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Welfare Consequences of Rising Wage Risk in the United States: Self-Selection into Risky Jobs and Family Labor Supply Adjustments*

Seonyoung Park† Donggyun Shin‡

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Abstract

Wages in the United States have become more volatile since the early 1970s. This paper quantitatively demonstrates that the welfare cost caused by this change is substantially overstated when heterogeneity in individual risk preferences and workers' risk choices are neglected. Family labor supply adjustments reduce the welfare cost and do so most effectively when borrowing and saving behavior is allowed. It is also found that wives increase their labor supply significantly in response to increases in the variance of husbands' permanent wage shocks, and this 'added-worker' effect is mostly accounted for by wives' labor supply adjustments on the extensive margin.

JEL Codes: D52, E21, E25, J22
Keywords: heterogeneity, insurance, wage risk, welfare costs, labor supply

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†Corresponding author. Department of Economics, Alfred Lerner College of Business & Economics, University of Delaware, 413 Purnell Hall, Newark, DE 19716, U.S.A. Email: ypark@udel.edu. Phone: 1-302-831-1909. Fax: 1-302-831-6968.

‡School of Economics and Finance, Victoria University of Wellington, Wellington, New Zealand. Email: don.shin@vuw.ac.nz.

1 Introduction

The last half-century has seen a large increase in men's earnings volatility in the United States (US). This change can be observed both in annual earnings[1] and hourly wages (e.g., Heathcote et al., 2010), both of which first increased in volatility markedly during the 1970s. These changes were accompanied by stagnation in men's hourly wages (e.g., Elsby et al., 2016). The welfare implications of these changes depend crucially on households' preferences and abilities to self-insure against greater wage risk. This paper examines the welfare cost of this increased wage volatility and finds that it is significantly overstated by existing homogeneous agent-job models of the economy.

More precisely, this paper evaluates the welfare costs of the increased wage volatility in the US from the early 1970s to the early 2000s. As in Blundell et al. (2016a), we take wage shocks as the primitive source of uncertainty faced by households, and view changes in labor hours as the household's optimal responses to the increased shocks. Crucially, we include the 1970s in the analysis period. As noted by Shin and Solon (2011), the dramatic increase in men's earnings volatility during the 1970s preceded the well-documented increase in long-run earnings inequality observed during the 1980s and 1990s. The time pattern of these two events suggests that they might have been driven by different causes, and consequently might command different welfare implications. As reliable consumption data are not available for the 1970s, the conventional empirical approach for investigating the relationship between income changes and consumption changes is unfeasible. We therefore estimate a model that describes the early 1970s economy and ask how much welfare costs

[1] See Moffitt and Zhang (Table 3, 2018) for a complete review of studies up through Spring 2018. For example, using the Panel Study of Income Dynamics (the workhorse data set in the literature), Shin and Solon (2011) find that, apart from cyclical fluctuations, men's earnings volatility increased significantly during the 1970s, stabilized afterwards until a new upward trend appeared after 1998. While existing studies produce a consistent result regarding trend movements of men's earnings volatility until around 2000, as well discussed by Carr et al. (2020) among others, there is a lack of consensus between PSID-based studies and those using administrative data regarding recent trends in men's earnings volatility especially during the Great Recession. Our welfare analysis, therefore, excludes the Great Recession and its aftermath from the analysis period. See Section 2 for a brief discussion of why the PSID is desirable for our research purposes.

(in lifetime consumption equivalent) a typical household of the 1970s economy would suffer in facing the higher variance of wage shocks of the 2000s.

A more appropriate welfare evaluation, however, requires simultaneous consideration of a number of issues raised by existing studies. As emphasized by Blundell et al. (2008) and Cunha et al. (2005), among others, a rise in earnings volatility forms only a necessary condition for welfare loss. Identifying economic risk associated with earnings changes will require further information on whether the changes were the results of (heterogeneous) agents' purposive choice and whether the affected individuals were insured against the changes. While several recent papers have investigated the effectiveness of various insurance measures by analyzing consumption data in conjunction with the income data,[2] little effort has been made to consider heterogeneity in individual risk preferences and individual choice of job-related wage risk in the welfare analysis of rising volatility. As we will demonstrate in this paper, the welfare cost of the rising wage volatility is significantly overstated by neglecting heterogeneous workers' self-selection into jobs with high wage risk.

Another shortcoming of the literature evaluating the welfare effects of the changing wage structure is the lack of study on the role played by gender differences in observed volatility trends. As documented by several studies (e.g., Congressional Budget Office, 2007; Dynan et al., 2012, and Ziliak et al., 2011), male earnings have become more volatile, while female earnings have become less volatile since the early 1970s, suggesting that exogenous changes in the variance of wage shocks in the economy might also be different between genders. Existing studies, however, often find rising trends in men's wage volatility and assume equality between genders in wage dynamics when evaluating welfare consequences of rising wage in-

[2]For example, Krueger and Perri (2006) document that rising income inequality has been accompanied by an increase in consumption inequality of a smaller degree because individuals mitigate their income fluctuation through an endogenous credit constraint. Heathcote et al. (2014) find that households are well-insured against sharp increases in wage inequality through private- and government-provided insurance. Gorbachev (2011) shows a smaller increase in consumption volatility compared to the increase in family income volatility. Using relatively old data (from 1980 to 1992), Blundell et al. (2008) show evidence of partial insurance on permanent income shocks and almost full insurance on transitory shocks through various insurance mechanisms, such as, taxes, transfers, family labor supply, and durable goods. Attanasio et al. (2005) and Blundell et al. (2016a) emphasize the importance of family labor supply as an insurance mechanism and find strong evidence of smoothing of permanent wage shocks.

equality and volatility (e.g., Heathcote et al., 2010; Hong et al., forthcoming). We believe that, due to limited insurability, welfare implications are different depending on how changes in the variance of wage shocks are distributed between genders.

This paper contributes to the literature by investigating the aforementioned issues jointly. First and most importantly, we allow, in a standard general equilibrium model with incomplete markets, heterogeneity in individual risk preferences, 'job' (or sector) heterogeneity in wage risks, and workers' self-selection into risky jobs. We focus on the bias in the measured welfare cost generated by the conventional approach. This is done by comparing the estimation results of our heterogeneous agent-job model to those of the otherwise-comparable homogeneous agent-job model. For the purpose of introducing heterogeneity in individual risk preferences into our augmented model, we adopt an empirical distribution of individual risk aversions that several existing studies[3] derive consistently using individuals' responses to the hypothetical gambles over lifetime income commonly fielded by various surveys, such as the Panel Study of Income Dynamics (PSID), National Longitudinal Study of Youth 1979 (NLSY79), and Health and Retirement Study (HRS). Because the survey questions are structured in a way that individuals reveal their risk preferences in choosing between a job with a certain lifetime income (safe job) and a job with random, but higher mean lifetime income (risky job), the derived distribution of individual risk aversion matches the unique features of our model economy where agents with heterogeneous risk preferences make optimal choices of job-related wage risk. To the best of our knowledge, this paper is the first to analyze the effects of heterogeneous workers' self-selection into risky jobs in the literature on the welfare evaluation of rising inequality/volatility.[4]

[3]See, among others, Barsky et al. (1997), Kimball et al. (2008, 2009), Sahm (2012), and Light and Ahn (2010).

[4]Preference heterogeneity has been an important ingredient of many macroeconomic models. For example, Krusell and Smith (1998) find that allowing heterogeneity in the discount factor across generations has substantial effects on wealth inequality. Cagetti (2003) uses a life-cycle model of wealth accumulation, estimates different discount factors and risk aversion parameters for three different education groups, and explains the importance of precautionary savings on wealth accumulation. In testing the theory of precautionary savings, Fuchs-Schündeln and Schündeln (2005) consider the importance of self-selection into occupations due to differences in risk preferences. Using the German reunification "experiment", they find that failing to control for risk aversion in the presence of self-selection could lead to a systematic underestimation of the

Second, as previously noted, an implicit assumption in the literature is that both males and females experience the same trend in wage volatility. This partly reflects the complication associated with identifying women's wage shocks in the presence of their endogenous selection into the labor force. The increasingly-prominent role played by women in the labor force further raises the importance of our ability to cope with this selection issue and account for gender differences in the trends and levels of wage shock variances when analyzing welfare implications. Of course, individuals make decisions on not only their desired level of wage risk associated with a 'job' but also labor market participation. This implies that the selection issue exists not only in the heterogeneous agent-job model but also in the conventional homogeneous agent-job model. We therefore estimate both the heterogeneous and homogeneous agent-job models to identify structural parameters of variances of job-specific and employment-specific wage shocks, respectively, individuals would experience *were* they assigned to jobs and employment *randomly*. In the estimation process, observed variances of wage shocks are used as empirical moments to be matched with corresponding model-generated moments.

Third, we conduct a quantitative assessment of the effectiveness of various insurance measures in mitigating the welfare cost caused by the increased variance of wage shocks. While this is not new to the literature, our analyses of interactive effects among self-insurance mechanisms and cross-responsiveness of precautionary labor supply with respect to an increase in spouses' wage uncertainty are somewhat new and deserve further attention. Lastly, we conduct various tests that demonstrate the robustness of the results.

Our analysis of Panel Study of Income Dynamics (PSID) data shows that the observed variance of permanent wage shocks increased similarly for both genders from the early 1970s to the 2000s. While the variance of male transitory shocks increased significantly, that of female ones declined to some degree. These are changes in the observed variance of wage shocks that mix changes in the true variance of wage shocks and changes in the composition

importance of precautionary saving.

of workforce. To estimate the structural parameters of gender-and-job-specific true variances of wage shocks in the heterogeneous agent-job model, we use two sets of empirical moments, among others, that are related to heterogeneous agents' risk and job choice. One set of moments are observed variances of wage shocks by risk-preference group for each gender, and the other are wage-related empirical moments, such as the wage gap between risky and safe jobs and within-job gender wage gaps. As the latter set of moments require definitions of the risky and safe jobs, we use the PSID to define "risky" and "safe" jobs empirically (see Section 4 for details of our estimation strategy). After estimating the structural parameters of gender-and-job-specific true variances of wage shocks that existed in the early 1970s and the 2000s, we provide a quantitative assessment of the welfare costs caused by these changes in true variances of wage shocks, using a two-earner model in which husband and wife choose their life-cycle labor supply, job types, consumption, and savings. A relevant question is how much welfare cost the 1970s households would have paid if they had received the gender-job-specific variance of wage shocks the 2000s households experienced. Heterogeneous risk preferences, job heterogeneity in wage risk, workers' self-selection into risky jobs along with gender differences in wage dynamics constitute unique features of the model. We also estimate an alternative model that assumes homogeneous risk preferences and only one type of job and compare the results to those from our augmented model.

Our most important result shows that the welfare cost of the changes in the variance of wage shocks is significantly exaggerated by neglecting heterogeneity in risk preferences and workers' risk choices. The measured welfare cost is approximately twice as great in the homogeneous agent-job model as in our heterogeneous model. Both heterogeneity in risk preferences and self-selection into job types are important in explaining the welfare cost gap between the models, although the former is somewhat greater than the latter. It is easy to understand the latter, 'job selection' effect: Allowing more options (risky vs. safe jobs) generally increases agents' utility compared to the case of having one type of job. The finding of the large 'preference effect,' however, may surprise some who consider that the distribution of risk aversion is right-skewed. The greater welfare loss of more risk-averse agents relative to an agent with the average risk aversion (an agent in the homogeneous model) may dominate the smaller welfare loss of less risk-averse agents, resulting in a greater welfare cost in the heterogeneous, relative to homogeneous, model. While this conjecture may be valid in a fictitious world where agents have no ability to cope with the increased variances of wage shocks, once the model is enriched to allow various self-insurance mechanisms (e.g., labor supply and borrowing/saving), it no longer holds. The welfare gain of allowing agents these insurance mechanisms is greater for the more risk-averse agents compared to the 'average' risk-averse agent; it is smaller for the less risk-averse agents; and the former dominates the latter. As emphasized previously, a more appropriate and realistic welfare evaluation should consider agents' insurability against the increased variance of wage shocks.

Our additional analyses, including examination of the effectiveness of insurance mechanisms, are qualitatively robust with respect to the underlying assumption about preference and job heterogeneity. To summarize briefly, while family labor supply adjustments, relative to borrowing and saving, are more effective in reducing the welfare costs caused by the increased variance of permanent shocks, the welfare costs of the increased variance of transitory shocks are more effectively mitigated by borrowing/saving. More interestingly, family labor supply adjustments can reduce the welfare costs of the increased variance of permanent shocks more effectively when borrowing and saving behavior is allowed. Our results further show that, when households are hit by the increased variance of male (female) permanent shocks, added-worker effects by wives (husbands) play a greater role in mitigating the welfare loss, relative to the effects of husbands' (wives') self-labor supply adjustment. We also find that about 60 percent of the wives' 'added-worker' effect in response to the increased variance of husbands' permanent shocks is accounted for by the extensive margin of wives' labor supply adjustments. Lastly, measured welfare costs remain similar whether or not prices are allowed to vary following the increased variance of wage shocks.

The rest of the paper is organized as follows. Section 2 describes some basic facts that motivate our research. Section 3 presents our models, and Section 4 discusses determination of model parameters. Using the estimated models from Section 4, Section 5 analyzes welfare consequences of the increased variances of wage shocks and conducts various robustness tests. Section 6 concludes.

2 The Facts

We begin our analysis by establishing the facts that motivate our exercise. As a key feature of our analysis is the different roles played by volatility in men's and women's wages, we focus on gender differences in the variances of permanent and transitory wage shocks in this exposition. Our empirical evidence is based on a sample of married, prime-age respondents in the PSID for the period 1970 to 2014. We choose the PSID for our analysis for two primary reasons. First, the PSID permits us to include the 1970s, a period omitted from many administrative-data studies, in our analysis. The inclusion of this period is essential because the large increase in men's earnings volatility during the 1970s preceded the well-documented increase in earnings inequality observed during the 1980s and 1990s. As such, these two events might have been driven by different causes and might command different welfare implications. Second, administrative data do not contain information on labor hours, which precludes studying the volatility of hourly wages. As our analysis will show, the PSID exhibits different trends in the volatilities of hours and hourly wages for both genders. On the other hand, a drawback of using the PSID is the potential for measurement error in earnings and hours variables. The literature on earnings inequality/volatility often emphasizes the importance of these errors in surveys like the PSID and the Current Population Survey (CPS). However, while these measurement errors may distort measured earnings/wage volatility, there is no reason to believe that they produce biased results in the trends in measured volatility, which is the focus of the current study. Furthermore, as our emphasis is on the difference in measured welfare costs between the homogeneous and heterogeneous agent models, our results are likely to be even less sensitive to the issue of measurement error. (See Appendix A for a brief introduction of various samples used in this paper.)

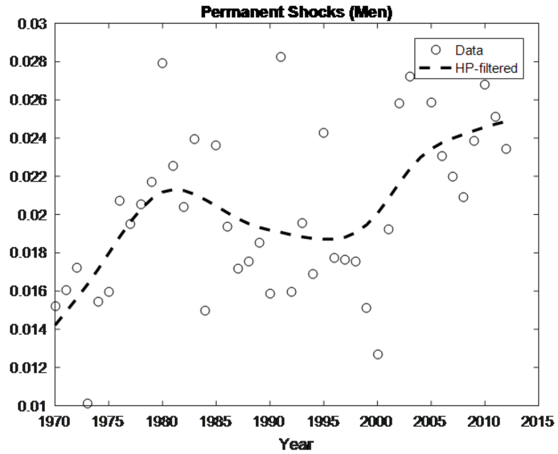
We first test whether the variances of wage shocks, both permanent and transitory, have increased during the last fifty years, and if so, by how much. The distinction of permanent from transitory is useful because it informs the welfare evaluation of wage changes. Not only do permanent and transitory wage shocks have different welfare consequences, but the effectiveness of potential insurance measures is different depending on the nature of the wage shocks. In response to the recent research (aforementioned) addressing the gender differentials in volatility, we model stochastic wage processes separately by gender. On the basis of the standard permanent-transitory decomposition method commonly adopted in the literature,[5] Figure 1 shows how the variances of male and female permanent and transitory wage shocks have changed over time using directly-observed moments that do not correct for participation.

Comparing both wage shocks across genders reveals that the variances of both have been greater for females than males, reflecting women's weaker labor market attachment. More importantly, while the variances of permanent shocks have similarly increased for genders, the variances of transitory shocks exhibit differing trends. While the variance of male transitory shocks has increased dramatically, the variance of female transitory shocks has generally decreased. As a result, the gender gap in the variance of transitory wage shocks had almost closed by the early 2010s. To quantify the magnitude of these changes, we compare the five-year average of each variance over the 1970–1974 period with that over the 2002–2006 period. These reference periods are selected to avoid the differential impacts of the recession of the mid-1970s and the Great Recession on wage volatility. The variances of male and female permanent shocks increased from the early 1970s to the mid-2000s by similar amounts, 0.0072 (from 0.0156 to 0.0228) and 0.0073 (from 0.0172 to 0.0245), respectively.

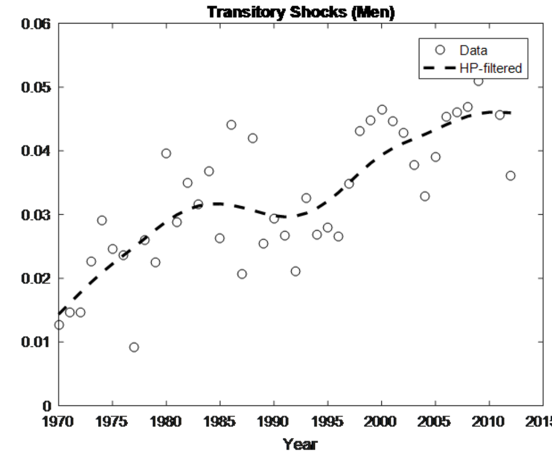
[5] The generalized method of moments (GMM) estimation procedure is also described in Sections 4 and Appendix B.

Figure 1: Observed Variances of Wage Shocks

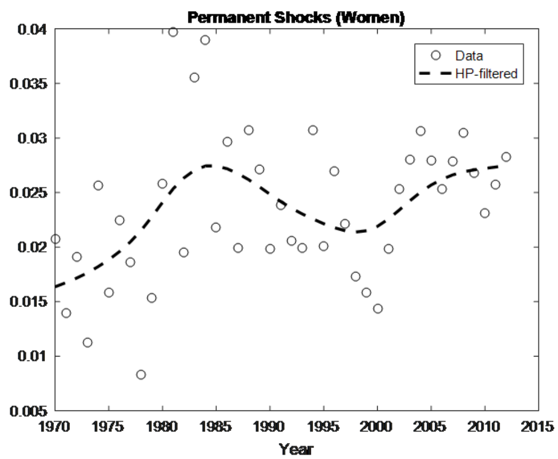
(A) Male Permanent Wage Shocks



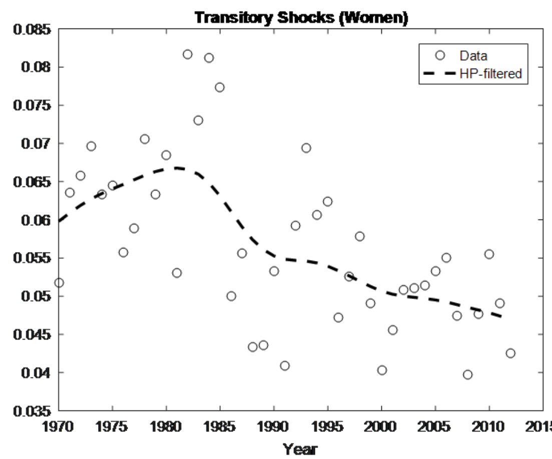
(B) Male Transitory Wage Shocks



(C) Female Permanent Wage Shocks



(D) Female Transitory Wage Shocks



Source: The Panel Study of Income Dynamics, 1970–2014.
Notes: See the text for estimation methodology, and Appendix A for sample restrictions and variable definitions.

But while the variance of male transitory shocks increased by 0.0249 (from 0.0177 to 0.0426), the variance of female transitory shocks *decreased* by 0.0120 (from 0.0617 to 0.0497).

