained unchanged.

In the 1995 sample, a large majority of jobs were assigned by the government in both sectors. In particular, nearly 90% of the GOV jobs and 80% of the SOE jobs were assigned by the government. In 2002, the share of government assigned jobs declined somewhat in both sectors (to about 76% in the GOV sector and 69% in the SOE sector), although they still constitute a majority of all jobs. When we estimate the importance of precautionary savings, we restrict our sample to government assigned jobs in both years to mitigate the self-selection bias.

The SOE reform in the late 1990s led to different income expectations between the two groups. In the 2002 survey, about 24% of the SOE workers expected to have lower income in the next five years, compared to 11% of GOV employees who expected income to decline.[12]

The Chinese government also started to reform the pension system in the early 1990s. Under the traditional system, retirement benefits were directly provided by the government and workers or employers were not required to contribute to pension funds. Starting in 1991, the Chinese government gradually pushed out a series of pension reform plans, aiming to establish a new pension system that combines a pay-as-you-go system and a mandatory individual retirement account. The new pension system requires both a worker and her employer to make contributions to a pension fund established on the worker's behalf, based on a fraction of the worker's salary. The new pension system was gradually implemented over time: it was started in 1991, broadened in scope in 1995, and accelerated in 1997.[13]

Following the pension reforms, the share of individuals who contributed to pension funds rose from 37.0% in 1995 to 67.8% in 2002, as shown by Table 2. The reforms were implemented more broadly in SOEs than in the government sector. Thus, a larger share of SOE workers had contributed to pension funds than GOV employees. As shown in Table 3, the share of SOE workers who reported pension contributions rose from 46% in 1995 to 82% in 2002, and the share of GOV workers with pension contributions also increased from 17% to 49% during that period.

As discussed in the theoretical model in Section 3, pessimistic income expectations and pension participation can change savings through PIH effects, but those changes do not reflect precautionary savings. Section 5.4 below discusses the estimation of precautionary savings, controlling for the effects of income expectations and pension participation.

5. Empirical results

The empirical model in Eq. (12) is estimated using the CHIP data. 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.

Table 5 shows the estimation results. Columns (i) and (ii) show the full-sample estimation for 1995 and 2002, respectively. Columns (iii) and (iv) show the estimation with the sample restricted to government assigned jobs.

5.1. Evidence of precautionary savings

First, consider the estimation results in the full sample (Columns (i) and (ii) in Table 5). The estimated value of β_1 in 1995 is slightly negative (at -0.047) and statistically insignificant, indicating that the savings behaviors of SOE and GOV workers were statistically and economically similar in 1995 when demographic characteristics are controlled for. In 2002, however, SOE workers accumulated significantly more wealth than GOV employees (reflected by a much large estimate of $\beta_1 = 0.366$). The Chow test rejects the null hypothesis that β_1 is identical between 1995 and 2002, with a p-value of .055. The difference between the two estimated values of β_1 $(0.366 - (-0.047) = 0.413)$ is not just statistically significant, but also economically large; it suggests that, all else equal, the extra savings of SOE workers relative to GOV workers after the reform were about 0.413 times of their annual permanent income, or about 5 months worth of permanent income.[14] Despite the potential downward bias caused by self-selection in the full sample, the evidence here suggests that increases in the relative income uncertainty for SOE workers after the reform led to significant precautionary savings.

The estimated coefficient β_2 on *RISK* suggests that idiosyncratic income risks conditional on being employed had positive and significant effects on savings in both 1995 and 2002. Thus, all households—working in SOE or GOV—responded to increases in idiosyncratic income risks by raising savings, consistent with the implication of the theoretical model presented in Section 3. This source of savings represents households' responses to variations in idiosyncratic income risks conditional

[12] The 1995 survey does not include a question about income expectations. Before the reform, since workers in both sectors all held an iron rice bowl, they should not expect their income to decline.

[13] The official policy announcements were made by China's State Council in three important documents (in Chinese): (1) "On Reforming the Old-Age Insurance System for Enterprise Employees" (State Council Document No. 33, 1991); (2) "On Deepening the Reform of the Old-Age Insurance System for Enterprise Employees" (State Council Document No. 6, 1995); and (3) "On the Establishment of a Unified Basic Old-Age Insurance System for Enterprise Employees" (State Council Document No. 26, 1997). See He et al. (2017) for further discussions about China's pension reforms.

