Wage Volatility and Changing Patterns of Labor Supply*

Jay H. Hong[†], Byoung Hoon Seok[‡], and Hye Mi You[§]

April 2018

Abstract

Over the past few decades, the skilled-unskilled hours differential for US men increased when the skill premium rose sharply. This contrasts with previous studies suggesting dominant income effects. Based on PSID data, we show that skilled men experienced much larger increases in wage volatility than unskilled men. With the rise in volatility in the wage processes, our general equilibrium incomplete markets model replicates the increased hours differential despite the increased skill premium. We find that hours adjustment is important for self-insurance in the short run, whereas precautionary savings play a crucial role in the long run.

Keywords: Skill Premium, Wage Volatility, Labor Supply by Skill Group, Precautionary Savings

JEL Classifications: E24, J22, J24, J31

* We thank editor Dirk Krueger and three anonymous referees for their extremly insightful comments and suggestions. We are also grateful to Mark Aguiar, Mark Bils, Yongsung Chang, and participants at Rochester, Midwest Macro Meetings, Canadian Economics Association, NY/Philadelphia Workshop on Quantitative Macroeconomics, Shanghai Macro Workshop, European Meetings of the Econometric Society, Yonsei, Cleveland Fed, Seoul National Univerity, Sogang , Korea, Sungkyunkwan, and West Virginia for their valuable comments. Kyooho Kwon and Jinhee Woo provide excellent research assistance. This work is partially supported by Research Resettlement Fund for the new faculty of Seoul National University. All errors remain ours.

† Department of Economics, Seoul National University, Email: jayhong@snu.ac.kr

‡ Department of Economics, Ewha Woman's University, Email: bhseok@ewha.ac.kr

§ Corresponding Author: College of Economics and Finance, Hanyang University, Email: hyemiyou@hanyang.ac.kr

1 Introduction

It has been well documented in the literature that over the past four decades the skill premium, the relative wages of U.S. men with a college degree (skilled) to those without a college degree (unskilled), increased substantially (Katz and Murphy (1992); Krusell et al. (1994); Autor et al. (2008)). The substitution effect from a rise in the skill premium predicts skilled men to increase hours worked relative to unskilled men, while the income effect works in the opposite direction. In a recent survey, Saez et al. (2012) writes: "At the margin, substitution possibilities...can be captured by a compensated labor supply elasticity. With some notable exceptions, the profession has settled on a value for this elasticity close to zero for prime-age males." This implies that the income effect possibly governs the response of hours worked of skilled men relative to unskilled men to a rise in the skill premium,[1] that is, skilled men who face higher relative wages likely reduce their work hours relative to unskilled men. However, difference in hours worked between skilled and unskilled men actually increased in the U.S. when the skill premium rose most rapidly.

This paper explores the changing patterns of the second moment of wages, i.e. increased wage volatility, to resolve this discrepancy. As wages become more volatile, individuals under incomplete markets work longer hours as well as accumulate more precautionary savings for self-insurance. Flodén (2006) shows that an increase in future wage uncertainty raises current labor supply due to a precautionary motive in a simple two-period model. Pijoan-Mas (2006) also emphasizes the impact of wage volatility on labor supply decisions, by showing that both the level of precautionary savings and labor supply are higher in an economy with incomplete markets than in a complete-markets economy. If skilled men were exposed to larger increases in wage volatility than unskilled men, then this channel can potentially explain the observed rise in skilled-unskilled hours differential concurring with a rising skill premium.

This paper begins by estimating the time path of wage volatility *separately* for skilled

[1] A recent paper by Boppart and Krusell (2016) also shows that a utility function that features a dominant income effect is consistent with the long-run reduction in hours worked and changes in the other aggregate variables in the U.S.

and unskilled men using the data from the Panel Study of Income Dynamics (PSID). We document that there are marked differences in the trends of the second moment–as well as the first moment–of wages *across* skill groups over the past few decades. Our estimation shows that i) wage volatility (measured by the variance of residual wages) has risen for both skill groups between 1967 and 2000, consistent with the findings in the related literature and ii) skilled men have experienced *much larger increases* in their wage volatility compared to the unskilled, which is our contribution to the literature.[2]

In order to quantify the effect of the changing wage structure in explaining the skilled-unskilled hours differential, we develop a general equilibrium incomplete markets model with heterogeneous agents of different skill levels. Agents face uninsurable idiosyncratic wage shocks, which consist of a persistent and a transitory component drawn from a skill-specific distribution. In particular, the initial steady state of the model is calibrated to the 1967 U.S. economy. We then feed in the estimated wage processes from the PSID data to quantify how much the absolute and relative increase in wage volatility can explain the widening gap between skilled and unskilled hours.

The model can generate an increase in the skilled-unskilled hours differential during the transition, qualitatively consistent with the data. The greater rise in the volatility of wage rates for skilled men causes them to increase their hours worked relatively more to build up a larger stock of precautionary savings. This effect is strong enough to swamp the wealth effect from the rise in the skill premium. The model quantitatively over-predicts the increase in the relative labor supply of skilled men, which we believe is due to household labor supply decisions missing in this study. We argue that hours adjustment is important

[2] Although examining the causes of the rising wage volatility is outside the scope of this study, we view the increasing adoption of performance-pay contracts and deunionization in the U.S. labor market as important factors behind the increases in wage volatility. Lemieux et al. (2009) show that about 20% of the rise in the variance of male log wages is associated with performance pay contracts and that performance-pay contracts have been widely adopted in occupations of managers and sales and in finance, insurance, and real estate industries, regarded as skill-intensive compared to others. Card et al. (2004) claim that deunionization is an important contributor to the rise in male wage inequality. Acikgoz and Kaymak (2014) explore skill-biased technical changes as a potential explanation behind the decline in U.S. unionization rate.

for self-insurance in the short run, whereas precautionary savings play a dominant role in the long run. Once precautionary savings are built up, the wealth effect causes skilled men to reduce their relative labor supply.

As a validity check, we assess the model's implications for consumption and wealth. Based on nondurables consumption data from the Consumer Expenditure Survey (CEX), we document that the relative consumption of skilled to unskilled households continued to increase since the early 1980s. This path is well in line with the relative consumption of skilled men generated by the model. The model can also replicate the increasing trend in net worth of skilled households relative to unskilled households in both the PSID and the CEX data. In addition, our model's implications for changes in the cross-sectional variation of hours worked and the correlation between wages and hours are broadly consistent with what we observe in the data.

This paper is closely related to Heathcote et al. (2010), who explore the implications of observed changes in the wage structure (including a rising skill premium, a declining gender gap, and the increasing overall residual wage dispersion) for cross-sectional inequality in hours, earnings, and consumption. Our work differs from theirs in that we focus on the differences in the evolution of wage volatility *across different skill groups.* We explore the quantitative effects of the differences in increased wage volatility on the evolution of the relative hours worked of skilled to unskilled men.

We can also relate our paper to Erosa et al. (2014) and Castro and Coen-Pirani (2008) who examine differences in male labor supply by skill groups. However, Erosa et al. (2014) focus on their life-cycle features to study aggregate labor supply responses to changes in the economic environment. Their main purpose is to quantify the aggregate responses of labor supply to wage changes and evaluate the importance of the extensive versus intensive margins in this response. Instead, our study aims to explain changing patterns of relative hours worked using differences in the evolution of wage volatility across skill type. Our work also differs from Castro and Coen-Pirani (2008) in that they explore demand-side factors to explain changes in business cycle fluctuations of hours by skill groups,[3] whereas we consider supply-side factors to address the trends in hours by skill levels.

Several recent studies, including Santos (2014), Elsby and Shapiro (2012), Michelacci and Pijoan-Mas (2008), and Michelacci and Pijoan-Mas (2012) have tried to explain phenomena related to our motivating facts using a different mechanism. These studies all explore the effects of current experience/hours on future labor market outcomes as potential explanations for the phenomena, although the details of the mechanism vary. Santos (2014) considers a model in which current work hours affect future productivity in order to explain the increase in the correlation between hours and wages for the last quarter of the 20th century in the U.S. He shows that in the CPS data, this dynamic effect has become stronger for higher wage quintiles, whereas it has weakened for lower wage quintiles, leading to the recent rise in the hours-wages correlation. Elsby and Shapiro (2012) focus on the extensive margin of labor supply and explore changes in the returns to experience as driving forces behind the changing patterns of employment/nonemployment rates by education. They find that the return to experience for high-school dropouts has fallen significantly since the 1970s, which contributed to a downward trend in their employment rate. Michelacci and Pijoan-Mas (2008) and Michelacci and Pijoan-Mas (2012) exploit a search-matching framework to explain the link between wage inequality and hours worked. They assume that current hours of work affect the future probability of getting outside job offers. In this framework, greater wage inequality makes outside offers more attractive, inducing agents to work longer hours. Michelacci and Pijoan-Mas (2008) use this mechanism to explain the divergence in hours worked between the U.S. and Continental Europe, while Michelacci and Pijoan-Mas (2012) show that this model is consistent with the positive correlation between the increase in wage inequality and the rise in hours worked across occupation and industry groups in the U.S. census data. Relative to these studies, we propose an alternative mechanism based on a self-insurance motive. Our work shows that the evolution of the second moment of wages

[3]Castro and Coen-Pirani (2008) attempt to explain increased cyclicality of skilled hours since the mid-1980s and propose changes in capital-skill complementarity as potential explanations.

plays a quantitatively important role in explaining the recent changes in U.S. male hours worked by skill groups.

The remainder of this paper is organized as follows. Section 2 describes the stylized facts on changes in the U.S. wage structure and the trends in hours worked by skill group. In section 3, we describe a general equilibrium model with heterogeneous agents and incomplete markets. We then describe the calibration procedure in section 4. Section 5 presents main quantitative results and Section 6 discusses several sensitivity analyses. Section 7 then concludes the paper.

2 Data

This section documents how the U.S. wage structure and hours worked have changed by skill type over the past few decades. Our data of wages and hours worked are drawn from the Panel Study of Income Dynamics (PSID) 1967 through 2011. We also use the Current Population Survey (CPS) March Supplements over this period to document the relative wages and hours worked of skilled to unskilled men for comparison. In order to avoid issues associated with changing selection into labor force participation among women, we restrict our attention to men. Detailed sample selection criteria are provided in the Online Appendix.

2.1 Changes in the Wage Structure

We measure skill (e) based on one's educational attainment: we define a skilled worker (s) as one with a college degree or higher and an unskilled worker (u) as one without a college degree. We assume that an individual wage follows a skill-specific process that depends on a time-varying skill price per efficiency unit of labor, years of experience, and persistent and transitory labor productivity shocks. Specifically, the hourly wage of individual i, aged a at time t and skill type $e \in \{s, u\}$ is assumed to follow

$$\ln w^e_{iat} = \beta^e_t + f^e(X_{it}) + y^e_{it}, \tag{1}$$

where β_t^e is a skill-specific time dummy, f^e is a skill-specific return to experience (assumed to be a cubic polynomial of X_{it}), $X_{it} \equiv a_{it} - S_i - 5$ is potential experience with S_i representing years of schooling, and y_{it}^e is the log wage residual. We assume that the log wage residual y_{it}^e consists of a persistent and a transitory component:

$$y_{it}^e = \mu_{it}^e + \upsilon_{it}^e + \theta_{it}^e,$$

where μ_{it}^e is a persistent component, $\upsilon_{it}^e \sim (0, \lambda_t^{e,\upsilon})$ is a transitory component, and $\theta_{it}^e \sim (0, \lambda^\theta)$ is measurement error. The variance $\lambda_t^{e,\upsilon}$ of the transitory component υ_{it}^e is allowed to vary over time, whereas the measurement error's variance is time-invariant and independent of skill type. The persistent component μ_{it}^e is modelled as an AR(1) process:

$$\mu_{it}^e = \rho^e \mu_{it-1}^e + \eta_{it}^e,$$

where ρ^e is the persistence and $\eta_{it}^e \sim (0, \lambda_t^{e,\eta})$ is the persistent wage shock that has a time-varying variance of $\lambda_t^{e,\eta}$. The initial value of persistent component is drawn from a time-invariant, skill-specific distribution: $\mu_0^e \sim (0, \lambda^{e,\mu})$. We assume that all four variables, υ_{it}^e, θ_{it}^e, η_{it}^e, and μ_0^e are orthogonal and i.i.d. across individuals. We focus on time effects in this specification because empirical evidence suggests that time effects rather than cohort effects have been important in explaining rising wage inequality in the U.S. for recent decades (see Heathcote et al. (2005)).

With this statistical model, we estimate the wage processes of skilled and unskilled men separately in two stages. In the first stage, we run an Ordinary Least Squares (OLS) regression of log hourly wages on a time dummy and a cubic polynomial of potential experience by skill type, as shown in Equation (1). The residuals from these first-stage OLS regressions are exploited to estimate the parameters governing persistent and transitory wage shocks. Since a transitory shock in wages disappears after one period, autocovariances of individual log wage residuals depend solely on the persistent component, which help us to identify a persistent component from a transitory one.[4] In implementing the estimation procedure, we

[4] A large literature in labor and macroeconomics has been devoted to estimating labor income processes.

compute sample autocovariances of the residual y^e_{it} from the first-stage regressions of all possible orders for each age group in every survey year. According to the specifications above, the autocovariances of log residual wages can be written as functions of model parameters including ρ^e, υ^e_{it}, θ^e_{it}, η^e_{it}, and μ^e_0. The parameters are then estimated by minimizing the equally weighted distance between the sample autocovariances and the model counterparts.[5] To disentangle a transitory component from a measurement error, we use an estimate of 0.02 for the measurement error taken from French (2004).

We define the skill premium as the mean wage of skilled men relative to the mean wage of unskilled men:

$$\text{skill premium}_t = \overline{w}^s_t / \overline{w}^u_t,$$

where $\overline{w}^e_t$ denotes the average hourly wage in period t for skill type e. Figure 1 plots the trends in the male skill premium since 1967, using data from the PSID and CPS. Both data show similar trends in the skill premium over the sample period, although the timings differ slightly. As many previous studies document, the skill premium declined from the beginning of the sample period to the mid-1970s and increased sharply thereafter.[6]

[Put Figure 1 here]

Table 1 reports the estimated variances of persistent and transitory wage shocks of each skill type from the second stage regression. To better present the trends in these variances,

---

As Krueger et al. (2010) summarize, these studies often find that the estimates of persistent and transitory variances, based on the covariances of the levels of log earnings, are significantly different from those based on their first differences. A recent paper by Daly et al. (2014) attempts to reconcile these different estimates and finds that "outlying" earnings observations either prior to or following a missing observation are misspecified in standard labor income processes used in this literature. However, they find that without dropping such "outlying" observations, one can use the estimates for persistent variances based on the levels of log earnings because the biases are quantitatively small. We find that a time-varying persistent wage component drives most of important quantitative results in our model while the impacts of a transitory wage component are quantitatively small.

[5] Details of the estimation procedure are provided in the Online Appendix.

[6] The differences in the skill premium between the two datasets began to increase around the mid-1990s, which is due to a reduction in the representativeness of the PSID pointed out in the literature (Gouskova (2014)).

we depict their paths in Figure 2. The estimated variance ($\lambda_t^{s,\eta}$) of a persistent wage shock for skilled men rose substantially for the sample period. In particular, they experienced sharp rises in the variance of their persistent wage component during the 1980s and 1990s. The estimated variance ($\lambda_t^{u,\eta}$) of a persistent wage shock for unskilled men also increased over the sample period, but by much smaller than skilled men. The variance for unskilled men rose mostly during the 1970s and stayed roughly constant thereafter until it rose again in the mid 2000s.[7]

[Put Table 1 here]

[Put Figure 2 here]

The estimated variances ($\lambda_t^{e,v}$) of a transitory wage shock also exhibited different patterns by skill groups. Skilled men experienced a mild increase in the variance of a transitory wage shock during the 1980s and faced a sudden increase in its variance during the 1990s. Conversely, the bottom right panel of Figure 2 indicates that the variance of a transitory wage shock of unskilled men increased mostly during the first two decades and stagnated thereafter.

[Put Figure 3 here]

These changes in the variances of persistent and transitory wage shocks ultimately raised the volatility of residual wages faced by both skill groups. Figure 3 reports the variance decomposition of log wage residuals by skill type, based on the estimated parameters. In order to show the trend more easily, we normalize the total residual wage variance of skilled men to 1 in 1967. The left panel shows that the total variance of residual wages of skilled men more than doubled over the sample period. The rise in wage volatility occurred mostly during the 1980s and 1990s. During the 1980s, the rise in the variance of the persistent

[7]Blundell et al. (2008) estimate time-varying variances of permanent income shocks during the 1980s using the PSID data, and find that college graduates experienced continued growth in the variance of their permanent income shocks, whereas the variance in the permanent income shocks declined for non-college graduates towards the late 1980s.

component caused the total variance to increase, while the increasing variances of both persistent and transitory shocks are responsible for the rise in the total variance of residual wages during the 1990s. Over the same period, unskilled men also experienced a large increase in the variance of their residual wages. Their total residual wage variance reached about 50% higher by 2000 but to a level lower than skilled men. Unlike skilled men, unskilled men's wage volatility increased most rapidly between 1967 and the mid 1980s. Comparing both skill groups, we conclude that skilled men experienced a much larger increase in their residual wage volatility (measured as the estimated variance of log wage residuals) for the sample period than unskilled men.

