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Why Did the Average Duration of Unemployment Become So Much Longer?

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Abstract

This paper examines the causes of the observed increase in the average duration of unemployment over the past thirty years. First we analyze whether changes in the demographic composition of the U.S. labor force, particularly the age and gender composition, can explain this increase. We then consider the contribution of institutional changes, such as the change in the generosity and coverage of unemployment insurance. We find that changes in the composition of the labor force and institutional changes can only partially account for the longer duration of unemployment. We construct a job search model and calibrate it to U.S. data. The results indicate that more than 70 percent of the increase in the duration of unemployment over the past thirty years can be attributed to an increase in within-group wage inequality.

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1 Introduction

It is commonly observed that unemployment duration moves together with the unemployment rate. Recent data show that this relationship is changing. Figure 1 shows the U.S. unemployment rate (left scale) and average unemployment duration (right scale) between 1948 and 2003.[1] We can observe a clear discrepancy between the two in recent years.

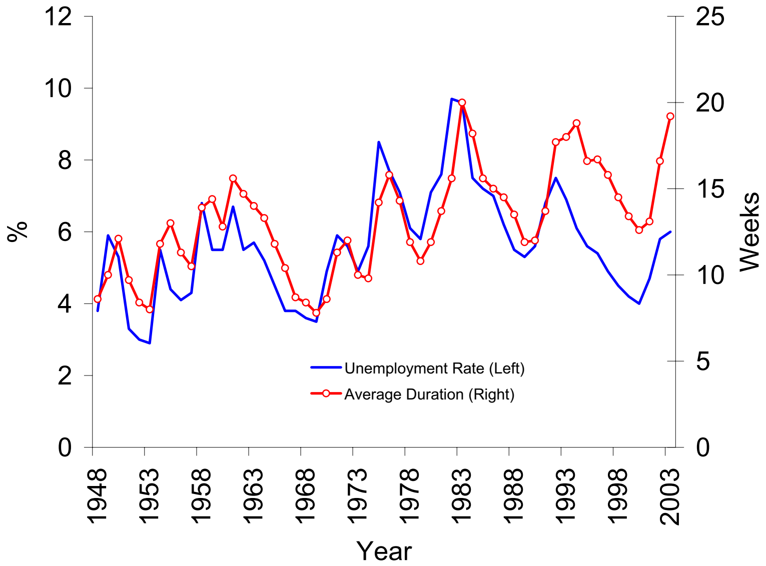

Figure 1: Unemployment Rate (Left Scale) and Average Unemployment Duration (Right Scale).
Data Source: Current Population Survey

This tendency is more apparent when we compare the trends of these two series. Figure 2 compares the Hodrick-Prescott (HP) trend of the two series.[2] Casual observation of Figure

[1]The data in this paper are taken from Bureau of Labor Statistics (BLS) website, `www.bls.gov`, except where noted otherwise.

[2]For HP-filtering, we use $\lambda = 6.25$ for yearly data following Ravn and Uhlig's (2002) suggestion.

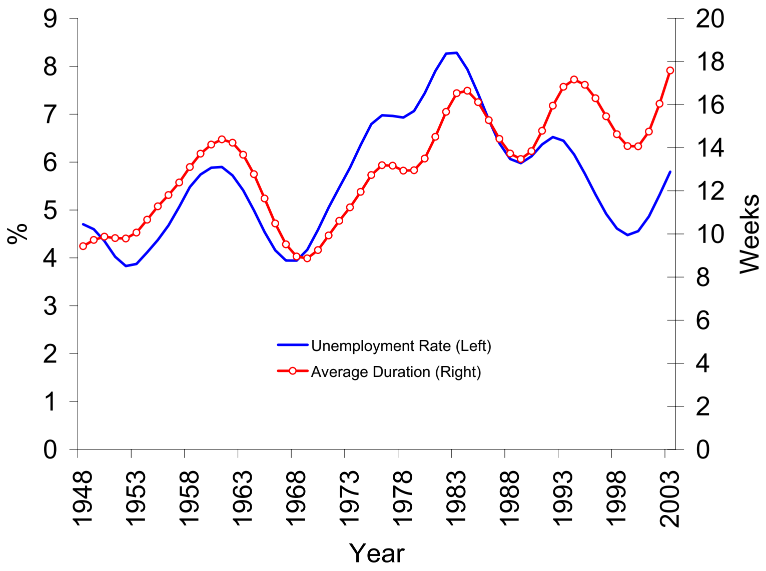

Figure 2: Trends of Unemployment Rate (Left Scale) and Average Unemployment Duration (Right Scale).

Data Source: Current Population Survey

2 suggests that the difference of the trend has been particularly pronounced in recent years. There has been a dramatic decline in the U.S. unemployment rate during the past twenty years. The average duration of unemployment, however, remained high during the 1990s. In higher frequency however, the two series seem to have maintained a stable relationship (Baker, 1992). Figure 3 shows the HP-detrended series. The two series are moving together, even in recent years.

There are other studies that document the increase in the average duration of unemployment. Baumol and Wolff (1998), Valletta (1998), Abraham and Shimer (2001), Juhn, Murphy, and Topel (1991, 2002) also point out that unemployment duration is becoming longer in recent years despite the low levels of the unemployment rate.