[14] The dependent variable in our model is the ratio of financial wealth to annual permanent income (W/P). Thus, an increase in W/P of 0.413 units implies an increase in W of an amount equivalent to $0.413 * 12 = 4.96$ months of permanent income.

on being employed, and it is different from the responses of savings behaviors to unemployment risks specific to the SOE households captured by β_1. Furthermore, the estimated values of β_2 for 1995 and 2002 are similar in magnitude and both are significant. In contrast, the value of β_1 was much larger in 2002 than in 1995 and turned from insignificant to significant. In other words, whereas β_2 stays roughly constant over time, β_1 has much larger time-variations that capture the effects of changes in unemployment risks for SOE workers caused by the reform.

The estimation also suggests that households with high permanent income tend to save more, consistent with the presence of non-homothetic preferences, although the coefficient on $\log(P)$ was insignificant in 1995 and became significant in 2002.

The occupation of the household head had mixed effects on savings. The household head occupations are partitioned into four groups: professionals, directors or managers, skilled workers, and unskilled workers and others. Professionals are our reference group. The estimation suggests that directors and managers saved more than professionals in 1995, although the differences in savings behaviors across occupation groups become insignificant in the 2002 sample.

The coefficients of both health care dummy variables are small and insignificant in 1995 but become significantly negative in 2002 (the reference group here includes those households who self financed health care expenditures). This result is consistent with China's health care system and its reform. In 1995, most workers were covered under a near-free public health care system, so that the health care status did not have significant impact on households' savings. However, after the health insurance reform that started in 1998, a significant fraction of health care spending was shifted to private households (Huang and Gan, 2017). 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 savings against potential health expenditure shocks.

To control for the effects of education reforms on households' savings behavior and potential competitive savings motive in the marriage market emphasized by Wei and Zhang (2011), our regression includes 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 leads to lower disposable income and reduced wealth accumulation. The number of boys contributes positively to savings in 1995, 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 1% 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).

To control for the effects of housing reform on savings, the regression includes a non-homeownership dummy. The coefficient for this variable is insignificant for both years, possibly reflecting that the housing market in China was still underdeveloped through 2002.

In addition, the regression also controls for other demographic variables such as age, age-squared, sex of the household head, marital status, and the household size. The full sample estimation suggests that households with female heads save significantly more than those with male household heads in 1995; and they save even more in 2002. Married households also saved more, although the effects of marital status on wealth accumulation were statistically significant only for 1995, not for 2002. The household size had little effects on savings in 1995, but larger households saved significantly more in 2002.

5.2. The self-selection bias

The literature shows that self-selection can lead to a substantial downward bias in the estimated magnitude of precautionary savings (Fuchs-Schündeln and Schündeln, 2005). An individual with high risk aversion has an incentive to choose a job with low income risk and, all else equal, she is also likely to save more. Without correcting self-selection biases, the estimation using the full sample may understate the true magnitude of precautionary savings.

To control for potential self-selection bias, we follow the approach in Fuchs-Schündeln and Schündeln (2005) and restrict our sample to workers whose jobs were assigned by the government. As discussed in Section 2, jobs were assigned primarily based on quotas and "needs" of the local governments, rather than the preferences of individual workers. Thus, under the regime with government job assignments, individual workers are less likely to self select into different sectors.

The estimation results using the subsample with government assigned jobs are shown in Table 5 (Columns (iii) and (iv)). Our estimation shows that self-selection indeed caused a significant downward bias in the estimated value of β_1 after the reform, but not before. In particular, the estimated value of β_1 in 1995 in the subsample with government assigned jobs is similar to that in the full sample (-0.012 vs. -0.047), both are statistically insignificant. In 2002, however, the estimate of β_1 for workers with assigned jobs becomes much greater and statistically more significant than that in the full sample

(0.539 vs. 0.366). As in the full sample, the Chow test for the SOE dummy in this subsample estimation strongly rejects the null hypothesis that the estimated value of β_1 in 2002 is identical to that in 1995, with a p-value of .049.[15]

To summarize, there are two important findings. First, even without controlling for self-selection biases, we find significant presence of precautionary savings caused by the large-scale SOE reform. Second, self-selection causes significant downward biases in estimating precautionary wealth accumulation. When self-selection is mitigated by focusing on government assigned jobs, the magnitude of precautionary savings rises significantly relative to that estimated from the full sample.

5.3. Quantitative importance of precautionary savings

Using the SOE reform as a natural experiment, we have identified the presence of precautionary savings. But to what extent can precautionary savings 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 contribution of precautionary savings to wealth accumulation. The idea is to compare 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.

This idea is implemented with three steps. The first step calculates the model's predicted wealth held by SOE workers in 1995 and in 2002 (denoted by $\hat{W}^{soe}_{1995}$ and $\hat{W}^{soe}_{2002}$, respectively) using the baseline estimation results after correcting self-selection biases based on the subsample with government assigned jobs (Columns (iii) and (iv) in Table 5).