One possible explanation behind the greater rise in wage volatility for skilled men than unskilled men is that the increasing share of college graduates caused more recent college graduates to be drawn from the bottom of the ability distribution.[8] In this case, the cross-sectional variation of residual wages may widen without any changes in wage volatility faced by individuals. However, we argue that the estimation results are not mainly driven by these changes in labor composition by education. Firstly, the estimation exploits a panel dimension to identify time-series variations of wage volatility from the cross-sectional variation of wages. Secondly, the fraction of college graduates increased most rapidly during the 1970s and grew only slowly thereafter, whereas wage volatility among skilled men changed little during the 1970s and rose sharply since the early 1980s.[9]

We also examine the possibility that wage volatility appears to have risen due to increased unemployment risks. Considering that wages tend to decline after a spell of unemployment, frequent unemployment spells may increase variability of observed wages. In the PSID data, we did not find any distinct trends in the transition probabilities from employment

[8]According to the CPS data, the share of college graduates was about 15% in 1967 and rose to 31% in 2006.

[9]Carneiro and Lee (2011) examine the effect of distributional changes in the context of the skill premium. Specifically, they computed the time path of skill premium adjusted for changes in quality of new college graduates and compare it with the raw skill premium. They found that the distance between the two increased the most during the 1970s when the share of college graduates rose most rapidly.

to unemployment for either skill group, particularly during the 1980s and 1990s. Thus, the estimated rises in residual wage volatility for both skill groups are not likely due to changes in unemployment risks.

[Put Figure 4 here]

To summarize, there have been substantial changes in the U.S. wage structure between 1967 and 2010. It is well known that the skilled premium rose sharply, in favor of skilled men. On top of that, wage volatility also increased by different magnitudes across skill groups. While we estimate wage processes using data up to 2010, we use the estimates up to 2000 for our quantitative evaluation. We make this choice because we have a smaller number of empirical covariances available to identify the variances of persistent wage shocks towards the end of the sample period; hence, the estimates for more recent survey years are relatively less accurate. Figure 4 depicts the Hodrick-Prescott (HP) filtered time series of the variances of persistent and transitory wage shocks by skill group as well as the skill premium used for our quantitative exercises.

2.2 Trends in Hours Worked

In this section, we document the trends in hours worked by U.S. men by skill group using the PSID. In particular, we focus on changes in labor supply at the intensive margin. We make this choice because the second moment of wages tends to have a larger impact on hours adjustment along the intensive margin rather than the extensive margin. Individuals facing more volatile wages may find staying in the labor market less attractive, yet they do not withdraw from the labor market unless they have very large wealth or are old enough to retire.[10] Most workers who do not fall in this category tend to increase their work hours to build up a larger stock of precautionary savings. Motivated by rising wage volatility

[10] Low et al. (2010) consider the impact of wage risk on employment rates by doubling the standard deviation of annual earnings growth and find that it had little effect on employment rates of workers younger than 55.

for U.S. male workers, this paper naturally focuses on the intensive margin of hours worked. Moreover, the participation margin of labor supply is closely related to unemployment shocks and retirement, which are outside the scope of this paper.[11]

[Put Figure 5 here]

The left panel of Figure 5 plots the trends in the average weekly hours (annual hours worked divided by 52 weeks) of U.S. men who worked at least 260 hours in the PSID and CPS over the 1967-2010 period. In order to clean out the effect of demographic changes on the trends in the average hours worked, we depict the trends in hours, holding the age and race composition constant.[12] The average weekly hours show similar patterns in both datasets. Although unskilled men's weekly hours in the PSID are about one hour less than those in the CPS throughout the sample period, both datasets share the same trends in unskilled men's hours. Both skill groups reduced their weekly hours during the 1970s[13] and began to increase their hours worked in the early 1980. This increasing pace continued until 2000 when both skill groups decreased labor supply again.

The right panel presents the differences in the average weekly hours between skilled and unskilled men over the same period. In the PSID, skilled men on average worked about 1

[11]There are many previous studies on changes in U.S. labor supply at the extensive margin (see Juhn (1992), Juhn et al. (2002), Juhn and Potter (2006), and Elsby and Shapiro (2012)). These studies document that nonemployment rate of U.S. men increased significantly over the past few decades with it concentrated among less-skilled men. They also find that the early retirement rate has increased and unemployment duration has become longer. Suggested explanations for the phenomena include a reduction in real wages of less-skilled men, an expansion of disability benefits program, and a decline in returns to experience.

[12]Specifically, we classify men in each skill group by age and race, and compute the average share of each age and race group for the first 5 years, and use the shares to obtain the weighted average of weekly hours in each skill group as if the age and race composition stayed constant over time.

[13]In contrast with our finding, **?** show that based on U.S. Census data, male hours worked per worker increased slightly during the 1970, while hours worked per person declined for the same period. This discrepancy is attributable to a difference in data treatment. In their study, workers are classified as employed based on a question about work status. Instead, we use the 260 hours cut for the employed. Those who were included in our sample because they worked more than 260 hours, but were classified as unemployed in **?** because they were not employed at the time of survey contribute to the decline in male hours worked during the 1970s in Figure 5.

hour more than unskilled men in the beginning of the sample period. The hours differential rose to about 2.5 hours in the early 1980s and declined only slightly thereafter. The hours differential around 2000 is about one hour larger than in the early 1970s. The similar pattern also appears in the CPS. In the CPS, the initial increase in the skilled-unskilled hours differential continued for a longer period until the early 1990s, and then began to decline. However, the hours differential around 2000 is still larger than it was in the beginning of the sample period in both datasets.

The increase in the hours gap between skilled and unskilled men in the last few decades of the 20th century has been documented in the literature. For example, based on five waves of American Time Use Surveys (ATUS), Aguiar and Hurst (2007) show that less-educated men reduced their time spent on total market work relative to those with a college degree or higher over 1965-2003.[14] Additionally, Costa (2000) and Santos (2014) find a similar pattern from the Current Population Survey (CPS) using wage as a measure of skill.[15]

Assuming that age effects are time-invariant, the rise in the relative hours of skilled to unskilled men may be due to either time effects (skilled-unskilled hours differential has increased over time within every age group) or cohort effect (hours gap between skilled and unskilled men is larger among younger cohorts than older cohorts). We claim that time effects rather than cohort effects have been more important in the rise in the skilled-unskilled hours differential, as is the case for increased wage volatility. To illustrate this point, consider a model in which the skilled-unskilled hours differential (hd) depends on age (a), time (t), and

[14]According to Aguiar and Hurst (2007), hours worked of all men (including both extensive and intensive margins) declined substantially between 1965 and 2003, while Figure 5 shows that hours worked of both skilled and unskilled men in the early 2000s are no less than those in the beginning of the sample period. This is because Aguiar and Hurst (2007) look at all men, while we focus on employed men. Despite this difference in the trends in hours between all men and employed men, the trends in the skilled-unskilled hours differential show the same pattern.

[15]Costa (2000) shows that the length of a work day by higher-wage earners has increased relative to that of lower-wage earners between 1973 and 1991 in the CPS. Santos (2014) confirms that this finding holds for the CPS data extended to 2006.

cohort ($c = t - a$), where these three components are additively separable:

$$hd(a, t, t - a) \equiv g_1(a) + g_2(t) + g_3(t - a).$$

As in Heathcote et al. (2005), we compute changes in the hours differential within cohort, within age, and between age groups to identify time and cohort effects. If we track changes in the hours differential of the same cohort $c = t - a$ between t and $t+1$, the change excludes cohort effects and captures age and time effects:

$$\Delta hd^c_{t,t+1} = g_1(a+1) - g_1(a) + g_2(t+1) - g_2(t) = \Delta g_1(a) + \Delta g_2(t).$$

The average of within-cohort changes in the hours differential across different cohorts is then given by $\overline{\Delta hd^c_{t,t+1}} = \Delta g_1(a) + \overline{\Delta g_2(t)}$. If the average varies by subperiods, it is attributable to time effects because age effects are time-invariant.

Instead, following the hours gap between skilled and unskilled men within the same age group a between t and $t+1$ sorts out age effects and reflects the sum of time and cohort effects:

$$\Delta hd^a_{t,t+1} = g_2(t+1) - g_2(t) + g_3(t+1-a) - g_3(t-a) = \Delta g_2(t) + \Delta g_3(t-a).$$

Note that the correlation between within-cohort changes and within-age changes is mainly determined by the strength of time effects. Taking the average of within-age changes across different age groups yields $\overline{\Delta hd^a_{t,t+1}} = \overline{\Delta g_2(t)} + \overline{\Delta g_3(t-a)}$.

Lastly, variations in the skilled-unskilled hours differential between ages a and $a+1$ in period $t+1$ indicate the difference between age and cohort effects, while removing time effects:

$$\Delta hd^{t+1}_{a,a+1} = g_1(a+1) - g_1(a) - (g_3(t+1-a) - g_3(t-a)) = \Delta g_1(a) - \Delta g_3(t-a).$$

The average between-age group change is given by $\overline{\Delta hd^{t+1}_{a,a+1}} = \Delta g_1(a) - \overline{\Delta g_3(t-a)}$. Any variations in the average between-age group changes across different subperiods can then be ascribed to time-varying cohort effects.

[Put Table 2 here]

Table 2 presents the average annual changes in the skilled-unskilled hours differential within cohort, within age, and between age groups (within-period) for various subperiods. The results show that within-cohort changes in the relative labor supply are statistically different across sub-periods. However, between-age groups variations are not statistically different from zero except for the first five years in the sample. Given that the age effects are time-invariant, whereas both time and cohort effects are time-varying, the results imply that time effects are strong and cohort effects are not.

In addition, since both within-cohort and within-age changes in the skilled-unskilled hours differential have time effects in common, the correlation between the two should be strong if the time effects are strong. The correlation is very close to one as shown in the bottom of Table 2. To the contrary, the correlation between within-age and between-age variations is close to zero, implying that cohort effects are weak. These results suggest that time effects rather than cohort effect have been more important in the trends of the skilled-unskilled hours differential. Based on these empirical evidences, our study focuses on how hours worked by skill level change over time in response to rising wage volatility and abstracts from life-cycle features of hours.

3 The Model

To understand how our empirical findings are connected, we develop a general equilibrium heterogeneous agent model with incomplete markets. Below we explain elements of our model in a greater detail.

3.1 Households

There are two types of single-person households–skilled and unskilled–in this economy.[16] We denote the skill type by e where $e \in \{s, u\}$ and the population share of skilled households at period t is given by π_t.[17] Each household is endowed with one unit of time in each period and faces a skill-specific process for his idiosyncratic productivity. In the beginning of a period t, a household of skill type e draws persistent and transitory components (μ_t^e and υ_t^e, respectively) of his productivity, x_t^e, from his type-specific distribution. After observing his productivity draw, the household decides whether to be employed or non-employed. If employed, he provides $h_t x_t^e$ efficiency units of labor and gets paid w_t^e per efficiency unit of labor, where h_t is hours worked chosen by this household. There is a minimum requirement for the number of hours such that each household should work at least $\underline{h}$ when employed. If this household is non-employed, then his hours of work h_t is zero, and he receives no wage income.

In this economy, only a risk-free asset is available for saving and no borrowing is allowed, which make markets incomplete. Households insure themselves against bad shocks through *both* precautionary savings and labor supply decisions.

At the end of the period, households face a survival probability of γ: a fixed fraction γ of the population survives to the next period and the remainder is replaced by newborn households. We make this choice because it allows the model to generate reasonable predictions

[16]By considering single-person households, we ignore intra-household transfers. In recent decades, labor force participation rates have risen more markedly among highly-educated women while the correlation between husband's and wife's education levels has become stronger (see Greenwood et al. (2014)). These trends may have affected the skilled-unskilled hours differential through income-pooling, intra-household bargaining, etc. While we understand the importance of these channels, we focus on the role of wage process in male hours worked in this paper. To our relief, trends in the skilled-unskilled hours differential among married men are very similar to those of non-married men in the data.

[17]We abstract from educational choice by households to focus on differences in labor supply by skill level. According to the CPS March supplements, the fraction of individuals with a college degree has increased rapidly until 1970s and the increase has slowed down afterwards. The extent that this change in acquiring a college education affects labor supply decisions is unaccounted for in our analysis, thus, our quantitative results may be biased.

on hours and saving without an explicit life-cycle, which adds age to the state space. Out of these newborn households, a $\tilde{\pi}_t$ fraction is assumed to be born as skilled, thus the evolution of total number of skilled workers is given by $\pi_{t+1} = \gamma\pi_t + (1-\gamma)\tilde{\pi}_t$. Unclaimed assets of deceased households are collected and redistributed to all households in a lump-sum transfer, T_t. Households pay income tax, $\Omega\left(r_t a_t + w_t^e h_t x_t^e\right) \equiv r_t a_t + w_t^e h_t x_t^e - \kappa\left(r_t a_t + w_t^e h_t x_t^e\right)^{1-\tau}$, where parameters τ and κ governs the progressivity and level of the tax schedule, respectively. A government fund its expenditures, Φ_t, using the tax revenue.

Each household maximizes the expected lifetime utility from a stream of consumption c_t net disutility from hours worked h_t:

$$\mathbb{E}\sum_{t=0}^{\infty}(\beta^e\gamma)^t u(c_t, h_t),$$
$$u(c_t, h_t) = \frac{c_t^{1-\sigma}}{1-\sigma} - \psi^e\frac{h_t^{1-\nu}}{1-\nu} \text{ and } \beta^e\gamma \in (0,1), e \in \{s, u\}$$

where β^e is a skill-specific time discount factor, ψ^e is a skill-specific constant, σ is the inverse of elasticity of intertemporal substitution with respect to consumption, and ν is the same but with respect to hours worked. Consumption and leisure are assumed to be additively separable in a constant relative risk aversion (CRRA) utility function, which is common for both skill types. In period t, households choose consumption c_t, hours of work h_t, and asset holdings a_{t+1} for the next period subject to

$$\begin{aligned} c_t + a_{t+1} &\leq a_t(1+r_t) + w_t^e h_t x_t^e + T_t - \Omega\left(r_t a_t + w_t^e h_t x_t^e\right), \\ c_t &\geq 0, a_{t+1} \geq 0, h_t \in \{0\} \cup [\underline{h}, \overline{h}], \end{aligned}$$

where r_t is real interest rate on the risk-free asset and w_t^e is real wage per efficiency unit of labor for skill type e. In order to incorporate a skill premium into the model, we consider two different real wages by skill types. The logarithm of the idiosyncratic productivity x_t^e of skill type e is specified as the sum of a persistent and a purely transitory component (μ_t^e and υ_t^e, respectively), given by

$$\ln x_t^e = \mu_t^e + \upsilon_t^e,$$

where $\upsilon_t^e \sim N(0, \lambda_t^{e,\upsilon})$. The persistent component is modelled as an AR(1) process:

$$\mu_t^e = \rho^e \mu_{t-1}^e + \eta_t^e,$$

where $\eta_t^e \sim N(0, \lambda_t^{e,\eta})$. These specifications allow for differences by skill types in the persistence of productivity shocks and in the variances of both persistent and transitory components. The variances of both persistent and transitory productivity shocks are time-varying in order to capture changing patterns in wage volatility seen from the PSID data. The recursive formulation of a household problem of skill type e is then given by:

$$V_t^e(a_t, \mu_t^e, \upsilon_t^e; G_t) = \max_{c_t \geq 0, a_{t+1} \geq 0, h_t \in \mathcal{H}} \left\{ u(c_t, h_t) + \beta^e \gamma \mathbb{E}_t V_{t+1}^e(a_{t+1}, \mu_{t+1}^e, \upsilon_{t+1}^e; G_{t+1}) \right\},$$

subject to

$$\begin{aligned}
c_t + a_{t+1} &\leq a_t(1 + r_t) + w_t^e h_t x_t^e + T_t - \Omega\left(r_t a_t + w_t^e h_t x_t^e\right), \\
\ln x_t^e &= \mu_t^e + \upsilon_t^e, \\
\upsilon_t^e &\sim N(0, \lambda_t^{e,\upsilon}), \\
\mu_t^e &= \rho^e \mu_{t-1}^e + \eta_t^e, \qquad \eta_t^e \sim N(0, \lambda_t^{e,\eta}), \\
\mathcal{H} &= \{0\} \cup [\underline{h}, \overline{h}], \\
G_{t+1} &= \mathbf{T}_t\left(G_t\right),
\end{aligned}$$

where G_t is a distribution of households over $(e, a_t, \mu_t^e, \upsilon_t^e)$ space and $\mathbf{T}_t$ is a mapping from G_t to G_{t+1}.

3.2 Production

We consider a representative firm that produces output using a standard Cobb-Douglas technology with competitive factor markets. The firm's production function is given by

$$F(K_t, H_t, z_t) = z_t K_t^\alpha H_t^{1-\alpha},$$

where K_t is aggregate capital which depreciates at a rate of δ and H_t is aggregate labor input. The aggregate labor H_t is a constant elasticity of substitution (CES) aggregator of efficiency hours worked by skilled S_t and unskilled U_t households:

$$H_t = \left\{ \chi_t U_t^{\phi} + (1-\chi_t) S_t^{\phi} \right\}^{\frac{1}{\phi}}, -\infty < \phi \leq 1,$$

where $\frac{1}{1-\phi}$ denotes the elasticity of substitution between the two types of labor inputs. The parameter χ_t is introduced to capture skill-biased technical changes over time. A decrease in χ_t makes efficiency units of hours worked by skilled workers more productive relative to those of unskilled workers, leading to a rise in the skill premium.