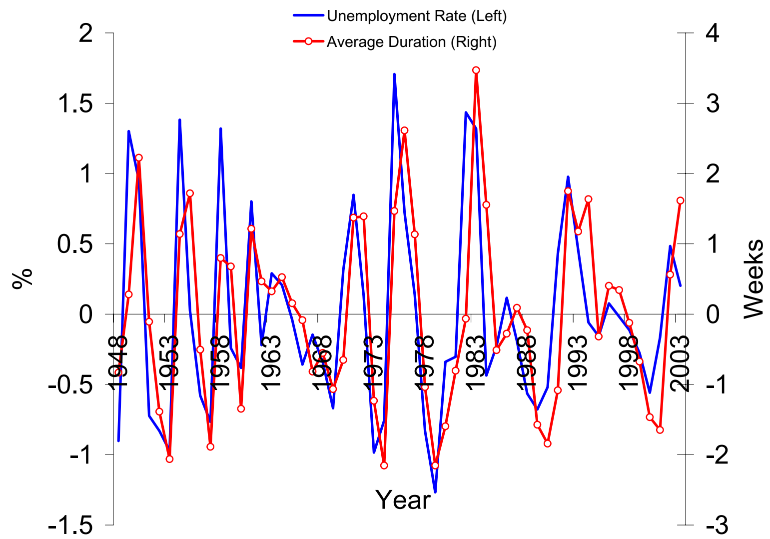

Figure 3: Detrended Unemployment Rate (Left Scale) and Average Unemployment Duration (Right Scale).

Data Source: Current Population Survey

Valletta (1998) examines how the incidence and duration of unemployment changed during 1967-1998 and shows that there is an upward trend in the duration of unemployment. He argues that rising incidence and duration of permanent job loss can account for most of the increase in the duration of unemployment. His study does not explicitly analyze the underlying economic reasons of this phenomenon. He speculates that a rapid technological change, changing job search strategies, and measurement issues can be possible explanations.

Abraham and Shimer (2001) focus on the changes in the measurement of unemployment duration and demographic changes as the sources of the observed increase in the duration. They find that the redesign of the Current Population Survey (CPS) in 1994 explains a half a week increase in the duration. They then examine two important demographic changes in the composition of the U.S. labor force which are likely to increase the duration of unemployment: the aging of the baby-boom cohort and the increase in women's labor-force attachment. They conclude that the first factor explains a half week increase in the duration. Therefore, in total, a one week increase can be attributed to the measurement issues and aging of the baby boom cohort. They attribute the remainder of the increase to the increase in women's unemployment duration.

There is no doubt that the increase in women's labor force attachment explains a part of the increase in the average duration. However, provided that a significant increase in average duration can be observed in the samples that consist of only males, a substantial amount of the increase in duration remains unexplained. For example, Juhn, Murphy, and Topel (1991, 2002) point out that for the samples of prime-aged men, unemployment duration is becoming longer in recent years (see Figure 3 in their 1991 paper and Figure 11 in their 2002 paper).

We investigate the economic reason of this phenomenon. In particular, we construct a search model of unemployment and examine the effect of the change in the economic environment on the worker's search activity. By evaluating the model quantitatively, we show that an increase in the dispersion of wages can have a large effect on the search length of an unemployed worker. Therefore, the observed recent increase in the wage dispersion can be a main economic force increasing the unemployment duration.

Juhn, Murphy, and Topel (1991) emphasize the labor-supply response to the change in the wage *level*. Their work can be considered as complementary to ours, which emphasize the change in wage *dispersion*. In her comments to Juhn, Murphy, and Topel (1991), Yellen (1991) argues that

> When there is wage dispersion, so that both good and bad jobs are available for workers with given skills, some workers will choose to remain unemployed, searching for good, rent-paying jobs, rather than work at the poor jobs that are readily available... The long-term unemployed are searching for work for which they are qualified. In this interpretation, unemployment is a response to

> wage dispersion rather than to wage levels, contrary to the authors' labor supply function, in which labor supply depends only on wage levels. (pp. 129–130)

Our hypothesis parallels her argument — the recent change in wage distribution may have had a significant impact on the duration of unemployment.

Baumol and Wolff (1998) argue that technological progress can explain the increase in the duration of unemployment. When technological progress is more rapid, the relative cost of hiring a worker whose training cost is higher (e.g. unskilled and/or old workers) increases. As a consequence, the supply of jobs available to these workers decreases, thereby increasing their duration of job search.

Our theory does not rule out technological change as a main cause of the longer unemployment duration. In fact, Violante (2002) shows that the recent increase in embodied technological progress can explain a large part of the rise in wage dispersion which is not explained by worker characteristics. Our theory, combined with Violante's story, can provide another channel through which technological progress can affect the duration of unemployment.

In the next section, we briefly examine the time-series properties of average unemployment duration. In Section 3, we examine if demographic change is responsible for the increase in the duration of unemployment. In Section 4, we discuss if institutional change can explain the change in unemployment duration. In Section 5, we present our model and analyze it quantitatively. Section 6 concludes.

2 Time-Series Properties

Before exploring possible explanations for the change in unemployment duration, we examine some time-series properties of the data. Let D_t be the average unemployment duration at year t. For the whole sample period (t = 1948 to 2003), the average of D_t is 13.2 weeks. From Figure 1, we clearly observe an increase in average D_t. To identify the timing of the regime shift, we apply the idea of the Sup F test by Andrews (1993). Let us denote the mean

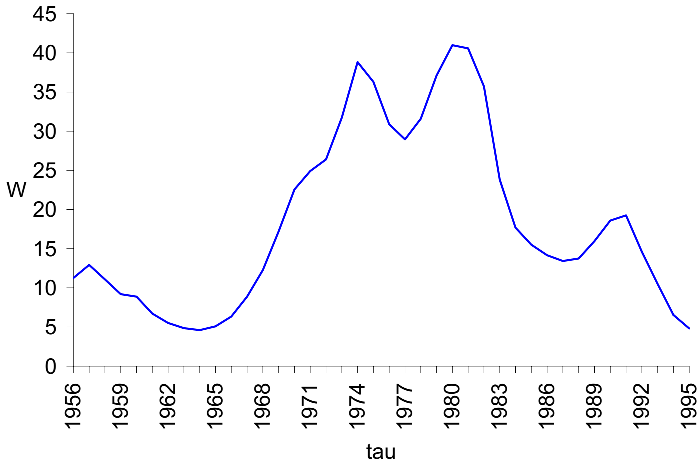

Figure 4: W Statistic.

of the whole sample as $m \equiv [\sum_{t=1948}^{2003} D_t]/T$. The sample size, T, is 56. Let S be the sum of the square mean-deviation : $S \equiv \sum_{t=1948}^{2003}(D_t - m)^2$.