The second step computes 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 step, except that the SOE dummy is set to zero. Denote by $\tilde{W}^{soe}_t$ the counterfactual wealth holdings of SOE workers in year $t \in \{1995, 2002\}$.

The third (and final) step computes the magnitude of precautionary wealth accumulation (denoted by W^{ps}) stemming from the large-scale SOE reform according to the relation

$$W^{ps} = (\hat{W}^{soe}_{2002} - \hat{W}^{soe}_{1995}) - (\tilde{W}^{soe}_{2002} - \tilde{W}^{soe}_{1995}). \tag{13}$$

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

Our calculation suggests that, with self-selection biases corrected, precautionary savings account for 43.1% of financial wealth accumulation for SOE households between 1995 and 2002, and this magnitude is statistically significant at the 5% level, with a standard error of 0.204.[16] Thus, the SOE reform in the late 1990s led to quantitatively important precautionary savings by SOE households.

In comparison, without correcting self-selection biases, the contribution of precautionary savings to SOE household wealth accumulation would have been lower at 31% (calculated based on Columns (i) and (ii) in Table 5), which is also significant with a standard error of 0.152. Therefore, self-selection leads to a downward bias of the estimated precautionary savings of about 28% ($(0.431 - 0.31)/0.431 \approx 0.28$).

Our findings on the quantitative importance of both precautionary savings and self-selection biases are consistent with those obtained by Fuchs-Schündeln and Schündeln (2005) using German data.[17]

5.4. The PIH effects

The large-scale SOE reform not only led to significant changes in the relative job security between SOE and GOV employees, it also produced potentially large differences in future income expectations between the two groups. As illustrated in the simple theoretical model in Section 3, a worker who expects declines in future income would like to increase savings, but such an increase in savings reflects a desire for intertemporal consumption smoothing (i.e., a PIH effect) rather than a motive of precautionary savings. Similarly, pension participation can also affect individuals' savings decisions through the PIH effect. As described in Section 4.3, the new pension system was implemented gradually over time and not all individuals in a given year in our sample had contributed to pension funds.

[15] Comparing the estimation results between the full sample and the subsample with government assigned jobs, we see that not only the coefficient on the SOE dummy changes, but some other coefficients, especially those on occupations changed significance. In particular, the coefficient on unskilled and other workers in 2002 turned from insignificant in the full sample to significant in the subsample. This difference partly reflects the fact that unskilled workers were less likely to obtain jobs through government assignments than other occupations. In our 2002 sample, the share of government assigned jobs for unskilled workers is about 52%, much lower than that for professionals and directors or managers (about 81%) or that for skilled workers (about 70%), suggesting that self-selection biases are likely more pronounced for unskilled workers.

[16] We calculate the standard error of the contribution of precautionary savings to financial wealth accumulation by using the STATA command "nlcom," which computes standard errors for nonlinear combinations of parameter estimates based on the delta method.

[17] Fuchs-Schündeln and Schündeln (2005) find that precautionary wealth estimated from a sample with former GDR households who chose their jobs before the German reunification amounts to 22% of total wealth, while that in the West German sample amounts to 13% of total wealth. They argue that, since the former GDR sample is not subject to self-selection biases while the West German sample is, self-selection causes a downward bias in the estimation of precautionary savings of about 41%, which is slightly larger than our estimate of 30% using Chinese data.

Please cite this article as: H. He et al., Breaking the "iron rice bowl:" Evidence of precautionary savings from the chinese state-owned enterprises reform, Journal of Monetary Economics (2017), https://doi.org/10.1016/j.jmoneco.2017.12.002

This section discusses our approach to controlling for PIH effects in estimating precautionary savings, focusing on the effects of both short-term income expectations and pension participation.

5.4.1. Short-term income expectations

The 2002 CHIP survey contains a question that asks households about their expectations of income changes over the next five years (increase, decrease, or no change). As Table 3 shows, a significant fraction of SOE workers (23.8%) surveyed in 2002 expected future income to decline, and a much smaller fraction of GOV workers (11.4%) expected income to decline. Thus, the reform has caused different income expectations in addition to different unemployment risks across the two groups of workers.[18]

To control for the effects from income expectations, we construct a dummy variable "income decline" that equals one if the household head expected income to decline in the next 5 years and zero otherwise. The "income decline" dummy is then added to the baseline model as a control, and the model is re-estimated using the subsample with government assigned jobs. The estimation results are reported in Table 6 (Columns (i) and (ii)). Since the 1995 survey does not contain information about income expectations, we are able to include the income decline dummy as a control only for the 2002 regression. The estimated coefficient on the dummy variable "income decline" is small and insignificant (0.002).