At first, our finding of the increased variance of female permanent wage shocks seems at odds with existing studies (e.g., Dynan et al., 2012; Ziliak et al., 2011), which report that female earnings have become less volatile. However, these studies measure different outcomes and employ different methodologies in that they examine earnings instead of hourly wages and use the standard deviation of residualized earnings changes as the measure of earnings volatility. As explained by Shin and Solon (2011), an earnings volatility measure based on the dispersion of year-to-year earnings changes reflects permanent shocks in addition to transitory ones. In order to explore the difference in these methods, we replicate directly[6] the finding that female earnings have become less volatile from the early 1970s to the 2000s using our PSID sample. Then we find that the reduction in hours volatility slightly 'over-explains' the reduction in earnings volatility for the same sample period. In short, the downward trend in female earnings volatility observed in several existing studies is primarily due to the decreasing trend in hours volatility, while female wage volatility increased slightly over the same time.[7] Despite the substantial increase in the variance of female permanent wage shocks during the sample period, overall female wage volatility has increased only slightly due to the decrease in the variance of transitory wage shocks. However, considering that permanent wage shocks are even more consequential than transitory ones, the associated welfare cost would be greater than what is suggested by the mild increase in the overall volatility of female wages.

A limitation of this initial analysis is that these changes in observed variances of wage shocks confound the effects of self-selection into employment with the effect of exogenous changes in wage shocks. Confounding these two effects is likely to be particularly important for female workers. For instance, if women are more likely to exit the labor market upon facing a negative transitory shock, this response affects observed wage changes and must be accounted for in the estimation of the wage process. Subsequent sections are devoted to model-based identification of the true variances of the four types of wage shocks and a quantitative assessment of the welfare costs caused by exogenous changes in their variances. The observed moments in Figure 1 will be used as empirical moments to be matched by

[6]More precisely, we follow Shin and Solon (2011) by computing the standard deviation of residuals obtained from the regression of a two-year change in log earnings on a quadratic in age each year.

[7]We find that the estimated covariance of hours and wage changes does not show a clear trend for both genders. For men, hours volatility increased only slightly from the early 1970s to the early 2000s, and a majority of the increase in earnings volatility is attributed to the increased volatility of the wage rate. All these results are available electronically from the authors upon request.

corresponding model-generated moments.

3 Model

We study the welfare effects of the *changes* in variances of the four types of wage shocks (henceforth the increased variances of wage shocks) in a general equilibrium model. The model features incomplete markets in which households choose their life-cycle labor supply, types of jobs (risky or safe), consumption, and savings. A representative firm employs different types of labor across genders, and the government runs a pension system. We estimate and compare two versions of the model: a baseline model which extends the conventional model by allowing individual heterogeneity in risk preferences, job heterogeneity in wage risk, and workers' self-selection into risky/safe jobs; and an alternative (or conventional) model which assumes homogenous risk preference and allows only one type of job. The economies that are described by the baseline and alternative models are called the heterogeneous and homogeneous economies, respectively.

3.1 Baseline Model

3.1.1 Economic environment

The economy is populated by a continuum of households in a single cohort, where each household consists of a married couple. Let $j \in J = \{j_1, j_2, ..., j_T\}$ denote an age of the life-cycle. Adults in the household start economic activity at age j_1 and retire at j_R, after which they receive a pension of b until j_T. In each period, the household makes decisions on the couple's labor supply, types of 'jobs' (or sectors),[8] and the household's consumption and savings. Job types are identified by a combination of the price per efficiency unit of labor and the associated wage risk: risky jobs are subject to a greater variance of wage shocks and pay

[8] In the current study, each 'job', risky or safe, consists of multiple jobs of the same type, and consequently has the meaning of 'sector'. Individuals are allowed to change jobs within each risky or safe 'job' (or sector) or between 'jobs' (or sectors). Jobs and sectors are used interchangeably.

a higher labor price compared to safe jobs. Job heterogeneity is introduced into the model to consider heterogeneous agents' risk choice in the welfare analysis of the rising wage risk. Let s^d denote an endogenously determined share of households choosing d-combination of jobs for a husband and a wife: s^1 represents the share of households that choose risky jobs for both the husband and wife; s^2 risky for husband and safe for wife; s^3 risky for husband and non-work for wife; and so on all the way up to s^7 non-work for husband and risky for wife; s^8 non-work for husband and safe for wife; and s^9 non-work for both. Obviously, $s^d \in [0, 1]$ and $\sum_{d=1}^{9} s^d = 1$.

The household faces uncertainty not only over wages but also life expectancy. They face a probability ξ_i of surviving from age $j-1$ to j. The household pays flat taxes (τ_k, τ_l) on labor and capital income, and the government uses taxes to finance a public pension system for retirees. In the spirit of Heathcote et al. (2010), once the pension system has been financed, any excess tax revenues are used for government spending.

3.1.2 Preferences

Each household maximizes expected lifetime utility over sequences of consumption and the couple's labor supply:

$$E_1 \sum_{j=1}^{j_T} \beta^{j-1} \phi_j u(c_{i,j}, l_{i,j}^m, l_{i,j}^f; \gamma_i), \tag{1}$$

where β is a discount factor with $\beta \in (0.1)$, ϕ_j is the unconditional probability of surviving up to age j with $\xi_j = \frac{\phi_j}{\phi_{j-1}}$, $c_{i,j}$ is consumption of household i in age j, $l_{i,j}^g$ is labor supply of gender $g \in \{m, f\}$ with m representing a male (or husband) and f a female (or wife), and γ_i is the risk aversion parameter, which varies across households. For various reasons to be stated in Section 4, the risk aversion parameter is assumed to be fixed over ages and common between husband and wife. The period utility function is represented by

$$u(c_{i,j}, l_{i,j}^m, l_{i,j}^f; \gamma_i) = \frac{c_{i,j}^{1-\gamma_i}}{1-\gamma_i} - \chi^m \frac{(l_{i,j}^m)^{1+\sigma_m}}{1+\sigma_m} - \chi^f \frac{(l_{i,j}^f)^{1+\sigma_f}}{1+\sigma_f}, \tag{2}$$

where χ^g is a weight on the disutility of labor supply of gender g, and σ_g concerns the Frisch elasticity of labor supply of gender g with the elasticity being determined by $1/\sigma_g$. We borrow the functional form from Heathcote et al. (2010), but extend it by allowing heterogeneous risk preferences. This is an important addition to the literature. As we will demonstrate in Section 5, the welfare cost caused by rising wage risk is significantly exaggerated by assuming homogeneous risk preference.[9]

3.1.3 Technology

Output Y is produced by a representative firm hiring aggregate capital K and aggregate labor H from competitive markets. The production technology takes a Cobb-Douglas form:

$$Y = ZF(K, H) = ZK^{\alpha}H^{1-\alpha}, \tag{3}$$

where α is the share of capital in the output, and Z is total factor productivity. Capital depreciates at the rate of δ.

The aggregate labor input H is assumed to be given by

$$H = [\lambda^{m,R}H^{m,R} + (1-\lambda^{m,R})H^{f,R}]^{\lambda_R}[\lambda^{m,S}H^{m,S} + (1-\lambda^{m,S})H^{f,S}]^{1-\lambda^R}, \tag{4}$$

where four types of input, $H^{g,n}$, are indexed by gender $g \in \{m, f\}$ and riskiness of job $n \in \{R, S\}$, with m, f, R, and S, respectively, representing male, female, risky job, and safe job. λ^R represents the share of risky job (or input) among the total, and $\lambda^{m,n}$ stands for the weight placed on men within each job, whose changes capture gender-biased demand shifts. Our specification therefore implies that, within each job (risky or safe), male and

[9] A recent study by Blundell et al. (2016a, 2018) considers nonseparability of consumption and labor supply in the utility function when analyzing consumption insurance against wage shocks. Because we focus on comparing the welfare costs of the rising wage risk between the homogeneous agent-job and heterogeneous agent-job models, our results are less sensitive to different specifications of the utility function. In addition, Blundell et al. (2016a) write, "Hence, while allowing for nonseparability is important if one wants to provide a correct specification of preferences, its role in explaining consumption insurance against wage shocks is modest."

female efficiency units of labor are perfect substitutes (as in Heathcote et al. (2010)). Our specification also allows different shares of gender-specific labor inputs between the risky and safe jobs.

3.1.4 Wages and job choice

Individuals are endowed with efficiency units of labor per market hour, which depend on age (or experience) and idiosyncratic wage shocks, both current and historical. As in Blundell et al. (2016b), we allow gender differences in the wage process and consider the covariation of spouses' wages.

$$w_{i,j,t}^{g,n} = \underbrace{p_t^{g,n}}_{\text{price per efficiency unit}} \times exp(\underbrace{\overbrace{f^g(j_{i,t})}^{\text{deterministic}} + \overbrace{z_{i,j,t}^{g,n}}^{\text{stochastic}}}_{\text{efficiency units}}). \tag{5}$$

Here, $w_{i,j,t}^{g,n}$ is the hourly wage rate for an adult in household i of gender g, age j at time t, which also depends on the type of job n. $p_t^{g,n}$ is the gender-job-specific price per efficiency unit of labor, $f^g(j_{i,t})$ is the deterministic component of the wage rate which is a function of age, and $z_{i,j,t}^{g,n}$ is the stochastic component. We assume the stochastic component consists of two orthogonal components: permanent and transitory. Following the literature, the permanent component is assumed to follow a first-order autoregressive (AR(1)) process, and the transitory one is assumed to be independently distributed. Unlike existing studies, however, we allow individuals to face different stochastic wage processes depending on the type of job they take. Thus, it follows that

$$z_{i,j,t}^{g,n} = v_{i,j,t}^{g,n} + \varepsilon_{i,j,t}^{g,n}, \tag{6}$$

where $v^{g,n}_{i,j,t}$ and $\varepsilon^{g,n}_{i,j,t}$ are permanent and transitory components, respectively, and

$$v^{g,n}_{i,j,t} = \rho^{g,n} v^{g,n}_{i,j-1,t-1} + \eta^{g,n}_{i,j,t},\ \ with\ (\eta^{m,n}_{i,j,t}, \eta^{f,n}_{i,j,t})' \sim N\Big(0, \Sigma_{perm}\Big)\ and\ (\varepsilon^{m,n}_{i,j,t}, \varepsilon^{f,n}_{i,j,t})' \sim N\Big(0, \Sigma_{trans}\Big). \tag{7}$$

Allowing job heterogeneity in wage risk is also an important addition to the literature. The measured welfare cost of the increased variance of wage shocks would be different depending on whether the affected individuals have the ability to choose their optimal risk levels associated with jobs. As we will demonstrate in Section 5, the welfare cost of rising wage volatility is significantly exaggerated by neglecting job heterogeneity in wage risk and preferences. Consequently, one of our main tasks is to identify, for each gender, how the variances of wage shocks are different between the risky and safe jobs and how they have changed from the early 1970s to the early 2000s.[10] We identify these differences in Section 4.

Finally, the husband's and wife's wage shocks are assumed to be correlated in the following way:

$$\Sigma_{perm} = \begin{pmatrix} \left(\sigma^{m,n}_{\eta_{j,t}}\right)^2 & \rho^P_{\eta^m,\eta^f}\sigma^{m,n}_{\eta_{j,t}}\sigma^{f,n}_{\eta_{j,t}} \\ \rho^P_{\eta^m,\eta^f}\sigma^{m,n}_{\eta_{j,t}}\sigma^{f,n}_{\eta_{j,t}} & \left(\sigma^{f,n}_{\eta_{j,t}}\right)^2 \end{pmatrix}, \tag{8}$$

and

$$\Sigma_{trans} = \begin{pmatrix} \left(\sigma^{m,n}_{\varepsilon_{j,t}}\right)^2 & \rho^T_{\varepsilon^m,\varepsilon^f}\sigma^{m,n}_{\varepsilon_{j,t}}\sigma^{f,n}_{\varepsilon_{j,t}} \\ \rho^T_{\varepsilon^m,\varepsilon^f}\sigma^{m,n}_{\varepsilon_{j,t}}\sigma^{f,n}_{\varepsilon_{j,t}} & \left(\sigma^{f,n}_{\varepsilon_{j,t}}\right)^2 \end{pmatrix},$$

where $\rho^P_{\eta^m,\eta^f}$ and $\rho^T_{\varepsilon^m,\varepsilon^f}$ represent the correlations of permanent and transitory shocks between husband and wife, respectively. To reduce the parameter space, for each type of wage shock, the correlation is assumed to be the same between risky and safe jobs and fixed over time

[10] Our model allows gender differences in the variances of permanent and transitory wage shocks, conditional on the same risky and safe job choice. While it enlarges the parameter space, it makes the model meaningfully more flexible. When the variance of wage shocks is assumed to be equal between genders within each job (sector), observed gender differences in wage dynamics are purely based on gender differences in composition changes, which is more or less restrictive. Instead, our model allows for behavioral differences in wage dynamics between genders within each sector.

(equivalently, age).

3.1.5 Decision problem of a household

In each period, a household makes decisions on consumption, saving, types of jobs (or sector) including non-work, and labor hours. A set of state variables for the household is denoted by $\Omega = \{a, j, \nu^m, \nu^f, \varepsilon^m, \varepsilon^f\}$, where $a \in \mathcal{A} \equiv [\underline{a}, \infty)$ is current asset holdings with $\underline{a}$ being the borrowing limit. Given the degree of risk preferences, $\gamma \in \Gamma \equiv [\underline{\gamma}, \bar{\gamma}]$, optimal decision rules are a set of functions for consumption, $c(\Omega)$, the couple's labor supply, $l^g(\Omega)$, and asset holdings, $a(\Omega)$, which solve the household problem: each household faces nine mutually exclusive alternatives, depending on the couple's choice of job types and labor market participation. The value function for the household's problem, V, is given by

$$V(\Omega; \gamma, d) = \begin{cases} \max_d \{V_d(\Omega; \gamma)\}_{d=1}^{9} & \text{if } j < j_R,\ d = \{1, 2, ..., 9\} \\ V_9(\Omega; \gamma) & \text{if } j \geq j_R \end{cases} \tag{9}$$

where d=1 if the husband and the wife choose risky jobs; d=2 if the husband chooses a risky job and the wife chooses a safe job; d=3 if the husband chooses a risky job and the wife chooses non-work; all the way up to d=7 if the husband chooses non-work and the wife chooses a risky job; d=8 if the husband chooses non-work and the wife chooses a safe job; and d=9 if the husband and the wife choose non-work. It should be clear that, conditional on risk preference, households reselect their jobs (sectors) at each stage of their life-cycle depending on current shocks and the entire history of shocks as reflected in assets holding. For example, if an agent was able to get very far away from the liquidity constraint, she/he is likely to make a different choice regarding the job type, among others, compared to an agent with the same risk preference close to the constraint. This explains not only why an individual can change job types over the course of the life-cycle even if risk preferences are fixed over ages but also why couples can choose different types of jobs even with a common risk preference. Note that our state variables are reduced to two (a, j) for the retired household's problem.

The value function of each case is defined by

$$V_d(\Omega;\gamma) = \begin{cases} \max_{a',c,l^m,l^f}\{u(c,l^m,l^f;\gamma) + \beta\xi' EV[(\Omega';\gamma)|\Omega,\gamma,d]\} & \text{if } j < j_R \\ \max_{a',c}\{u(c;\gamma) + \beta\xi' EV[(\Omega';\gamma,d=9)|\Omega,\gamma,d=9]\} & \text{if } j_R \le j < j_T \\ \max_c u(c;\gamma) & \text{if } j = j_T \end{cases}, \tag{10}$$

subject to

$$c + \xi' a' = [1 + (1-\tau_k)r)]a + I_{j<j_R}(1-\tau_l)\sum_{g\in\{m,f\}} p^{g,n}\epsilon^{g,n}(j,\nu^{g,n},\varepsilon^{g,n};\gamma)l^g + I_{j\ge j_R}b(1-\tau_l) \tag{11}$$

$$0 \le l^g \le 1,\ c \ge 0,\ a' \ge \underline{a},$$

where β represents the discount factor; τ_k is the tax rate on capital income; τ_l is the tax rate on labor income; $p^{g,n}$ is the price of an efficiency unit of labor of gender g working in job n; $\epsilon^{g,n}$ is the gender-job-specific efficiency units of labor, which consists of both deterministic and stochastic components of wages; b is a lump-sum transfer taxed at τ_l; and I is an indicator function.

3.1.6 Recursive stationary equilibrium

A recursive stationary equilibrium is defined as a collection of value functions and decision rules, endogenous shares of households choosing different types of jobs and labor market participation, prices, aggregate capital, aggregate gender- and job-specific labor inputs, government spending, and stationary distribution that satisfy the following conditions: The decision rules and value functions solve the household problem in (9) - (11); the prices are determined competitively; capital, labor, and good markets clear; the government budget constraint is satisfied; and the measure of households is consistent with household decision rules. (A detailed definition of the equilibrium is provided in Appendix C.)

3.2 Alternative Model

The alternative model assumes homogenous risk preferences and only one type of job in terms of wage risk. While this trivializes the workers' selection between risky and safe jobs, the wage processes for men and women must still be estimated within the model, as workers self-select into employment. All other economic environments and model assumptions remain identical between the two models.

4 Determination of Model Parameters

For the purpose of assessing welfare consequences of the increased variances of wage shocks from the early 1970s to the early 2000s, our strategy is to estimate a model that describes the early 1970s economy and ask how much welfare costs (in lifetime consumption equivalent) the average household optimizing under the economic environment of the 1970s would suffer from when the same household faced the higher variances of wage shocks that existed in the early 2000s. In the evaluation process, *all the model parameters are fixed at the early-1970s level*. Otherwise, all economic agents re-optimize their behavior in response to changes in the variances of wage shocks, and markets clear both before and after the arrival of new wage shocks. We estimate the model that matches the empirical moments for the early 2000s only for the purpose of obtaining the true variances of wage shocks that existed in that economy.

4.1 Externally determined model parameters

With the exception of the distribution of risk preferences, the externally-determined model parameters are set commonly in both the baseline and alternative models. These parameters are summarized in Table 1.

Demography: We assume that households start their economic activities at age j_1, which represents an actual age of 25. They work for forty years, retire at age j_{41} (actual age of

65), and die at age j_{61} (actual age of 85). Survival probabilities for the 1970s and 2000s are obtained from the Vital Statistics of the United States for 1972 and 2004.

Technology: For the 1970s and the 2000s, respectively, the capital's share of income, α, is set to 0.31 and 0.34; total factor productivity, Z, is set to 0.88 and 1; and the depreciation rate of capital, δ, is set to 0.062 and 0.067. These values are provided by the *Bureau of Economic Analysis*. Each pair of values represent annual averages within the periods of 1970–1974 and 2002–2006.

Tax: As summarized by the *Economic Report of the President*, the tax rate on labor income, τ_l, increased from 0.28 in the early 1970s (annual average of 1970–1974) to 0.31 in the 2000s (annual average of 2002–2006). In contrast, the tax rate on capital income, τ_k, was reduced from 0.35 to 0.26.

Preferences: In the period utility function, the parameters, σ_m and σ_f, which correspond to the Frisch elasticities of labor supply for males and females, respectively, are set to 2.08 and 0.57 for the 1970s and 2.08 and 0.80 for the 2000s.[11] Consequently, female labor supply is less elastic in the 2000s relative to the early 1970s.