Here, we would like to identify the timing of the increase in average D_t. To achieve this, we divide the sample period into two different subperiods. We denote the break point by τ. Then, we calculate the mean of the first subperiod $m_1 \equiv [\sum_{t=1948}^{\tau} D_t]/T_1$. Here, the sample size $T_1 = \tau - 1948 + 1$. Let S_1 be the sum of the square mean-deviation: $S_1 \equiv \sum_{t=1948}^{\tau}(D_t - m_1)^2$. Similarly, for the second subperiod, $m_2 \equiv [\sum_{t=\tau+1}^{2003} D_t]/T_2$, where $T_2 = 2003 - \tau$. Let S_2 be the sum of the square mean-deviation: $S_2 \equiv \sum_{t=\tau}^{2003}(D_t - m_2)^2$.

We calculate a Wald-type statistic, $W \equiv T(S - S_1 - S_2)/(S_1 + S_2)$, for each τ. Following convention (Maddala and Kim 1998, p.395), we consider $\tau = 1956, ..., 1995$ (disregarding the first 15% and the last 15%). A large value of this statistic indicates a structural change.

Figure 4 plots the series of W. The value of W exceeds 25 between 1972 and 1982.[3] It seems likely that structural change occurred during this period. W reaches its maximum at

[3]The asymptotic critical value for 1% significance (against the null hypothesis of no structural change during the period) is 12.35 (Andrews 1993, Table 1).

1980.[4] The average of D_t for $t = 1948, ..., 1980$ (m_1) is 11.5 weeks, while the average of D_t for $t = 1981, ..., 2003$ (m_2) is 15.5 weeks. Therefore, the difference before and after the change ($m_2 - m_1$) when $\tau = 1980$ is 4.0 weeks. When we select a τ from the period between 1972 and 1982, the difference in average duration, $m_2 - m_1$, takes a value between 3.4 weeks to 4.0 weeks. In the next section, we explore whether various explanations can account for this magnitude of increase in average unemployment duration.

3 Demographic Adjustment

In this section, we examine if the change in demographic composition can account for the recent increase in average unemployment duration. In particular, we focus on the composition in age and gender.

First we divide the unemployed people into two gender groups: men and women. Then each group is divided into seven age groups: $A_m = \{$ men 16-19, men 20-24, men 25-34, men 35-44, men 45-54, men 55-64, men 65+ $\}$ and $A_w = \{$ women 16-19, women 20-24, women 25-34, women 35-44, women 45-54, women 55-64, women 65+ $\}$.[5] Therefore, we have fourteen demographic groups. Let $f_t(i)$ be the fraction of unemployed workers who are in group i at time t, and let $D_t(i)$ be the average duration of unemployment for workers who are in group i at time t. Then, by definition, the average duration of unemployment at time t is

$$D_t = \sum_{i \in A_m \cup A_w} f_t(i) D_t(i). \tag{1}$$

Equation (1) implies that if the fraction of the group who experiences longer unemployment spells increases, then the average unemployment duration increases.

Table 1 shows that in general, older workers experience a substantially longer unemployment duration than younger workers (except for the 65+ age group). It also shows that women's duration tends to be shorter than men's (except for the 25-34 age group). Figure

[4] When we specify the process of D_t as an AR(1) process, the W statistic calculated from the residuals reaches its peak at 1982.

[5] These age categories coincide with the age groups used by the BLS.

	16-19	20-24	25-34	35-44	45-54	55-64	65+
Men	12.0	16.7	17.9	22.5	24.8	26.8	24.8
Women	11.2	15.1	18.0	20.3	23.2	25.2	21.3

Table 1: Average Duration of Unemployment for Each Demographic Group (Weeks, Year 2003)

Data Source: Current Population Survey

5 shows the average age of unemployed workers from 1948-2003.[6] Since the 1970s, there has been some increase in the average age, mainly due to the aging of the baby-boom cohort. Figure 6 shows the fraction of women in the unemployment pool. Here we observe some increase over time. In the following, we explore the effect of demographic changes on the duration of unemployment.

We consider two thought-experiments. The first is "What if the people in each group behaved the same but the demographics (the composition of the population) changed as in reality?" Second, "What if the demographics (the composition of the population) remained the same but the people in each group changed their behavior as in reality?" The first experiment captures the effect of the change in composition of the population, and the second experiment captures the effect of people's changing behavior.

3.1 First Experiment — Pure Composition Effect

We compute how much the unemployment duration has increased because of the demographic change. We examine the *demographic* change in unemployment duration[7] by computing

$$D^d_{t_1,t_0} = \sum_{i \in A_m \cup A_w} f_{t_1}(i) D_{t_0}(i). \tag{2}$$

For robustness, we examine two different base years (t_0): 1967 and 2003.

Figure 7 shows $D^d_{t_1,t_0}$ for $t_0 = 1967$ and $t_0 = 2003$. Neither of the series have an apparent trend. In fact, if we calculate the average of $D^d_{t_1,t_0}$ before 1980 and after 1981, there is a

[6] This figure is computed by assigning the middle value to each group and taking weighted sum. We assigned 70 to the 65+ group.