The coefficient on the SOE dummy (β_1) under this model specification captures the extra savings by an SOE household relative to a comparable GOV household in the group that did not expect income to decline (i.e., with the "income decline" dummy set to zero). The estimated β_1 is similar to that from the baseline model (0.559 vs. 0.539), and it remains significant at the 5% level. This finding suggests that controlling for changes in short-term income expectations does not have a large impact on our baseline estimation of precautionary savings by SOE households following the reform.

5.4.2. Pension participation

To control for the effects of pension participation, we construct a dummy variable "no-pension" that takes a value of one if a household head *did not* make contributions to pension funds and zero otherwise. The no-pension dummy is included in the baseline empirical model as an additional control. The estimation results are shown in Columns (iii) and (iv) of Table 6.

The estimated coefficient on the no-pension dummy is small and insignificant in the 1995 sample, but turns positive and significant at the 5% level (0.350) in the 2002 sample. This positive coefficient implies that, all else equal, workers who did not participate in the pension system saved significantly more than those who did.

When the no-pension dummy is included in the model as a control, the coefficient on the SOE dummy (β_1) measures the extra savings by an SOE household relative to a comparable GOV household, conditional on that they both participated in the new pension system. In the 1995 regression, the estimated β_1 is small and insignificant, similar to that obtained in the baseline regression (−0.016 vs. −0.012). However, the 2002 estimation of β_1 becomes large and significant at the 5% level, and its magnitude is modestly greater than that obtained from the baseline regression (0.621 vs. 0.539). These findings suggest that, controlling for the PIH effects from pension participation, the estimated precautionary savings are quantitatively more important than in the benchmark case without such controls.

Our findings here suggest that, although changes in short-term income expectations did not have significant impact on savings, pension participation was relatively more important.

5.5. SOE firm sizes

There is evidence that the impact of the large-scale SOE reform on SOE workers in large firms was very different from that on workers in small- or medium-sized 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 energy, telecommunications, and heavy manufacturing) were corporatized or consolidated into large state-owned conglomerates, while smaller and loss-making SOE firms were shut down or privatized (Hsieh and Song, 2015). Evidence suggests that those large SOEs that survived the reorganization gained even more government protections for their monopoly power, leading to higher profits than before the reform (Li et al., 2015).

Since the government policy explicitly favored large SOEs, workers in large SOEs faced smaller increases in unemployment risks than those in small SOEs (Appleton et al., 2002). Therefore, one should expect to see stronger precautionary savings motives for workers in smaller (and riskier) SOEs.

To examine this issue, we divide the SOE firms into two groups: central or provincial SOEs (CSOE) vs. local SOEs (LSOE).[19] Consistent with the evidence provided by Appleton et al. (2002), LSOE workers in the 2002 CHIP sample reported much more layoff experience than CSOE workers. The benchmark model in Eq. (12) is also modified by replacing the SOE dummy

[18] The question on income expectations is not available in the 1995 CHIP survey. In the 2002 sample, the fact that SOE workers are more likely to expect an income decline than GOV workers could be driven by increased unemployment risks. In that sample, the correlation between the SOE dummy variable and the expected income decline dummy is small but positive (at 0.158 and significant at the 1% level). Since the survey does not provide information on the size of the expected income declines, it is hard to completely disentangle precautionary savings from income expectations conditional on staying employed. Still, the qualitative information about expected income changes provided in the CHIP survey helps to control for the PIH effects on household savings.

[19] LSOE also includes urban collective enterprises.

Table 5
Baseline IV-Tobit regressions.