Regarding the individual risk aversion parameter, γ_i, what is needed for the baseline model is the distribution of risk aversion that is appropriate for the focus of our research question: how *heterogeneous workers' self-selection into risky jobs* affects the measured welfare cost of rising wage uncertainty. The distribution can be obtained from individuals'

[11] We borrow the parameter values from Heathcote et al. (2010) whose model generates Frisch elasticities of labor supply of 0.48 for males for 1967–2005 and 1.77 and 1.25 for females in 1967 and 2005, respectively. Using these values, Heathcote et al. (2010) replicate some important empirical facts on the ratio of average female to average male hours and the correlation between year-to-year growth rates of a husband's and a wife's wages. Also, these values are generally consistent with empirical evidence from micro econometric analysis. For example, using the March Current Population Survey, Blau and Kahn (2007) document a decline in the elasticity of labor supply for married females from 1980 to 2000. (See Reichling and Whalen (2012) for a review of estimates of the Frisch elasticity of labor supply.) Using a structural model, a more recent study by Attanasio et al. (2018) also estimate the Frisch elasticity at 0.87 for the median household for 1980 to 2012. This estimate is roughly equal to the average of 0.48 (for male) and 1.25 (for female) Heathcote et al. (2010) obtained for 2005. For women, however, the Frisch elasticity estimate suggested by Heathcote et al. (2010) is larger than the estimate obtained by Blundell et al. (2016a): For 1999–2009, their estimated model implies Frisch elasticities of 0.528 and 0.850 for men and women, respectively. At the end of Section 5, we conduct a further robustness test of our main results by adopting a smaller estimate of Frisch elasticity suggested by Blundell et al. (2016a).

responses to the hypothetical gambles over lifetime income commonly fielded by various surveys such as the Panel Study of Income Dynamics (PSID), National Longitudinal Study of Youth 1979 (NLSY79), and Health and Retirement Study (HRS). The survey questions are structured in a way that individuals reveal their risk preferences in choosing between a job with a certain lifetime income (safe job) and a job with random, but higher mean lifetime income (risky job). Precisely, the above surveys commonly ask their respondents to choose between a safe job and a risky job. With equal chances, the risky job will double lifetime income or cut lifetime income by a specific fraction (or downward risk). Depending on their responses to the initial question, respondents are directed to subsequent questions where they choose between a safe job and another risky job with a different downside risk. While individuals accepting the risky job in the initial question then consider another risky job with a higher downside risk, those initially declining the risky job consider one with a lower downside risk. Individuals' responses to these questions can be formed into four ordinally-ranked categories of risk preferences from the most risk averse group (those who say 'no' to both the initial and subsequent questions) to the most tolerant group ('yes' to both questions). A series of empirical studies (e.g., Barsky et al., 1997; Kimball et al., 2008, 2009; Light and Ahn, 2010; and Sahm, 2012) exploit this information from various surveys and estimate the risk aversion parameter values that maximize the likelihood of those four categories being observed in their respective samples. The resulting measure of individual-level risk aversion is a cardinal measure of aversion that can be compared in a meaningful way across individuals. Because individuals reveal their risk preferences through choices between a safe job and a risky one, the derived distribution of individual risk aversion matches the unique features of our model economy where agents with heterogeneous risk preferences make optimal choices of job-related wage risk.

Precisely, we adopt an empirical distribution of individual risk aversions that is consistently estimated by several micro data based studies (aforementioned). For example, assuming log-normality of the distribution of risk aversion, Kimball et al.'s (2009) analysis of 20–69 year-olds in the 1996 PSID finds risk aversion distributed $\hat{\gamma}_i \sim lognormal$ (1.05, 0.76). To cross-validate the PSID-based empirical distribution with another data set and derive the distribution for the age range of 25–84 (the life-cycle in our model), we use the parameter values of individual risk aversion that are estimated by Light and Ahn (2010) based on the same income-gambling questions addressed to the NLSY79 respondents in 1993, 2002, and 2004 (for details of the estimation procedure, see p.917 of their paper). The original set of estimates of individual risk aversion supplied by Light and Ahn (2010) are balanced in the sense that everyone has valid estimates from 1979 through 2004, even though ages vary across individuals in a year. For *each* individual, we apply Ordinary Least Squares (OLS) to the regression of the logarithm of the estimated risk aversion level on a constant, age and age squared. Then, for each individual, we calculate the simple average of the predicted values over the 25–84 stages of the life-cycle and allow heterogeneity in age-fixed individual-specific (innate) risk preferences in the model.[12] The results show that the estimated parameters of individual relative risk aversions ($\hat{\gamma}_i$) are lognormal (1.33, 0.79).[13] Then we reproduce the empirical distribution based on 20–69 year-olds and obtain $\hat{\gamma}_i \sim lognormal$ (1.07, 0.71). This is virtually identical to Kimball et al.'s (2009) PSID-based result of *lognormal* (1.05, 0.76) for the same age range.

As will be demonstrated, under the empirical distribution of $\hat{\gamma}_i \sim lognormal$ (1.33, 0.79), our heterogeneous agent-job model matches some important data moments not directly targeted in the estimation, such as the overall dispersion of hours worked over the life-cycle and the employment size of the risky sector (to be defined subsequently). For the conventional homogeneous agent-job model, everyone is assumed to have the same risk aversion of $log\hat{\gamma}$

[12] Sahm (2012) quantifies both changes in risk tolerance over age and differences across individuals. Using panel data on hypothetical income gambles over lifetime income in the Health and Retirement Study (HRS), she finds that, although risk tolerance changes with age and macroeconomic conditions, 73% of the systematic variation is explained by persistent differences across individuals. In light of this work, we assume that risk preferences are heterogeneous, but age-invariant over the life-cycle. In addition, with age-varying risk preferences, the value function becomes discontinuous over the life-cycle, and global concavity is not guaranteed.

[13] The current paper does not consider gender-differences in risk aversion, as information on individual risk aversion is available only for the respondents, not for their spouses. Alternatively, we pool males and females and estimate the average risk aversion for a couple in the household.

=1.33 (equivalently, the mean risk aversion level of 5.62). Importantly, because we focus on the *difference* between the heterogeneous model with a non-degenerate distribution of risk aversion and the homogeneous model with a degenerate risk aversion in the measured welfare cost of the rising wage risk and the *relative* effectiveness of self-insurance mechanisms (e.g., family labor supply vs. borrowing/saving), our main results may not be sensitive to the adopted empirical distribution. In addition, a forthcoming discussion will include a robustness test of our main results with respect to changing the distribution of individual risk aversion.[14]

Endowment: deterministic component: Using the PSID for the period 1970 to 2014, we first estimate the following wage equation for each gender:

$$\log w^g_{i,j,t} = \beta^g_0 + \beta^g_1 j_{i,t} + \beta^g_2 j^2_{i,t} + z^g_{i,j,t}, \tag{12}$$

where $w^g_{i,j,t}$ is the real average hourly earnings of individual i of gender g at age j in year t, calculated by the ratio of total annual labor income divided by annual hours (deflated by CPI-U-RS, 2013=100), $j_{i,t}$ is age, and $z^g_{i,j,t}$ represents the stochastic wage component. The β coefficients are externally determined by applying Ordinary Least Squares (OLS) estimation to equation (12). They are then used to predict the individual deterministic wage rate in the same way for both the risky and safe sectors. The stochastic wage process is, however, modeled separately by sector and by gender, as described in the following subsection.

[14]The mean and median of the measured risk aversion levels are 5.62 and 3.78, respectively. These values are located in the middle of the spectrum of the estimates of risk aversion adopted by existing studies. They are somewhat higher than those commonly adopted in the consumption literature, where estimates of relative risk aversion are commonly set between 1 and 3. For example, with the reciprocal of the degree of risk aversion ($1/\gamma$) viewed as the elasticity of intertemporal substitution (EIS), Attanasio and Weber (1993, 1995) find a high elasticity (as high as 0.8). Heathcote et al. (2010) also assume a mean risk aversion of 1.5 in their model (so about 0.67 of EIS). On the contrary, on the basis of the log-linearized Euler equation, Hall (1988) finds that the magnitude of EIS is very small (less than 0.1), implying a mean risk aversion of greater than 10.

Table 1: Summary of Externally Determined Parameters

Parameter	Name (Source)	1970s	2000s
Demography			
j_R	Age of Retirement (assumption)	41	41
j_T	Terminal Age (assumption)	60	60
$\{\xi\}$	Conditional survival probability (U.S. Life Tables, Centers for Disease Control and Prevention, 1972, 2004)	See text	See text
Technology			
α	Capital's share of income (Bureau of Economic Analysis, 1970–74, 2002–06)	0.31	0.34
Z	Total factor productivity (Bureau of Economic Analysis, 1970–74, 2002–06)	0.88	1
δ	Depreciation rate (Bureau of Economic Analysis, 1970–74, 2002–06)	0.062	0.067
Tax			
τ_l	Labor income tax rate (Economic Report of the President, 1970–74, 2002–06	0.28	0.31
τ_k	Capital income tax rate (Economic Report of the President, 1970–74, 2002–06	0.35	0.26
Preferences			
σ_m, σ_f	Frisch elasticity of labor supply parameters, male/female (Heathcote et al. (2010))	2.08/0.57	2.08/0.80
γ_i for heterogeneous	Risk preferences (NLSY79, see text)	$\ln N(1.33,\ 0.79)$	
γ for homogeneous	Mean of γ_i's (NLSY79, see text)	5.62	
Endowment			
$f^g(j)$	Deterministic wages (PSID, NLSY, see text)	See text	See text

4.2 Model estimation and internally determined model parameters

4.2.1 Estimation strategy

The remaining model parameters are internally determined by estimating the model. For the baseline model, they include the share of total labor input from the risky sector (λ^R), the male share of the total labor input in the risky sector ($\lambda^{m,R}$), the male share among the total labor input in the safe sector ($\lambda^{m,S}$), the marginal disutility of work by gender (χ^m, χ^f), the discount factor (β), the borrowing limit ($\underline{a}$), the retirement benefit (b), and gender-job-specific variances of wage shocks. For the alternative model, three job-related parameters (λ^R, $\lambda^{m,R}$, $\lambda^{m,S}$) are dropped, and men's share of the total labor input (λ^m) is included in the list of structural parameters.[15] In addition, the alternative model focuses on the variances of wage shocks that are gender-specific but common to both sectors.

One important set of ingredients in the current life-cycle model are the set of parameters in equations (7) and (8): the job (or sector)-specific profiles of variances of permanent and transitory wage shocks and the persistence of permanent shocks each individual faces over the course of her or his career. Identification of these parameters is tricky. A required input of the baseline model is the job-specific profiles of variances of wage shocks an individual would expect over the job career if she/he *were* randomly assigned to the job, risky or safe. These structural parameters are not directly obtained from observed data, as the true profiles of variances of wage shocks are possibly different between risky and safe jobs, and workers self-select into jobs in the process of maximizing their lifetime expected utility. The observed variances of wage shocks for heterogeneous risk preference groups, therefore, mix the effects of the true variance of job-specific wage shocks with the effects of self-selection into job types. As we will show, these observed variances together with other relevant empirical moments

[15] Precisely, $H = \lambda^m H^m + (1-\lambda^m)H^f$. As in the baseline model, male and female labor inputs are perfect substitutes in production.

will be used to estimate the model's structural parameters.

Another challenging aspect of the estimation process lies in the number of structural parameters to be estimated. In the current two-sector model, the variances of wage shocks are allowed to vary even when workers stay in one sector (risky or safe), reflecting the age-varying nature of the variances of wage shocks (e.g., Baker and Solon, 2003; Gordon, 1984). For example, wages are relatively more volatile at the early stage of the work career due to more frequent job changes within sector and become relatively stable in the middle stage as the worker accumulates more job-specific human capital. This involves identification of 320 variances of both permanent and transitory shocks from age 25 to age 64 of the work-cycle for both genders and both sectors plus four parameters of persistence of permanent shocks by gender and by sector. To reduce the parameter space, we adopt a parametric form of the shock profile. We set the gender-job-specific variance of wage shocks (permanent or transitory) as a quartic function of age.

$$\left(\sigma^{g,n}_{\omega_j}\right)^2 = \delta^{g,n,\omega}_0 + \delta^{g,n,\omega}_1 j + \delta^{g,n,\omega}_2 j^2 + \delta^{g,n,\omega}_3 j^3 + \delta^{g,n,\omega}_4 j^4, \tag{13}$$

where $g \in \{m, f\}$, $n \in \{R, S\}$, and $\omega \in \{\eta, \varepsilon\}$, where m, f, R, S, η, and ε, respectively, denote male, female, risky, safe, permanent, and transitory. This assumption reduces the number of parameters to be estimated by 280 (from 320 to 40).

For the same reason, the profiles of wage shock variances have to be estimated even in the homogeneous agent-job model for both genders. For instance, if women are more likely to exit the labor market after a negative transitory shock, this would affect observed wages and has to be taken into account in the estimation of the wage process. Specifically, for each gender, the variance of the wage shock (permanent or transitory) one would expect at age j is determined by the following function:

$$\left(\sigma^{g}_{\omega_j}\right)^2 = \delta^{g,\omega}_0 + \delta^{g,\omega}_1 j + \delta^{g,\omega}_2 j^2 + \delta^{g,\omega}_3 j^3 + \delta^{g,\omega}_4 j^4. \tag{14}$$

Finally, the correlations between husband and wife in both permanent and transitory wage shocks are also important model ingredients. Other things being equal, a household would suffer from a greater welfare loss if permanent shocks are more highly correlated between spouses. These correlations are also internally determined within the model as the observed correlations between spouses' wage shocks are also affected by self-selection.

We target a number of empirical moments aimed primarily at pinning down the gender-sector-specific profiles of variances of wage shocks in equation (13). First, we use various four measures of relative wages within- and between-sectors. These include the hourly wage rate in the risky (relative to safe) sector, the ratio of female-to-male wages within each of the sectors, and the ratio of female to male wages overall. In order to obtain these empirical moments, we must define the risky and safe sectors. We determine the risky and safe sectors empirically using the PSID sample from 1976 through 1997.[16] We take the hourly wage rate of household heads collected from his/her main job, classify each respondent's main job according to the 20 occupation groups classified by the 2000 Census Occupational Classification System, pull all the residualized wage changes across individuals and years within each occupation group,[17] and compute the standard deviation of the changes for each occupation group. Appendix D displays estimated occupation-specific wage volatilities in descending order. The results show that the difference in estimated volatility between the construction and sales occupation groups is more distinct compared to any other pair of neighboring occupations below the sales occupation group. We, therefore, group the first four occupations in the 'risky' (precisely, relatively risky) sector, and the rest in the 'safe' (relatively safe) sector. When all observations are pooled across occupations within each sector, the standard deviation of residualized wage changes are 0.25 and 0.15 in the risky and safe sectors, with the difference being statistically significant at any conventional

[16]This is the longest sample period from which we can obtain information on the hourly wage rate from the main job held during the survey week for both genders on a yearly basis. The average hourly earnings variable defined as the ratio of annual earnings to annual hours is not appropriate for this purpose, as annual earnings often pertain to multiple jobs and job codes are available only for main jobs.

[17]To obtain residualized wage changes, we previously applied an OLS estimation to the regression of year-to-year changes in the logarithm of the hourly wage rate against a constant, age, and age squared.

significance level.[18] To reduce the sensitivity of the results coming from misclassification of the sectors, we do not use the employment size of the risky sector as a target moment directly.[19] Instead, we use Current Population Survey (CPS) data to compute the various types of relative wages mentioned at the beginning of this paragraph and include them in the list of empirical moments. The data moments displayed in Table 2 suggest that not only the female-to-male wage ratio (overall or within each sector) but also the hourly wage rate in the risky, relative to safe, sector increased from the early 1970s to the early 2000 (rows 5 through 8).

Second, in our effort to pin down the wage-shock-variance profiles in equation (13), we also include in the list of empirical moments the life-cycle profiles of wage shock variances observed by gender and by various preference groups and by gender. Combined with the above empirical moments of various types of relative wages, these observed moments will help the model identify both the size of the risky sector and the structural parameters of gender-job-specific profiles of wage shock variances in equation (13), as they are the results of heterogeneous individuals' wage-risk/job choices. We use the NLSY79 for 1985 to 2014 to derive the life-cycle profiles of observed variances of wage shocks (permanent and transitory) for each gender-preference group. Using the individual value of risk aversion previously described, we first classify individuals among four preference groups: those whose risk tolerance level is within the top 25% of the entire distribution, between 25% and 50%, between 50% and 75%, and the rest (the most risk averse group).[20] Then, for each gender-preference group, we apply OLS to equation (12) and, using the OLS residuals, estimate the

[18] The 95 percent confidence intervals for the true volatilities in the risky and safe sectors are (0.250, 0.254) and (0.152, 0.153), respectively. These are obtained based on the asymptotic distribution of the sample mean squared residual under classical regression assumptions (Schmidt, 1976).

[19] It should be noted that using the ratio of female to male wages within the risky sector, the ratio of female to male wages within the safe sector, and the ratio of female to male wages does not impose the employment size of the risky sector. In a previous version, we also tried an alternative definition of the risky (or volatile) sector: Including manufacturing and construction industries in the risky sector and the rest in the safe sector (e.g., Barlevy, 2001), and found little difference in the main result.

[20] The PSID is not suitable for this purpose, as it addresses the income gambling questions to household heads only in 1996. When the sample is divided into four preference groups, the resulting sample size is too small to estimate the life-cycle profiles of wage shock variances for each gender-preference group, which is particularly true for female heads.

life-cycle profiles of variances of wage shocks using the standard GMM estimation described in Appendix B.[21] Then, for each gender-preference group, we apply OLS to the regression of the estimated variance of wage shocks (permanent or transitory) against quartic polynomials in age, and use the resulting estimated coefficients as target moments for the coefficients in equation (13). For the alternative (homogeneous) model, we follow a similar procedure except that we focus on the empirical moments that are gender-specific but common to all preference groups.[22] The empirical moments for correlations between husband and wife in the permanent and transitory wage shocks are common for both models.

Panel A in Figure 2 presents a visualization of the empirical moments of the life-cycle profiles of variances of wage shocks by gender and by preference group for the early 1970s. The four aforementioned preference groups are denoted by Q1 (alternate long and short dashed line), Q2 (dashed line), Q3 (dotted line), and Q4 (solid line connecting circular points), respectively, from the most tolerant to the most risk averse group. In each case, the solid line without a data point represents the life-cycle profile of variances of wage shocks for all workers, which is used as the set of empirical moments for the homogeneous agent-job model. In each case, the solid line without a data point represents the life-cycle profile of variances of wage shocks for all workers, which is used as the set of empirical moments for the homogeneous agent-job model. In all figures, observed variances are systematically higher for more tolerant groups at all stages of the life-cycle. Interestingly, only for the most tolerant

[21] Since risk tolerant individuals are more likely to choose risky jobs, differences in the observed variance of wage shocks across different preference groups contain information on how many workers are employed in the risky sector and how variances of wages shocks are different between the two sectors. For example, other things being constant, if the share of the risky sector is relatively small (say 20%), and the difference between the two sectors in the variance of wage shocks is large, we would tend to observe a large difference between the most risk tolerant group and the other three groups in the size of observed variance of wage shocks. If the risky sector accounts for 50% of the total employment, and the difference between the two sectors in the variance of wage shocks is also large, we would tend to observe a large gap between the second and the third preference group in the observed variance of wage shocks, compared to either between the first and second groups or between the third and fourth groups.

[22] The resulting estimated coefficients, however, represent the life-cycle profiles of observed wage shock variances that apply for the NLSY79 cohort, not for the cohort of, say, the early 1970s economy. Under the assumption that the life-cycle *shape* of the variances of wage shocks is time-invariant, Appendix E explains how to shift the life-cycle profiles observed from the NLSY sample and derive the life-cycle profiles for the cohort of the early 1970s (or 2000s) economy.

group (Q1) is the estimated variance profile located well above the profile for all individuals. As previously noted, this is informative in identifying the size of the risky sector.[23]

For each of the early 1970s and the early 2000s, a total of 102 empirical moments are employed for the baseline model. Among them, 90 empirical moments concern the wage process. The additional 12 target moments are listed in the upper part of Table 2. Finally, the 54 structural parameters in the baseline model (listed in Table 3) are estimated by the Simulated Method of Moments (SMM, McFadden, 1989), which chooses parameter values that minimize the distance between a set of empirical moments and corresponding model-generated moments. For the alternative (homogeneous) model, 30 model parameters (24 wage shock-related moments plus 6 additional moments) are internally determined by the SMM that uses 33 empirical moments.[24]

[23]For both genders and all preference groups, estimated profiles are approximately U-shaped over the 25 to 64 stages of the life-cycle, which is particularly pronounced for transitory shocks. Individuals experience larger variances of wage shocks in the earlier and later stages of the life-cycle as compared to the middle stage. This is quite consistent with Baker and Solon (2003) and Gordon (1984), among others, except that the latter studies focus on men's earnings. Observed variances of both permanent and transitory shocks are generally greater for females than males for most of the work career. See panel A of Appendix F for the corresponding empirical moments from the early 2000s.

[24]See Appendix G for details of the SMM procedure and Appendix H for the computation algorithm. We also employ the U.S. social security benefit system in the model, as described in Appendix I.