The assumption of competitive factor markets imply

$$r_t + \delta = F_K(K_t, H_t, z_t) = z_t \alpha (K_t/H_t)^{\alpha-1},$$

$$w_t^s = F_S(K_t, H_t, z_t) = z_t(1-\alpha)(1-\chi_t)(K_t/H_t)^{\alpha} H_t^{1-\phi} S_t^{\phi-1},$$

$$w_t^u = F_U(K_t, H_t, z_t) = z_t(1-\alpha)\chi_t(K_t/H_t)^{\alpha} H_t^{1-\phi} U_t^{\phi-1}.$$

Therefore, the ratio between the rental rate of skilled labor and the rental rate of unskilled labor, ω_t, is given by:

$$\omega_t \equiv w_t^s / w_t^u = \frac{1-\chi_t}{\chi_t} (S_t/U_t)^{\phi-1}.$$

The skill premium defined by the mean relative hourly wages of skilled to unskilled men can be written as

$$\text{skill premium}_t = E(w_t^s x_t^s)/E(w_t^u x_t^u).$$

3.3 Competitive Equilibrium

A competitive equilibrium consists of value functions, individuals' optimal policies, aggregate inputs, prices, and a law of motion for the distribution $G_{t+1} = \mathbf{T}_t(G_t)$ such that:

given sets of parameters $\{z_t, \chi_t, \pi_t, \lambda_t^{s,\eta}, \lambda_t^{u,\eta}, \lambda_t^{s,\upsilon}, \lambda_t^{u,\upsilon}\}_{t=0}^{\infty}$,

i) Given $\{w_t^s, w_t^u, r_t\}_{t=0}^{\infty}$, households of skill type e optimally choose $c_t(e, a_t, \mu_t^e, \upsilon_t^e; G_t)$, $a_{t+1}(e, a_t, \mu_t^e, \upsilon_t^e; G_t)$, and $h_t(e, a_t, \mu_t^e, \upsilon_t^e; G_t)$ that are consistent with households problem specified above,

ii) The firm chooses K_t, S_t, and U_t to maximize profits,

iii) The labor market for each skill type clears:

$$\begin{aligned} S_t &= \int h_t(s, a_t, \mu_t^s, \upsilon_t^s; G_t) \exp(\mu_t^s + \upsilon_t^s)\ dG_t(s, a_t, \mu_t^s, \upsilon_t^s); \\ U_t &= \int h_t(u, a_t, \mu_t^u, \upsilon_t^u; G_t) \exp(\mu_t^u + \upsilon_t^u)\ dG_t(u, a_t, \mu_t^u, \upsilon_t^u) \end{aligned}$$

iv) The capital market clears:

$$K_t = \sum_e \int a_t\ dG_t(e, a_t, \mu_t^e, \upsilon_t^e)$$

v) The goods market clears:

$$\sum_e \int c_t(e, a_t, \mu_t^e, \upsilon_t^e; G_t)\ dG_t(e, a_t, \mu_t^e, \upsilon_t^e) + \Phi_t = F(K_t, H_t, z_t) + (1-\delta)K_t - K_{t+1},$$

where $H_t = \left\{\chi_t U_t^\phi + (1-\chi_t) S_t^\phi\right\}^{\frac{1}{\phi}}$

vi) The government budget constraints are satisfied:

$$\sum_e \int T_t\ dG_t(e, a_t, \mu_t^e, \upsilon_t^e) = (1-\gamma) \sum_e \int a_t\ dG_t(e, a_t, \mu_t^e, \upsilon_t^e);$$

$$\Phi_t = \sum_e \int \Omega\left(r_t a_t + w_t^e h_t x_t^e\right)\ dG_t(e, a_t, \mu_t^e, \upsilon_t^e)$$

vii) Individual and aggregate behaviors are consistent.

4 Calibration

This section describes our calibration strategy. We first begin with a few parameters set based on related literature. We then turn to parameters whose values are chosen so that the model can match relevant data moments. Table 3 summarizes the values of our model parameters and the data moments used for calibration.

The model period is one year. The common estimates for risk aversion fall between 1 and 2 in the literature (see Attanasio (1999)): we set the parameter σ to 1.5. This value for the relative risk aversion implies that the intertemporal elasticity of subsitution is less than

1 (i.e., the income effect is greater than the subsitution effect). According to Keane (2011), most micro estimates for the Frisch elasticity of hours worked range between 0.1 and 0.7. The parameter ν is set to match this elasticity of 0.4. Following Aiyagari (1994) and Prescott (1986), we assign 0.36 and 0.08 to, respectively, the capital share in the production function, α, and the depreciation rate of capital, δ. There is a large literature that attempts to estimate the elasticity of substitution between skilled and unskilled workers and most estimates fall between 1.4 and 2.[18] We take an intermediate value of 1.5 by setting the parameter ϕ to 1/3. The maximum hours of work, $\overline{h}$, is 1, which corresponds to 112 hours per week, while the minimum hours of work, $\underline{h}$, is 0.04, equivalent to 5 hours per week, consistent with how we treat the data.[19] Using estimates in Chen and Guo (2013), the parameters τ and κ, which govern the income tax schedule, are set at 0.1679 and 0.8081, respectively.

Another set of parameters is calibrated so that the initial steady state of the model can replicate target data moments. The discount factor for skilled men, β^s, is picked so that the equilibrium real interest rate matches an annual real interest rate of 0.04. The discount factor for unskilled men, β^u, is chosen so that the relative wealth of skilled to unskilled households from the initial steady state of the model is consistent with its data counterpart. Due to the lack of wealth data by skill type available in the late 1960s, we refer to the 1962 Survey of Financial Characteristics of Consumers (SFCC). According to this data set, the relative net worth of skilled households is 2.5 in 1962. We choose the value of the parameter β^u by targeting the relative wealth of skilled households in the initial steady state of 2.5. The preference parameters ψ^s and ψ^u are chosen to match the average weekly hours worked of skilled and unskilled men for the first 5 years in our PSID sample, which are 43.9 and 42.7 hours, respectively. We set the survival probability γ to 0.972 by targeting the average

[18] Katz and Murphy (1992) find that this elasticity is about 1.4 based on CPS March Supplements for the period 1967-1973. Autor et al. (2008) show that the elasticity is estimated to be about 1.6 using the CPS data extended to 2005. Heckman et al. (1998) and Card and Lemieux (2001) obtain the estimates of 1.5 and 2, respectively.

[19] We solve our model using different values of $\overline{h}$ and $\underline{h}$ and find that our main results are not sensitive to these upper and lower bounds of labor supply.

life span of workers of 35 years because the data covers workers of ages 25 through 59.

[Put Table 3 here]

The remaining parameters include variables whose values vary over time. We set the total factor productivity, z_t, to 1 in every period.[20] The population share of skilled men, π_t, is taken from the PSID. The parameter χ_t governing the skill-biased technical change is chosen to match the Hodrick-Prescott (HP) trend of the skill premium in the PSID over the sample period. In regards to the parameters that determine the idiosyncratic productivity, x_t^e, we take the HP trend of our point estimates for the variances of persistent and transitory wage shocks from the PSID data. Figure 4 depicts the trends in the skill premium and in the variances of both type of wage shocks by skill level.

5 Results

This section presents our main quantitative results. We assume that the 1967 U.S. economy is in the steady state and that the economy experiences unexpected changes in the variances of wage shocks as well as in the skill premium between 1967 and 2000. Specifically, in the beginning of 1967, workers in the economy receive a new set of complete information about the first and second moments of their wage process (*i.e.* the skill premium and wage volatility) over the 1967-2000 period. After 2000, the economy gradually converges to a new steady state.

This section begins with the initial steady state results. It is then followed by transitional dynamics of hours worked, precautionary savings, and consumption by skill group.[21] Lastly, we also present a couple of counterfactual experiments which highlight the model's mechanism.

[20] As an alternative, we solve the model by normalizing the output y_t to 1 in every period and find that the short-run quantitative results differ little. In the long-run, there are some level effects at work, yet the long-run skilled-unskilled hours differential is the same as the benchmark outcome.

[21] The computation algorithm to solve the model is described in the Online Appendix.

5.1 Initial Steady State

We solve for the initial steady state allocations of the model economy by targeting the 1967 U.S. economy. Table 4 summarizes the results. In the initial steady state, the wage rate per efficiency unit of labor (w) for skilled men is 65% larger than that for unskilled men. We do not have a direct data counterpart for the wage rate. Instead, hourly wages are observable in the data. Since we target the skill premium defined by the average hourly wage of skilled men relative to unskilled men in the initial steady state, the model replicates the ratio of hourly wages (wx) between skilled and unskilled men from the data. Skilled men also work 3% longer hours (or 1.1 hours more per week) than unskilled men, replicating what we observe in the data. Skilled men on average earn 63% more wage income than unskilled men, consistent with the data. In the model, skilled men's hours are also more dispersed, displaying a larger standard deviation compared to unskilled men. This is because skilled men derive less disutility from hours worked and therefore adjust their hours by a larger amount in response to wage shocks than unskilled men do. We observe the same pattern in the data, although the difference in the standard deviation of hours between the two skill groups in the data is not as large as that in the model.

[Put Table 4 here]

In the initial steady state, skilled men are exposed to less volatile wages than unskilled men. Skilled men face smaller wage volatility through both persistent and transitory wage components. These differences in wages and hours between the two skill groups affect their consumption, resulting skilled men's average consumption to be 41% larger than that of unskilled men.

5.2 Transitional Dynamics

5.2.1 Hours Worked

The main experiment in this study is the comparison of the transitional dynamics generated by the model in response to exogenous changes in the wage structure with what we observe in the U.S. economy since 1967. In our benchmark economy, we consider gradual exogenous changes from 1967 to 2000 in the following three factors: i) the skill premium; ii) the variances of persistent and transitory wage shocks; iii) the skill composition in the labor force.[22] Households are assumed to have perfect foresight about the changes in the wages and the skill composition. After 2000, the wage structure stays unchanged and the economy gradually converges to a new steady state.

[Put Figure 6 here]

According to the model, the exogenous changes considered have a big impact on the evolution of hours worked by both skill groups. Figure 6 presents how hours worked by skill group evolve along the transition to the new steady state in our benchmark model.[23] As the wage structure changes, skilled men's hours worked initially jump to 45.2 hours per week. Skilled men then reduce their work hours for the first decade while the skill premium declines. This pattern is reversed in the late 1970s when skilled men begin to increase their labor supply. The weekly hours of skilled men peak in the early 1990, reaching 45.6 hours. It is worth noting the evolution of skilled men's hours worked resembles the time path of the skill premium. This phenomenon is attributable to the perfect foresight assumption. In the benchmark model, skilled men have perfect foresight about the changes in the wage structure from 1967 to 2000. Anticipating a rise in wage volatility, skilled men have incentives to increase a buffer stock of precautionary wealth by working longer hours. The most efficient time allocation to this aim is to increase hours worked when the skill premium is high and

[22] The skill premium and the skill composition may not be completely separated. Changes in the skill premium are due to changes in both the skill biased technology χ and the skill composition π in the labor force. Moreover, the skill composition in the labor force has an indirect effect on the relative wages through the general equilibrium effect because aggregate savings are affected by changing shares of skilled and unskilled men.

[23] Due to the lack of business cycle features, the levels of hours generated by the model do not mimic those in the data. However, taking the difference in hours between the two skill groups eliminates the business cycle component that affects both skill groups evenly, justifying a comparison between the evolution of the skilled-unskilled hours differential in the model and its data counterpart.

work less when the skill premium is low, consistent with the benchmark model prediction. If skilled men believe the current wage structure to stay unchanged instead, they would rather work shorter (longer) hours when the skill premium is higher (lower) due to dominant wealth effects. We confirm this in Section 5.3 where we run a counterfactual under an alternative assumption about the formation of households' expectation. We find that the upward trend in skilled men's hours caused by incresed wage volatility does not last in the long run. As shown in Figure 7, skilled men increase their wealth rapidly beginning in the 1980s. As skilled men accumulate more wealth, they can afford a reduction in their work hours. Due to this wealth effect, skilled men begin to decrease their hours in the early 1990s. The model predicts that skilled men's work hours reach 42.8 hours per week by 2100, which is 1.1 hours smaller compared to the initial steady state level.

[Put Figure 7 here]

Unskilled men show a different pattern of time allocation to labor in terms of both timing and magnitude, compared to skilled men. Anticipating a rise in wage volatility, unskilled men also have incentives to increase their labor supply for a precautionary saving motive. The initial decline in the skill premium provides unskilled men with a good opportunity to pursue this scheme. Thus, unskilled men maintain longer work hours for the first decade since 1967 than in the initial steady state. However, this rise in unskilled men's work hours is much smaller than that of skilled men because they face a smaller increase in wage volatility and they derive larger disutility from work than skilled men. By working more when the skill premium is low, unskilled men can increase a buffer stock of wealth in the beginning of the sample period as shown in Figure 7. Due to wealth effects from this accumulated wealth, unskilled men begin to reduce their work hours sooner than skilled men. However, unskilled men's hours worked bounce back around 2000. The substantial amount of savings accumulated by skilled men lowers the real interest rate, discouraging unskilled men from saving in the long run. This general equilibrium effect causes unskilled men to decumulate their savings and increase their labor supply after after 2000. Nonetheless, unskilled men work shorter hours per week in the new steady state compared to the initial steady state.

[Put Figure 8 here]

Figure 8 depicts the model's predictions for the hours difference between skilled and unskilled men in the short run (left) vs. in the long run (right) along with their data counterparts. The skilled-unskilled hours differential closely follows the trend in hours worked of skilled men. The difference in weekly hours between skilled and unskilled men instantly rises to 2.3 hours as soon as the news regarding the exogenous wage changes arrives and then declines until the mid-1970s. Since then, the hours differential between the two groups increases sharply, reaching its peak in 2000. This short-run increase in the skilled-unskilled hours differential implies that the impact of wage volatility is strong enough to dominate the wealth effect from the rise in the relative wages for skilled men. Such model predictions are broadly consistent with their data counterparts, although the turning points occur slightly sooner in the data compared to the model. According to the PSID, the skilled-unskilled hours differential continued to increase from the early 1970s to the early 1980s and maintained its level from the early 1980s to the early 1990s. Let $\Delta(\overline{h^s} - \overline{h^u})_{SR}$ denote the short-run change in the hours differential, defined by the change in the skilled-unskilled hours differential between the initial steady state and the peak year. Table 5 shows that the short-run change in the hours differential in the model is 2.68 hours, while it is 1.36 hours in the data. This overstatement may be attributable to some missing factors that actually reinforced income effects relatively more for skilled men such as an increase in secondary earnings within household, a rise in asset prices, to name a few.[24]

[Put Table 5 here]

[24]Previous studies find that over the past few decades, the correlation between a husband's and a wife's education increased, while female labor force participation increased more sharply among more educated women. This implies that the share of skilled men with a working spouse increased. Given the dominant income effect, this may have caused skilled men to work less relative to unskilled men. Missing this channel is a potential reason behind the over-prediction problem of our model.

We find that the rise in the skilled-unskilled hours differential in response to a rising wage volatility is a short-run phenomenon. In the long run, precautionary savings play an important role in the evolution of hours worked. In the model, the skilled-unskilled hours differential continues to decline after 2000 until it reaches the new steady state. Analogous to the short-run change in the hours differential, we define $\Delta(\overline{h^s} - \overline{h^u})_{LR}$ as the long-run change in the hours differential between 1967 and 2100. The long-run change in the hours differential is 0.07 hours in the benchmark model, as shown in Table 5.

A useful check for the model's validity is to compare the model's implications for the second moments of hours worked with what we observe in the data. Table 6 summarizes how the coefficients of variation (CV) of hours worked by skill group changed over the sample period both in the model and the data. Admittedly, the model cannot generate as much variation in hours as we observe in the data because it abstracts from various factors that contribute to cross-sectional dispersion in hours, for instance, heterogeneity in disutility from work and fixed costs of work. However, the model is in line with a distinct feature of the observed trend in this second moment of hours worked: the coefficient of variation of hours increased more for unskilled men than for skilled men in both the model and the data. Although unskilled men experienced a smaller increase in their wage volatility, their lower wealth to income ratio appears responsible for the larger increase in the cross-sectional variation of hours, compared to skilled men.

[Put Table 6 here]

The changing wage structure also affects the correlation between wages and hours. With our parameterization for the risk aversion ($\sigma > 1$) where the wealth effect is dominant, permanent wage changes such as an increase in the skill premium cause households to adjust their hours in the opposite direction of the wage changes, reducing the wage-hour correlation. In contrast, rising wage volatility causes households to work more in response to a rise in their wages to accumulate more precautionary savings, which leads to a rise in the wage-hour correlation. As Table 6 shows, the model cannot match the levels of the correlation coefficients largely due to the measurement error in the data.[25] However, the direction of the changes in the correlation coefficients from the model is consistent with that in the data. In both the model and data, the wage-hour correlation increased for both skill groups with the rise more pronounced for skilled men than for unskilled men. Many previous studies including Costa (2000), Santos (2014), and Heathcote et al. (2010) find that the wage-hour correlation increased sharply over the past few decades. In this paper, we document that the pattern holds within each skill group as well.