Dep. variable:	Full sample		Job assigned	
W/P	(i) 1995	(ii) 2002	(iii) 1995	(iv) 2002
SOE	−0.047	0.366*	−0.012	0.539**
	(0.093)	(0.194)	(0.094)	(0.264)
RISK	0.197***	0.180***	0.170***	0.145***
	(0.048)	(0.043)	(0.052)	(0.049)
log(P)	1.253	2.473***	0.846	2.840**
	(0.905)	(0.854)	(1.010)	(1.261)
Director/manager	0.164**	−0.039	0.190**	0.122
	(0.078)	(0.157)	(0.080)	(0.177)
Skilled worker	−0.027	−0.009	−0.07	0.148
	(0.095)	(0.158)	(0.104)	(0.188)
Unskilled/others	−0.003	0.362	−0.127	0.798*
	(0.164)	(0.319)	(0.180)	(0.485)
Public health care	0.042	−0.831**	0.024	−0.976**
	(0.168)	(0.340)	(0.189)	(0.454)
Public med insurance	0.090	−0.594**	−0.009	−0.654*
	(0.145)	(0.299)	(0.165)	(0.396)
Child age	0.005	−0.002	0.006	0.000
	(0.005)	(0.009)	(0.006)	(0.011)
Num. of boys	0.022	−0.283***	0.045	−0.281**
	(0.044)	(0.102)	(0.047)	(0.120)
Children at school	−0.050	−0.274**	−0.097	−0.221
	(0.061)	(0.120)	(0.064)	(0.143)
Non−homeowner	0.083	−0.055	0.018	−0.095
	(0.065)	(0.148)	(0.068)	(0.179)
Age	−0.000	0.156	0.033	0.038
	(0.048)	(0.106)	(0.050)	(0.125)
Age^2*100	−0.007	−0.166	−0.046	−0.033
	(0.056)	(0.125)	(0.057)	(0.145)
Male	−0.452***	−0.862***	−0.364***	−0.807***
	(0.091)	(0.131)	(0.098)	(0.169)
Married	0.461***	0.228	0.503***	0.385
	(0.158)	(0.324)	(0.191)	(0.357)
Household size	−0.012	0.389***	−0.039	0.270*
	(0.047)	(0.125)	(0.050)	(0.152)
Log-likelihood	−8711.71	−7855.88	−7045.37	−5519.87
p-value				
(Chow test for SOE)		0.055		0.049
Sample size	4390	3027	3627	2170

Notes: Columns (i) and (ii) show the estimation results using the full sample. Columns (iii) and (iv) show those using the subsample with government assigned jobs. All regressions include controls for fixed effects of locations (provinces of current residence) and industries. Robust standard errors are in parentheses. ***, **, and * indicate p-values of less than 1%, 5%, and 10%, respectively.

variable with the two dummy variables, indicating whether the household head works in a CSOE or an LSOE. The regression model is now

$$W_i/P_i = \beta_0 + \beta_1^{CSOE} CSOE_i + \beta_1^{LSOE} LSOE_i + \beta_2 RISK_i + \beta_3 \log(P_i) + \beta_4' Z_i + v_i \tag{14}$$

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

Table 7 (Panel A) reports the regression results from the sample with government assigned jobs. From 1995 to 2002, β_1^{CSOE} increased from −0.157 to 0.343, although it is not significant in both years. In contrast, β_1^{LSOE} was estimated to be 0.075 and insignificant in 1995, but it rose sharply to 0.769 in 2002 and became significant at the 5% level. The Chow test rejects the null hypothesis that β_1^{LSOE} has not changed between 1995 and 2002, with a p-value of .08. This finding is consistent with the view that workers in LSOEs had stronger precautionary savings motives than those in CSOEs because they faced significantly higher unemployment risks.

5.6. Lifecycle effects

Households' consumption and savings behaviors vary significantly over the lifecycle. In an important contribution, Gourinchas and Parker (2002) estimate a structural lifecycle model using U.S. data and provide evidence that young households save for precautionary reasons whereas old households save mainly for retirement. To examine the lifecycle patterns

Table 6
Regressions controlling for PIH effects.

Dep. variable	Income expectation		Pension participation	
W/P	(i) 1995	(ii) 2002	(iii) 1995	(iv) 2002
SOE	−0.012	0.559**	−0.016	0.621**
	(0.094)	(0.268)	(0.102)	(0.277)
RISK	0.170***	0.150***	0.172***	0.145***
	(0.052)	(0.049)	(0.053)	(0.048)
log(P)	0.846	2.961**	0.896	2.816**
	(1.010)	(1.279)	(1.002)	(1.264)
Income decline		0.002		
		(0.161)		
No-pension			−0.025	0.350**
			(0.084)	(0.154)
Log-Likelihood	−7045.37	−5505.65	−7031.84	−5515.49
p-value of Chow test for SOE		0.044		0.031
Sample size	3627	2164	3627	2170

Notes: IV-Tobit regression results using the sample with government assigned jobs. Columns (i) and (ii) show the regression results controlling for income expectations. Since the question on income expectations is not available in the 1995 CHIP survey, we add the "income decline" dummy in the 2002 regression only. Columns (iii) and (iv) show the results controlling for pension participation. All other control variables shown in Table 5 are included. Robust standard errors are in parentheses. ***, **, and * indicate *p*-values of less than 1%, 5%, and 10%, respectively.

of precautionary savings for Chinese households, we split our sample into two cohorts: a young cohort (aged 25–44) and an old cohort (aged 45–55). We estimate the benchmark model in Eq. (12) for each age cohort using the sample with government assigned jobs.