Table 2: Data vs. Model-Generated Moments

Moments	1970s			2000s		
	Data	Hetero	Homo	Data	Hetero	Homo
Average household market hours	0.229[1]	0.233	0.229	0.266[1]	0.271	0.265
Ratio of female to male market hours	0.373[2]	0.394	0.382	0.687[2]	0.689	0.691
Male per capita employment	0.896[3]	0.898	0.889	0.852[3]	0.854	0.869
Female per capita employment	0.431[4]	0.438	0.426	0.675[4]	0.682	0.681
Ratio of average female to male wage in the risky sector	0.723[5]	0.714	N/A	0.911[5]	0.905	N/A
Ratio of average female to male wage in the safe sector	0.588[6]	0.588	N/A	0.756[6]	0.752	N/A
Hourly wage rate in risky relative to safe sector	1.311[7]	1.229	N/A	1.453[7]	1.391	N/A
Ratio of female to male wages (Gender wage gap)	0.592[8]	0.600	0.587	0.788[8]	0.794	0.781
Capital-to-output ratio	3.00[9]	3.00	2.96	3.20[9]	3.23	3.12
Gini coefficient of income	0.397[10]	0.418	0.408	0.466[10]	0.512	0.502
Negative asset share	0.155[11]	0.158	0.171	0.155[11]	0.166	0.179
Wage-Shock-Related Empirical Moments						
Life-cycle profiles of variances of permanent and transitory wage shocks by gender and by preference group	See panel A Figure 2[12]	See panel B, Figure 2[12]	N/A	See panel A Appendix F[12]	See panel B, Appendix F[12]	N/A
Life-cycle profiles of variances of permanent and transitory wage shocks by gender	See panel A, Figure 2[12]	N/A	See panel B, Figure 2[12]	See panel A, Appendix F[12]	N/A	See panel B, Appendix F[12]
Persistence of male permanent wage shocks by preference group (Q1)	0.941[13]	0.943	N/A	0.941[13]	0.942	N/A
(Q2)	0.961[13]	0.963	N/A	0.961[13]	0.963	N/A
(Q3)	0.966[13]	0.968	N/A	0.966[13]	0.970	N/A
(Q4)	0.970[13]	0.973	N/A	0.970[13]	0.974	N/A

Table 2: Data vs. Model-Generated Moments (Cont'd)

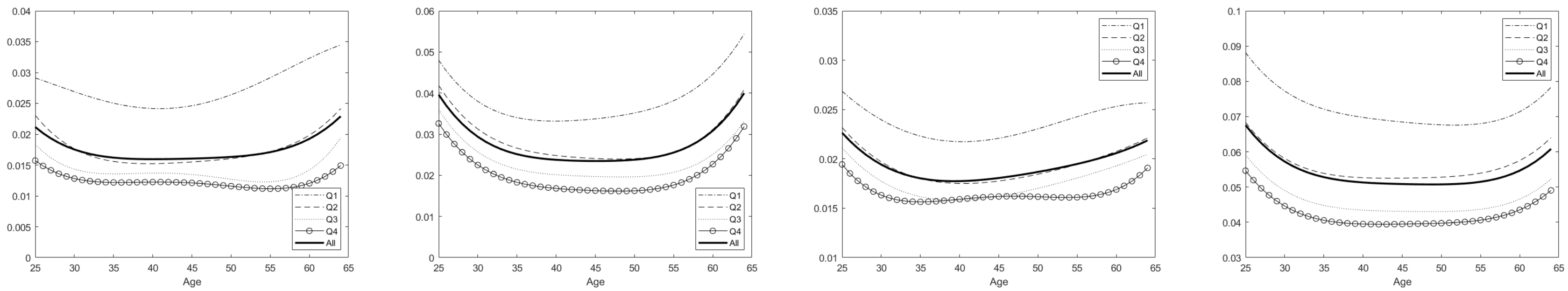
Moments		1970s			2000s		
		Data	Hetero	Homo	Data	Hetero	Homo
Persistence of female permanent wage shocks by preference group	(Q1)	0.943[13]	0.946	N/A	0.943[13]	0.946	N/A
	(Q2)	0.961[13]	0.954	N/A	0.951[13]	0.956	N/A
	(Q3)	0.953[13]	0.957	N/A	0.953[13]	0.960	N/A
	(Q4)	0.957[13]	0.960	N/A	0.957[13]	0.965	N/A
Persistence of male permanent wage shocks (all workers)		0.965[13]	N/A	0.968	0.965[13]	N/A	0.970
Persistence of female permanent wage shocks (all workers)		0.953[13]	N/A	0.957	0.953[13]	N/A	0.960
Covariance of male and female permanent shocks		0.0034[14]	0.0032	0.0032	0.0040[14]	0.0038	0.0041
Covariance of male and female transitory shocks		0.0077[14]	0.0073	0.0075	0.0119[14]	0.0117	0.0121

Sources: (1) The authors' calculation using the PSID. Male average market hours are normalized to 1/3 (0.3333). Female market hours are calculated accordingly. (2) The authors' calculation using the PSID. (3) The authors' calculation using the March CPS. (4) The authors' calculation using the March CPS. (5) The authors' calculation using the March CPS. (6) The authors' calculation using the March CPS. (7) The authors' calculation using the March CPS. (8) The authors' calculation using the March CPS. (9) Bureau of Economic Analysis. (10) Gini coefficient of income: Census Bureau. (11) Wolff (2000). (12) The authors' calculation using the NLSY and the PSID. See Section 4 and Appendix E for the derivation. (13) The authors' calculation using the NLSY. (14) The authors' calculation using the PSID.
Notes:(1) Each pair of data moments represent annual averages within periods of 1970–1974 and 2002–2006, respectively. (2) The (relatively) risky sector includes legal occupations, Healthcare Practitioners and Technical Occupations, Management Occupations & Business Operations Specialists & Financial Specialists, and Construction Trades & Extraction Workers. The (relatively) safe sector includes all the other occupations. See Section 4 and Appendix D for a detailed discussion of the definitions of the risky and safe sectors. (3) Q1, Q2, Q3, and Q4 represent, respectively, those whose risk tolerance level is within the top 25% of the entire distribution (most risk tolerant group), between 25% and 50%, between 50% and 75%, and the rest (the most risk averse group).

Figure 2: Life-Cycle Profiles of Variances of Permanent and Transitory Wage Shocks in the 1970s Economy

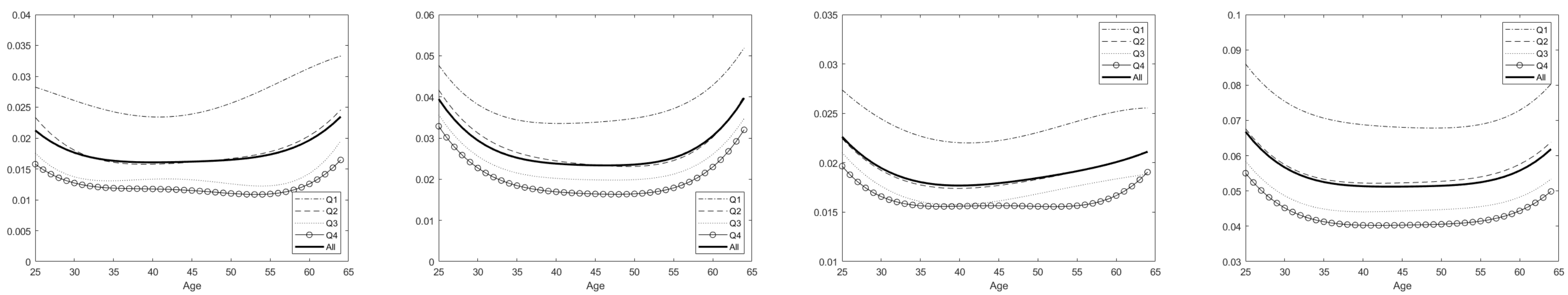
A. Empirical Moments

(A) Male Permanent (B) Male Transitory (C) Female Permanent (D) Female Transitory



B. Model-Generated Moments

(A) Male Permanent (B) Male Transitory (C) Female Permanent (D) Female Transitory



Sources for pane A: The authors' estimation using both the PSID (1970–2014) and NLSY79 (1979–2014).

Notes: See the text for the derivation of empirical moments and corresponding model-generated moments. The solid line represents the life-cycle profile of variance of wage shocks in the homogeneous economy. In the heterogeneous economy, Q1 (alternate long and short dashed line), Q2 (dashed line), Q3 (dotted line), and Q4 (solid line connecting circular points) represent, respectively, the life-cycle profiles of those whose risk tolerance level is within the top 25% of the entire distribution, those between 25% and 50%, those between 50% and 75%, and the rest (the most risk averse group).

4.2.2 Estimation results

Targeted moments: Comparisons of the empirical moments and corresponding model-generated moments appear in Table 2, Figure 2, and Appendix F. These show that the current models match the data moments to a reasonable degree in both the heterogeneous or homogeneous economies for both the early 1970s and the early 2000s.

Non-targeted moments: Table 3 and Figure 3 illustrate how the estimated model performs in dimensions not directly targeted in the estimation. First, for each gender and each time period (early 1970s and 2000s), Table 3 compares the employment share of the risky sector generated by the heterogeneous model with the corresponding empirical moment. These are calculated by the number of workers instead of man-hours. The empirical moments are obtained based on the CPS with the March CPS supplement weights being used. When both genders are combined in columns (3) and (6), the model-generated moments match the corresponding empirical moments almost perfectly. For women, while the share of the risky sector is slightly over-predicted by the model for each of the early 1970s and 2000s periods, the change in the share between the two periods appears very similar between the empirical and model-generated moments. For men, both the shares and the change remain very similar between the model-predicted and data moments. Second, Figure 3 shows how the estimated heterogeneous and homogeneous models imply for the overall dispersion of hours worked over the life-cycle. The March CPS Supplement is used to generate the data moments. For each time period, the average annual labor hour among the entire male population is normalized as 0.333. The takeaway result from Figure 3 is that, for each gender and each time period, the heterogeneous model matches the life-cycle variation of hours worked better than the homogeneous model does. Precisely, at each stage of the life-cycle, the data moment is closer to the moment generated by the heterogeneous, relative to homogeneous, model. As will be demonstrated subsequently, labor adjustments, extensive and intensive, play a crucial role in evaluating the welfare cost of the increased variances of wage shocks.

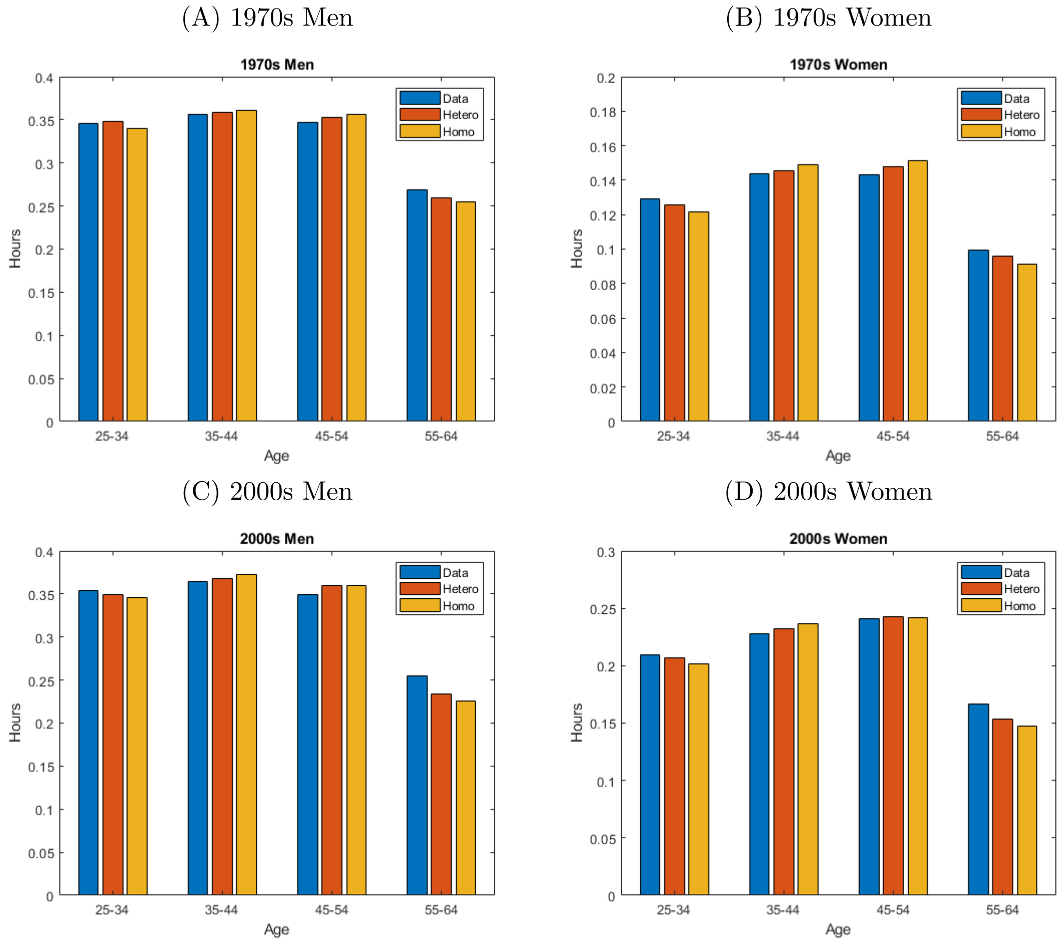
Table 3: Matching Non-Target Moments

	Data			Model		
	Men (1)	Women (2)	All (3)	Men (4)	Women (5)	All (6)
Employment share of the risky sector (1970–1974)	0.2671	0.1178	0.2186	0.2602	0.1388	0.2204
Employment share of the risky sector (2002–2006)	0.2945	0.2493	0.2745	0.2892	0.2623	0.2773

Sources: The authors' calculation using the CPS. March CPS supplement weights are used.
Notes: Heterogeneous model. The employment share is calculated by the number of workers instead of man-hours. See the text for definitions of the risky sector.

Estimated model parameters: Estimated structural parameters are presented in Table 4. The numbers in parentheses represent standard errors of the SMM estimator. They suggest that all parameters are estimated precisely. For brevity, the estimated model parameters in equations (13) and (14) are reported in Figure 4. Although not presented in values, all the coefficients in the quartic functions are estimated precisely. How did the variances of wage shocks change in the heterogeneous economy from the early 1970s to the 2000s? In Figure 4, thicker and thinner lines represent variances of permanent and transitory shocks, respectively, and solid and dashed lines correspond to the early 1970s and the 2000s, respectively. As shown in panel A, the variances of permanent shocks have increased in both risky and safe jobs and for both genders. While there have been dramatic increases in the variance of male transitory shocks in both risky and safe jobs, the variances of female transitory shocks have decreased in all jobs. Panel B shows how variances of wage shocks have changed for each gender in the homogeneous economy from the early 1970s to the early 2000s. As in the heterogeneous model, these changes are separated from the changes in the observed variances of wage shocks implied by the solid lines in Figure 2 and Appendix F, and represent exogenous changes in wage shock variances that are free of the effect of workers' selection into employment. When averaged over the life-cycle, the variance of permanent wage shocks increased by 0.0073 (from 0.0185 to 0.0258) for male and by 0.0071 (from 0.0203 to 0.0274)

Figure 3: Matching Non-Targeted Moments: Overall Dispersion of Hours Worked Over Life-Cycle



Data Source: Current Population Survey March Supplement.
Notes: For each time period, men's annual average hours are normalized by 1/3.

for female. While the variance of male transitory shock increased by 0.0242 (from 0.0298 to 0.0540), the variance of female transitory shock decreased by 0.0119 (from 0.0584 to 0.0465). A careful comparison of these estimates with those in Figure 1 suggests that, due to the workers' self-selection, the true variance of wage shocks recovered from the model is greater than corresponding observed variance of wage shocks reported in Figure 1 for each of the early 1970s and 2000s economies. Section 5 analyzes the welfare costs caused by these exogenous changes in the variance of wage shocks described in Figure 4.[25]

[25] Additional calculation shows that the estimated correlation of husbands' and wives' permanent wage shocks declined from 0.21 (1970s) to 0.17 (2000s). These correlations are somewhat lower than 0.57 suggested by Hyslop (2001), but more close to Voena (2015) and Blundell et al. (2016a).

Table 4: Summary of Internally Determined Parameters

Parameter	Data	1970s		2000s	
		Hetero	Homo	Hetero	Homo
λ^m	Male share among total labor input	N/A	0.665 (0.068)	N/A	0.570 (0.061)
$\lambda^{m,R}$	Male share among total labor input in risky sector	0.786 (0.059)	N/A	0.593 (0.054)	N/A
$\lambda^{m,S}$	Male share among total labor input in safe sector	0.631 (0.055)	N/A	0.562 (0.046)	N/A
λ^R	Share of risky input among total labor input	0.231 (0.024)	N/A	0.289 (0.029)	N/A
χ^m	Disutility of male work	39.313 (4.039)	345.975 (21.352)	7.964 (0.545)	48.394 (4.119)
χ^f	Disutility of female work	27.394 (3.168)	255.647 (18.334)	5.267 (0.737)	35.946 (3.554)
β	Discount factor	0.951 (0.069)	0.935 (0.056)	0.949 (0.065)	0.937 (0.059)
$\underline{a}$	Borrowing limit	-0.187 (0.018)	-0.172 (0.015)	-0.226 (0.025)	-0.216 (0.020)
b	Retirement benefits	0.305 (0.018)	0.291 (0.018)	0.396 (0.028)	0.375 (0.025)
Variance of wage shocks: Structural parameters in equations (13) and (14)					
$\delta_0^{g,n,\omega}$ through $\delta_4^{g,n,\omega}$: parameters in equation (13)		See panel A, Figure 4	N/A	See panel A, Figure 4	N/A
$\delta_0^{g,\omega}$ through $\delta_4^{g,\omega}$: parameters in equation (14)		N/A	See panel B, Figure 4	N/A	See panel B, Figure 4

Table 4: Summary of Internally Determined Parameters (Cont'd)

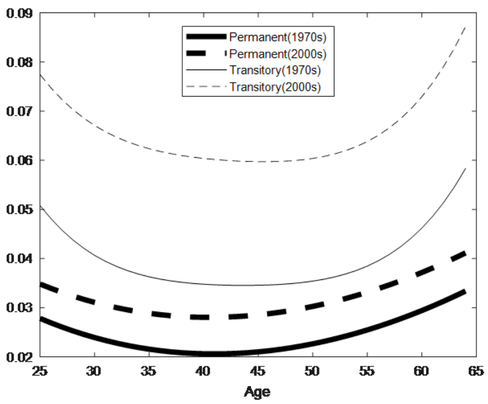
Parameters	Data	1970s		2000s	
		Hetero	Homo	Hetero	Homo
$\rho^{m,R}, \rho^{m,S}$	Persistence of permanent wage shocks for male in risky and safe sectors	0.955 (0.075) 0.971 (0.066)	N/A	0.957 (0.077) 0.973 (0.067)	N/A
$\rho^{f,R}, \rho^{f,S}$	Persistence of permanent wage shocks for female in risky and safe sectors	0.956 (0.072) 0.967 (0.067)	N/A	0.960 (0.078) 0.970 (0.071)	N/A
ρ^{m}, ρ^{f}	Persistence of permanent wage shocks for male and female	N/A	0.963 (0.061) 0.954 (0.065)	N/A	0.965 (0.063) 0.960 (0.067)
$\sigma^{p}_{\eta^m,\eta^f}$	Covariance of couple's permanent wage shocks	0.0031 (0.0004)	0.0032 (0.0004)	0.0037 (0.0005)	0.0038 (0.0005)
$\sigma^{t}_{\epsilon^m,\epsilon^f}$	Covariance of couple's transitory wage shocks	0.0070 (0.0006)	0.0074 (0.0006)	0.0112 (0.0008)	0.0115 (0.0008)

Notes: Numbers in parentheses are estimated standard errors obtained by the Simulated Methods of Moments estimation. See notes to Table 2. See equations (13) and (14) for definitions of δ coefficients.