5.2.2 Precautionary Savings and Consumption

In a class of model where households face uninsurable idiosyncratic shocks and markets are incomplete, households have strong incentives to increase wealth in response to a rising wage volatility. Figure 7 displays the evolution of wealth by skill level generated by the model. Exposed to a large wage volatility, skilled men accumulate a substantial amount of wealth in the long run, but this path is not monotone. For the initial decade while the skill premium declines, skilled men reduce their work hours and thus deplete their wealth to smooth consumption fluctuations. It takes another decade before skilled men begin to increase their asset holdings not only because the skill premium has not recovered enough but also because their wages become more volatile. Skilled men's wealth in the mid-1980s is about 25% lower than that in the initial steady state. Skilled men begin to increase their wealth in the late 1980s and this trend continues until the economy converges to the new steady state.

Unskilled men, on the other hand, gradually increase their wealth from the beginning. As the skill premium declines, unskilled men have sufficient resources to increase wealth without reducing consumption much. The wealth effect from this increased asset holdings allows unskilled men to reduce their hours worked beginning in the early 1980s. However,

[25] The smaller correlation coefficients in the data compared to their model counterpart are attributable to the measurement error in the data. Note that hourly wage in the data is annual labor income divided by annual hours worked and weekly hours are annual hours divided by 52 weeks. Thus, weekly hours and hourly wages are negatively correlated by construction.

unskilled men do not maintain this large wealth in the long run. The real interest rate continues to decline from the beginning as the aggregate capital stock increases due to large savings by unskilled men. As skilled men accumulate wealth more rapidly beginning in the late 1980s, the real interest rate declines even further,[26] which discourages unskilled men from accumulating more wealth.

[Put Figure 9 here]

In order to compare the model implications for wealth with their data counterparts, we use wealth data from the PSID and the CEX. The PSID publishes data on various categories of household wealth in every few years from 1984 to 2000. We define wealth as net worth and determine the skill type of a household based on the skill level of the male head of a household. Household wealth from the CEX is the sum of house value and financial assets, available yearly beginning in 1981. Figure 9 depicts the paths of relative household wealth of skilled to unskilled men in the model together with their data counterparts from the PSID and the CEX. The relative wealth of skilled households from both data sets displays a significant increase over the sample period, particularly during the 1990s that coincides with the longest expansion in the U.S. Due to the lack of business cycle fluctuations, our model cannot generate as strong relative wealth growth for skilled households during the 1990s as we observe in the data. However, both the level and the trend of the relative wealth of skilled to unskilled households from the model are broadly consistent with their data counterparts.

[Put Figure 10 here]

As an additional validity check of the model, we turn to the model's implications for consumption trends. Figure 10 compares the trends in the relative consumption of skilled to unskilled households generated by the model with their data counterparts. As for consumption in the data, we employ two measures of household consumption, noondurables and

[26] During the transition, the composition of skilled workers has monotonically increased from 15.4% in 1967 to 30.4% in 2000. Combined with this, the increase in the average skilled men's wealth accumulation has much stronger effect on the real interest rate over time.

nondurables+, based on the CEX as in Krueger and Perri (2005). Nondurables indicates household nondurables consumption, while nondurables + includes not only nondurables but also services, small durables, and imputed service flows from houses and cars. These time series are available beginning in 1981.

In the model, the ratio of household consumption between skilled and unskilled households declines as soon as the wage structure changes. This ratio then bounces back around the early 1980s and continues to increase until the economy converges to the new steady state. This time path of the relative consumption from the model is consistent with continued increases in the consumption ratio between the two skill groups since the early 1980s in the CEX data.[27]

5.3 Understanding the Mechanism

The main exogenous changes in the wage structure considered in this paper can be summarized by an increase in the relative wages of skilled men and a larger increase in wage volatility for skilled men relative to unskilled men. The benchmark results reflect the combined effects of these two driving forces on skilled men's relative hours worked. In order to separately identify the roles of these two exogenous forces in explaining the patterns of the skilled-unskilled hours differential, we implement two counterfactual experiments. In the first counterfactual, we unplug any changes in the skill premium from the benchmark model. In the other counterfactual, we have skilled men face the same change in the total wage variances as unskilled men.

Effect of Skill Premium

Figure 11 presents the results from a counterfactual experiment where the skill premium is held constant at its initial steady state level. With a constant skill premium, skilled men's work hours increase significantly as soon as the wage structure changes. Skilled men continue to increase their weekly hours until the mid-1980s, which enables them to increase

[27] Attanasio and Davis (1999) also document significant growth in the relative household consumption of more educated groups during the 1980s.

their asset holdings sooner than in the benchmark model. This rapid wealth accumulation causes skilled men to reduce their hours worked from the mid-1980s, while this reversal in hours worked occurs in the early 1990s in the benchmark model. Without positive wealth effects from a rising skill premium, skilled men work longer hours in the long run, compared to the benchmark result, but the difference is fairly small.

[Put Figure 11 here]

On the other hand, unskilled men show a monotonically declining pattern of work hours with a constant skill premium. Unskilled men increase their weekly hours immediately as the wage structure changes and reduce their work hours monotonically afterwards, in contrast with the benchmark result. The decreasing pace of unskilled men's hours worked slows down in the mid-1990s. In the long run, unskilled men work shorter hours than in the benchmark model because their lifetime wealth increases without a rise in the skill premium.

These patterns of hours worked of both skill groups lead to a substantial increase in the skilled-unskilled hours differential in the short-run. There is no longer a significant drop in the hours differential followed by a sharp increase, mimicking the path of the skill premium. With a constant skill premium, hours difference between the two skill groups continues to increase until the early 1990s and decline afterwards as skilled men reduce their weekly hours. The hours differential in this counterfactual rises by 2.9 hours in the short run, 0.3 hours more than it does in the benchmark model. Without wealth effects from the rise in the skill premium, skilled men work longer hours while unskilled men work less, compared to the benchmark economy. Consequently, the skilled-unskilled hours differential in the new steady state is 0.3 hours larger than its benchmark counterpart.

The Importance of Wage Volatility

By having both skill groups face the same increase in wage volatiltiy, we attempt to understand the role of the extra increase in wage volatilty for skilled men in explaining the relative labor supply of skilled men. In this experiment, skilled men experience only 36% of the actual rise in their total wage variances.

Figure 12 depicts the evolution of hours and wealth by skill level with the path of the real interest rate in this experiment. Exposed to a smaller increase in wage volatility, skilled men increase their labor supply in the short run by a smaller amount, compared to the benchmark result. Skilled men also accumulate a smaller buffer stock of precautionary savings in the long run, working more in the long run than in the benchmark model. Due to the smaller wealth accumulation by skilled men, the real interest rate does not decline much and hench unskilled men maintain a larger level of wealth, compared to the benchmark model. The skilled-unskilled hours differential in this experiment increases up to 2.9 hours in the short run, 0.8 hours less than in the benchmark model. This highlights the quantitative importance of the larger increase in wage volatility for skilled men in explaining the short-run increase in their relative labor supply. On the other hand, the long-run hours differential differs little between this experiment and the benchmark model.

[Put Figure 12 here]

We also find that the extra increase in wage volatility for skilled men contributes to significant growth in their wealth. By having skilled men face the same increase in their wage variances as unskilled men, skilled men's wealth in 2000 is reduced by 24% and in 2100 by 33%, relative to their levels in the benchmark model. This quantitative result can be related to many empirical studies attempting to measure the importance of precautionary savings motive in the observed wealth accumulation. These studies use a variety of econometric methods to estimate the contribution of the precautionary savings motive to wealth, and their estimates fall in a wide range.[28] Among these studies, Carroll and Samwick (1998) estimate 32% to 50% of wealth in the PSID sample to be attributable to the extra uncertainty some households face compared to the lowest uncertainty group. Our quantitative result is

[28] The lack of the consensus about the size of precautionary savings and wealth is mainly because it is challenging to identify exogeneous variations in income risks. In addition, income risks and the time path of income are often correlated, making it hard to disentangle the precautionary motive from the intertemporal subsitution motive in saving behavior. See Browning and Lusardi (1996) and Díaz et al. (2003) for a survey of the estimates in the literature.

broadly in line with their estimates.

5.4 Welfare Analysis

Changes in the wage structure considered in this study lead to significant adjustments in households' hours worked, consumption, and wealth. We assess how much households would gain or lose if they went through the transition to the new steady state. Specifically, we compute the percentage change in lifetime consumption households in the initial steady state should receive to give them the same utility they would enjoy from period t onward. It also indicates how much lifetime consumption households in the initial steady state would give up to avoid the changes in the wage structure from period t onward. This consumption equivalent (CEV) measure is the value ζ_t that solves

$$\int \sum_{\tau=t}^{\infty} (\beta^e \gamma)^{\tau-t} u([1+\zeta_t]c_\tau^*, h_\tau^*) dG^* = \int \sum_{\tau=t}^{\infty} (\beta^e \gamma)^{\tau-t} u(c_\tau, h_\tau) dG_\tau, e \in \{s, u\}$$

where the superscript $*$ denotes the initial steady state and $G_t(e, a, \mu, \upsilon)$ is the distribution of households across all possible states in period t.

[Put Figure 13 here]

Figure 13 plots the average welfare changes by skill type. The structural changes in wages turn out to be welfare-deteriorating for both skill groups. While skilled workers gain from the rise in the skill premium, a negative welfare effect from more volatile wages is dominating. Skilled men face about a 4% decline in their lifetime consumption in the long run. This long-run welfare loss of skilled men accompanies even larger welfare losses in the short run. During the transition to the new steady state, skilled men reduce their consumption and increase their hours worked to build up a larger buffer stock of precautionary wealth. The average welfare loss of skilled men is more than 10% of lifetime consumption in the late 1970s.

Unskilled men also experience large welfare losses both in the short run and in the long run, due to an increase in wage volatility and a decline in their relative wages. Unskilled men's long-run welfare loss is more or less the same as skilled men's. In the short run, unskilled men's welfare moves closely together with the evolution of the skill premium. During the 1970s when the skill premium moves in favor of them, the average welfare loss of unskilled men declines. It then reverses its trend in the late 1970s and continues to increase until the economy converges to the new steady state.

6 Discussion

In this section, we would like to show that our main results prevail even if we change certain assumptions of the model. First, we consider the effect on the hours worked of U.S. taxation changes around the mid-1980s that reduced the progressivity of labor income taxes. Second, we examine how important the perfect foresight assumption is in our quantitative results by considering an alternative assumption about expectation formation. It is then followed by a set of sensitivity analysis where the degree of relative risk aversion or the elasticity of intertemporal substitution varies.

6.1 Taxation Changes

Labor income tax affects the effective real wage faced by households and distorts their labor supply. It is then natural to ask whether the observed changes in the relative hours worked of skilled to unskilled men are caused by important changes in the U.S. tax system. One of the most important changes in the U.S. tax system in recent decades is the tax reform during the Raegan administration. The Tax Reform Act of 1986 devised to simplify the tax code reduced the tax rates for high-income workers significantly, lowering the progressivity of the U.S. tax system. In this section, we present the transitional dynamics in the presence of this tax change in 1986 and compare the results with our benchmark one. We parametrize the tax change based on the estimates in Chen and Guo (2013). We keep the values of τ and κ at the initial steady state until 1986. After 1986, the values of τ and κ are set at 0.0634 and 0.7973, respectively, which implies that the post-1986 tax code is less progressive, compared to the pre-1986 one.

[Put Figure 14 here]

Figure 14 shows that the tax change that occurred in 1986 leads to a substantial jump in the hours worked by both skill groups and hence in the hours difference between the two skill groups after 1986. The less progressivity implied by the tax reform makes leisure increasingly costly for households who draw better productivity shocks. Once the new tax system is adopted, households increase their work hours much more than in the benchmark model and this effect is stronger for households with more labor income. This, in turn, raises the skilled-unskilled hours differential in the short run compared to the benchmark result. While this effect is quantitatively large, the tax code change does not appear to be one of the main driving forces behind the observed changes in the skilled-unskilled hours differential. This increase in the hours differential caused by the tax code change occurs mostly after 1986 in the model, whereas in the data, the skilled-unskilled hours differential rose mainly until the early 1980s.

With the new tax system, individual savings also increase significantly. The increased labor supply of skilled men helps them build up an even larger buffer stock of wealth in the long run, compared to the benchmark model. Despite the larger wealth, skilled men do not reduce work hours much in the long run. Under this less progressive tax system, skilled men choose to substitute consumption for leisure and maintain longer working hours in the long run compared to the benchmark model. This substitution also occurs among unskilled men. Under the new tax regime, unskilled men also work longer hours in the new steady state than in the benchmark model. With these effects combined, the skilled-unskilled hours differential in the new steady state converges to the same level as its benchmark counterpart.

6.2 Expectations

In our benchmark model, we assume that households have perfect foresight about the evolution of the wage structure. Given the anticipated changes in the wage structure, households choose the optimal path of hours worked. To the extent that it is difficult to forecast the observed changes in wages, the actual evolution of hours may be different from the benchmark result. In order to address this concern, we consider an alternative assumption about the formation of households' expectations and examine how critical the perfect foresight assumption is in our main quantitative results for hours worked.[29]

Suppose that households do not know the evolution of the skill premium and the wage shock variances. Households believe that the current wage structure remains unchanged forever. In every period they observe a new wage structure and believe such a wage structure stays unchanged. Figure 15 displays the evolution of hours worked by skill level under such information updating. The evolution of skilled men's hours worked in this exercise is different from what we obtain from the benchmark model. From the late 1960s to mid-1970s when the skill premium declines, skilled men increase their hours worked in contrast with the benchmark result. Without foreseeing the rise in the skill premium since the 1970s, skilled men choose to work longer hours in response to a decline in the skill premium due to negative wealth effects. As the skill premium turns to an increasing pattern around the mid-1970s, skilled men reduce their work hours slighlty. However, as skilled men are exposed to larger wage volatility since the early 1980s, they increase their labor supply more rapidly than in the benchmark model. This pattern of hours adjustment continues until the mid-1990s when skilled men begin to decrease labor supply due to the wealth effect from the increased precautionary wealth. Despite the differences in the short run, the long-run path of skilled men's hours converges to the benchmark prediction.

[Put Figure 15 here]

Unskilled men's hours worked also show a different path in the short run compared to the benchmark counterpart. As soon as the wage structure changes, unskilled men face a decline

[29]It is also worth noting that the estimated variances of persistent and transitory shocks have different precisions at different periods. The imprecision of the estimates for certain periods implies a large forecasting error, suggesting that the perceived trends of wage variances may not be the same as the observed ones, which is another reason to relax the assumption on perfect foresight.

in the skill premium and a rise in the variance of persistent wage shocks. The declining skill premium provides unskilled men incentives to reduce work hours due to wealth effects, while the larger wage volatility causes them to work longer hours to increase precautionary wealth. The right panel of Figure 15 shows that unskilled men reduce their weekly hours during this period, suggesting that the gain from the declining skill premium was a dominant factor. Since the skill premium begins to pick up in the mid-1970s, unskilled men's work hours reverses to an increasing pattern. However, as the variance of persistent wage shocks for unskilled men decline in the 1980s, unskilled men resume to reduce their hours worked. After the early 1990s, the skill premium rises above its level in the initial steady state, inducing unskilled men to increase their hours slightly. This short-run change in unskilled men's hours contrasts with the benchmark prediction where unskilled men's hours increase only slightly for the first decade and decline thereafter.

[Put Figure 16 here]

Figure 16 depicts the skilled-unskilled hours differential based on this partial information updating. In the benchmark model, both skill groups have perfect foresight about the wage structure, which allows them to efficiently adjust their time allocation by increasing labor supply when the wage moves in favor of them. In the absence of the perfect foresight, this dynamic consideration is missing, causing both skill groups to increase their work hours when they face a decline in their relative wages or a rise in wage volatility. As a result, the skilled-unskilled hours differential increases from the beginning as opposed to the benchmark result, while it increases more in the short run than in the benchmark model. While the assumption about the expectation formation is important in determining the exact timing of the short-run increase in the skilled-unskilled hours differential, the quantitative importance of the changing wage structure in explaining the short-run changes in hours worked by skill level prevails regardless of the assumption.

6.3 Sensitivity Analysis

This subsection presents a sensitivity analysis with respect to two key model parameters governing households' intertemporal substitution of consumption and leisure.