In 1995, the estimated coefficient β_1 for the SOE dummy variable is small and statistically insignificant for both age groups, as in the baseline sample. In contrast, in 2002, the estimated value of β_1 is very different for the two age cohorts. Table 7 (Panel B) shows that, in 2002, the estimated value of β_1 for the young cohort is much greater than that for the baseline sample (0.942 vs. 0.539), and both are significant at the 5% level. The estimated value of β_1 for the old cohort is much smaller (0.292) and statistically insignificant. This evidence is consistent with the finding obtained by Gourinchas and Parker (2002) that young households behave as buffer-stock agents and old households behave more like certainty equivalent consumers.

5.7. Other demographic factors

During the periods of the SOE reform, there is evidence that specific demographic groups including female, less skilled, and less educated workers are more likely to be laid off (Appleton et al., 2002). We now examine the precautionary savings behaviors of these specific demographic groups.

Table 7 (Panel C) shows the estimation results using a few different subsamples of the data (again, focusing on government assigned jobs in both years). Evidently, if a household head is female, then the household has a stronger precautionary savings motive. The value of β_1 for this group in 1995 is insignificant, but becomes significant in 2002 and is indeed larger than that in the baseline sample (0.931 vs. 0.539). We obtain qualitatively similar results when we consider the samples with female or less skilled; female or less educated; or female, less skilled, or less educated. These results, putting together, suggest that SOE workers who faced higher unemployment risks accumulated more precautionary wealth in response to the reform.

6. Robustness

In this section, we examine the sensitivity of our estimation of precautionary savings. In particular, we consider the implications of sample selections, spouse occupations, and some alternative sampling and measurement methods. Although these factors change the particular estimates of the contributions of precautionary savings to SOE household wealth accumulation, we show that the quantitative importance of precautionary savings that we have obtained in the benchmark model controlling for self-selection biases remains robust.

Table 8 shows the estimated values of β_1 for 1995 and 2002 under these alternative specifications, as well as the implied contributions of precautionary savings to overall wealth accumulations for SOE households (the last column). For comparison, we also display the baseline estimation (Panel A).[20]

[20] To conserve space, we report the detailed estimation results from the alternative models and with alternative measurements in a Supplemental Appendix available online at http://www.frbsf.org/economic-research/files/wp2014-04_appendix.pdf.

Table 7
Precautionary savings and demographic factors.

Sample group	Variable	Coefficient	
		1995	2002
Panel A: *firm size effects*			
Baseline sample	CSOE	−0.157	0.343
		(0.128)	(0.237)
	LSOE	0.075	0.769**
		(0.150)	(0.367)
		[$n = 3627$]	[$n = 2170$]
Panel B: *lifecycle effects*			
Age 25–44	SOE	−0.015	0.942**
		(0.144)	(0.380)
		[$n = 2349$]	[$n = 1123$]
Age 45–55	SOE	0.042	0.292
		(0.145)	(0.594)
		[$n = 1278$]	[$n = 1047$]
Panel C: *other demographic factors*			
Female	SOE	−0.130	0.931*
		(0.193)	(0.526)
		[$n = 1305$]	[$n = 585$]
Female or less skilled	SOE	−0.126	1.365*
		(0.160)	(0.777)
		[$n = 1572$]	[$n = 756$]
Female or less educated	SOE	−0.043	0.871*
		(0.125)	(0.459)
		[n=2063]	[n=984]
Female, or less educated, or less skilled	SOE	−0.063	1.227*
		(0.126)	(0.697)
		[$n = 2157$]	[$n = 1060$]

Notes: Results are from the IV-Tobit regressions. All household heads in the samples had government assigned jobs. Each regression includes the same set of control variables shown in Table 5. Robust standard errors are in parentheses. Sample sizes are in squared brackets. ***, **, and * indicate p-values of less than 1%, 5%, and 10%, respectively.

6.1. Sample selection biases

Since the CHIP surveys do not keep track of individual households over time, the post-treatment group observed in 2002 includes only those workers who survived the SOE reform and who chose not to quit from their SOE jobs. There is evidence that workers with lower educational attainment or lower skills were more likely to be laid off (Appleton et al., 2002). In addition, during the period from 1995 and 2002, some workers who were not laid off chose to quit from SOE firms for private-sector jobs. The difference of worker characteristics before and after the reform can potentially cause biases in the estimation of precautionary savings.