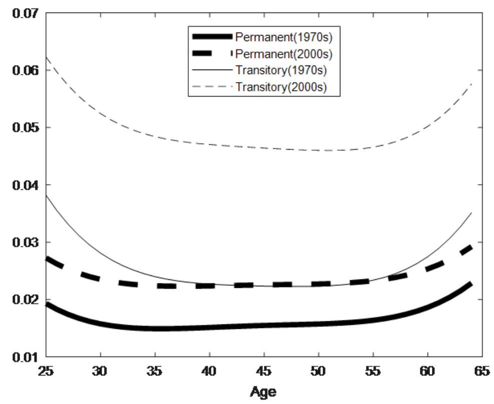
Figure 4: Changes in Gender-Job-Specific Profiles of Variances of Wage Shocks from the 1970s to the 2000s

A. Heterogeneous Agent-Job Model

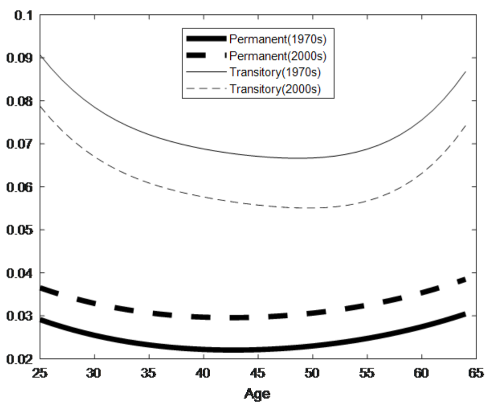
(A) Male risky job



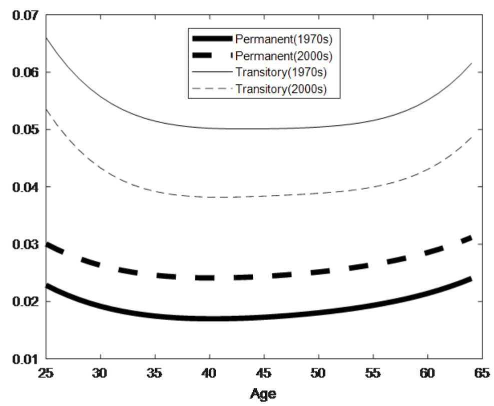
(B) Male safe job



(C) Female risky job

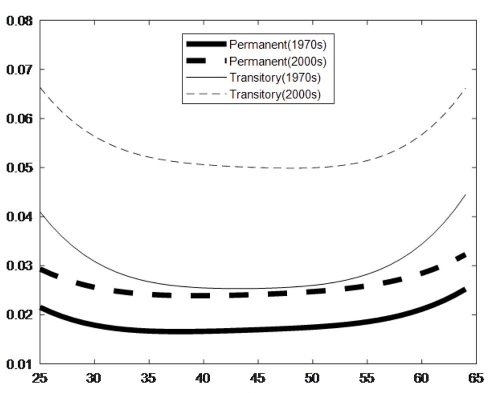


(D) Female safe job

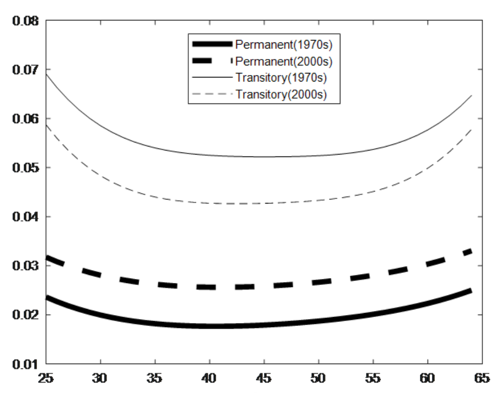


B. Homogeneous Agent-Job Model

(A) Male



(B) Female



Notes: Thicker and thinner lines are variances of permanent and transitory shocks, respectively, and solid and dashed lines are for the early 1970s and the 2000s, respectively.

Equilibrium configuration: For the heterogeneous economy, Table 5 describes an equilibrium configuration of worker-sector matching in the 1970s economy and how the equilibrium would shift following changes in the variance of gender-job-specific wage shocks. Panel A shows how much wage shock variance and unit price of labor a worker would expect from each sector in the 1970s economy. The price of female efficiency unit of labor in the safe sector as of the early 1970s economy is normalized as 1. To facilitate the comparison of the variances of wage shocks between the risky and safe sectors and between genders, the variances of wage shocks are simply averaged over the work-cycle. The variances of both permanent and transitory wage shocks are higher in the risky sector for both genders. The price per efficiency unit of labor is also higher in the risky sector for both genders. The equilibrium price differential between the two sectors (as measured by the ratio of the risky price to the safe price) is higher for men (1.84=2.65/1.44) than for women (1.39). As noted previously, due to lack of information on matching between husband and wife in terms of risk preference, we assign a common risk preference parameter to a household. Consequently, the higher price differential for men reflects that the ratio of variances of wage shocks between the risky and the safe sectors is higher for men, particularly for permanent wage shocks: 1.47 (=0.0243/0.0165) for men vs. 1.29 (=0.0247/0.0191) for women. Panel B of Table 5 displays the joint population distribution of gender and sector for the early 1970s economy. With employment represented by the number of workers, the share of the risky sector among the total workers is estimated at 22.0% $\left(=\frac{14.7}{14.7+52.1}\times 100\%\right)$. This is slightly smaller than the estimated share of the risky input ($\widehat{\lambda^R}$=0.231) based on efficiency unit of labor (see Table 4).[26]

How does the equilibrium change in response to exogenous changes in the variances of wage shocks? Panel C summarizes information contained in panel A of Figure 4: how variances of wage shocks change in each sector-gender cell. Once again, the variances of

[26]The estimated share of the risky sector (approximately 25%) is informed by the observation that, when the sample is divided into four preference groups with an equal size, the most risk tolerant group shows a much greater variance of wage shocks compared to the other groups, while the other three groups are less discernable in the variance of wage shocks they experience.

the wage shocks are simply averaged over the work-cycle. As shown in the last column, for each gender, changes in the variance of wages shocks (permanent or transitory) are similar between the risky and safe sectors. These results are quite consistent with existing evidence that increases in the variance of wage shocks are spread across all subgroups or sector-neutral (e.g., Gottschalk and Moffitt, 1994; Feigenbaum and Li, 2015).[27] Following these changes in wage shock variances, households re-optimize their behavior, while input demand and output supply functions of risk-neutral firms remain the same, and markets clear at a new level. As a result, the employment share of the total population increases from 66.8% (14.7% + 52.1% in the last column of panel B) to 70.2% (17% + 53.2% in the last column of panel D), partly reflecting the increase in precautionary labor supply in response to the increased uncertainty. Due to the nature of 'sector-neutral increase' in the variance of wage shocks, the share of the risky sector among the total workers increases slightly from 22% to 24.2% $\left(=\frac{17}{17+53.2}\times 100\%\right)$. At least two factors are responsible for this. First, the wide-spread increases in the variance of wage shocks make the two sectors less distinct from each other: for example, for men (women), the ratio of the variance of the permanent shock in the risky sector to that in the safe is reduced from 1.47 (1.29) to 1.34 (1.23), which makes the risky sector relatively more appealing after the change. Second, the equilibrium price differential between the risky and safe sectors increases slightly from the old equilibrium (1.84 for men and 1.39 for women) to the new one (1.86 for men and 1.41 for women).[28]

[27] Using the PSID data, Gottschalk and Moffitt (1994) find that all major industries experienced an increase in transitory fluctuations during the 1970s and the 1980s. Feigenbaum and Li's (2015) analysis of the PSID from 1970 to 2005 concludes that the rise in household income uncertainty during the sample period is widespread across all sub-populations.

[28] Our calculation shows that the conditional probability of switching from the safe to risky sector is slightly greater than the conditional probability of switching from the risky to safe sector (1.5% vs. 1% for men, and 2.6% vs. 1.2% for women).

Table 5: Equilibrium Configuration in Heterogeneous Economy

(A) Variance of Wage Shocks and Unit Price of Labor by Sector and by Gender in the 1970s Economy

	Male			Female		
	Wage shocks		Price of	Wage shocks		Price of
	Permanent	Transitory	labor	Permanent	Transitory	labor
Risky Sector	0.0243	0.0399	2.65	0.0247	0.0731	1.39
Safe Sector	0.0165	0.0268	1.44	0.0191	0.0535	1

Notes: Wage shocks are simply averaged over the work-cycle. The female unit price of labor in the safe sector is normalized to 1.

(B) Joint Population Distribution of Gender and Sector (%) in the 1970s Economy

	Male	Female	Total
Risky Sector	11.7	3.0	14.71
Safe Sector	33.2	18.9	52.1
Non-market Sector	5.1	28.1	33.2
Total	50	50	100

(C) Changes in Variance of Wage Shocks by Sector and by Gender

			1970s	2000s	Change
Permanent Wage Shocks	Risky Sector	Male	0.0243	0.0318	0.0075
		Female	0.0247	0.0323	0.0076
	Safe Sector	Male	0.0165	0.0237	0.0072
		Female	0.0191	0.0262	0.0071
Transitory Wage Shocks	Risky Sector	Male	0.0399	0.0658	0.0260
		Female	0.0756	0.0646	-0.0110
	Safe Sector	Male	0.0268	0.0509	0.0241
		Female	0.0525	0.0398	-0.0127

Notes: Variances of wage shocks are simply averaged over the work-cycle.

Table 5: Equilibrium Configuration in Heterogeneous Economy (Cont'd)

(D) New Equilibrium Distribution of Gender and Sector and Unit Price of Labor by Sector and by Gender: Share (%)/Price

	Male	Female	Total
Risky Sector	12.5/2.70	4.5/1.42	17.0
Safe Sector	33.4/1.45	19.8/1.01	53.2
Non-market Sector	4.2	25.7	29.8
Total	50.0	50.0	100.0

Notes: The female unit price of labor in the safe sector as of the early 1970s is normalized to 1. In panels A and C, the estimated variance of wage shocks for each gender-sector cell represents the estimated variance averaged over the work-cycle.

5 Welfare Consequences of Changing Wage Shocks

5.1 Welfare Cost Calculations

On the basis of the estimated models in Section 4, this section evaluates the welfare costs of the changing variances of wage shocks from the early 1970s to the early 2000s. This welfare analysis is based on the estimated models for the 1970s economy. The models for the 2000s economy were estimated to obtain the structural parameters of (gender-job-specific) variances of wage shocks that existed in the 2000s. Our central research questions are *how much welfare loss/gain the 1970s households would experience if they faced the variance of wage shocks of the 2000s (with all the other parameter values/economic conditions remaining the same at the 1970s level) and how the measured welfare costs are different between the conventional homogeneous agent-job model and our heterogeneous agent-job model*. We adopt two approaches initially. The first one is to allow output and input prices to vary in response to changes in the variances of wage shocks. In this full general equilibrium approach, all economic agents optimize their behaviors, and markets clear both before and after the changes in the variances of wage shocks. In the second approach, we start with the same general equilibrium state of the 1970s economy, but output and input prices stay fixed at the 1970s equilibrium level when variances of wage shocks increase to the 2000s level. Otherwise, households still re-optimize their behavior in response to the arrival of new wage shocks. This partial analysis, therefore, does not require market clearing after changes in the variances of wage shocks. A comparison of the results from the two approaches tells us the importance of the general equilibrium effects in our welfare analysis.

For each approach, the welfare cost is calculated as follows. First, for each household, the welfare costs of the increased variance of wage shocks are the values of ω and ω' in the heterogeneous and homogeneous economies, respectively, which solve the following equations (see Heathcote et al. (2010), among others, for a reference):

$$\begin{aligned} E\Big(\sum_{j=1}^{j_T}\beta^j\phi_j u(c^*_{i,j}, l^{m*}_{i,j}, l^{f*}_{i,j};\gamma_i\Big) &= E\Big(\sum_{j=1}^{j_T}\beta^j\phi_j u((1+\omega)c^{**}_{i,j}, l^{m**}_{i,j}, l^{f**}_{i,j};\gamma_i\Big), \\ E\Big(\sum_{j=1}^{j_T}\beta^j\phi_j u(c^*_{i,j}, l^{m*}_{i,j}, l^{f*}_{i,j}\Big) &= E\Big(\sum_{j=1}^{j_T}\beta^j\phi_j u((1+\omega')c^{**}_{i,j}, l^{m**}_{i,j}, l^{f**}_{i,j}\Big), \end{aligned} \tag{15}$$

where $\{c^*_{i,j}, l^{m*}_{i,j}, l^{f*}_{i,j}\}_{j=1}^{j_T}$ and $\{c^{**}_{i,j}, l^{m**}_{i,j}, l^{f**}_{i,j}\}_{j=1}^{j_T}$ are the equilibrium allocations of a 1970s household facing the 1970s and the 2000s wage shocks, respectively. The left-hand-side of each equation represents the maximized lifetime utility the 1970s household can expect under the assumption that all the model parameters of the 1970s economy remain unchanged over the life-cycle. The right-hand-side stands for the maximized lifetime utility the same 1970s household can expect under the assumption that all the other model parameters still remain unchanged at the 1970s level over the life-cycle, except that they face the new life-cycle wage shock profiles that existed in the 2000s economy. Then, measured welfare costs are averaged across all households to produce the welfare cost caused by the exogenous changes in the variances of various wage shocks as described in Figure 4.

5.2 Why a heterogeneous agent-job model?

Table 6 compares measured welfare costs between the homogeneous and heterogeneous economies. Panels A and B show results from the general equilibrium and partial equilibrium analyses, respectively. Several findings emerge immediately from this table. First, the results are quite similar between the two panels, implying little general equilibrium effects when assessing the welfare costs of the increased variances of wages shocks. This reflects the fact that prices do not change much in response to the increased variance of wage shocks.[29] The insignificance of the general equilibrium effect is also observed in a different, but related literature. For example, Storesletten et al. (2001) find that welfare gains from removing business cycle variation in idiosyncratic shocks remain similar whether or not the general equilibrium effect is considered.

Second, and more importantly, welfare costs are overstated by assuming homogeneous risk preferences and neglecting workers' self-selection into risky, as opposed to safe, jobs. For example, in panel A, when the 1970s households face the four types of wage shocks that exist in the 2000s economy, the measured welfare cost is 14.07% and 7.56% (in lifetime consumption equivalent) in the homogeneous and heterogeneous economies, respectively. Estimates in the last column suggest that the welfare cost is overstated by about 90% in the homogeneous economy. This estimate remains robust whether the general equilibrium effects are included in the analysis, as can be seen through a comparison of panels A and B.

Third, in both models, the increase in the variance of the male permanent shock is a dominant contributor to the overall welfare cost. Even though the variances of permanent shocks increased similarly for both genders, the welfare cost caused by the increased variance of the female permanent shock is about a half of the amount generated by the increased variance of the male permanent shock. This is mainly attributed to women's lower wages and hours, and consequently, women's smaller contribution to household income. Nevertheless,

[29] As noted at the end of Section 4, the price of the efficiency unit of labor increases only slightly in response to the changes in the variances of wage shocks. Concurrently, the interest rate decreases by a very small amount from 2.47% to 2.4%.

Table 6: Welfare Costs of Changes in Variances of Wage Shocks from the Early 1970s to the 2000s

(A) General Equilibrium Effects Included

Type of Shocks	Homogeneous Model ω' (%) (a)	Heterogeneous Model ω (%) (b)	Overstatement of Welfare Cost (a)/(b)
Δ All Types of Shocks	14.07	7.56	1.86
Δ Male permanent	7.73	3.74	2.07
Δ Male transitory	1.89	1.01	1.88
Δ Female permanent	3.98	2.02	1.99
Δ Female transitory	–0.74	–0.34	2.20

(B) Partial Analysis: Prices Are Fixed at the Equilibrium Level of the 1970s Economy

Type of Shocks	Homogeneous Model ω' (%) (a)	Heterogeneous Model ω (%) (b)	Overstatement of Welfare Cost (a)/(b)
Δ All Types of Shocks	13.33	7.03	1.90
Δ Male permanent	7.35	3.49	2.11
Δ Male transitory	1.82	0.96	1.89
Δ Female permanent	3.79	1.88	2.02
Δ Female transitory	–0.71	–0.32	2.21

Notes: The welfare costs are generated by changes in the variances of various types of wage shocks from the early 1970s to the 2000s, under the assumption that the model parameters/economic conditions remain unchanged at the 1970s levels. (See equation (15) for our working definition of the welfare cost.) For the homogeneous economy, variances of male permanent, male transitory, female permanent, and female transitory shocks have changed by 0.0073 (from 0.0185 to 0.0258), 0.0242 (from 0.0298 to 0.0540), 0.0071 (from 0.0203 to 0.0274), and –0.0119 (from 0.0584 to 0.0465), respectively, from the early 1970s to the early 2000s. See panel C of Table 5 for changes in the gender-job-specific variance of wage shocks for the heterogeneous economy.

the increase in the variance of female permanent shocks remains as the next-most important contributor to the overall welfare cost. Although the variance of male transitory shocks increased the most among the four types of wage shocks, the resulting welfare cost is much smaller compared to even the welfare cost of the increased variance of female permanent shocks. As expected, permanent shocks are even more consequential than transitory shocks. Little welfare gain is generated by the reduced variance of the female transitory shock.

Fourth, in each economy, the total welfare cost is greater from the combined effect of all four types of shocks, relative to the sum of the four measured welfare costs that would be generated by the changes to the four types of shocks occurring individually. In the partial analysis, for example, the combined effect is 13.33% vs. 12.25% in the homogeneous economy, and 7.03% vs. 6.01% in the heterogeneous economy. This suggests that there exist some interactive effects among the four types of wages shocks (due to limited insurability) in generating their welfare costs.

Why is the measured welfare cost higher for the homogeneous, relative to heterogeneous, economy? This question is of particular interest because the distribution of risk aversion is right-skewed. One might expect that the greater welfare loss of more risk-averse agents relative to an agent with the average risk aversion dominates the smaller welfare loss of less risk-averse agents, resulting in the greater welfare cost in the heterogeneous agent-job model. While this conjecture may be valid in a world where agents are unable to cope with the increased variances of wage shocks, it does not apply once the model is enriched to encompass various self-insurance mechanisms. The welfare gain of allowing agents these insurance mechanisms is greater for the more risk-averse agents compared to the 'average' risk-averse agent, and while it is smaller for the less risk-averse agents, the former dominates the latter. As emphasized in the introduction, a more appropriate and realistic welfare evaluation should consider agents' insurability against the increased variance of wage shocks.

Table 7 provides a quantitative explanation of why the measured welfare cost is overstated in the homogeneous agent-job model. *With little general equilibrium effect, for brevity, we* *adopt the partial equilibrium approach for the rest of the analysis.* As previously noted, even in this partial analysis, households still re-optimize their behavior in response to the arrival of new wage shocks. In panel A, we examine a source of the greater welfare cost for the homogeneous economy by teasing out the effects of assuming homogeneous risk preference (henceforth 'preference effects') from the effects of assuming homogeneous jobs (henceforth 'selection effects'). As observed in panel B of Table 6, the gap between the homogeneous and heterogeneous economies in estimated welfare cost is 6.3%P (=13.33%–7.03%). Because the estimated model parameters are not the same between the two economies, a portion of this total gap is attributable to differences between the two economies in model parameters aside from preference parameters. In the first row of panel A in Table 7, we use the heterogeneous model (that is, fix the other parameters as those in the heterogeneous model) in estimating the two effects. Column (1) represents the welfare cost of the increased variance of wage shocks for the heterogeneous economy, which is borrowed from panel B of Table 6. In column (2), we recalculate the welfare cost by eliminating the job selection process from the model. This is done by replacing the gender-job-specific wage shock variances of the estimated heterogeneous model (panel A of Figure 4) with the gender-specific wage shock variances of the homogeneous model (panel B of Figure 4). In column (3), we recalculate the welfare cost by imposing homogeneous risk preference in the estimated heterogeneous model. In column (4), we recalculate the welfare cost by imposing both homogeneous preference and homogeneous job in the estimated heterogeneous model. In all cases, households re-optimize their behavior in response to the increased variances of wage shocks.