Relative Risk Aversion

The relative risk aversion determines the curvature of utility from consumption. A higher risk aversion implies that the marginal utility of consumption diminishes more rapidly. Thus, risk averse agents choose to increase leisure (and reduce hours) relatively more in response to a rise in real wage than otherwise the same agents. In other words, a higher risk aversion is associated with a larger income effect of higher wages on leisure and a smaller (uncompensated) wage elasticity of labor supply. A simple derivation of the elasticity illustrates this point. Recall that the utility function in this study is given by $u(c,h) = \frac{c^{1-\sigma}}{1-\sigma} - \psi^e \frac{h^{1-\nu}}{1-\nu}, \sigma > 0, \nu < 0$. In a static case, the utility function implies that:

$$\frac{h^s}{h^u} = \left(\frac{\psi^u}{\psi^s} \frac{w^s}{w^u} \left(\frac{c^s}{c^u} \right)^{-\sigma} \right)^{\frac{-1}{\nu}}.$$

Taking logarithm of both sides of the above equation and taking the total derivative, we obtain:

$$\frac{d \ln \frac{h^s}{h^u}}{d \ln \frac{w^s}{w^u}} = \frac{-1}{\nu} + \frac{\sigma}{\nu} \frac{d \ln \frac{c^s}{c^u}}{d \ln \frac{w^s}{w^u}}.$$

Based on Attanasio and Davis (1999), assume that $d \ln \frac{c^s}{c^u} \approx d \ln \frac{w^s}{w^u}$. Then, we have:

$$\frac{d \ln \frac{h^s}{h^u}}{d \ln \frac{w^s}{w^u}} = \frac{\sigma - 1}{\nu}.$$

This implies that a rise in the skill premium causes the relative hours worked of skilled to unskilled men to decline if $\sigma > 1$, increase if $\sigma < 1$, and remain unchanged if $\sigma = 1$.

On the other hand, the parameter σ affects households' precautionary savings motives. More risk averse households attempt to accumulate more precautionary wealth when exposed to a rise in wage volatility. Suppose that skilled men face an increase in both the skill premium and wage volatility as observed in the U.S. data. A higher degree of risk aversion implies that the larger income effect causes skilled men to reduce hours worked more while the stronger precautionary motive induces them to increase hours worked more in the short run. Depending on which effect is larger, the short-run increase in skilled men's hours worked may vary by the degree of risk aversion.

In our benchmark calibration, we use $\sigma = 1.5$, which implies the dominant income effect in a static case. In order to examine how our main quantitative results hinge on this parameter value, we consider alternative values for the risk aversion parameter, $\sigma = 1$ (log utility) and $\sigma = 3$. For both values of σ, we recalibrate time discount factors β^s and β^u, and preference parameters ψ^s and ψ^u governing utility from non-working time. The recalibrated parameter values are summarized in Table 7.

[Put Table 7 here]

[Put Figure 17 here]

Figure 17 presents hours worked by skill level generated by a model with log utility. While the log utility reduces the income effect from a rising skill premium, it weakens the precautionary savings motive of skilled men compared to the benchmark model. The smaller increase in skilled men's weekly hours in the short run compared to the benchmark result implies that the latter effect is dominant. For unskilled men, both the smaller income effect and the weaker precautionary motive decreases the short-run rise in unskilled men's weekly hours relative to the benchmark result. However, these effects are quantitatively small. Consequently, the skilled-unskilled hours differential increases by a smaller amount in the short run, peaking at 3.1 hours in 2000, 0.6 hours smaller than in the benchmark model. In the long run, households can afford a larger reduction in their work hours because they are willing to bear larger fluctuations in consumption compared to the benchmark model. This effect is pronounced for skilled men, causing the hours differences between skilled and unskilled men to converge to an even smaller value in the long run relative to the benchmark value.

[Put Figure 18 here]

Figure 18 displays the transition dynamics of hours worked for $\sigma = 3$. While the benchmark model's qualitative implications for the skilled-unskilled hours differential prevail with more risk averse households, quantitative changes in the hours differential are more drastic. The higher degree of risk aversion causes skilled men to reduce their weekly hours relative to unskilled men more in response to a rising skill premium. However, this effect is more than offset by the strengthened precautionary motives of households, leading to substantial increases in both skilled men's hours worked and the skilled-unskilled hours differential in the short run. The hours differential rises above 6 hours at the peak with $\sigma = 3$, 2.3 hours more than in the benchmark model. Skilled men accumulate much larger wealth in the long run than in the benchmark model, yet they work longer hours in the new steady state to reduce fluctuations in consumption even more. Unskilled also save more while working shorter hours in the long run than in the benchmark model. Consequently, the hours differential between the two skill groups converges to 4.4 hours in the new steady state, about 3 hours larger than its initial steady state level.

Frisch Elasticity of Labor Supply

Our main quantitative results may vary by households' willingness to substitute hours worked across time and state. In our benchmark calibration, we use a value of 0.4 for the Frisch labor supply elasticity standard in the literature. However, according to Keane (2011), the estimates for the Frisch elasticity range from 0 to 0.7. Thus, we consider two alternative values (0.2 and 0.6) of the elasticity for a sensitivity check. Table 7 presents the recalibrated parameters for this analysis.

[Put Figure 19 here]

Figure 19 presents the trends in hours worked by skill level with the Frisch labor supply elasticity of 0.2. The main qualitative features of hours worked by skill level are the same as in the benchmark model. Quantitatively, the lower elasticity reduces the magnitude of changes in hours worked by both skill groups compared to the benchmark model, due to a stronger incentive to smooth hours worked across time and state. This effect reduces the short-run increase in the skilled-unskilled hours differential. In this exercise, the hours differential increase up to 2.6 hours in 2000, about 1 hour less than its largest value in the benchmark model. Households achieve these smoother profiles of hours at the expense of their long-run precautionary wealth. Consequently, both skill groups work longer hours in the new steady state than in the benchmark model. This leaves the long-run hours differential unchanged compared to its benchmark level.

[Put Figure 20 here]

In contrast, the larger Frisch labor supply elasticity causes households to adjust their hours by a larger amount as the relative wages change than in the benchmark model, as shown in Figure 18. This results in a larger increase in the skilled-unskilled hours differential in the short run. The hours differential at the peak is 4.7 hours, 1 hour larger than its benchmark counterpart. By substituting hours more flexibly across time and state, households accumulate more precautionary wealth in the long run and thus can afford to reduce their hours worked by a larger amount. These effects are more or less the same for both skill groups, so the skilled-unskilled hours differential converges to the same value as its benchmark counterpart.

7 Conclusion

In the past few decades in the U.S., skilled men increased their hours worked relative to unskilled men while the skill premium rose sharply. This fact contradicts the predictions of the previous literature suggesting a dominant income effect. This paper attempts to explain this discrepancy using wage volatility. In an incomplete markets framework, more volatile wages cause households to work longer hours to accumulate a buffer stock of precautionary wealth. Using the PSID, we estimate the wage process by skill level, and find that skilled men experienced larger increases in their wage volatility compared to unskilled men. This can potentially explain the rise in skilled-unskilled hours differential concurring with a rising skill premium observed in the data.

To quantify the effect of increased wage volatility, we develop a general equilibrium incomplete markets model where workers receive idiosyncratic labor productivity shocks. These productivity shocks are drawn from a skill-specific distribution whose processes follow our estimates from the PSID. The model can replicate the observed increase in hours worked of skilled men relative to unskilled men during the transition, while the model predicts a larger increase in the hours differential than actually seen in the data. However, as skilled men accumulate a large buffer stock of precautionary savings, they can afford to reduce their labor supply relative to unskilled men, reducing the hours gap between skilled and unskilled men in the long run. These results imply that hours adjustment is important for self-insurance in the short run while the effect of precautionary savings are dominant in the long run.

What caused the wage volatility of skilled men to increase more is outside the scope of this study. Exploring potential explanations behind the phenomenon may help improve our understanding of male labor supply, its evolution, and its macroeconomic implications. We leave this for future work.

References

Acikgoz, O. and B. Kaymak, "The Rising Skill Premium and Deunionization," *Journal of Monetary Economics* 63 (2014), 37–50.

Aguiar, M. and E. Hurst, "Measuring Trends in Leisure," *The Quarterly Journal of Economics* 122 (2007), 969–1006.

Aiyagari, S. R., "Uninsured Idiosyncratic Risk and Aggregate Saving," *The Quarterly Journal of Economics* 109 (1994), 659–684.

Attanasio, O. and S. J. Davis, "Relative Wage Movements and the Distribution of consumption," *Journal of Political Economy* 104 (1999), 1227–1262.

Attanasio, O. P., "Consumption," *Handbook of Macroeconomics* (1999), 741–812, john B. Taylor and Michael Woodford (Eds.).

Autor, D. H., L. F. Katz and M. S. Kearney, "Trends in U.S. Wage Inequality: Revising the Revisionists," *Review of Economics and Statistics* 90 (2008), 300–323.

Blundell, R., L. Pistaferri and I. Preston, "Consumption Inequality and Partial Insurance," *American Economic Review* 98 (2008), 1887–1921.

Boppart, T. and P. Krusell, "Labor Supply in the Past, Present, and Future: a Balanced-Growth Perspective," (2016), nBER Working Paper No. 22215.

Browning, M. and A. Lusardi, "Household Saving: Micro Theories and Micro Facts," *Journal of Economic Literature* 34 (1996), 1797–1855.

Card, D. and T. Lemieux, "Can Falling Supply Explain the Rising Return to College for Younger Men? A Cohort-Based Analysis," *The Quarterly Journal of Economics* 116 (2001), 705–745.

Card, D., T. Lemieux and C. Riddell, "Unions and Wage Inequality," *Journal of Labor Research* 25 (2004), 519–559.

Carneiro, P. and S. Lee, "Trends in Quality Adjusted Skill Premia in the United States," *American Economic Review* 101 (2011), 2309–2349.

Carroll, C. D. and A. A. Samwick, "How Important Is Precautinoary Saving?," *Review of Economics and Statistics* 80 (1998), 410–419.

Castro, R. and D. Coen-Pirani, "Why Have Aggregate Skilled Hours Become So Cyclical since the Mid-1980s?," *International Economic Review* 49 (2008), 135–185.

Chen, S.-H. and J.-T. Guo, "Progressive Taxation and Macroeconomic (In)Stability with Productive Government Spending," *Journal of Economic Dynamics and Control* 37 (2013), 951–963.

Costa, D., "The Wage and the Length of the Work Day: From the 1890s to 1991," *Journal of Labor Economics* 18 (2000), 156–181.

Daly, M., D. Hryshko and I. Manovskii, "Reconciling Estimates of Earnings Processes in Growth Rates and Levels," (2014), working paper.

Díaz, A., J. Pijoan-Mas and J. V. Ríos-Rull, "Precautionary Savings and Wealth Distribution under Habit Formation Preferences," *Journal of Monetary Economics* 50 (2003), 1257–1291.

Elsby, M. W. and M. D. Shapiro, “Why Does Trend Growth Affect Equilibrium Employment? A New Explanation of an Old Puzzle,” *American Economic Review* 102 (2012), 1378–1413.

Erosa, A., L. Fuster and G. Kambourov, “Towards a Micro-Founded Theory of Aggregate Labor Supply,” (2014), working Paper.

Flodén, M., “Labour Supply and Saving under Uncertainty,” *The Economic Journal* 116 (2006), 721–737.

French, E., “The Labor Supply Response to (Mismeasured but) Predictable Wage Changes,” *Review of Economics and Statistics* 86 (2004), 602–613.

Gouskova, E., “Parameter Estimates Comparison of Earnings Funcations in the PSID and CPS Data, 1976-2007,” *Economics Letters* 122 (2014), 353?57.

Greenwood, J., N. Guner, G. Kocharkov and C. Santos, “Marry Your Like: Assortative Mating and Income Inequality,” *American Economic Review* 104 (2014), 348–353.

Heathcote, J., K. Storesletten and G. L. Violante, “Two Views of Inequality over the Life-Cycle,” *Journal of the European Economic Association* 3 (2005), 765–775.

———, “The Macroeconomic Implications of Rising Wage Inequality in the United States,” *Journal of Political Economy* 118 (2010), 681–722.

Heckman, J. J., L. Lochner and C. Taber, “Explaining Rising Wage Inequality: Explanations with a Dynamic General Equilibrium Model of Labor Earnings with Heterogeneous Agents,” *Review of Economic Dynamics* 1 (1998), 1–58.

Juhn, C., “The Decline in Male Labor Market Participation: The Role of Declining Market Opportunities,” *Quarterly Journal of Economics* 107 (1992), 79–121.

Juhn, C., K. M. Murphy and R. H. Topel, “Current Unemployment, Historically Contemplated,” *Brookings Papers on Economic Activity* 33 (2002), 79–116.

Juhn, C. and S. Potter, “Changes in Labor Force Participation in the United States,” *Journal of Economic Perspectives* 20 (2006), 27–46.

Katz, L. F. and K. M. Murphy, “Changes in Relative Wages, 1963-1987: Supply and Demand Factors,” *Quarterly Journal of Economics* 107 (1992), 35–78.

Keane, M. P., “Labor Supply and Taxes: A Survey,” *Journal of Economic Literature* 49 (2011), 961–1075.

Krueger, D. and F. Perri, “Does Income Inequality Lead to Consumption Inequality? Evidence and Theory,” *Review of Economic Studies* 73 (2005), 163–193.

Krueger, D., F. Perri, L. Pistaferri and G. L. Violante, “Cross-sectional Facts for Macroeconomists,” *Review of Economic Dynamics* 13 (2010), 1–14.

Krusell, P., L. E. Ohanian, J. V. Ríos-Rull and G. L. Violante, “Capital-skill Complementarity and Inequality: A Macroeconomic Analysis,” *Econometrica* 68 (1994), 1029–1053.

Lemieux, T., W. B. MacLeod and D. Parent, “Performance Pay and Wage Inequality,” *Quarterly Journal of Economics* 124 (2009), 1–49.

Low, H., C. Meghir and L. Pistaferri, “Wage Risk and Employment Risk over the Life Cycle,” *American Economic Review* 100 (2010), 1432–1467.

Michelacci, C. and J. Pijoan-Mas, “The Effects of Labor Market Conditions on Working Time: The US-EU Experience,” (2008), working paper.

———, “Intertemporal Labor Supply with Search Frictions,” *Review of Economics Studies* 79 (2012), 899–931.

Pijoan-Mas, J., “Precautionary Savings or Working Longer Hours?,” *Review of Economic Dynamics* 9 (2006), 326–352.

Prescott, E. C., “Theory Ahead of Business Cycle Measurement,” *Federal Reserve Bank of Minneapolis Quarterly Review* (1986), 9–22.

Saez, E., J. Slemrod and S. H. Giertz, “The Elasticity of Taxable Income with Respect to Marginal Tax Rates: A Critical Review,” *Journal of Economics Literature* 50 (2012), 3–50.

Santos, R., “Dynamic Effects of Labor Supply: A Mechanism Explaining Cross-Sectional Differences in Hours,” *Review of Economic Dynamics* 17 (2014), 630–653.

Table 1: Estimated Variances of Persistent and Transitory Wage Shocks

Year	Persistent Shock Variance		Transitory Shock Variance	
	Skilled	Unskilled	Skilled	Unskilled
1967	0.0079 (0.0023)	0.0037 (0.0016)	0.0182 (0.0211)	0.0289 (0.0118)
1968	0.0000 (0.0033)	0.0114 (0.0079)	0.0049 (0.0123)	0.0097 (0.0085)
1969	0.0000 (0.0039)	0.0051 (0.0041)	0.0000 (0.0054)	0.0105 (0.0098)
1970	0.0000 (0.0038)	0.0017 (0.0043)	0.0128 (0.0123)	0.0132 (0.0091)
1971	0.0107 (0.0055)	0.0040 (0.0032)	0.0220 (0.0153)	0.0264 (0.0110)
1972	0.0125 (0.0075)	0.0104 (0.0073)	0.0333 (0.0155)	0.0333 (0.0104)
1973	0.0134 (0.0099)	0.0000 (0.0007)	0.0000 (0.0060)	0.0323 (0.0117)
1974	0.0026 (0.0023)	0.0000 (0.0018)	0.0205 (0.0131)	0.0306 (0.0102)
1975	0.0060 (0.0043)	0.0048 (0.0035)	0.0141 (0.0165)	0.0256 (0.0091)
1976	0.0000 (0.0020)	0.0170 (0.0060)	0.0347 (0.0160)	0.0481 (0.0135)
1977	0.0000 (0.0016)	0.0088 (0.0052)	0.0223 (0.0121)	0.0358 (0.0103)
1978	0.0090 (0.0045)	0.0075 (0.0050)	0.0246 (0.0128)	0.0331 (0.0107)
1979	0.0000 (0.0015)	0.0071 (0.0048)	0.0064 (0.0129)	0.0322 (0.0118)
1980	0.0067 (0.0043)	0.0115 (0.0063)	0.0522 (0.0226)	0.0499 (0.0113)
1981	0.0043 (0.0041)	0.0147 (0.0049)	0.0499 (0.0213)	0.0492 (0.0113)
1982	0.0109 (0.0061)	0.0057 (0.0060)	0.0321 (0.0162)	0.0511 (0.0129)
1983	0.0100 (0.0049)	0.0129 (0..0056)	0.0623 (0.0174)	0.0424 (0.0118)
1984	0.0149 (0.0071)	0.0208 (0.0071)	0.0322 (0.0162)	0.0558 (0.0133)
1985	0.0128 (0.0069)	0.0132 (0.0047)	0.0170 (0.0146)	0.0656 (0.0123)
1986	0.0363 (0.0089)	0.0091 (0.0050)	0.0516 (0.0234)	0.0691 (0.0123)
1987	0.0048 (0.0043)	0.0026 (0.0036)	0.0320 (0.0177)	0.0659 (0.0131)
1988	0.0141 (0.0067)	0.0095 (0.0046)	0.0409 (0.0143)	0.0891 (0.0134)
1989	0.0258 (0.0104)	0.0118 (0.0043)	0.0416 (0.0125)	0.0563 (0.0124)
1990	0.0032 (0.0054)	0.102 (0.0045)	0.0352 (0.0164)	0.0665 (0.0122)
1991	0.0261 (0.0112)	0.0043 (0.0044)	0.0515 (0.0183)	0.0648 (0.0136)
1992	0.0128 (0.0094)	0.0052 (0.0046)	0.0339 (0.0138)	0.0730 (0.0143)
1993	0.0439 (0.0132)	0.0151 (0.0053)	0.0614 (0.0142)	0.0779 (0.0133)
1994	0.0064 (0.0065)	0.0089 (0.0049)	0.0720 (0.0224)	0.0659 (0.0124)
1995	0.0183 (0.0161)	0.0114 (0.0055)	0.0657 (0.0228)	0.0558 (0.0136)
1996	0.0207 (0.0171)	0.0041 (0.0043)	0.0651 (0.0304)	0.0438 (0.0129)
1998	0.0250 (0.0109)	0.0164 (0.0042)	0.0720 (0.0248)	0.0518 (0.0127)
2000	0.0426 (0.0091)	0.0093 (0.0048)	0.1030 (0.0308)	0.0463 (0.0119)
2002	0.0120 (0.0094)	0.0072 (0.0054)	0.1381 (0.0329)	0.0973 (0.0138)
2004	0.0112 (0.0069)	0.0189 (0.0075)	0.1178 (0.0233)	0.0844 (0.0125)
2006	0.0383 (0.0088)	0.0114 (0.0059)	0.0729 (0.0211)	0.0500 (0.0095)
2008	0.0235 (0.0126)	0.0163 (0.0048)	0.0562 (0.0187)	0.0735 (0.0120)
2010	0.0295 (0.0166)	0.0193 (0.0062)	0.0728 (0.0207)	0.0645 (0.0117)

Note: Numbers in parentheses are standard errors computed from block bootstrapping of 200 replications.