[7] We follow Shimer's (1998) terminology.

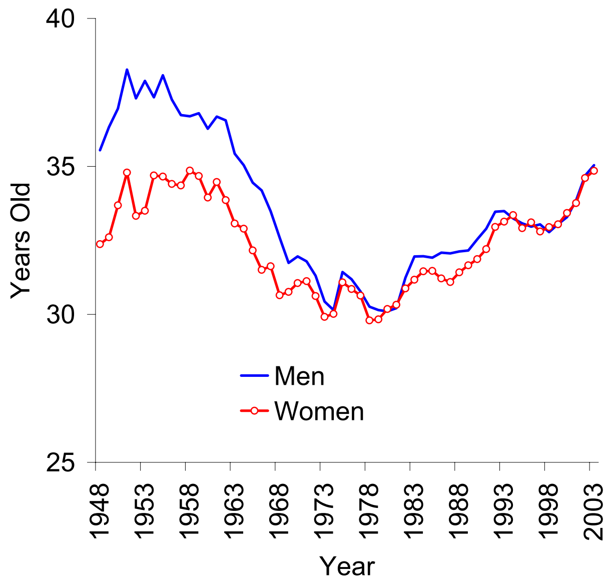

Figure 5: Average Age of Unemployed Workers

slight decline. Even when we take the difference between the smallest $D^d_{t_1,t_0}$ ($t_1 = 1972$ or 1973) and the largest $D^d_{t_1,t_0}$ in recent years ($t_1 = 2003$), it is less than two weeks, which is much smaller than the observed increase in Section 2.

To see the effect of the adjustment more clearly, we subtract $D^d_{t_1,t_0}$ (with $t_0 = 2003$) from the trend of average duration (shown in Figure 2). Figure 8 shows this series. The structural change is even more apparent than in previous figures – there is a clear shift in the mean between the late 1970s and the early 1980s.

3.2 Second Experiment — Changing Behavior

Next, we compute how much of the increase in the duration of unemployment would have occurred if the demographics were the same. This part (*genuine* change in unemployment

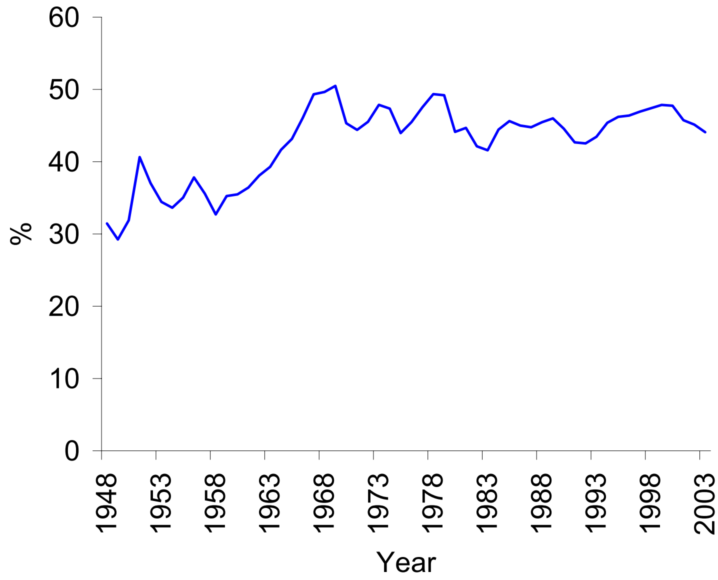

Figure 6: The Fraction (%) of Women in Unemployment Pool

duration) measures the change in average duration that is not accounted for by the demographic change. Here, we hypothetically assume that demographics remained the same from period t_0 to t_1, and that the duration of unemployment for each group $(D_t(i))$ followed the same pattern as the data. Then, the unemployment duration would be

$$D^g_{t_1,t_0} = \sum_{i \in A_m \cup A_w} f_{t_0}(i) D_{t_1}(i). \tag{3}$$

Figure 9 shows $D^g_{t_1,t_0}$ for $t_0 = 1967$ and $t_0 = 2003$.[8] They both show a very similar trend to Figure 1. In fact, the differences between the average before 1980 and after 1981 are 3.4 weeks ($t_0 = 1967$) and 3.9 weeks ($t_0 = 2003$). These are slightly smaller than the value in Section 2, but still substantial.

[8]The graph starts from 1967 due to the data limitation.

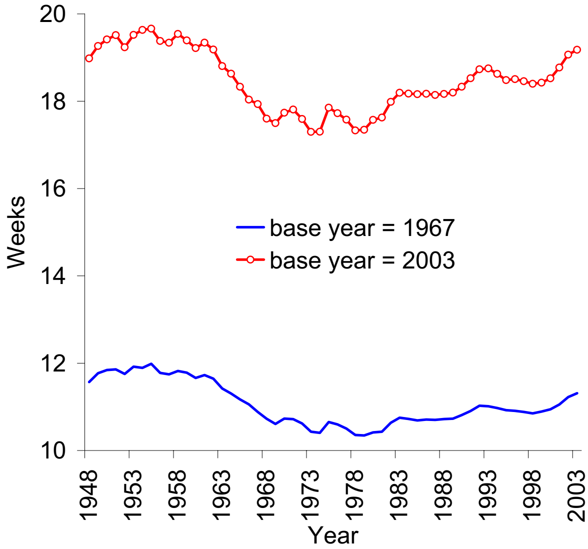

Figure 7: $D^d_{t_1,t_0}$ for $t_0 = 1967$ and $t_0 = 2003$

4 Institutional Change

An alternative candidate for an explanation is that this increase in unemployment duration might have been caused by the change in institutions. In fact, there have been some changes in the unemployment insurance system during the post-war period. Theoretically, if unemployment insurance became more generous, it may have an effect which lengthens the duration of unemployment. Baicker, Goldin, and Katz (1997) describe the changes of the unemployment insurance system since the 1930s. Although there have been increases in coverage in the 1970s (mainly including public sector workers), they argue that the generosity of the unemployment insurance has remained almost constant, and that the ratio of unemployment insurance claims to total unemployment has been declining during the post-war era (see their Figure 7.2). Consistent with this observation, Burtless (1983) argues that the