Table 7: Decomposition of Welfare Cost Gap between Homogeneous and Heterogeneous Economies

(A) Decomposition

	Hetero Pref Hetero Jobs (1)	Hetero Pref Homo Job (2)	Homo Pref Hetero Jobs (3)	Homo Pref Homo Job (4)	Decomposition by (2): Preference effects (4)–(2)	Decomposition by (2): Selection effects (2)–(1)	Decomposition by (3): Preference effects (3)–(1)	Decomposition by (3): Selection effects (4)-(3)
Using Hetero Model Parameters	7.03	8.62	10.34	12.06	3.44 (68%)	1.59 (32%)	3.31 (66%)	1.72 (34%)
Using Homo Model Parameters	8.21	9.96	12.34	13.33	3.37 (66%)	1.75 (34%)	4.13 (81%)	0.99 (19%)

(B) More on Effects of Preference Heterogeneity

		Decomposition by (2): Homo Pref Homo Job (4)	Decomposition by (2): Hetero Pref Homo Job (2)	Decomposition by (2): (4)–(2)	Decomposition by (3): Homo Pref Hetero Jobs (3)	Decomposition by (3): Hetero Pref Hetero Jobs (1)	Decomposition by (3): (3)–(1)
Using Hetero Model Parameters	Labor Supply & Saving/Borrowing Adjusted (1)	12.06	8.62	3.44	10.34	7.03	3.31
	No adjustments allowed (2)	14.67	16.44	-1.77	13.36	15.11	-1.75
Using Homo Model Parameters	Labor Supply & Saving/Borrowing Adjusted (3)	13.33	9.96	3.37	12.34	8.21	4.13
	No adjustments allowed (4)	21.41	24.92	-3.51	19.79	23.25	-3.46

Notes: See the text for definitions of the preference effect and the selection effect. In case of using the parameters of the heterogeneous (homogeneous) model, the sum of the job selection and the preference effects is 5.03 (5.12), which accounts for 79.8% (81.3%) of the total gap between the homogeneous and heterogeneous economies in the measured welfare cost (6.3%P=13.33%–7.03%). The residual gap that is not explained by the sum of the two effects is attributed to differences between the two estimated models in terms of other model parameters (see Table 4). In panel A, numbers in parentheses represent percentages of the job selection and the preference heterogeneity effects among the explained gap.

The difference between columns (4) and (1) in the welfare cost is 5.03%P which 'explains' about 80% of the total welfare cost gap between the two models, 6.3%P. As noted above, the 'residual' gap (1.27%P=6.3%P–5.03%P) is attributable to differences between the two estimated models in the other model parameters. This corresponds to 20% of the observed total welfare cost gap. In the second row, we conduct a similar exercise using the estimated homogeneous model. As shown in the last four columns, the decomposition results are somewhat different depending on which column is used ((2) or (3)). The takeaway from these results is that the majority of the overstated welfare cost in the homogeneous model is attributable to the preference effects which explains about 66% to 81% of the 'explained' gap (5.03%P) depending on which model and which column are used in the decomposition. The non-negligible minority (19% to 34%) is attributable to the selection effect.[30]

Understanding the job selection effect is easy, as individuals can reduce the welfare costs of the increased variances of wage shocks by a larger extent when they have more (job) choices. Why is it that the measured welfare cost is smaller in the world of heterogeneous, relative to homogeneous, preferences? It is important to notice that, in estimating the size of the 'preference effect' in panel A, we allowed each household to adjust all self-insurance measures (e.g., labor supply and borrowing/saving) following changes in the variance of wage shocks. These preference effects are restated in the first and third rows in panel B. In the second and fourth rows of panel B, we re-estimate the preference effects, assuming that households do not have any ability to adjust their labor supply and borrowing/saving when facing new variances of wage shocks, that is, they are forced to maintain their previous optimal levels of labor supply and borrowing/saving at the early 1970s levels. These 'would-be' preference effects appear universally negative, confirming our conjecture that the welfare gain of introducing self-insurance mechanisms to the world of no insurance is greater

[30] The relatively smaller role played by the job selection effect is partly related to the fact that the equilibrium share of the risky (or safe) sector does not change much in response to changes in the variances of wage shocks. Precisely, we find in Section 4 that, when both input and output prices change following the increased variance of wage shocks, the equilibrium share of the risky sector among the total workers increases by 2.2%P from 22% to 24.2%. With prices being fixed at the 1970s level, the share increases by 1.9%P from 22% to 23.9%.

when agents' risk preferences are heterogeneous compared to the homogeneous case. We can further illustrate this using the heterogeneous model, focusing on the preference effect derived by (4) – (2). With no ability to cope with the increased variances of wage shocks (the second row), the measured welfare costs are 14.67 and 16.44, respectively, under homogeneous and heterogeneous risk preferences. After allowing households to adjust labor supply and borrowing/saving behaviors in response to the increased wage uncertainty, the estimated welfare cost is reduced by a larger amount for the heterogeneous (7.82%P=16.44%–8.62%) than homogeneous (2.61%P=14.67%–12.06%) risk preferences, resulting in a smaller welfare cost in the heterogeneous risk preference economy. As emphasized previously, a more appropriate and realistic welfare evaluation should consider agents' insurability against the increased variance of wage shocks.

5.3 Effectiveness of self-insurance measures

In this subsection, we evaluate the relative effectiveness of self-insurance mechanisms in reducing the welfare costs caused by the increased wage uncertainty. We continue our focus on the partial analysis for this purpose. In panel B of Table 6, the estimated welfare cost is 7.03% in the heterogeneous agent-job model. This estimate is obtained by allowing households to adjust both of their labor supply and borrowing/saving behaviors in response to the increased variance of wage shocks. The estimate would be larger if households were not allowed to adjust their self-insurance behaviors in response to the increased wage uncertainty. Our strategy is, therefore, to force the simulated households to maintain their previous labor supply or borrowing/saving choices (made in the general equilibrium environment of the 1970s economy) even after they face the new levels of wage shocks generated by the increased variances. We then re-estimate the (higher) welfare cost by comparing the maximized utility level in the 1970s economy and the new counterfactual utility level, and compare the counterfactual welfare costs to those in panel B of Table 6. We conduct a parallel analysis using the homogeneous agent-job model. This is to test if, aside from the overstatement of the welfare cost, the conventional homogeneous agent-job model leads to a biased result regarding the relative effectiveness of self-insurance mechanisms.

5.3.1 Self-insurance from family labor supply adjustment vs. borrowing and saving

Table 8 compares family labor supply adjustment and borrowing/saving in terms of their effectiveness as a self-insurance mechanism. More specifically, they are compared in their abilities to reduce the welfare cost caused by the exogenous increase in the variances of wage shocks. Estimates in the first two columns are imported from Table 6, representing the welfare costs estimated with all insurance measures used by households. In columns (3) and (4), we re-compute the welfare costs with only the family labor supply fixed at the early 1970s level. In columns (5) and (6), we repeat the analysis with only borrowing and saving fixed at the early 1970s level. Then, we fix both measures at the early 1970s level, estimate the welfare costs, and report the results in columns (7) and (8). Numbers in parentheses represent the additional welfare *costs* of not using either family labor supply or borrowing/saving or both, relative to the case of allowing both mechanisms as in columns (1) and (2). For example, in the heterogeneous economy, an additional 3.83%P (=10.86%-7.03%) of the welfare cost is associated with not being able to adjust family labor supply to changes in the variances of the four types of wage shocks but still allowing adjustments of borrowing and saving. Numbers in brackets stand for additional welfare *gains* of using either or both measures, relative to the case of not allowing any of the insurance mechanisms, that is, those estimates in columns (7) and (8). For example, focusing on the heterogeneous economy, allowing the 1970s households to adjust only family labor supply in response to changes in the variances of the four types of wage shocks (estimate in column (6)) creates an additional welfare gain of 2.38%P (=12.47%-10.09%), relative to the case of no self-insurance (12.47% in column (8)).

Table 8: Effectiveness of Family Labor Supply and Borrowing/Saving as Insurance Mechanisms against Welfare Reduction

	Both Insurance Measures Are Adjusted %, (%P), [%P]		Family Labor Supply Fixed at the Early 1970s Level %, (%P), [%P]		Borrowing/Saving Fixed at the Early 1970s Level %, (%P), [%P]		Both Measures Are Fixed at the Early 1970s Level %, (%P), [%P]		Interactive Insurance Effects of Two Measures %P	
	Homo-geneous (1)	Hetero-geneous (2)	Homo-geneous (3)	Hetero-geneous (4)	Homo-geneous (5)	Hetero-geneous (6)	Homo-geneous (7)	Hetero-geneous (8)	Homo-geneous (9)	Hetero-geneous (10)
Δ All Types	13.33	7.03	21.10	10.86	19.31	10.09	24.69	12.47	2.39	1.45
of Shocks	–	–	(7.77)	(3.83)	(5.98)	(3.06)	(11.36)	(5.44)	–	–
(1)	[11.36]	[5.44]	[3.59]	**[1.61]**	[5.38]	**[2.38]**	–	–	–	–
Δ Male	7.35	3.49	11.49	5.66	10.19	4.94	12.98	6.46	1.35	0.65
Permanent	–	–	(4.14)	(2.17)	(2.84)	(1.45)	(5.63)	(2.97)	–	–
(2)	[5.63]	[2.97]	[1.49]	**[0.80]**	[2.79]	**[1.52]**	–	–	–	–
Δ Male	1.82	0.96	2.42	1.24	3.24	1.78	3.45	1.88	0.39	0.18
Transitory	–	–	(0.60)	(0.28)	(1.42)	(0.82)	(1.63)	(0.92)	–	–
(3)	[1.63]	[0.92]	[1.03]	**[0.64]**	[0.21]	**[0.11]**	–	–	–	–
Δ Female	3.79	1.88	5.91	2.77	5.25	2.53	6.68	3.16	0.69	0.25
Permanent	–	–	(2.12)	(0.89)	(1.46)	(0.65)	(2.89)	(1.28)	–	–
(4)	[2.89]	[1.28]	[0.77]	**[0.39]**	[1.42]	**[0.63]**	–	–	–	–
Δ Female	–0.71	–0.32	–0.57	–0.23	–0.43	–0.16	–0.37	–0.11	0.08	0.04
Transitory	–	–	(0.14)	(0.09)	(0.28)	(0.16)	(0.34)	(0.21)	–	–
(5)	[0.34]	[0.21]	[0.20]	[0.11]	[0.06]	[0.05]	–	–	–	–

Notes: See notes to Table 6. Numbers in parentheses represent additional welfare costs of not using an insurance measure (or measures), relative to the case of all insurance measures being fully adjusted. Those in brackets stand for additional welfare gains of using an insurance measure (measures), relative to the case of no adjustment in any insurance measure.

Several important findings emerge from Table 8. First, in all cases, measured welfare costs are higher for the homogeneous, relative to heterogeneous, agent-job model, re-confirming our previous findings in Table 6. The extent of overstated welfare costs by the representative worker-job model is independent of the availability of insurance measures. In most cases, the estimated welfare cost is approximately 2 times greater in the homogeneous economy than in the heterogeneous economy, with little variation in the ratio across different cases in Table 8. The results imply that all of the following observations are preserved in a qualitative sense even in the homogeneous agent-job model which is frequently adopted by existing studies. Subsequent discussions focus on our preferred model, the heterogeneous agent-job model.

Second, as shown in row (1), family labor supply adjustments are somewhat more effective than borrowing and saving in mitigating the welfare cost caused by changes in the variances of the four types of wage shocks. This can be seen in the greater welfare gain (or greater reduction in the welfare cost) from adjusting family labor supply than from borrowing and saving.

Third, more interestingly, there exists some interactive effect between the two insurance mechanisms in mitigating the welfare loss of the increased wage uncertainty. For example, when the economy is hit by the four types of wage shock changes, 5.44%P (=12.47%–7.03%) of additional welfare cost is associated with not using any insurance measure, relative to the case of using both insurance mechanisms. Of the total welfare gain (5.44%P), the 'direct' (or pure) effect of family labor supply adjustment is 2.38%P, as shown in the bracket of column (6). In contrast, borrowing and saving behavior alone creates 1.61%P of the welfare gain, as revealed in the bracket of column (4). The sum of these two direct effects amounts to 3.99%P, and the difference between the total gain (5.44%P) and the sum of the two direct effects (3.99%P) is called an 'indirect' or 'interactive' effect, which is 1.45%P, as shown in the last column.[31] That a joint introduction of the two mechanisms creates additional welfare gains

[31] Alternatively, this indirect effect of 1.45%P can be obtained by adding the additional welfare costs (3.83%P and 3.06%P) of fixing the two measures one by one at the 1970s level (as shown in the parentheses of columns (4) and (6)) and subtracting the total welfare gain of having both insurance mechanisms (5.44%P) from the sum (6.89%P).

of 1.45%P (or the insurance effect of each mechanism is greater when the other mechanism is allowed to be adjusted) is qualitatively consistent with Attanasio et al.'s (2005) finding of complementarity of the two measures.

Fourth, the relative effectiveness of an insurance measure is different depending on the type of wage shock under consideration. While family labor supply adjustment is more effective in mitigating the welfare costs caused by the increased variance of permanent wage shocks, the welfare cost of the increased variance of (male) transitory shocks is more effectively reduced by borrowing and saving. For example, when only family labor supply is allowed to be adjusted in the state where both measures are fixed at the early 1970s level, the welfare cost caused by the increased *male permanent* shock is reduced from 6.46% to 4.94%, producing 1.52%P of additional welfare gain. When only borrowing and saving are allowed, the welfare cost of this same shock is reduced only to 5.66%, producing a smaller relative welfare gain of 0.80%P. In contrast, the welfare cost of the increased variance of *male transitory* shock is reduced from 1.88% to 1.78% and 1.24% when family labor supply and borrowing/saving are introduced to the world of no insurance, respectively, resulting in an additional welfare gain of 0.11%P and 0.64%P.

5.3.2 Details on effectiveness of family labor supply adjustments

In Table 8, we analyze the effects of family labor adjustments as a whole without distinguishing an individual's own labor supply adjustment from his or her spouse's labor supply adjustment. To understand how labor supply adjustments are jointly made within each household, we decompose the total welfare effect of family labor adjustments between the effect of self-adjustment and the 'added' worker effect. For brevity, results are reported only for the heterogeneous model. We analyze only the case of increased variance of permanent shocks, as they are most effectively mitigated by family labor supply adjustments.

Table 9: Details on Effectiveness of Family Labor Supply Adjustment as an Insurance Mechanism in Heterogeneous Model

	Borrowing/Saving Adjusted					Borrowing/Saving Fixed				
	Both Insurance Measures Are Adjusted (%)	Self Labor Supply Adjustment Not Allowed (%)	Family Labor Supply Fixed (%)	Added Worker Effects (%P)	Effects of Self Labor Supply Adjustment (%P)	Added Worker Effect (%P)	Effects of Self Labor Supply Adjustment (%P)	Family Labor Supply Adjusted (%)	Only Spouse Labor Supply Adjusted (%)	Both Measures Are Fixed at the 1970s Level (%)
	(a)	(b)	(c)	=(c)–(b)	=(b)–(a)	=(f)–(e)	=(e)–(d)	(d)	(e)	(f)
Δ Male Permanent	3.49	4.37	5.66	**1.29**	**0.88**	**0.95**	**0.57**	4.94	5.51	6.46
Δ Female Permanent	1.88	2.19	2.77	**0.58**	**0.31**	**0.46**	**0.17**	2.53	2.70	3.16

Notes: See notes to Table 6.

Table 10: More on Wives' 'Added-Worker' Effects

(A) Borrowing/Saving Allowed

		Wives' Labor Supply Adjusted (%) (b)	Family Labor Supply Fixed at the 1970s (%) (c)	Added-Worker Effects (%P) (c)-(b)	Population Share	Added-Worker Effects × Population Share	Contribution to Added-Worker Effects (%)
Total Added-Worker Effect (in Table 9)	(1)	4.37	5.66	1.29 (12%Δ in labor supply)	1	1.29	100 (100)
Intensive	(2)	5.39	6.74	1.35	0.436	0.59	45.6 (35.9)
Extensive	(3)	14.9	33.83	18.9	0.037	0.70	54.3 (64.1)
Non-Work	(4)	2.80	2.80	0	0.525	0	0 (0)

(B) Borrowing/Saving Fixed

		Wives' Labor Supply Adjusted (%) (e)	Family Labor Supply Fixed at the 1970s (%) (f)	Added-Worker Effects (%P) (f)-(e)	Population Share	Added-Worker Effects × Population Share	Contribution to Added-Worker Effects (%)
Total Added-Worker Effect (in Table 9)	(1)	5.51	6.46	0.95 (17.5%Δ in labor supply)	1	0.95	100 (100)
Intensive	(2)	6.41	7.24	0.83	0.436	0.36	37.9 (37.7)
Extensive	(3)	22.99	34.56	11.57	0.05	0.59	62.1 (62.3)
Non-Work	(4)	3.00	3.00	0	0.512	0	0 (0)

Notes: See notes to Table 6. Intensive: Those wives who supply labor both before and after the increase in the variance of their husbands' permanent wage shocks. Extensive: Those wives who supply labor only after the increase. Non-Work: Those wives who always stay out of the labor force.

Table 9 summarizes the results. As in Table 8, we report the results under two scenarios: when borrowing/saving adjustment is allowed so that the effect of family labor adjustment includes both direct and indirect effects (estimates in the first five columns), and when adjustment of borrowing/saving is not allowed so that only the direct effect of family labor supply adjustment is analyzed (those in the last five columns).[32] We find that, when households are hit by increased variance of male permanent shocks, added-worker effects by wives play a greater role in mitigating the welfare loss relative to the effects of husbands' own labor supply adjustment. With or without the borrowing/saving mechanism, about 60% of the total welfare-improving effect of family labor supply adjustments is explained by the added-worker effect. Still, husbands' labor supply adjustment also plays some role in reducing the welfare loss. Estimates in the second row show that, when households are subject to the increased variance of female permanent shocks, for reasons stated previously, both estimated 'added-worker' effects by husbands and effects of wives' self-labor supply adjustments become generally smaller compared to the case when the households are hit by increased male permanent shocks. (It is worth noting that variances of male and female permanent shocks increased in similar magnitudes during our sample period.) Even in this case, however, added-worker effects by husbands play a greater role in reducing the welfare loss, relative to wives' self-labor supply adjustment.

5.3.3 More on added-worker effects

In Table 10, we further decompose the total added-worker effect observed in Table 9 into the effects of the extensive and intensive margins of spouses' labor supply adjustments. We discuss wives' added-worker effects only, as they are much greater than husbands' added-worker effects. In each panel, the first three estimates in the first row are imported from Table 9.

[32] Readers may verify that, in each scenario, the sum of the added worker effect and the effect of self-labor supply adjustment in Table 9 is identical to the total effect of family labor supply adjustments observed in Table 8. In Table 8, when the household is hit by an increased variance of male permanent shocks, the sum of the two effects (that is, the effect of family labor adjustment) is 2.17%P and 1.52%P with and without adjustment of borrowing/saving, respectively. These numbers are, respectively, in the parenthesis of row (2) and column (4) of Table 8 and the bracket of row (2) and column (6).

Focusing on the case of allowing the borrowing and saving behavior (panel A), these estimates show that wives' added-worker effect reduces the welfare cost of the increased variance of husband's permanent wage shocks by 1.29%P (from 5.66% to 4.37%). Rows (2) through (4) present how this additional welfare gain is generated by three groups of wives: those who work both before and after the increase in the variance of their husbands' permanent wage shocks (intensive); those who work only after the increase (extensive); and those who stay out of the labor force even after the increase (non-work). Estimates in the third column suggest that the welfare gain of allowing wives' labor supply adjustments is much larger for a typical household in the extensive group (18.9%P) than in the intensive group (1.35%P). Obviously, the welfare gain is zero for each household in the non-work group. The group-specific welfare gain, when multiplied by the corresponding group population share in the fourth column, produces the amount of each group's contribution to the total added-worker effect (1.29%P) in the fifth column. The last column converts these contributions to percentages. The results show that the extensive group contributes to the total added-worker effect by 54.3%.