Table 2: Average Annual Changes in Skilled-Unskilled Hours Differential

	Within Cohort	Within Age	Between Age
	$\Delta g_1(a) + \Delta g_2(t)$	$\Delta g_2(t) + \Delta g_3(t-a)$	$\Delta g_1(a) - \Delta g_3(t-a)$
1967–1971	-0.0045	-0.2110	0.2146^{**}
	(0.1661)	(0.1817)	(0.0981)
1972–1976	0.2986^{*}	0.3224^{*}	-0.0097
	(0.1599)	(0.1738)	(0.0955)
1977–1981	0.2838^{*}	0.2285	0.0622
	(0.1483)	(0.1613)	(0.0884)
1982–1986	-0.1076	-0.1548	0.0241
	(0.1402)	(0.1529)	(0.0837)
1987–1991	-0.0059^{*}	0.1029	-0.0903
	(0.1336)	(0.1463)	(0.0795)
1992–1996	-0.0833	-0.1038	0.0034
	(0.1252)	(0.1370)	(0.0747)
1997–2001	0.4945^{**}	0.5113^{**}	0.0194
	(0.1875)	(0.2032)	(0.1111)
2002–2006	-0.1360	-0.1266	0.0206
	(0.1497)	(0.1618)	(0.0879)
2007–2012	-0.1748	-0.0670	-0.0197
	(0.1864)	(0.2012)	(0.1084)
Correlation with Within Age	0.9882	—	0.0156

Note: Numbers in parentheses represent standard errors. Coefficients with a * and a ** are statistically significant at a 10% and a 5% level, respectively.

Table 3: Parameters

Parameter	Target
	Parameters taken from the literature
$\sigma = 1.5$	Relative risk aversion of 1.5
$\nu = -2.5$	Frisch elasticity of hours of 0.4
$\alpha = 0.36$	Capital share (e.g. Aiyagari (1994))
$\delta = 0.08$	Capital depreciation rate (e.g. Prescott (1986))
$\phi = 1/3$	Substitution elasticity of 1.5 b/w skill groups (e.g. Heckman et. al. (1998))
$\overline{h}= 1$	112 hours per week
$\underline{h}= 0.04$	5 hours per week
$\tau = 0.1679$	Progressivity of income tax (Chen and Guo (2013))
$\kappa = 0.8081$	Level of income tax (Chen and Guo (2013))
	Parameters specific to model economy
$\beta^s= 1.0036$	Real interest rate of 0.04
$\beta^u = 0.9946$	$K^s/K^u = 2.5$
$\psi^s= 22.0225$	Avg. weekly hours of skilled men in PSID: 43.9
$\psi^u= 26.6494$	Avg. weekly hours of unskilled men in PSID: 42.7
$\gamma = 0.972$	Average work life of 35 years
π_t	Skilled share from PSID
χ_t	Time series of skill premium in PSID data
$\rho^s= 0.9834$	Own estimate for the skilled from PSID
$\rho^u= 0.9859$	Own estimate for the unskilled from PSID
$\lambda^{s,\mu}= 0.1172$	Own estimate for the skilled from PSID
$\lambda^{u,\mu}= 0.1488$	Own estimate for the unskilled from PSID
$\lambda_t^{s,\eta}, \lambda_t^{u,\eta}, \lambda_t^{s,\upsilon}, \lambda_t^{u,\upsilon}$	Own estimates from PSID for 1967 through 2000

Table 4: Initial Steady State Results

	Model			Data
	Skilled (s)	Unskilled (u)	Ratio (s/u)	Ratio (s/u)
Wage Rate: w	2.21	1.34	1.65	..
Hourly Wage: wx	2.41	1.53	1.58	1.58
Weekly Hours: h	0.39	0.38	1.03	1.03
Labor Income: wxh	0.94	0.58	1.63	1.63
Standard Deviation of Hours: $SD(h)$	0.04	0.03	1.28	1.03
Persistent Volatility: $\lambda^{\eta}/\left(1-\rho^{2}\right)$	0.12	0.18	0.68	0.68
Transitory Volatility: λ^{υ}	0.01	0.02	0.56	0.56
Consumption: c	0.83	0.59	1.41	..

Source: PSID 1967-1971, authors' calculation.

Table 5: Changes in the Skilled-Unskilled Hours Differential

	Short-Run Change $\Delta(\overline{h^s}-\overline{h^u})_{SR}$	Long-Run Change $\Delta(\overline{h^s}-\overline{h^u})_{LR}$
Data	1.36	
Benchmark	2.68	0.07

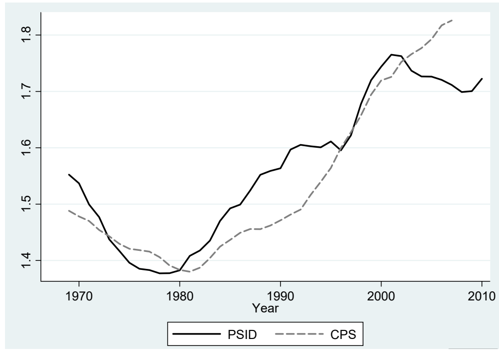
Table 6: Dispersion in Hours and Wage-Hour Correlation: Model vs. Data

	$CV(h)$				$Corr(wx,h)$			
	Model		Data		Model		Data	
	S	U	S	U	S	U	S	U
1967	0.0972	0.0745	0.1830	0.1821	−0.0206	−0.0835	−0.1959	−0.1735
2000	0.1269	0.1077	0.1964	0.2170	0.2736	0.1338	−0.0597	−0.1308
Change	0.0297	0.0332	0.0134	0.0349	0.2942	0.2173	0.1362	0.0427

Table 7: Recalibrated Parameter Values for Sensitivity Analysis

	Benchmark	$\sigma=1$	$\sigma=3$	Frisch elasticity=0.2	Frisch elasticity=0.6
β^s	1.0036	1.0035	1.0034	1.0041	1.0032
β^u	0.9946	0.9963	0.9900	0.9943	0.9947
ψ^s	22.023	19.555	30.077	233.34	10.010
ψ^u	26.649	19.787	63.553	300.91	11.870

Figure 1: Trends in Skill Premium for U.S. Men



Note: The skill premium is the ratio between skilled men's average hourly wage and unskilled men's.

Figure 2: Estimated Variances of Persistent and Transitory Wage Shocks

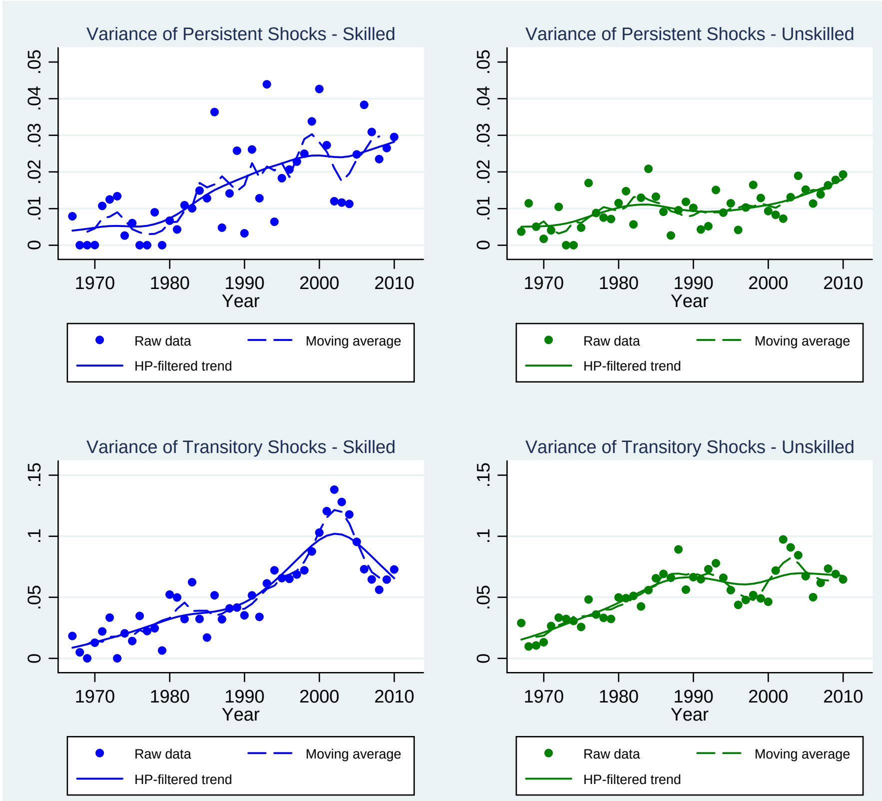
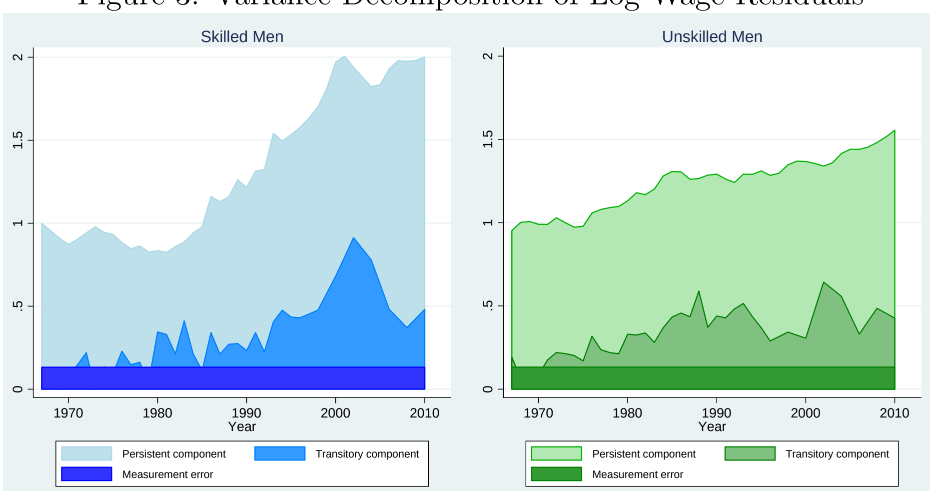
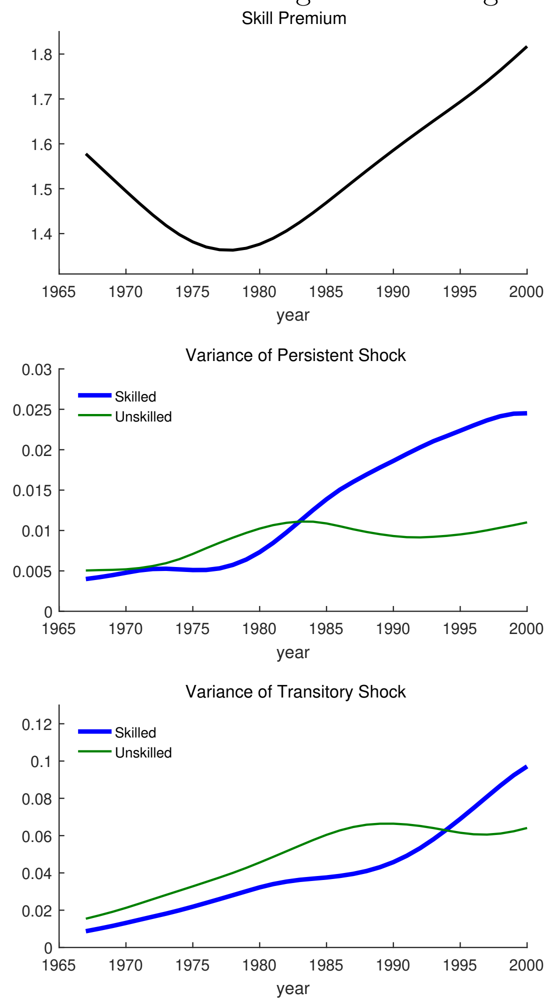


Figure 3: Variance Decomposition of Log Wage Residuals



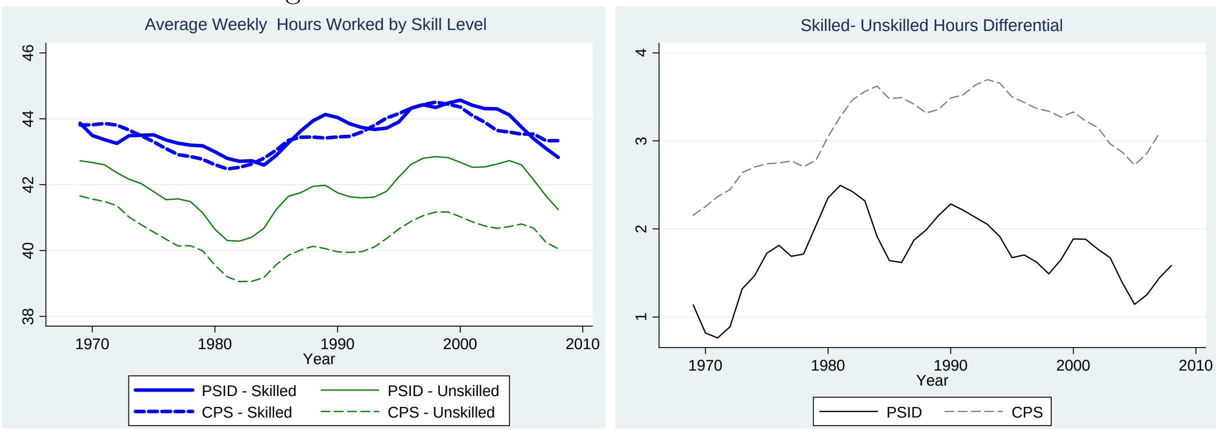
Note: The estimated variance of total wage residual of skilled men in 1967 is normalized to 1.

Figure 4: Summarized Changes in the Wage Structure



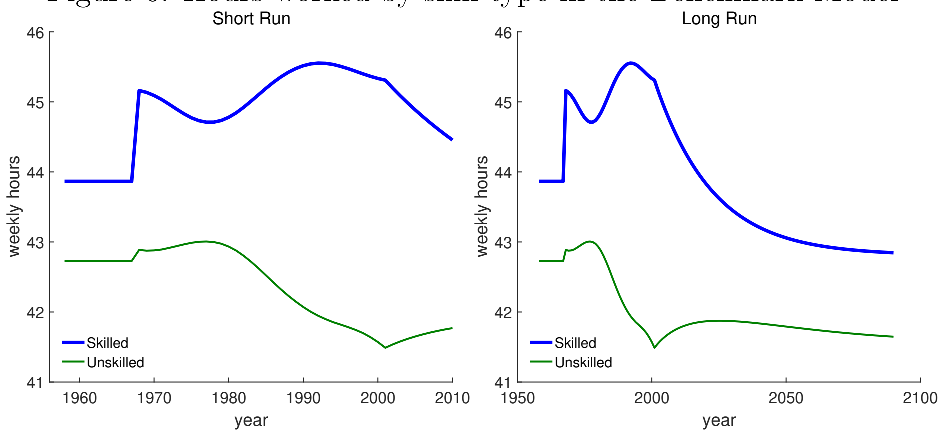
Note: This figure displays trends in the skill premium (top), the variance of persistent wage shocks by skill level (middle), and the variances of transitory wage shocks by skill level (bottom). The trends are extracted by the HP filter.

Figure 5: Trends in Male Hours Worked in the U.S.



Note: These trends are based on U.S. men who worked at least 260 hours in the past year. Both panels depict 5-year moving averages.

Figure 6: Hours worked by skill type in the Benchmark Model



Note: The left and right panels are from the same exercise. They are drawn in different scales to emphasize the different pattern of hours in the short run vs. long run.