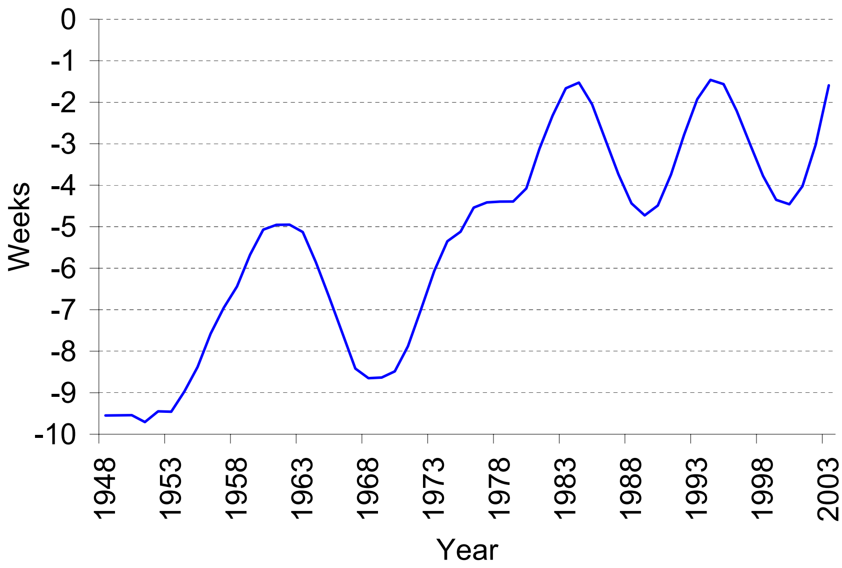

Figure 8: The deviation of actual average duration (HP trend) from $D^d_{t_1,t_0}$ ($t_0 = 2003$).

insured unemployment rate has been stable throughout the 1950s to the 1970s.[9] Therefore, change in the unemployment insurance system is not likely to be an explanation for the longer unemployment duration.[10] Baumol and Wolff (1998) also support this view. They examine the effect of institutional changes on the duration of unemployment and conclude that institutional factors like changes in the coverage and generosity of unemployment insurance, changes in the rate of unionization, and changes in the minimum wage cannot account for the observed increase in the duration of unemployment.

[9]Baumol and Wolff (1998) report that the insured coverage rate (the percent of unemployed workers receiving benefits) has been dropping if the data is extended until 1990s.

[10]Ehrenberg and Oaxaca (1976) estimate that moving from a welfare system without an unemployment insurance to a system similar to the U.S. unemployment insurance system increases the duration of unemployment by one week for old (age 45-59) male, and by less than one week for the other demographic groups. This result suggests that even if the increase in coverage increases the unemployment duration, its quantitative impact is small.

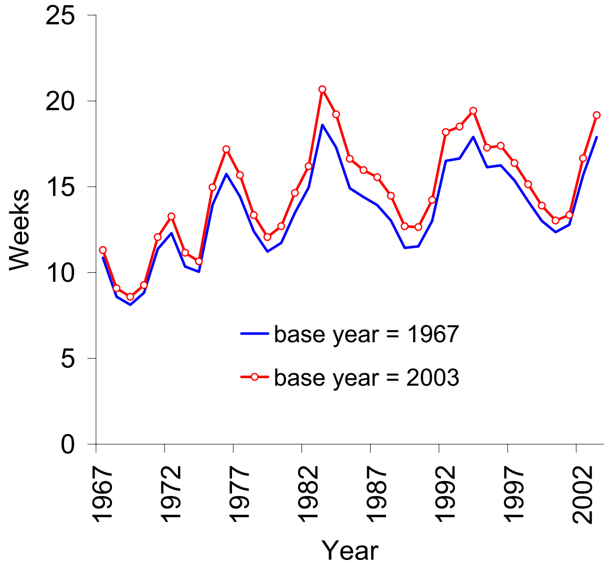

Figure 9: $D^{g}_{t_1,t_0}$ for $t_0 = 1967$ and $t_0 = 2003$

5 Model

In Sections 3 and 4 we explored if demographic changes and institutional changes can account for the observed increase in average unemployment duration. Both of these factors, however, have not had a large enough effect to drive the larger change in average duration. In this section, we construct a search model and investigate several possible explanations.

5.1 Setup

We employ a variant of the search model by McCall (1970). Consider a worker who is unemployed and searching for a job. We assume that there is no borrowing or saving, and that the period utility for an unemployed worker is u_s. For an employed worker receiving wage w, the momentary utility is $u_e(w) \equiv \ln(w)$.

An unemployed worker receives one wage offer each period. He decides whether or not to accept it. Once he accepts, he will work at that wage until he is separated from the job. If he rejects, the search continues in the subsequent period. The separation occurs exogenously at the probability $\alpha \in [0,1)$ every period. After the separation, the worker is unemployed for at least one period.