Numbers in parentheses in the third and last columns represent corresponding labor supply adjustments. The total added-worker effect of 1.29%P comes from wives increasing their total hours (employment times average hours) by 12%. This amounts to an increase in per capita employment from 0.436 to 0.473 and an increase in the average working hours from 0.240 to 0.249. We previously noted that the estimated variance of male permanent wage shocks increased from 0.0185 to 0.0258. A back-of-the-envelope calculation shows that, when the percentage is calculated at the mean of each variable obtained before and after the increase in the variance of male permanent shocks, the cross-elasticity of the wife's precautionary labor supply with respect to the husband's permanent wage risk is estimated at 0.73. The last column shows that a 64.1% of the increase in wives' total hours (0.013) is accounted for by the extensive margin of wives' labor supply adjustments.[33] These findings

[33]Precisely, among the 'extensive' group of wives, per capita employment increases from 0 to 0.037, and the average hours increases from 0 to 0.231.

are qualitatively preserved even when borrowing/saving is not allowed in panel B. Most importantly, in terms of both the welfare cost and the amount of labor supply adjustment, more than 60 percent of wives' added worker effects is accounted for by the extensive margin of wives' labor supply adjustments. When borrowing/saving is not allowed, households depends more heavily on wives' added worker effects, increasing the cross-elasticity to 1.06.[34]

5.4 Changing Economic Environment

Economic environments have changed concurrently with the increased variances of wage shocks, including technological changes, gender-biased demand shifts, and changes in tax codes. We expect that these changes also affect households' welfare levels. For example, the rapid growth of information technology not only enhances average total factor productivity, but also reduces the durability of existing skills and knowledge, which may in turn result in greater wage volatility. In Table 11, we investigate how the welfare evaluation changes when these factors are considered in our preferred model. All the estimates represent the estimated welfare costs generated by changes in the four types of wage shocks. The estimate in the first row is borrowed from our previous analysis in panel B of Table 6. To consider the welfare effects of technological change, for example, we repeat the same analysis in the first row except that the externally-determined TFP value of the 1970s is replaced with that of the 2000s at the time the wage-shock-variances of the 1970s are replaced by those of the 2000s.[35] For the effects of gender-biased demand shifts, we replace the estimated values of

[34] Wu and Krueger (forthcoming) adopt a similar model as in the current study (but without preference and job heterogeneity) to quantify the welfare losses from idiosyncratic wage risk and find that both the cost as well as the demand for public insurance through progressive income taxation is reduced significantly in the presence of spousal labor supply adjustments. The current results are also in line with Pistaferri (2003) who adopts a *single-agent male* earner model without participation decision and finds that the effect of increased wage uncertainty on *intensive* labor supply adjustment is negligible. Even in our two-earner model, when households are hit by increased variance of male permanent shocks, the effects of husbands' own labor supply adjustment (including both intensive and extensive margins) play a smaller role in mitigating the welfare loss relative to wives' added-worker effects (see Table 9).

[35] Acemoglu (2002), for example, argues that technological change has been skill-biased, induced by the rapid increase in the supply of skilled workers. The current study, however, considers exogenous skill-neutral technological change in the analysis. Sharma et al. (2007) provide a groundwork for using a change in the total factor productivity as a proxy for skill-biased technological change by finding that the technological

$\lambda^{m,R}$ and $\lambda^{m,S}$ of the 1970s with those of the 2000s. Similar exercises are conducted using externally-determined capital and labor tax rates in Table 1.

Table 11: Consideration of Changing Other Economic Conditions

	Welfare costs (%)
Initial result: Panel B in Table 6	7.03%
All the economic conditions are changed to the 2000s levels	–1.08% (∇ 8.11%P)
Increase in TFP	1.85% (∇ 5.18%P)
Gender-biased demand shifts	5.08% (∇1.95%P)

Notes: Numbers in parentheses represent percentage point changes in the welfare cost relative to the initial estimate (7.03%) See notes to Table 5.

Our results show that the changing economic conditions generally create welfare gains. In the second row, we allow all four aforementioned economic conditions to vary at the same time the variances of the wage shocks change, which reduces the measured welfare cost by about 8.11%P. Consequently, despite the increased variances of wage shocks, a typical household of the 1970s economy would have enjoyed welfare gains due to concurrent changes in various economic conditions. The largest part of the welfare gain comes from the increased TFP, which alone contributes to the reduction of the welfare cost by 5.18%P. Gender-biased demand shifts also contribute to the reduction of the welfare costs to some degree, although to a lesser extent than the increased TFP. Although not reported for brevity, the reduced capital income tax rate and the increased labor income tax rate make little difference in the measured welfare cost of the increased variances of wage shocks.

progress explains 74% of the average TFP growth for the sample period.

5.5 Further Robustness Tests

5.5.1 Distribution of individual risk aversions

All our previous analyses are based on the distribution of individual risk aversion, $\widehat{\gamma}_i \sim$ *lognormal* (1.33, 0.79), that is derived from the NLSY79 sample. The derived distribution is based on predicted individual values for 25–84 stages of the life-cycle. Because the measured welfare cost depends on the distribution, we examine how the results change when a different distribution is adopted.

Although the level of risk aversion affects saving behavior even after retirement, in this exercise, we derive the distribution of individual risk aversion for 25–64 and apply it for the whole life-cycle stages up to 84. The resulting distribution is $\widehat{\gamma}_i \sim$ *lognormal* (1.08, 0.64). Excluding the most risk-averse stages of the life-cycle reduces the mean risk aversion from 5.62 to 4.08 and the variance from 0.79 to 0.64. We then reproduce the results in panel B of Table 6 using this new distribution. Although not reported in a separate table for brevity, the measured welfare costs are generally reduced. For example, when the 1970s households face the new variances of the four types of wage shocks that exist in the 2000s economy, the measured welfare costs are 8.63% and 4.91% in the homogeneous and heterogeneous economies, respectively. These are compared to respective estimates of 13.33% and 7.03% that are based on the entire distribution of risk aversion for 25–84. However, the extent of the overstated welfare cost by the conventional homogeneous agent-job model (the focus of the current research) survive this new exercise. For example, when all four types of wage shocks are considered, the ratio of the two welfare costs between the homogeneous and heterogeneous economies is approximately 1.8. This ratio is just marginally smaller than the corresponding estimate (1.9) in Table 6.[36] All the other results are also preserved in this new exercise.

[36]The ratio is reduced slightly from the ratio in Table 6 based on the full distribution, as the variance of risk aversion shrinks from 0.73 to 0.64, and therefore, agents become relatively more homogeneous in risk aversion in the heterogeneous economy.

5.5.2 Frisch elasticities

As noted in Section 4, we borrow the parameter values of Frisch elasticities from Heathcote et al. (2010) whose model generates Frisch elasticity of labor supply of 0.48 for males for 1967–2005; 1.77 for females in 1967 and 1.25 in 2005. While these values are generally consistent with empirical evidence from micro econometric analysis, the estimate for women is larger than the estimate obtained by Blundell et al. (2016a). For 1999–2009, their estimated model implies Frisch elasticities of 0.528 and 0.850 for men and women, respectively. Other things being held constant, the welfare cost of the increased variances of wage shocks would be larger if (women's) labor supply were less elastic. We therefore adopt Blundell et al.'s (2016a) smaller estimate for female labor supply elasticity and repeat the welfare analysis. Because Blundell et al.'s (2016a) estimate of 0.850 is derived for 1999–2009, it corresponds to 1.25 of female labor supply elasticity in Heathcote et al. (2010) for 2005. We therefore take the ratio of the two estimates (0.850/1.25); multiply Heathcote et al.'s estimate of 1.77 for 1967 by the ratio; and use this adjusted female Frisch elasticity (1.19) for the early 1970s. With the smaller estimate, we replicate all the welfare analyses from Table 6 through Table 11. Although not reported in additional tables, all our previous observations survive this new exercise. To repeat just a few, in Table 6, adopting a smaller Frisch elasticity for women slightly increases the measured welfare cost from 13.33% and 7.03% in the homogeneous and heterogeneous economies to 14.99% and 7.7%. The extent of overstated welfare cost by the homogeneous agent-job model remains virtually identical as 1.9 times. When Table 9 is reproduced with the smaller elasticity, the added-worker effects still appear greater than the effects of self-adjustment. In the analysis of Table 10, with the smaller elasticity estimate for women, the cross-elasticity of wives' precautionary labor supply with respect to the variance of husbands' permanent wage shocks is slightly reduced to 0.65 and 0.92 with and without adjustments of borrowing and saving, respectively. The changes are empirically unimportant. Also, with or without borrowing/saving, a majority (about 70%) of wives' added worker effect is accounted for by the extensive margin of wives' labor supply adjustments.

5.5.3 Perfect Foresight vs. myopic beliefs about future wages shocks

Up to this point, we have assumed that the 1970s households foresee perfectly how the variances of wage shocks will change in the future. In this subsection, we examine the other polar case: The 1970s households have no ability to predict the future wage shocks, and therefore, mistakenly believe that the variances of wage shocks they experience will remain unchanged in the future. This is done by making the 1970s households expect that life-cycle profiles of variances of wage shocks will remain at the 1970s level, but face the realized wage shocks that are generated under the 2000s wage structure. At each stage of their life-cycle, households are surprised by the new level of wage shocks generated by the wage shock variances that existed in the 2000s, but still optimize their behavior assuming that the variances of wage shocks from the next life-cycle stage will remain at the 1970s level. With these extremely myopic beliefs about the future wage shocks, the 1970s households would suffer from a greater welfare loss, compared to the case where the variances of new wage shocks were correctly anticipated. The true welfare cost will be in between the two polar cases.

Appendix J compares measured welfare costs/gains between the two polar cases. As we conjecture, surprising the 1970s households with higher variances of wage shocks than they expect makes the measured welfare cost greater for both the homogeneous and the heterogeneous economies. As the increased wage uncertainty is not anticipated, the households are unable to fully utilize their self-insurance mechanisms. For example, although not reported in a table for brevity, when we reproduce estimates in Table 10 under the assumption of myopic beliefs about the future shocks, the estimated cross-elasticity of the wife's precautionary labor supply with respect to the husband's wage uncertainty becomes 0.41 when borrowing/saving is allowed, compared to the estimate of 0.73 obtained from Table 10. Otherwise, all previous findings generally survive this new exercise in a qualitative sense. In particular, even in this myopic world, the welfare cost is significantly overstated by neglecting heterogeneous workers' risk choices, with the ratio of the measured welfare costs between the homogeneous and heterogeneous agent-job model being slightly greater than 2.

6 Conclusion

Using a general equilibrium model with incomplete markets in which married couples choose their life-cycle labor supply, consumption, and savings, we provide a quantitative assessment of the welfare costs caused by the increased wage uncertainty in the United States from the 1970s to the 2000s. Heterogeneous risk preferences, workers' self-selection into risky jobs, and gender differences in wage dynamics constitute unique features of the model.

The most important finding is that the welfare cost of the increased variances of wage shocks is significantly exaggerated by neglecting heterogeneity in risk preferences and workers' risk choices. The measured welfare cost is approximately twice as large in the homogeneous agent-job model as in our heterogeneous agent-job model. Both preference heterogeneity and self-selection into risky jobs play important roles in explaining the welfare cost gap between the two models, although the preference effect is somewhat greater than the job selection effect.

The welfare cost remains substantial at 7% (in lifetime consumption equivalent) even when increases in wage shocks are anticipated, heterogeneous workers self-select into risky jobs, and self-insurance mechanisms (family labor supply adjustments and borrowing and saving) are allowed in the model. Even in this 'idealistic' world, self-insurance mechanisms alone are insufficient for absorbing the entirety of the increased wage shocks, even though these within-household insurance mechanisms are well-functioning, which suggests that inter-family insurance mechanisms are necessary to fully insure against these shocks.

The good news is that most of our main results, including the relative effectiveness of insurance mechanisms, are robust in a qualitative sense with respect to the underlying assumptions about preference and job heterogeneity, expectations of future wage shock changes, and important model parameters. To summarize briefly, the welfare-improving effects are generally greater from family labor supply adjustments than borrowing and saving adjustments. While, compared to borrowing and saving, family labor supply adjustments are more effective in reducing the welfare costs caused by the increased variance of permanent wage shocks, the increased variance of transitory shocks are more effectively mitigated by borrowing/saving. What interests us more is the finding of the interactive effect of family labor supply adjustments and borrowing/saving in mitigating the welfare costs: Family labor supply adjustments can reduce the welfare costs of the increased variance of permanent shocks more effectively when borrowing and saving behavior is allowed. Evidence shows that when households are hit by increased male (female) permanent shocks, added-worker effects by wives (husbands) play a greater role in mitigating the welfare loss, relative to the effects of husbands' (wives') self-labor supply adjustment. We also find that a majority of the wives' 'added-worker' effect in response to the increased variance of husbands' permanent shocks is accounted for by the extensive margin of wives' labor supply adjustment. These insurance mechanisms become less effective when changes in permanent wage shocks are unanticipated.

Current findings suggest that labor and credit/financial market policies should be interacted to effectively reduce the welfare loss of households caused by the increased variance of wage shocks. The result that wives increase their labor supply substantially (participation in particular) in response to the increased variance of husbands' anticipated permanent wage shocks suggests that policies to enhance the predictability of permanent wage shocks, such as lengthening the advance notice period of mass layoffs, could be welfare-improving.

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Appendix

Appendix A Data

We use multiple sources of data, including the Panel Study of Income Dynamics (PSID), the National Longitudinal Study of Youth 1979 (NLSY79), and the Current Population Survey (CPS). First, the PSID is used to estimate trends in cross-sectional variances of permanent and transitory wage shocks by gender from 1970 to 2012. They are initially estimated up to 2014, but the last two estimates are dropped following the convention in the literature. Second, the NLSY79 for the period 1985 to 2014 is used to construct life-cycle profiles of variances of wage shocks by gender and by risk preference group. As explained in Appendix E in detail, the two sets of estimation results are carefully combined to generate observed life-cycle profiles of variances of wage shocks by gender and by preference group that hold for each of the early 1970s and the 2000s economies. Third, the CPS is used to obtain various empirical moments on wages and employment, such as per capita employment by gender, hourly wage rate in the risky, relative to the safe, sector, and the ratio of average female to male wage in each sector, that are used in estimating our model.

We use the data from the nationally-representative component of the PSID sample, the Survey of Research Center. We do not use the Survey of Economic Opportunity component of the PSID mainly because of the serious irregularities in the sample's selection (see the lengthy footnote 11 in Shin and Solon (2011)). We restrict our PSID sample to those who are married, between the ages of 25 and 64, and who work at least 100 hours per year. Following existing studies (e.g., Heathcote et al., 2010), we define the hourly wage rate as the ratio of total annual labor income to annual hours. Total labor income is a comprehensive earnings measure, which includes, in addition to wage and salary income, bonuses, overtime, tips, commissions, and the labor parts of farm and business income. We exclude, however, business and farm income from total labor income, as the PSID's treatment of business and farm income in total labor income has changed over the years, resulting in inconsistency in the total labor income variable over time. We also exclude imputations for missing values. For the NLSY79, we use the national random sample, excluding supplemental and military samples. As in the PSID, we use the average hourly earnings variable defined as total annual labor income divided by annual hours. Total labor income includes wages, salary, overtime pay, commissions, and tips from all jobs. Since the NLSY79 selected respondents who were 14 to 22 as of 1979, and since we use the NLSY up to the 2014 survey, respondents in our NLSY sample are at most 57 years old. We restrict our NLSY sample to those who are the ages of 25 and 57, and who work at least 100 hours per year. Wage rates are deflated by CPI-U-RS in 2013 dollars.

Finally, for the purpose of obtaining various empirical moments by job (or sector), we estimate wage volatility by detailed occupation group and determine the risk and safe sectors empirically. For this purpose, we use the PSID, but analyze the hourly wage rate of household heads collected from his/her main job. The average hourly earnings variable defined as the ratio of annual earnings to annual hours is not appropriate for this purpose, as the annual earnings often pertain to multiple jobs, and job codes are available only for the main jobs. See Section 4 for a detailed discussion of how to define the risky and safe sectors empirically.

Appendix B Estimation of Wage Processes

We first define the parameter vector, $\mathbf{\Theta} = \{\rho^g, \rho^P_{\eta^m,\eta^f}, \rho^T_{\varepsilon^m,\varepsilon^f}, \{(\sigma^g_{\eta t})^2, (\sigma^g_{\varepsilon_t})^2\}^{2014}_{t=1970}\}_{g\in\{m,f\}}$, where ρ^g represents the persistence of permanent shocks for gender g; $\rho^P_{\eta^m,\eta^f}$ and $\rho^T_{\varepsilon^m,\varepsilon^f}$ are the correlations of permanent and transitory shocks between husband and wife, respectively; and $(\sigma^g_{\eta t})^2$ and $(\sigma^g_{\varepsilon_t})^2$ are the year-specific variances of transitory and permanents shocks, respectively. These are the model parameters that appear in equations (7) and (8). As discussed in Section 2, these parameters are not directly estimated from the data, as estimated variances of wage shocks, for example, confound the effects of self-selection into employment with the effect of exogenous wage shocks. Our strategy is to estimate these parameters by

GMM and include the estimates in the set of target moments to be matched by the corresponding moments generated by our estimated model. Figure 1 displays the GMM estimates of $(\sigma^g_{\eta_t})^2$ and $(\sigma^g_{\varepsilon_t})^2$.

We use the PSID for the period of 1970–2014. Since the PSID switched the survey from annual to biennial beginning in 1997 (1996 income year), wage shock variances cannot be estimated for 1997, 1999, 2001, 2003, 2005, 2007, 2009, 2011, and 2013 income years. To resolve this issue, following Heathcote et al. (2010), we assume that the cross-sectional variance in a missing year is a weighted average of the two variances in neighboring years. To obtain empirical moments for GMM for each year t and each gender g, we group individuals in the sample into 10-year adjacent age cells. For example, the first age cell includes individuals between 25–34, the second 35–44, ..., all the way up to the last age cell including individuals ages 55-64. Then, using residual wages from equation (12), we compute the sample covariance between wages for individuals in an age group at time t and wages of the same individuals at time s for all t and s. Because not all individuals contribute to each moment, sample sizes vary across different moments. We also group households in a similar way to compute the empirical covariance of husbands' and wives' wages.

The GMM estimator solves the following minimization problem:

$$\widehat{\Theta}_{GMM} = arg\min_{\Theta} g(\Theta)' W g(\Theta), \tag{B.1}$$

where $g(\Theta) = M_d - M_m(\Theta)$; M_d is a vector of the stacked empirical moments; $M_m(\Theta)$ is the population moments; and W is a weighting matrix. Following Altonji and Segal (1996), we use an identity matrix for W. This procedure is typical in the literature (see Appendix 2 of Heathcote et al. (2010), among others, for a reference). Unlike existing studies, however, we use the GMM estimates as target moments to be matched with corresponding model moments, instead of using them as the estimated structural parameters.

The life-cycle profiles of the variances of wage shocks are estimated in a similar fashion except that all of the time subscripts (t) are replaced by age subscripts (j) in Θ, and the entire sample is regrouped into 1-year age cells.