Figure 7: Transition Dynamics in the Benchmark Model

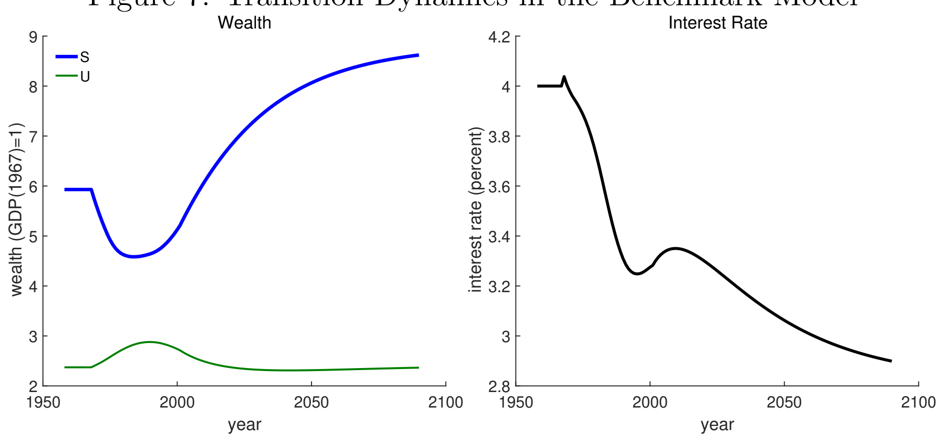
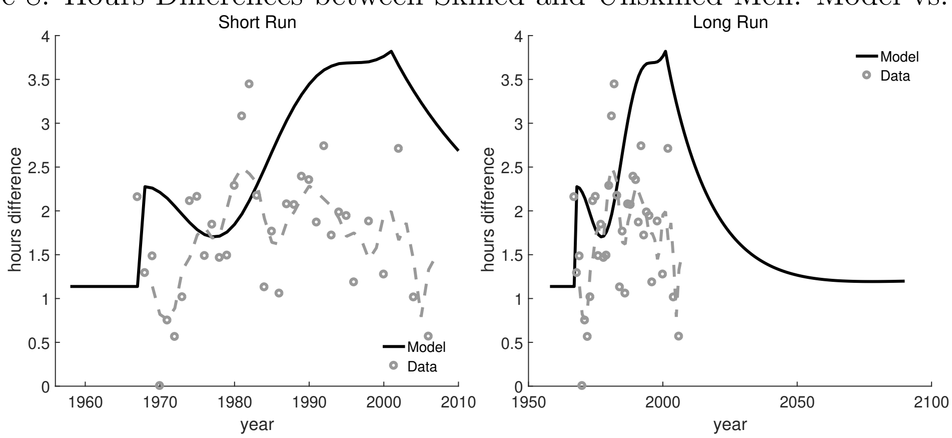
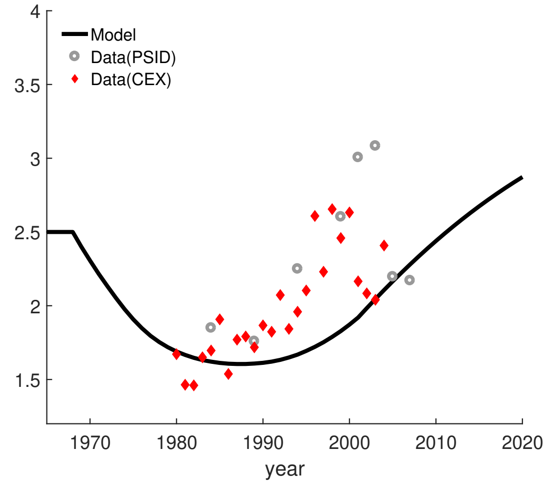


Figure 8: Hours Differences between Skilled and Unskilled Men: Model vs. Data



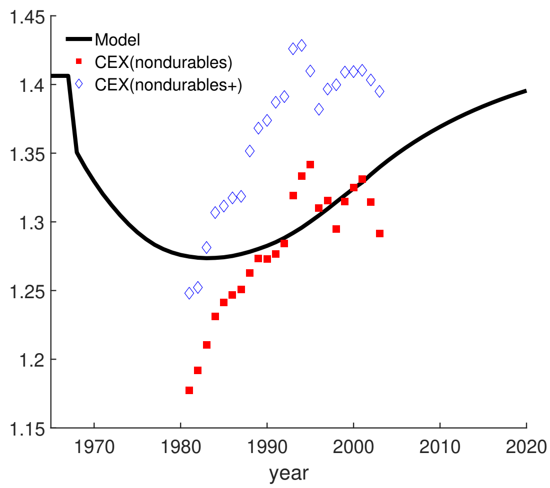
Note: The left and right panels are from the same exercise. They are drawn in different scales to emphasize the different pattern of hours in the short run vs. long run.

Figure 9: Relative Wealth of Skilled to Unskilled Households: Model vs. Data



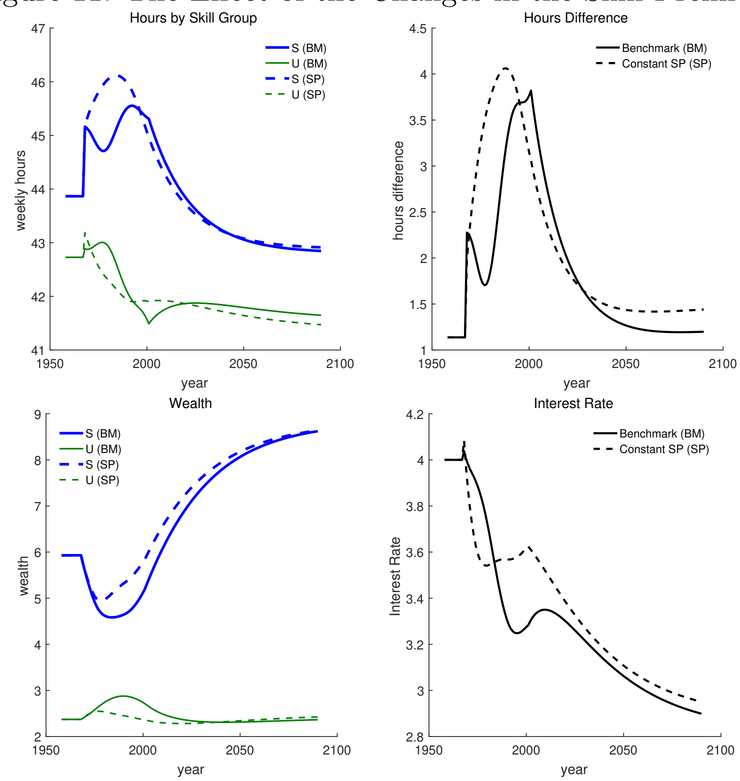
Note: The wealth ratio from the PSID is based on household net worth excluding net farm and business assets. Wealth from the CEX is the sum of house value and financial assets.

Figure 10: Consumption of Skilled Households Relative to Unskilled: Model vs. Data



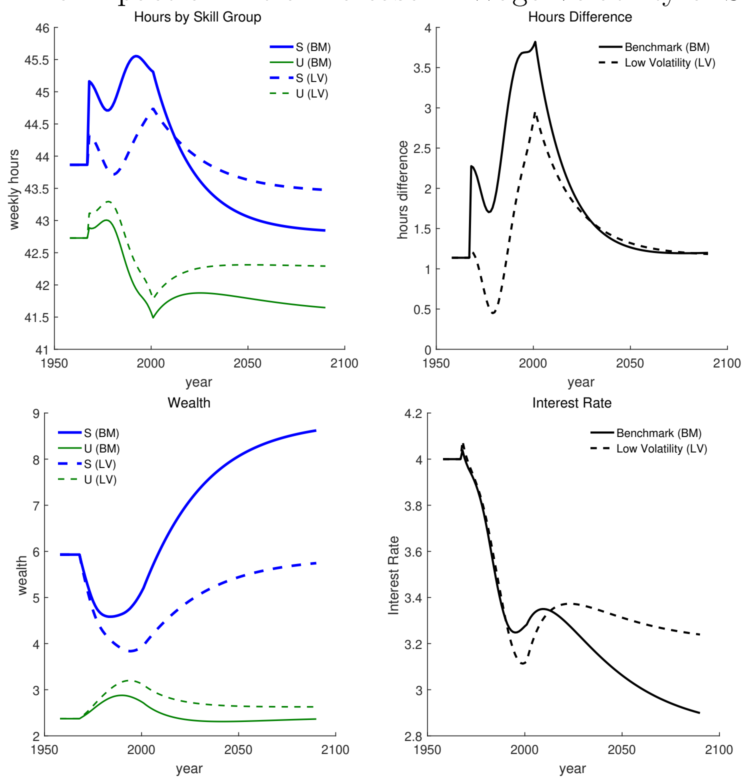
Note: Nondurables indicates household nondurables consumption and nondurables+ includes nondurables, services, small durables, and imputed service flows from houses and cars. Both series are from the CEX.

Figure 11: The Effect of the Changes in the Skill Premium



Note: The counterfactual labeled as 'SP' removes any changes in the skill premium from the benchmark model.

Figure 12: The Impact of Extra Increase in Wage Volatility of Skilled Men



Note: In the counterfactual labeled as'LV', both skilled and unskilled men face the same increase in the variance of each component of wages.

Figure 13: Welfare Changes from the Changing Wage Structure

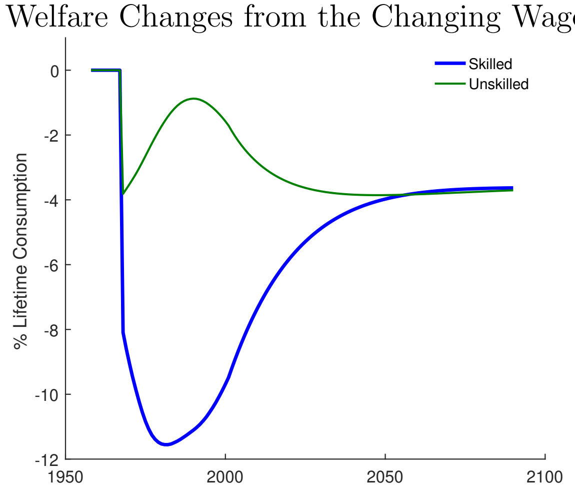
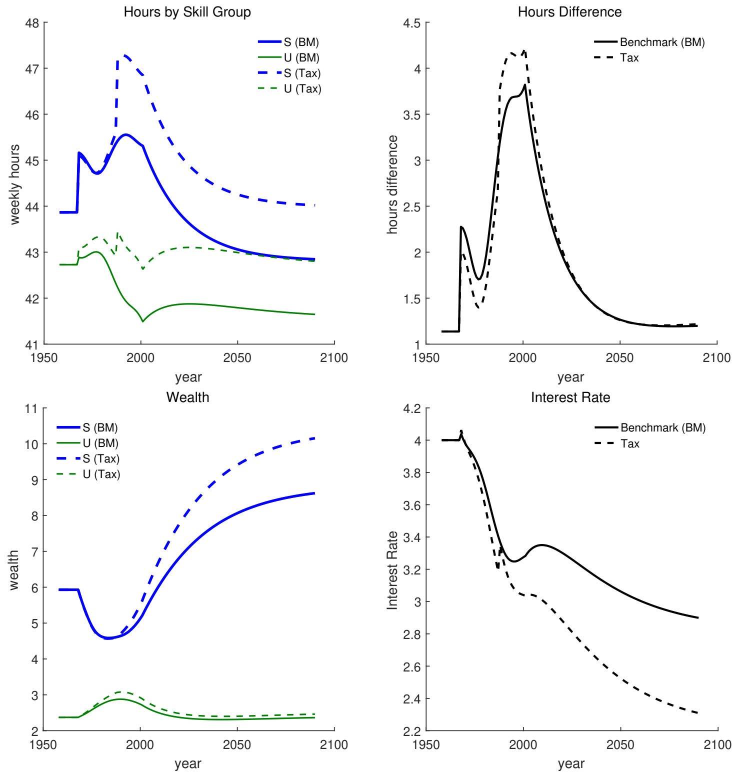
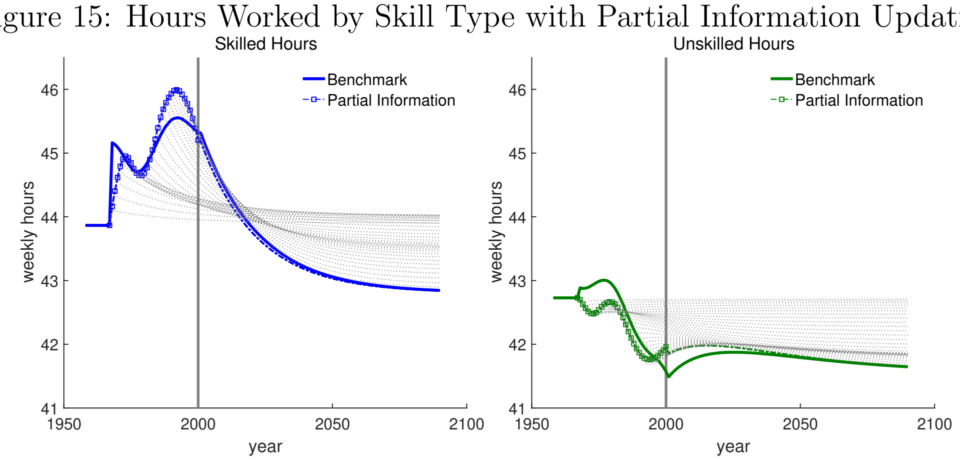


Figure 14: The Effects of the Tax Reform in 1986 in Transitional Dynamics



Note: In this experiment, households are surprised by the tax change effective in 1986 as well as other exogenous driving forces in 1967.

Figure 15: Hours Worked by Skill Type with Partial Information Updating



Note: Each square indicates one-period information update. Dotted lines indicate hypothetical paths without any additional information update.

Figure 16: Skilled-Unskilled Hours Differential with Partial Information Updating