To balance the 1995 and 2002 samples, we use the standard propensity score weighting approach in the spirit of Rosenbaum and Rubin (1983). We first estimate the propensity score for each individual in the pooled 1995 and 2002 samples, using a Logit model. The Logit model specifies the probability $\hat{p}$ that an individual belongs to the 2002 (post-treatment) sample as a function of a number of individual characteristics, including age, gender, education, occupation, industry, and geographic location. If an individual is observed in the 2002 sample, then that would indicate that she was still working in either SOE or GOV, and had not been laid off or quit. The estimated probability $\hat{p}$ is the propensity score. Following the approach in Hirano and Imbens (2001) and Hirano et al. (2003), we weigh each observation in the actual samples in 1995 and 2002 by the inverse propensity scores. In particular, we assign a weight of $\frac{1}{\hat{p}}$ to the 2002 sample and a weight of $\frac{1}{1-\hat{p}}$ to the 1995 sample. Finally, we estimate the baseline regression model in Eq. (12) for each year (1995 and 2002) using the weighted sample.

As shown in Panel B of Table 8, the estimation results using the samples weighted by the propensity scores are similar to our benchmark estimation. In particular, the point estimate for β_1 remains insignificant and very close to zero in 1995 (0.0003), and the 2002 estimate of β_1 becomes significant, with a magnitude similar to that in the benchmark case (0.571 vs. 0.539). With the sample selection biases mitigated, precautionary savings account for about 44.5% of total SOE household wealth accumulation, which is close to that obtained in the benchmark (43.1%).

6.2. Spouse effects

The precautionary savings that we have estimated are based on the regression model in Eq. (12), through a dummy variable indicating whether the household head works in an SOE. However, a large fraction of households in our sample

Table 8
Robustness.

Cases	1995	2002	Contributions
A. Benchmark	−0.012	0.539**	43.1%**
	(0.094)	(0.264)	(0.204)
B. Sample selection bias	0.0003	0.571**	44.5%**
	(0.101)	(0.278)	(0.210)
C. Spouse effects	−0.006	0.464*	36.9%*
	(0.099)	(0.265)	(0.207)
D. Eliminating zero wealth	0.034	0.372*	24.7%*
	(0.086)	(0.216)	(0.141)
E. Alternative risk measure	−0.022	0.521**	42.2%**
	(0.094)	(0.260)	(0.202)
F. Very liquid asset	0.003	0.475*	40.2%*
	(0.091)	(0.251)	(0.210)
G. Non-housing Non-business wealth	0.111	0.851**	48.9%**
	(0.131)	(0.357)	(0.197)

Notes: We use the standard IV (2SLS) regression for the case that eliminates zero-wealth observations (Panel D) and the case with non-housing non-business wealth (Panel G). We use IV-Tobit regressions for the other cases. Panel A shows the benchmark estimation results. Panels B-G report the estimation results under alternative model specifications or variable measurements. Each regression uses the sample with government assigned jobs and includes the same set of control variables shown in Table 5. The last column shows the contributions of precautionary savings to total wealth accumulation by SOE workers. Robust standard errors are in parentheses. ***, **, and * indicate *p*-values of less than 1%, 5%, and 10%, respectively.

are dual-income families. In particular, about 70% of SOE households and 76% of GOV households have dual income earners. The spouse working status affects the overall income uncertainty for a family in the post-reform period. For example, it is plausible that a household whose head works at the SOE sector but whose spouse works at the government sector is not as exposed to the reform as a household in which both the head and the spouse work for an SOE.

To control for the effects of the working status of the spouse, we add a dummy variable SOE^{sp} that indicates whether or not the spouse works for an SOE in our regressions. All else equal, we should expect a family with the spouse working for an SOE to have more precautionary savings than an average household after the SOE reform. Thus, the coefficient for SOE^{sp} in the 2002 sample should be positive. This turns out to be true. In particular, for the 2002 sample, the estimated coefficient for SOE^{sp} is significantly positive at 0.237 (with a *p*-value less than 10%, not shown in the table).

With the spouse effects controlled, the coefficient β_1 for the SOE dummy captures the marginal impact for the household head to work in an SOE when the unemployment risks rose relative to GOV workers. The point estimate for β_1 is 0.464, which is smaller than the benchmark value of 0.539, but it remains statistically significant at the 10% level. As shown in Panel C of Table 8, the implied contribution of precautionary savings to the observed increase in total savings is about 36.9% for SOE households.

6.3. Excluding zero wealth observations

The empirical results summarized above are obtained based on the sample that includes zero-wealth observations. To examine whether these results are driven by zero-wealth observations, we exclude those observations from the sample and re-estimate the benchmark model in Eq. (12) using the standard IV (2SLS) approach (instead of the IV-Tobit approach used for estimating the benchmark model). With the zero-wealth observations excluded, the sample size reduces to 3190 and 1977 observations for 1995 and 2002, respectively. Table 8 summarizes the estimation results (in Panel D).