We assume that the wage offer is independently and identically distributed, and follows a lognormal distribution

$$\ln(w) \sim \mathcal{N}(m - \sigma^2/2, \sigma^2).$$

Therefore,

$$E[w] = e^m \text{ and } Var[w] = e^{2m}(e^{\sigma^2} - 1).$$

The worker's problem in each period is characterized by the following Bellman's equation:

$$v(w) = \max \left\langle u_e(w) + \beta \left\{ (1-\alpha)v(w) + \alpha \left[u_s + \beta \int v(w')dF(w') \right] \right\}, u_s + \beta \int v(w')dF(w') \right\rangle. \tag{4}$$

where $F(\cdot)$ is the distribution function of the wage offer.

This problem has a simple reservation-wage property: the worker accepts the wage offer if w is above the reservation wage, and rejects if it is below the reservation wage. The reservation wage $\bar{w}$ solves (derivation is in the Appendix)

$$u_e(\bar{w}) - u_s = \frac{\beta}{1 - \beta(1-\alpha)} \int_{\bar{w}}^{\infty} \left[u_e(w') - u_e(\bar{w}) \right] dF(w'). \tag{5}$$

Let $\lambda \equiv F(\bar{w})$ be the probability that an unemployed worker is still unemployed next period.

5.2 Calibration

Equation (5) is scale-free in the sense that if m is replaced by $m + \mu$ and u_s is replaced by $u_s + \mu$, the reservation wage $\bar{w}$ will become $\bar{w} \cdot e^{\mu}$ and λ remains the same. Therefore, we can normalize m. Here, we set $m = 0$. We set one month as one period. Therefore, $\beta = 0.947^{1/12}$.[11]

[11] The value for the annual discount rate β is taken from Cooley and Prescott (1995).

In the following sections, we conduct two comparative statics to evaluate how the change in parameters affects the duration of unemployment. There have been two dramatic changes in the labor market in recent years:

1. Change in the incidence of unemployment.

2. Change in the dispersion of wages.

Here, as a benchmark, we set the parameters so that the model matches the data before 1970. As is examined in the next section, the incidence of unemployment (the probability of becoming unemployed during each month) was stable around 0.025 before 1970. Therefore, we set $\alpha = 0.025$.

The other parameters, u_s and σ, are set so that the following two conditions are satisfied.

1. The average duration of unemployment, $1/(1-\lambda)$, is approximately 2.5 (10 weeks).

2. The $90\% - 10\%$ log wage difference of employed workers is 0.9 (From Katz and Autor 1999, Figure 5d).

These two conditions provide $u_s = -6.2$ and $\sigma = 0.65$. Note that when we compute the $90\% - 10\%$ log wage difference, we only look at the accepted wage offers (which can be observed in the data).

5.3 Change in the Incidence of Unemployment

The dynamics of the unemployment rate u_t is governed by

$$u_{t+1} = \alpha(1-u_t) + \lambda u_t,$$

where the first term in the right-hand-side is the number of workers separated at time t, and the second term in the right-hand-side is the number of workers who are unemployed at time t and rejected the time-t job offer (as before, λ is defined as $\lambda \equiv F(\bar{w})$). In steady-state, u_t is constant (call it $\bar{u}$), and

$$\bar{u} = \frac{\alpha}{1+\alpha-\lambda}. \tag{6}$$

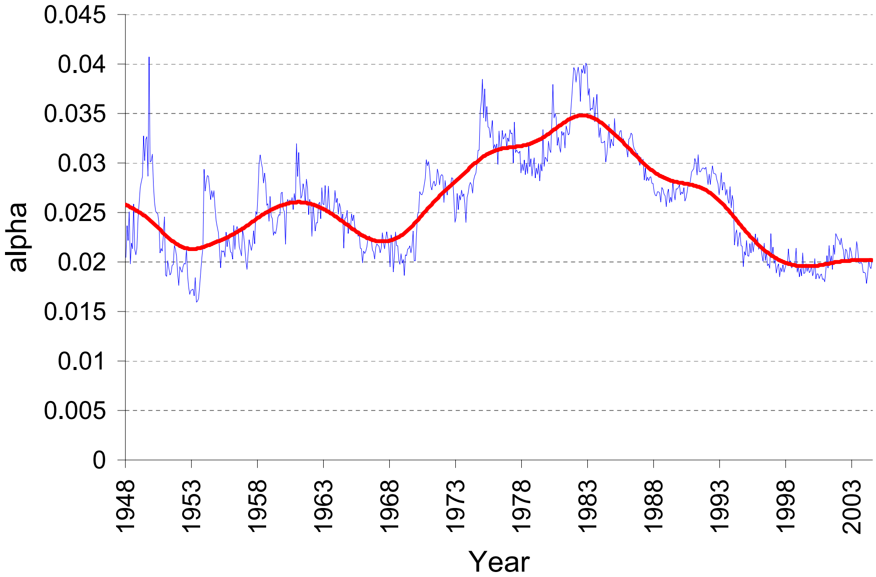

Figure 10: Incidence of unemployment

As is apparent from (6), $\bar{u}$ is increasing in both α and λ. When λ increases (longer duration), $\bar{u}$ goes down only if α decreases by a sufficient amount.

Indeed, this can be seen in the data. Figure 10 plots the incidence of unemployment (the number of the workers unemployed less than five weeks divided by the total employment) using monthly BLS series. The thick line depicts the HP-trend.[12] Before 1970, the incidence was stable around 0.025. It went up to 0.035 in the 1980s, and then went down to 0.020 in the 1990s. The very small value in recent years reflects the coexistence of a low unemployment rate and a long unemployment duration.[13]

[12]We use $\lambda = 129600$ for monthly data following Ravn and Uhlig (2002).

[13]In fact, Juhn, Murphy, and Topel (2002) argue that the decrease in the unemployment rate observed in the 1990s is driven almost entirely by the decreased incidence of unemployment. Even though the duration of unemployment remained as high as its level in the 1980s, the lower rate of separation dragged the unemployment rate down to very low levels.