Appendix C Definition of a Recursive Stationary Equilibrium

Let $S_\Omega \equiv \mathcal{A} \times \mathcal{J} \times \mathcal{V}^m \times \mathcal{V}^f \times \mathcal{E}^m \times \mathcal{E}^f$ be the state space, $B(S_\Omega)$ be the Borel sigma algebra on S_Ω, and $(S_\Omega, B(S_\Omega))$ be the measurable space, respectively. The probability measure of household over the measure space is μ, which is consistent with household behavior. Then, a recursive stationary equilibrium is a collection of decision rules, $\{c(\Omega;\gamma,d), a'(\Omega;\gamma,d), l^m(\Omega;\gamma,d), l^f(\Omega;\gamma,d)\}$ endogenous shares of households choosing different types of jobs and labor market participation, $\{s^1, s^2, ..., s^9\}$; value functions, $\{V(\Omega;\gamma,d)\}$; prices, $\{r, p^{m,R}, p^{f,R}, p^{m,S}, p^{f,S}\}$; aggregate capital, K; aggregate gender- and job-specific labor inputs, $\{H^{m,R}, H^{f,R}, H^{m,S}, H^{f,S}\}$; government spending, G; and stationary distribution $\mu(\Omega;\gamma,d)$ such that

1. The decision rules and value functions solve the household problem.
2. $\sum_{d=1}^{9} s^d = 1$ ($s^d = \int_{S_\Omega I_d=1} d\mu(\Omega;\gamma,d)$).
3. Factor prices are determined competitively:

$$\begin{aligned}
r &= \alpha Z\Big(\frac{H}{K}\Big)^{1-\alpha} - \delta, \\
p^{m,R} &= (1-\alpha)Z\Big(\frac{K}{H}\Big)^{\alpha} \frac{H\lambda^R\lambda^{m,R}}{\lambda^{m,R}H^{m,R} + (1-\lambda^{m,R})H^{f,R}}, \\
p^{f,R} &= (1-\alpha)Z\Big(\frac{K}{H}\Big)^{\alpha} \frac{H\lambda^R(1-\lambda^{m,R})}{\lambda^{m,R}H^{m,R} + (1-\lambda^{m,R})H^{f,R}}, \\
p^{m,S} &= (1-\alpha)Z\Big(\frac{K}{H}\Big)^{\alpha} \frac{H(1-\lambda^R)\lambda^{m,S}}{\lambda^{m,S}H^{m,S} + (1-\lambda^{m,S})H^{f,S}}, \\
p^{f,S} &= (1-\alpha)Z\Big(\frac{K}{H}\Big)^{\alpha} \frac{H(1-\lambda^R)(1-\lambda^{m,S})}{\lambda^{m,S}H^{m,S} + (1-\lambda^{m,S})H^{f,S}}.
\end{aligned} \tag{C.1}$$

4. Markets clear:

(a)

$$
\begin{aligned}
H^{m,R} &= H_1^m + H_2^m + H_3^m,\ H^{m,S} = H_4^m + H_5^m + H_6^m, \\
H^{f,R} &= H_1^f + H_4^f + H_7^f,\ H^{f,S} = H_2^f + H_5^f + H_8^f, \\
\textit{where } H_d^g &= \int_{S_\Omega \times \Gamma, I_d=1} \epsilon^g(j, \nu^g, \varepsilon^g; \gamma, d) l^g(\Omega; \gamma, d) d\mu(\Omega; \gamma, d),
\end{aligned} \tag{C.2}
$$

(b)

$$
K = \sum_{d=1}^{9} \int_{S_\Omega \times \Gamma, I_d=1} a(\Omega; \gamma, d) d\mu(\Omega; \gamma, d), \tag{C.3}
$$

(c)

$$
\begin{aligned}
&C + K' + G = ZK^\alpha H^{1-\alpha} + (1-\delta)K, \\
&\textit{where } C = \sum_{d=1}^{9} \int_{S_\Omega \times \Gamma, I_d=1} c(\Omega; \gamma, d) d\mu(\Omega; \gamma, d) \text{ is aggregate consumption.}
\end{aligned} \tag{C.4}
$$

5. The government budget constraint is satisfied:

$$
\begin{aligned}
&G + (1-\tau_l) b \sum_{d=1}^{9} \int_{S_\Omega \times \Gamma, j \geq j_R, I_d=1} d\mu(\Omega; \gamma, d) = \\
&\tau_k r \sum_{d=1}^{9} \int_{S_\Omega \times \Gamma, I_d=1} a(\Omega; \gamma, d) d\mu(\Omega; \gamma, d) + \tau_l (\sum_{g \in \{m,f\}, n \in \{R,S\}} p^{g,n} H^{g,n}).
\end{aligned} \tag{C.5}
$$

6. The probability measure is consistent with household decision rules:

$$
\mu'(\mathbf{S_\Omega}; \gamma, d) = \int_{S_\Omega \times \Gamma, I_d=1} Q(\Omega, \mathbf{S_\Omega}) d\mu(\Omega; \gamma, d) \text{ for all } \gamma \in \Gamma \text{ and } d \in \mathbf{D}, \tag{C.6}
$$

where $\mathbf{S_\Omega} \equiv (\mathbf{A} \times \mathbf{J} \times \mathbf{V^m} \times \mathbf{V^f} \times \mathbf{E^m} \times \mathbf{H^f})$ is the typical subset in $B(S_\Omega)$ and $Q(\Omega, \mathbf{S_\Omega})$ is the transition function such that

$Q(\Omega, \mathbf{S_\Omega}) = I_{\{j+1 \in \mathbf{J}, a(\Omega;\gamma,d) \in \mathbf{A}\}} Pr\{(\nu^m)' \in \mathbf{V^m}, (\nu^f)' \in \mathbf{V^f}, (\varepsilon^m)' \in \mathbf{E^m}, (\varepsilon^f)' \in \mathbf{E^f} | \nu^m, \nu^f, \varepsilon^m, \varepsilon^f)\xi.$

Appendix D Standard Deviation of Year-to-Year Residualized Wage Changes by Occupation Groups

Table D.1: Volatility by Occupation

Occupation	
Legal Occupations	0.2866
Healthcare Practitioners and Technical Occupations	0.2577
Management Occupations & Business Operations Specialists & Financial Specialists	0.2477
Construction Trades & Extraction Workers	0.2200
Sales Occupations	0.1980
Transportation and Material Moving Occupations	0.1971
Community and Social Services Occupations	0.1951
Personal Care and Service Occupations	0.1837
Arts, Design, Entertainment, Sports, and Media Occupations	0.1797
Education, Training, and Library Occupations	0.1687
Life, Physical, and Social Science Occupations	0.1539
Computer and Mathematical Occupations	0.1517
Protective Service Occupations	0.1498
Healthcare Support Occupations	0.1483
Architecture and Engineering Occupations	0.1386
Office and Administrative Support Occupations	0.1288
Installation, Maintenance, and Repair Workers	0.1277
Building and Grounds Cleaning and Maintenance Occupations	0.1271
Production Occupations	0.1239
Food Preparation and Serving Occupations	0.1196

Sources: The authors' calculation using the PSID for 1976 to 1997.
Notes: We use the hourly wage rate of household heads collected from his/her main job. The average hourly earnings variable is not suitable for this purpose, as annual earnings often pertain to multiple jobs, and job codes are available only for the main job held during the survey week. Using the hourly wage rate of household heads collected from his/her main job, we classify each respondent's main job according to the 20 occupation groups classified by the 2000 Census Occupational Classification System; pull all the year-to-year changes in residualized wages across individuals and years within each occupation group; and compute the standard deviation of the changes for each occupation group. To obtain residualized wage changes, we previously applied ordinary least squares estimation to the regression of changes in the logarithm of the hourly wage rate against a constant, education, age, and age square each year. The first four occupation groups are included in the risky sector, and the rest in the safe sector. For the risky sector as a whole, the estimated standard deviation is 0.2497. The corresponding figure for the safe sector is 0.1515.

Appendix E Derivation of Life-Cycle Profiles of Observed Variances of Wage Shocks by Gender and by Preference Group for the Early 1970s (2000s) Economy

To obtain the gender-preference-group-specific life-cycle profiles of observed variances of wage shocks for the early 1970s, we take the following two steps.

(**Step 1**)

Using the NLSY79 for 1985 to 2014, we derive the life-cycle profiles of variances of wage shocks by gender and by preference group. First, using the individual value of risk aversion previously described in the text, we classify individuals among four preference groups: those whose risk tolerance level is within the top 25% of the entire distribution, between 25% and 50%, between 50% and 75%, and the rest (the most risk averse group). Then, for each gender-preference group, we apply OLS to equation (12) and, using the OLS residuals, estimate the life-cycle profiles of variances of wage shocks for each gender-preference group by the standard generalized method of moments estimation described in Appendix B. Then, for each gender-preference group, we apply OLS to the regression of the estimated variance of wage shocks (permanent or transitory) against quartic polynomials in age, and (with some adjustments described in the second step) use the resulting estimated coefficients as target moments for the coefficients in equation (13). The process also produces empirical moments of persistence of permanent shocks and covariance of husband's and wife's permanent (as well as transitory) shocks.

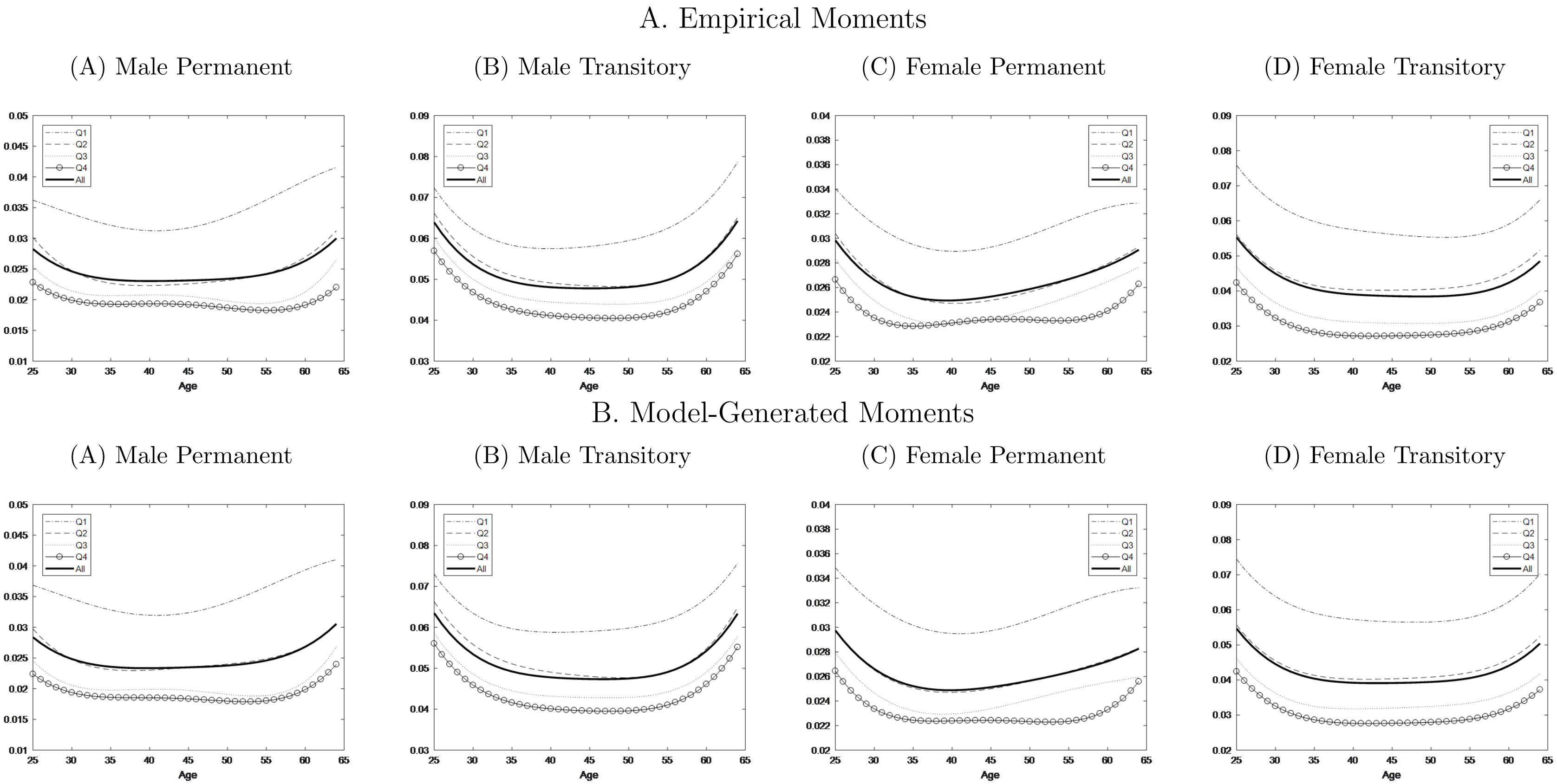
(**Step 2**)

The resulting estimated coefficients from the first step, however, represent the life-cycle profiles of observed wage shock variances that apply for the NLSY79 cohort, not for the cohort of the early 1970s economy. In Figure 1, we already obtained the observed variance of wage shocks for the early 1970s based on the PSID, but these results do not contain life-cycle information on the variance of wage shocks nor deliver variance of wage shocks by preference group. For each gender and type of shocks, Figure 1 displays a yearly estimate of variance of wage shocks that can be interpreted as the life-cycle *average* of observed variances for an agent with the *average* risk aversion. To derive the life-cycle profiles of observed variances of wage shocks for the early 1970s economy, we assume that the life-cycle *shape* of the variances of wage shocks is time-invariant and identical between the NLSY79 and the PSID.[37] Under the assumption, we shift the life-cycle profiles of wage shock variances obtained from the NLSY79 by the difference between the life-cycle average variance of wage shocks from the NLSY79 and the variance obtained from the PSID averaged for 1970 to 1974. To obtain the life-cycle average variance from the NLSY79, we pool all observations across preference groups, estimate a common life-cycle profile for each gender, and calculate the simple average of the estimated variances across different stages of the life-cycle. Practically, this process implies shifting only the estimated constant in the quartic function derived from the NLSY79. A similar procedure is followed to derive the observed wage shock profiles by gender and by preference group for the 2000s.

For the alternative (homogeneous) model, we follow the above two steps except that we focus on the empirical moments that are gender-specific but common to all preference groups.

[37]Indeed, Shin (2012) demonstrates that the life-cycle pattern of observed earnings volatility remains very similar between the PSID and the NLSY79. The gap between the two data sets in the estimated earnings volatility remains virtually identical across different stages of the life-cycle.

Appendix F Life-Cycle Profiles of Variances of Permanent and Transitory Wage Shocks by Gender and by preference Group, the 2000s Economy



Sources for panel A: The authors' estimation using both the PSID (1970–2014) and NLSY79 (1979–2014).

Notes: See the text for the derivation of empirical moments and corresponding model-generated moments. The solid line represents the life-cycle profile of variance of wage shocks in the homogeneous economy. In the heterogeneous economy, Q1 (alternate long and short dashed line), Q2 (dashed line), Q3 (dotted line), and Q4 (solid line connecting circular points) represent, respectively, the life-cycle profiles of those whose risk tolerance level is within the top 25% of the entire distribution, those between 25% and 50%, those between 50% and 75%, and the rest (the most risk averse group).

Appendix G Simulated Method of Moments

The model estimation processes for the 1970s and the 2000s are the same. Thus, we only describe the processes for the 1970s economy in this appendix.

Let M_d represent a vector of empirical moments that are computed from various data sources in the US, as described in Table 2. There are 33 and 102 empirical moments for the homogeneous and heterogeneous economy, respectively. Let $\mathcal{B}' \equiv \{\lambda^R, \lambda^{m,R}, \lambda^{m,S}, \chi^m, \chi^f, \beta,$ $\underline{a}, b, \delta_0^{g,\omega,n} - \delta_4^{g,\omega,n}\}$ represent the 54 structural parameters to be estimated internally. For the homogeneous economy, there are 30 structural parameters, $\mathcal{B}' \equiv \{\lambda^R, \chi^m, \chi^f, \beta,\ \underline{a}, b, \delta_0^{g,\omega} - \delta_4^{g,\omega}\}$. Given the externally determined parameter values, we obtain 100,000 households. We use the model to simulate their life-cycle labor hours, savings, and job choices, and generate the moments analogous to the empirical moments, denoted by $M_m(\mathcal{B})$ (again, 33 and 102 moments for the homogeneous and heterogeneous economies, respectively). Obviously, each of these model-generated moments is a function of a set of deeper structural parameters, $\mathcal{B}$. We define the column vector of deviations of the empirical moments from corresponding model-generated moments by $g(\mathcal{B}) = M_d - M_m(\mathcal{B})$. The simulated Method of Moments (SMM) estimator chooses the value of $\mathcal{B}$ that minimizes the weighted sum of the squared deviations between the empirical and model-generated moments:

$$\widehat{\mathcal{B}}_{SMM} = arg \min_{\mathcal{B}} g(\mathcal{B})' W g(\mathcal{B}), \tag{G.1}$$

whereW is some optimal weighting matrix. The variance-covariance matrix of $\widehat{\mathcal{B}}_{SMM}$ is estimated by

$$\widehat{\Sigma}_{\widehat{\mathcal{B}}} = (\widehat{G}' W \widehat{G})^{-1} \widehat{G}' W \Omega W \widehat{G} (\widehat{G}' W \widehat{G})^{-1}, \tag{G.2}$$

where $\widehat{G} = \frac{\partial}{\partial \mathcal{B}} g(\mathcal{B})|_{\mathcal{B}=\widehat{\mathcal{B}}}$, and Ω is the variance-covariance matrix of the empirical moments. We estimate the model with an optimal weighting matrix $= \Omega^{-1}$.

Appendix H Solution Algorithm (Heterogeneous Economy)

Given a set of model parameters, $\mathcal{B} \equiv \{\lambda^R, \lambda^{m,R}, \lambda^{m,S}, \chi^m, \chi^f, \beta, \ \underline{a}, b, \delta_0^{g,\omega,n} - \delta_4^{g,\omega,n}\}$ for $g \in \{m, f\}$, $\omega \in \{R, S\}$, we

1. Generate a discrete grid over the state space along with the risk aversion space and couple's choices of job types (9 cases), where we discretize a support of γ into 15, a into 50, j into 60, ν^m into 10, ν^f into 10, ε^m into 2, and ε^f into 2 points.

2. Guess the shares of males and females in the risky job and the safe job ($s_0^{r,m}, s_0^{r,f}, s_0^{s,m}, s_0^{s,f}$), interest rate ($r_0$), and gender-job-specific labor inputs ($H_0^{m,R}, H_0^{m,S}, H_0^{f,R}, H_0^{f,S}$); and we compute the other four prices using equation (C.1).

3. Solve the job-specific value functions and saving functions.

4. Compute the stationary distributions (μ) by simulating a large sample of 100,000 households.

5. Compute aggregates for capital and (gender-specific) labor as in equations (C.2)-(C.4).

6. Check asset and labor market clearing conditions and get ($s_1^{r,m}, s_1^{r,f}, s_1^{s,m}, s_1^{s,f}$, $H_1^{m,R}$,$H_1^{m,S}$, $H_1^{f,R}$, $H_1^{f,S}$), and r_1.

7. Update ($s_0^{r,m}, s_0^{r,f}, s_0^{s,m}, s_0^{s,f}$, $H_0^{m,R}, H_0^{m,S}, H_0^{f,R}, H_0^{f,S}$) and r_0 until markets clear and $s_1^{r,m} \approx s_0^{r,m}$, $s_1^{r,f} \approx s_0^{r,f}$, $s_1^{s,m} \approx s_0^{s,m}$, $s_1^{s,f} \approx s_0^{s,f}$, $H_1^{m,R} \approx H_0^{m,R}$, $H_1^{m,S} \approx H_0^{m,S}$, $H_1^{f,R} \approx H_0^{f,R}$, $H_1^{f,S} \approx H_0^{f,S}$, and $r_1 \approx r_0$.

8. Get the consumption functions.

9. Check the final goods market clearing condition and the government budget constraint.

For the homogeneous economy, the algorithm is much simpler in that we only solve for one type of value function and decision rule with a smaller state space.

Appendix I US Pension system

Following Heathcote et al. (2010), we employ the US social security benefit system in the model, which is referred to as the "primary insurance amount" (PIA). The PIA is the benefit a person would receive if he/she elects to begin receiving retirement benefits at his/her normal retirement age. At this age, the benefit is neither reduced for early retirement nor increased for delayed retirement. The PIA will be the sum of (a) 90% of the first 0.221 of his/her average indexed annual earnings, (b) 32% of his/her average indexed annual earnings over 0.221 and through 1.331, and (c) 15% of his/her average indexed annual earnings over 1.331.

Appendix J Perfect Foresight vs. Myopic Beliefs about Future Wage Shocks

	Perfect Foresight		Myopic Beliefs	
	Heterogeneous (%)	Homogeneous (%)	Heterogeneous (%)	Homogeneous (%)
Δ All Types of Shocks	7.03	13.33	9.22	19.68
Δ Male permanent	3.49	7.35	4.57	10.88
Δ Male transitory	0.96	1.82	1.25	2.70
Δ Female permanent	1.88	3.79	2.43	5.57
Δ Female transitory	–0.32	–0.71	–0.43	–0.91

Notes: See notes to Tables 6. Estimates in the first two columns are obtained by assuming that the 1970s households foresee perfectly how the life-cycle profiles of variances of wage shocks will change in the future. Those in the last two columns are derived under the assumption that the 1970s households mistakenly believe that life-cycle profiles of variances of wage shocks will remain at the 1970s level, but face the realized wage shocks that are generated under the 2000s wage structure.