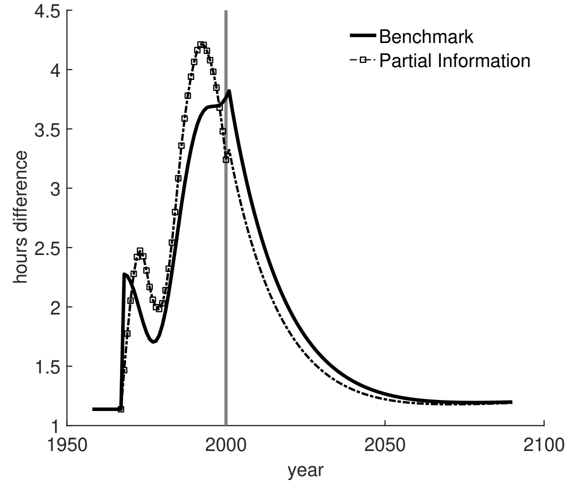


Figure 17: Sensitivity Analysis with Log Utility

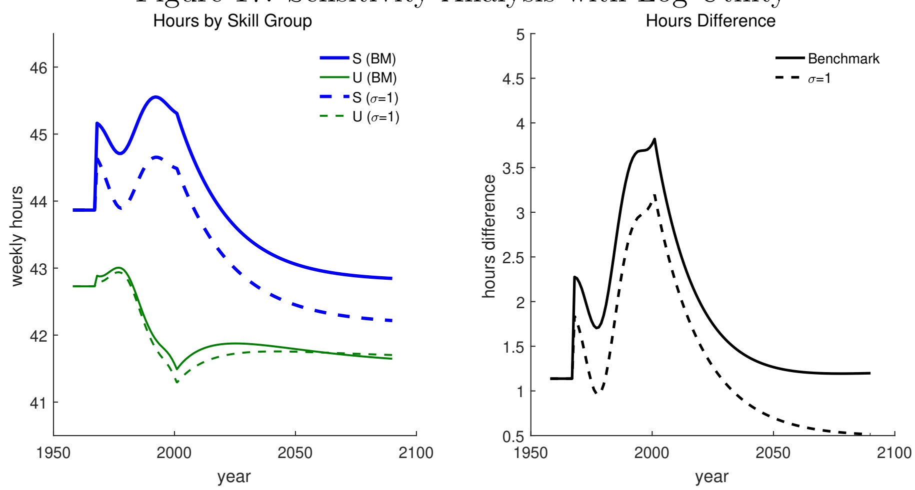


Figure 18: Sensitivity Analysis with $\sigma = 3$

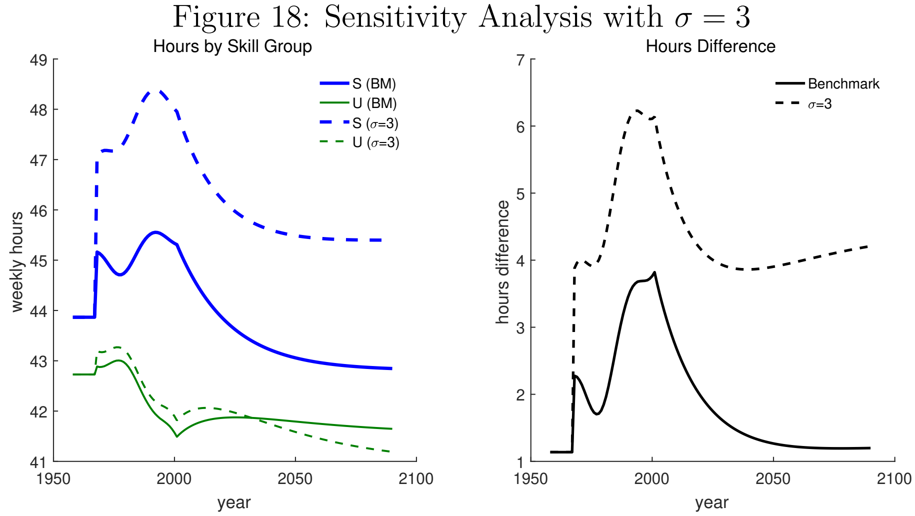


Figure 19: Sensitivity Analysis with with Frisch Elasticity of 0.2

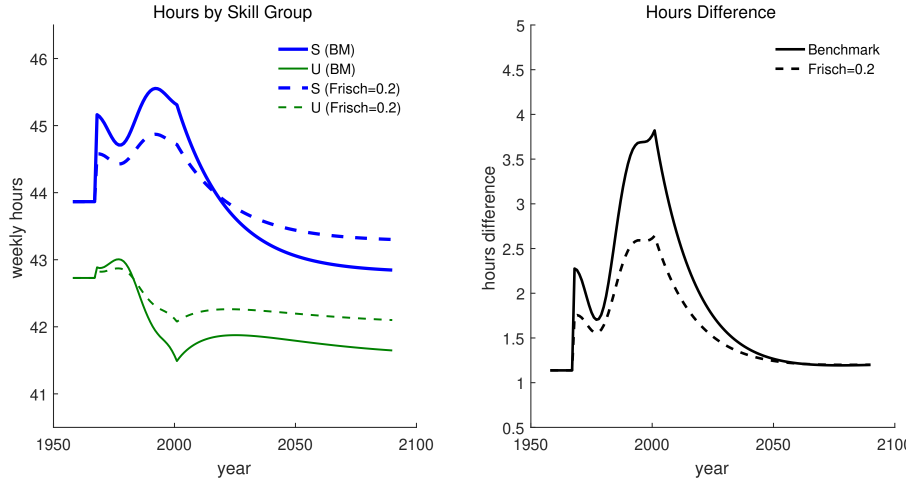
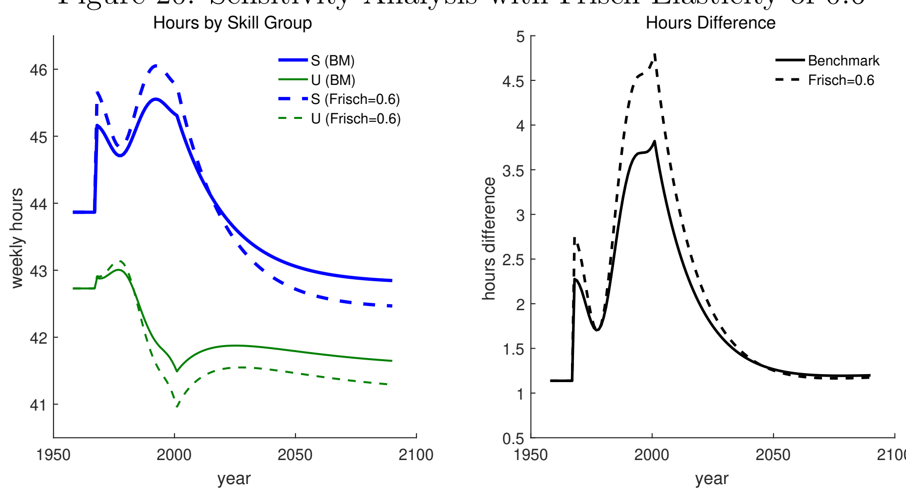


Figure 20: Sensitivity Analysis with Frisch Elasticity of 0.6



Appendix

(For Online Publication)

Appendix I. Data

We use data from Panel Study of Income Dynamics (PSID) in 1968 through 2011 to document how wages, hours worked, consumption, and wealth by skill level have evolved. Since the survey is conducted biennially beginning 1997, we exploit a total of 37 surveys from PSID. We also exploit data from Current Population Survey (CPS) March Supplements to document the trends in the skill premium and the skilled-unskilled hours differential for comparison purposes.

PSID: We begin with the core Survey Research Center (SRC) sample, which represents the U.S. population in 1968. We restrict our sample to male heads of household who participated in the labor market last year (i.e., worked at least 260 hours).We only include men aged between 25 and 59 with non-missing information about their educational attainment, total annual work hours, and labor income in the past calendar year. Hourly wage is obtained by dividing annual labor income by annual work hours. Extreme outliers are excluded from our sample by dropping those whose reported annual hours worked in the past year are more than 5840 hours, or whose hourly wage is less than half the federal minimum wage. Self-employed men are also removed because it is difficult to distinguish between labor and capital shares out of their income. The resulting sample is an unbalanced panel. We also extract information on household wealth and consumption for this sample. Household wealth is defined by net worth based on farm and business assets, checking/savings accounts, stocks, IRA/private annuities, net worth of vehicles, home equity, net worth of other real estate, other assets, and other debts. These variables are available for years 1984, 1989, 1994, 1999, 2001, 2003, and 2007. Household consumption is defined by the sum of food at home and food away from home, available for all survey years except for 1973, 1988, and 1989. This food expenditure is divided by the number of adult equivalents, where adult equivalent is defined by (number of adults $+0.7$(number of children))$^{0.7}$ according to the census equivalence scale. We use the consumption per adult equivalent for our analysis.

CPS: We apply the same sample selection criteria to data from the CPS. We include men aged between 25 and 59 with reported educational attainment. We obtain annual hours worked by multiplying previous calendar year's weeks worked by usual hours worked per week. In surveys before 1976, usual hours worked per week in the past year are not available and weeks worked in the past year are coded in intervals. Therefore, we impute both variables for previous surveys using the average weeks worked and the average usual hours worked per week in the same education and weeks worked interval cells in the 1976 survey. Annual earnings in the CPS are income from wages and salaries. We multiply top-coded earnings by 1.5, following Katz and Murphy (1992). Hourly wage is annual earnings divided by annual hours worked. As with the PSID, we exclude those who worked less than 260 hours and more than 5840 hours in the past year, who earned less than half the federal minimum wage per hour, or who is currently self-employed.

Appendix II. Estimation of Wage Processes

As described in section 2.1, following Heathcote et al. (2010), we model the log wage residual y^e_{it} as the sum of persistent and transitory shocks with time-varying variances:

$$y^e_{it} = \mu^e_{it} + \upsilon^e_{it} + \theta^e_{it}$$

where μ^e_{it} is a persistent component, $\upsilon^e_{it} \sim (0, \lambda^{e,\upsilon}_t)$ is a transitory component, and $\theta^e_{it} \sim (0, \lambda^\theta)$ is measurement error. The persistent component μ^e_{it} is assumed to follow an AR(1) process:

$$\mu^e_{it} = \rho^e \mu^e_{it-1} + \eta^e_{it}$$

where ρ^e is the persistence and $\eta_{it}^e \sim (0, \lambda_t^{e,\eta})$ is a persistent wage shock whose variance $\lambda_t^{e,\eta}$ varies over time. The initial value of the persistent component is drawn from a skill-specific distribution: $\mu_0^e \sim (0, \lambda^{e,\mu})$. We assume that all four variables, υ_{it}^e, θ_{it}^e, η_{it}^e, and μ_0^e are orthogonal and i.i.d. across individuals.

As we mention in 2.1, we take the estimate of 0.02 from French (2004) for the variance, λ^θ, of measurement error. Then, we estimate a parameter vector Φ^e, which includes two time-invariant parameters ρ^e and $\lambda^{e,\mu}$ and a set of time-varying parameters $\{\lambda_t^{e,\eta}, \lambda_t^{e,\upsilon}\}_{t=1967}^{2006}$ for each skill group $e \in \{s, u\}$. Since the PSID are available biennially beginning in 1997, we do not have empirical moments for transitory shocks for years 1997 and 1999. In order to resolve this issue, we assume that the cross-sectional variance of log residual wages in these missing years is the average of that in the previous year and in the subsequent year, and identify the variances of transitory wage shock for missing years as is done in Heathcote et al. (2010).

For each sample year t, we construct 10-year adjacent age cells from ages 29 to 54 such that, for instance, the age group 29 consists of those aged 25 to 34 years. We then compute the empirical autocovariance, $\widehat{g}_{a,t,n}^e$, of all possible orders for each age/year (a, t) cell in our PSID sample using log wage residuals $\widehat{y}_{it}^e$ from the first-stage regressions:

$$\widehat{g}_{a,t,n}^e = \frac{1}{I_{a,t,n}^e} \sum_{i=1}^{I_{a,t,n}^e} \widehat{y}_{it}^e\big|_{a_{it}=a} \cdot \widehat{y}_{i,t+n}^e\big|_{a_{it}=a}, \quad n \geqslant 0,$$

where $I_{a,t,n}^e$ is the number of observations for nth order autocovariance for age/year (a, t) cell in skill group e. We then pick the parameters $\widehat{\Phi}^e$ that minimize the equally weighted distance between this empirical autocovariance matrix and its theoretical counterpart:

$$\widehat{\Phi}^e = \arg\min_{\Phi^e} \left[\widehat{G}^e - G^e(\Phi^e)\right]' I \left[\widehat{G}^e - G^e(\Phi^e)\right],$$

where $\widehat{G}^e$ is a stacked vector of empirical autocovariances as well as cross-sectional variances for missing years, $G^e(\Phi^e)$ is the theoretical counterpart, and I is an identity matrix. Our estimation strategy is closest to that in Heathcote et al. (2010). As a robustness check, we estimate the variances of persistent and transitory wage shocks of the whole sample as Heathcote et al. (2010) do and confirm that our estimates are consistent with theirs.

Appendix III. Algorithm

A set of parameters β^s, β^u, ψ^s, and ψ^u is calibrated to match targets in the initial steady state. We allow the parameters $\mathbf{\Lambda}_t = \{\chi_t, \pi_t, \lambda_t^{s,\eta}, \lambda_t^{s,\upsilon}, \lambda_t^{u,\eta}, \lambda_t^{u,\upsilon}\}$, governing the skill premium, the population share of the skilled, and persistent and transitory idiosyncratic productivity, to vary over time.

Solving for the steady state

Under a set of parameters for the steady state (including $\mathbf{\Lambda}_*$),

1. Guess price r. Given this guess, compute w^u and w^s (given $\mathbf{\Lambda}_*$).

2. Solve the value function and get $h(e, a, \mu^e, \upsilon^e), a'(e, a, \mu^e, \upsilon^e), c(e, a, \mu^e, \upsilon^e)$.
3. Generate a sample of population N over $(e, a, \mu^e, \upsilon^e)$ space. That is, $N_S = \pi_* N$ sample over (a, μ^s, υ^s) and $N_U = (1 - \pi_*)N$ sample over (a, μ^u, υ^u).
4. Compute aggregate statistics:

$$
\begin{aligned}
K &= \sum_e \int a dG(e, a, \mu^e, \upsilon^e) \\
S &= \int h(s, a, \mu^s, \upsilon^s)\ \exp(\mu^s + \upsilon^s)\ dG(s, a, \mu^s, \upsilon^s) \\
U &= \int h(u, a, \mu^u, \upsilon^u)\ \exp(\mu^u + \upsilon^u)\ dG(u, a, \mu^u, \upsilon^u) \\
H &= \left\{\chi U^\phi + (1 - \chi) S^\phi\right\}^{\frac{1}{\phi}}
\end{aligned}
$$

   And define

$$\tilde{r} = z\alpha \left(K/H\right)^{\alpha - 1} - \delta$$

5. Check if $|r - \tilde{r}| < \epsilon$. If not, update r and go back to step 1.

Solving transition economy (Changing parameters over time)

The economy was originally at the initial steady state $(*)$. There is a gradual change for $\mathbf{\Lambda}$ from $\mathbf{\Lambda}_*$ to $\mathbf{\Lambda}_{**}$ for periods $t = t_1, \cdots, t_\tau$, and then the economy converges to the new steady state $(**)$ at $t_T (> t_\tau)$.

1. A sequence of parameters is given: $\{\omega_t, \lambda_t^{s,\eta}, \lambda_t^{s,\upsilon}, \lambda_t^{u,\eta}, \lambda_t^{u,\upsilon}\}_{t=t_1}^{t_\tau}$.
2. Solve two (initial and final) steady states and find a set of parameters $\{\chi_t, r_t\}_{t=t_0=*}$, $\{\chi_t, r_t\}_{t=t_T=**}$, where χ_t is adjusted so that the model matches the observed skill premium (sp_t) in each steady state. Record the stationary distribution $G_*(e, a, \mu^e, \upsilon^e)$ and the value function $V_{**}^e(a, \mu^e, \upsilon^e)$.
3. Guess the sequences of $\{\chi_t, r_t\}_{t=t_1}^{t_T - 1}$. Then, the other prices are given by

$$
\begin{aligned}
w_t^s &= w^s(\chi_t, r_t; \alpha, \delta, \phi) \\
w_t^u &= w^u(\chi_t, r_t; \alpha, \delta, \phi)
\end{aligned}
$$

$$
\begin{aligned}
w_t^u &= z(1 - \alpha)\chi_t \left(\frac{r_t + \delta}{z\alpha}\right)^{\frac{\alpha}{\alpha - 1}} \left(\chi_t + (1 - \chi_t)\left(\frac{\chi_t}{1 - \chi_t} \cdot \omega_t\right)^{\frac{\phi}{\phi - 1}}\right)^{\frac{1 - \phi}{\phi}} \\
w_t^s &= w_t^u \cdot \omega_t
\end{aligned}
$$

4. Solve for workers decisions backwards and get the decision rules during the transition periods:

(a) Using the value function at the new steady state: $V^e_{**}(a, \mu^e, \upsilon^e; \mathbf{\Lambda}_{**})$,

(b) In each period during the transition ($t = t_1, \cdots, t_{T-1}$), solve

$$V^e_t(a_t, \mu^e_t, \upsilon^e_t) = \max_{c_t, a_{t+1}, h_t} \{u(c_t, h_t) + \beta^e \gamma \mathbb{E}_t V^e_{t+1}(a_{t+1}, \mu^e_{t+1}, \upsilon^e_{t+1})\}$$

subject to constraints with $p_t = \{r_t, w^u_t, w^s_t\}$ and $\mathbf{\Lambda}_t$ for $t = t_1, \cdots, t_{T-1}$.
Solve the problem above backwards from $t = t_{T-1}, \cdots, t_1$ to obtain the decision rules: $c_t(e, a_t, \mu^e_t, \upsilon^e_t)$, $h_t(e, a_t, \mu^e_t, \upsilon^e_t)$, $a'_t(e, a_t, \mu^e_t, \upsilon^e_t)$ and the value function, $V^e_t(a_t, \mu^e_t, \upsilon^e_t)$.

5. Given a distribution at t, $G_t(e, a_t, \mu^e_t, \upsilon^e_t)$, simulate $G_{t+1}(\cdot)$ by applying $c_t(\cdot), h_t(\cdot)$, and $a'_t(\cdot)$. Start from $t = t_1$where we know $G_{t_1}(\cdot)$ from the distribution of the initial steady state $G_*(\cdot)$

   (a) First calculate the aggregate statistics:

$$\begin{aligned} K_t &= \sum_e \int a_t dG_t(e, a_t, \mu^e_t, \upsilon^e_t) \\ S_t &= \int h_t(s, a_t, \mu^s_t, \upsilon^s_t) \exp(\mu^s_t + \upsilon^s_t) dG_t(s, a_t, \mu^s_t, \upsilon^s_t) \\ U_t &= \int h_t(u, a_t, \mu^u_t, \upsilon^u_t) \exp(\mu^u_t + \upsilon^u_t) dG_t(u, a_t, \mu^u_t, \upsilon^u_t). \end{aligned}$$

   (b) Using these aggregate statistics, compute

   - an updated share $\tilde{\chi}_t$ (which matches the observed skill premium (sp_t)) such that

$$\tilde{\chi}_t = \frac{(S_t/U_t)^{\phi-1}}{\omega_t + (S_t/U_t)^{\phi-1}},$$

     where $\omega_t = w^s_t / w^u_t$.
     Note that $sp_t = [w^s_t \times E(\exp(\mu^s_t + \upsilon^s_t))]/[w^u_t \times E(\exp(\mu^u_t + \upsilon^u_t))]$. Hence, $\omega_t = sp_t \times E(\exp(\mu^u_t + \upsilon^u_t))/E(\exp(\mu^s_t + \upsilon^s_t))$.
   - and an implied price $\tilde{r}_t$ such that

$$\tilde{r}_t = z_t \alpha \left(K_t / \tilde{H}_t\right)^{\alpha-1} - \delta,$$

     where

$$\tilde{H}_t = \left\{\tilde{\chi}_t U^\phi_t + (1 - \tilde{\chi}_t) S^\phi_t\right\}^{\frac{1}{\phi}},$$

   (c) Get an updated distribution G_{t+1} and go back to (5a) until we get G_{t_T}.

6. Check if the guessed sequences $\{\chi_t, r_t\}^{**}_{t=*}$ are close enough to the model-implied sequences of $\{\tilde{\chi}_t, \tilde{r}_t\}^{**}_{t=*}$ from (5b). If not, update the guesses and go back to step 3.