α	$\bar{w}$	Avg. Duration
0.020	1.04	11.4 weeks
0.025	0.96	10.0 weeks
0.030	0.90	9.2 weeks
0.035	0.84	8.4 weeks

Table 2: Effect of Change in α

In the following, we conduct comparative statics by comparing the model outcome in steady-states. Focusing on the steady-state simplifies our calculation, and considering that we analyze a long-run trend which is much longer than the unemployment spells (average 10 weeks), we belive that it provides a fairly good approximation.

Qualitatively, it is clear from (5) that $\bar{w}$ is decreasing in α, so a higher α leads to a shorter duration. The intuition is that a worker becomes less selective to the job when the probability that the job will be terminated is high. Quantitatively, Table 2 summarizes the average duration of unemployment for different values of α. The difference between $\alpha = 0.025$ (before the 1970s[14]) and $\alpha = 0.020$ (recent years) is 1.4 weeks, which is smaller than half of the change in unemployment duration that we identified in Section 2 (3.4 weeks to 4.0 weeks). Moreover, in the 1980s the duration was long even though α was large. This deepens the puzzle. In the next section, we turn to the next possible factor.

5.4 Change in the Dispersion of Wages

Many labor economists have observed that there are substantial wage differentials for observationally equivalent workers. Mincer-style wage equations typically explain less than 30% of the overall wage variation. The remaining variation, which is more than 70%, is often called within-group (or residual) wage inequality.

There is a large body of theoretical literature which attempts to explain the existence of wage dispersion among workers of the same characteristics. The most popular approach

[14] This value is the average for 1948-1979. If we take the average for 1948-1970, the value is somewhat lower at $\alpha = 0.024$.

is to utilize a model of search and matching in a frictional labor market (see, for example, Burdett and Mortensen 1998). In the following analysis, we do not attempt to explain the source of the within-group wage inequality. Instead, we take the wage dispersion as given for the workers.

Beginning with Katz and Murphy (1992) and Juhn, Murphy, and Pierce (1993), researchers have noticed that there has been a significant increase in within-group wage inequality during recent years. Some recent studies investigate the causes of this increase in within-group wage inequality. For example, Violante (2002) argues that the rapid investment-specific technological change in recent years is the major cause of the increase in within-group wage inequality. In our model, we take this increase in within-group wage inequality as given, and analyze how this increase can change the behavior of the workers who are searching for a job.

Here we evaluate the impact of an increase in σ. Katz and Autor (1999, Figure 5d) show that male within-group wage inequality increased from 0.9 to 1.2 over the past 30 years. Thus, we set a new value of σ, so that the $90\% - 10\%$ log wage difference is 1.2.

It is well known that in McCall-style search models, a mean-preserving spread in the wage offer distribution increases the reservation wage. An unemployed worker tends to wait longer since the option value of a job opportunity (opportunity cost of accepting a job) increases with the variance of the wage offer. When the variance increases, the possibility of receiving a very low wage and very high wage both increase. An increase in the probability of a very low-wage offer does not affect the value of waiting, since those offers are always rejected anyway. Higher probability of a very high-wage offer, however, increases the value of waiting since those are the offers that are accepted and affect the worker's welfare. Therefore, the increase in variance makes the job opportunity better.

Table 3 summarizes how $\bar{w}$ and the average duration change when σ changes. When σ changes from 0.65 to 0.93 so that the $90\% - 10\%$ log wage difference matches the data (0.9 and 1.2, respectively), $\bar{w}$ increases from 0.96 to 1.02 and the average duration of unemployment

90% − 10% log wage difference	σ	$\bar{w}$	Avg. Duration
0.9	0.65	0.96	10.0 weeks
1.0	0.75	0.99	11.1 weeks
1.1	0.84	1.01	12.0 weeks
1.2	0.93	1.02	12.8 weeks

Table 3: Effect of Change in σ

increases by 2.8 weeks.[15] This value is between 70% to 82% of the amount of increase in data.[16] Here we are keeping the value of $\alpha = 0.025$. When we include the effect of the change in α by changing to $\alpha = 0.02$, this model "over-explains" the data — the average duration becomes 4.8 weeks longer.[17] Considering that all of the increase in the within-group wage inequality may not correspond to the increase in dispersion of wage offer, it is reasonable to think that these two factors explain a large part (but perhaps not all) of the increase in the average duration.

To see the robustness of the result, we carried out the same experiment for different period lengths (which coincides with the frequency of the wage offers). The calibrated values of u_s, σ (when the log wage difference is 0.9), and σ' (when the log wage difference is 1.2) are shown in Table 4. (β and α are adjusted as the period length changes.) Table 4 also shows the reservation wages under σ (denoted as $\bar{w}$), the reservation wages under σ' (denoted as $\bar{w}'$), and the increase in the unemployment duration. We consistently observe an increase in the unemployment duration of more than two weeks. In each case, the impact of increasing wage inequality is quantitatively substantial in explaining the longer unemployment duration.

Not all the within-group wage inequality may be the pure "luck" factor (good or bad wage draw in the model) from the viewpoint of the workers. It is possible that some part of the within-group wage inequality is due to the unobservable characteristics of the workers,

[15] Also, in this model, higher $\bar{w}$ implies a larger dispersion in the unemployment duration, since the coefficient of variation in the search length is $\lambda^{1/2}$.

[16] In Section 2, we estimated the increase to be between 3.4 weeks and 4.0 weeks. The above numbers follow since $(2.8/4.0) \times 100 = 70\%$ and $(2.8/3.4) \times 100 = 82\%$.