The estimated value of β_1 is 0.034 (insignificant) in 1995 and 0.372 (significant at the 10% level) in 2002. Thus, excluding zero-wealth observations from the sample modestly reduces the estimated magnitude of precautionary savings. Nonetheless, precautionary savings still account for about 24.7% (s.e. = 0.141) of total wealth accumulations for SOE workers.

6.4. Alternative risk measure

In our benchmark model, household idiosyncratic risks (*RISK*) are measured by the log variance of log income over the current and recent past years. To examine the sensitivity of our results, we consider the risk measure used by Carroll and Samwick (1998), which is the logarithm of the variance of log income for 16 different educational and occupational groups.[21] Unlike our measure *RISK*, which reflects a household's income variations across time, this alternative risk measure is com-

[21] The 16 groups correspond to the cross products of the 4 occupation categories and 4 education categories described in Section 4.3 and Table 2.

puted based on cross-sectional variations of income in the current year. The estimation results are shown in Panel E of Table 8.

Our main results are not sensitive to using the alternative risk measure. The estimated values of β_1 are similar to those obtained from the baseline model. In particular, β_1 increases from −0.022 (insignificant) in 1995 to 0.522 (significant at the 5% level) in 2002. These estimates imply that precautionary savings account for about 42.2% (s.e. = 0.202) of the increases in financial wealth for SOE workers from 1995 to 2002.

6.5. Alternative wealth measures

Some alternative measures of wealth such as very liquid assets (VLA) and non-housing non-business wealth (NHNBW) are also commonly used in the literature (Carroll and Samwick, 1998). We now examine the sensitivity of our empirical results to these alternative measures of wealth (see Table 1 for the definition of these variables).

Panel F of Table 8 presents the results using very liquid assets as wealth measure to construct the dependent variable in Eq. (12). The estimated value of β_1 increases from 0.003 (insignificant) in 1995 to 0.475 (significant at the 10% level) in 2002. These estimates imply that precautionary savings contribute about 40.2% (s.e. = 0.210) to the observed increases in wealth accumulation of SOE households following the reform.

Panel G of Table 8 shows that, when we use the non-housing non-business wealth to replace financial wealth, the estimated value of β_1 is 0.111 and insignificant in 1995 and it increases substantially to 0.851 (significant at the 5% level) in 2002. In this case, precautionary savings account for about 48.9% (s.e. = 0.197) of the increases in wealth accumulation for SOE workers from 1995 to 2002.

7. Conclusion

China's large-scale reform of the state-owned enterprises (SOEs) in the late 1990s provides a natural experiment for identifying and quantifying the importance of precautionary savings in a rapidly growing transition economy. With self-selection biases mitigated, we find significant evidence of precautionary savings stemming from sudden increases in unemployment risk for SOE workers relative to that for government employees. Our estimation suggests that precautionary savings can account for about 40 percent of the actual increase in wealth accumulation by urban SOE households in China for the period from 1995 to 2002. Furthermore, demographic groups more vulnerable to unemployment risks following the reform also accumulated more precautionary wealth. These findings suggest that precautionary savings associated with large structural changes in the Chinese economy is quantitatively important.

Appendix A. A case study: massive lay-off in Fushun, Liaoning

Smyth et al. (2001) present a case study of massive lay-off happened in Fushun, Liaoning. Fushun is a medium sized city located 45 km 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 that had the largest number of layoffs were coal, textiles, light industry, electronics, machinery, and chemicals. For example, of the 71,000 workers in SOEs in the coal sector in Fushun, 35,000 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 ties with their former SOEs employers. In practice, there are four different types of layoffs from an SOE firm: (1) *fang jia*: a worker is put on a temporary leave; (2) *xia gang*: a worker is put on a long-term leave; (3) *tui yang*: a worker takes voluntary early retirement. (4) *mai duan*: a firm pays a lump-sum amount (usually not exceeding three years of salary) to buy out or terminate the labor contract with a worker. In our sample, we include all four types of layoffs.

Allowances were paid to laid-off 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 that hindered the effectiveness of government-sponsored re-employment institutions. A large proportion of laid-off workers were middle-aged, female, less educated, or low skilled. 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 because they were concerned that seeking employment in non-state sectors would cut their ties with their former SOE employers. Among the

laid-off workers who have registered at re-employment centers in Fushun, 50% were middle-aged. Among these middle-aged workers, only half of them were successfully re-employed.[22]

Supplementary material

Supplementary material associated with this article can be found, in the online version, at 10.1016/j.jmoneco.2017.12.002.

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[22] 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 workers found new jobs. See Lee (2000) for details.