[17] In this case, σ changes to 0.98 to match the 90% − 10% log wage difference of 1.2.

Period Length	u_s	σ	σ'	$\bar{w}$	$\bar{w}'$	Increase in Duration
2 weeks	-5.6	0.79	1.10	1.42	1.57	1.9 weeks
4 weeks	-6.2	0.65	0.93	0.96	1.02	2.8 weeks
6 weeks	-7.0	0.55	0.81	0.75	0.76	2.7 weeks

Table 4: Results for Different Period Lengths

which is known by the workers but not by the econometricians. Gottschalk and Moffitt (1994, Table 2) show that during their sample period (both subperiods of 1970-78 and 1979-87), the permanent component explains 2/3 of the log-wage dispersion, and the temporary component accounts for 1/3. Their permanent component estimate includes the returns to education, so it is a plausible estimate that 1/2 of the within-group inequality is transitory, i.e. due to the "luck" factor. Therefore, we re-set the initial value of σ to match $90\% - 10\%$ log wage difference of 0.45 (which makes $\sigma = 0.32$ and $u_s = -3.0$), and see how the change in σ changes the average duration of unemployment.

$90\% - 10\%$ log wage difference	σ	$\bar{w}$	Avg. Duration
0.45	0.32	1.03	10.0 weeks
0.50	0.37	1.07	11.2 weeks
0.55	0.42	1.10	12.1 weeks
0.60	0.47	1.13	12.9 weeks

Table 5: Effect of Change in σ

Table 5 shows the result. The change in the average duration is vary similar to Table 3. In particular, when the $90\% - 10\%$ log wage difference increases from 0.45 to 0.6, the average duration increases by 2.9 weeks.

6 Conclusion

In this paper, we examined the causes of the increase in the U.S. average unemployment duration in recent years. We showed that demographic change can account for only a small fraction of the observed increase in the average duration of unemployment. Institutional change does not seem to be a major factor, either. By quantitatively evaluating a search model, we showed that both the decrease of separation and the increase in wage dispersion (within-group wage inequality) can make the average duration longer. The effect of the wage dispersion is particularly large, and it alone accounts for 70-82% of the observed increase in duration.

Clearly, wages and unemployment (incidence and duration) are determined simultaneously in the labor market. Our analysis is a first step towards a better understanding of the interaction between them in the context of the recent U.S. economy. More detailed and complete analysis of this interaction is an important future research agenda.

Appendix

A Derivation of (5)

First, we establish that the solution to (4) has a reservation-wage property. The second term of max does not depend on w. Suppose that $v(\cdot)$ is a continuous and nondecreasing function. Since $\lim_{x\to 0} u_e(x) = -\infty$ and $\lim_{x\to\infty} u_e(x) = \infty$ (and continuity), there exists at least one value of $\bar{w}$ which satisfies

$$u_e(\bar{w}) + \beta\left\{(1-\alpha)v(\bar{w}) + \alpha\left[u_s + \beta\int v(w')dF(w')\right]\right\} = u_s + \beta\int v(w')dF(w').$$

With such a $\bar{w}$, the optimal choice for the unemployed worker is to accept when $w \geq \bar{w}$ and to reject when $w < \bar{w}$.

In (4), with the supposition that $v(\cdot)$ is continuous and nondecreasing, the right-hand-side is also continuous and nondecreasing. It is also straightforward to see that (4), seen as a mapping, is a contraction mapping. Therefore, by the standard argument, it can be established that $v(\cdot)$ is in fact continuous and nondecreasing.

When $w \geq \bar{w}$, solving (4) for $v(w)$ yields

$$v(w) = \frac{u_e(w) + \beta\alpha\left[u_s + \beta\int v(w')dF(w')\right]}{1-\beta(1-\alpha)}. \tag{7}$$

Since this is equal to $u_s + \beta\int v(w')dF(w')$ when $w = \bar{w}$,

$$\frac{u_e(\bar{w}) + \beta\alpha\left[u_s + \beta\int v(w')dF(w')\right]}{1-\beta(1-\alpha)} = u_s + \beta\int v(w')dF(w'),$$

and therefore

$$\frac{u_e(\bar{w})}{1-\beta} = u_s + \beta\int v(w')dF(w'). \tag{8}$$

This can be rewritten as

$$\begin{aligned}\frac{u_e(\bar{w})}{1-\beta} &= u_s + \beta\int_0^{\bar{w}} v(w')dF(w') + \beta\int_{\bar{w}}^{\infty} v(w')dF(w') \\ &= u_s + \beta\int_0^{\bar{w}} \frac{u_e(\bar{w})}{1-\beta}dF(w') + \beta\int_{\bar{w}}^{\infty} \frac{u_e(w') + \beta\alpha\left[u_e(\bar{w})/(1-\beta)\right]}{1-\beta(1-\alpha)}dF(w'),\end{aligned} \tag{9}$$

where the second equality used the fact that $v(w) = u_s + \beta \int v(w')dF(w')$ when $w \leq \bar{w}$, (7), and (8). The left-hand-side can be decomposed as

$$\frac{u_e(\bar{w})}{1-\beta} = \int_0^{\bar{w}} \frac{u_e(\bar{w})}{1-\beta} dF(w') + \int_{\bar{w}}^{\infty} \frac{u_e(\bar{w})}{1-\beta} dF(w').$$

Using this to the left-hand-side of (9), subtracting $\beta \int_{\bar{w}}^{\infty} u_e(\bar{w})dF(w')/(1-\beta)$ from both sides, and rearranging yields (5). Since the left-hand-side of (5) is strictly increasing in $\bar{w}$ and the right-hand-side of (5) is nonincreasing in $\bar{w}$, the reservation wage $\bar{w}$ is unique.

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