"Rising Unemployment Duration in the United States: Causes and Consequences"

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May 2005

ABSTRACT

Since the mid-1970s, the duration of a typical unemployment spell in the United States has increased substantially relative to the unemployment rate. Using microdata on unemployment from the complete set of monthly Current Population Survey files for the period 1976-2004, I investigate the causes and consequences of rising unemployment duration. The duration of completed unemployment spells is formed using the synthetic cohort approach, and the data are adjusted for major changes in survey design that occurred in 1994. The empirical analysis focuses on two primary explanations for rising unemployment duration: changes in women's labor force attachment (Abraham and Shimer 2002) and changes in the incidence and duration of permanent job loss that relate to declining job security (Valletta 1999). The results provide support for both explanations, although the link to changing labor force attachment for women is less clear than the link to permanent job loss. These results bolster recent findings (Campbell and Duca 2004) suggesting that rising unemployment duration has lowered the aggregate unemployment rate that is consistent with stable wage and price inflation (the "NAIRU").

* I thank Jaclyn Hodges and Geoffrey MacDonald for outstanding research assistance. The views expressed in this paper are those of the author and should not be attributed to the Federal Reserve Bank of San Francisco or the Federal Reserve System.

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1. Introduction

The duration of unemployment spells has increased significantly relative to the unemployment rate over the past three decades (Figure 1). Several explanations have been offered for this trend. Abraham and Shimer (2002) argue that rising unemployment duration is concentrated among women and largely is a consequence of the increase in women's labor force attachment, which has reduced the incidence of short-term unemployment associated with transitions in and out of the labor force. By contrast, Valletta (1998) argued that rising unemployment duration is linked to declining job security through the key contribution of rising permanent job loss.

In this paper, I attempt to distinguish between these competing explanations for rising unemployment duration. I use microdata on unemployment from the complete set of monthly Current Population Survey files for the period 1976-2004, and I apply a "synthetic cohort" approach to estimate the expected completed duration of unemployment. The data and methods are described in Sections 2 and 3, including adjustment for major changes in survey methodology implemented in 1994 that altered measurement of key unemployment variables. Using these duration estimates and related covariates, including data on individual characteristics and reasons for unemployment, I directly test the competing explanations for rising unemployment duration. The tabulations and regression results presented in Section 4 reveal trends in unemployment by reason and unemployment duration. The trend toward rising duration has been evident throughout the sample frame and in fact strengthened after 1994. Group-specific regressions and formal decompositions uncover little or no direct support for the explanation based on women's labor force attachment, although evidence is uncovered suggesting that rising duration for labor force entrants has made an important contribution to the overall increase. Support for the job loss explanation is stronger, with the results largely confirming earlier findings of the key contribution of rising incidence and duration of permanent job loss to rising unemployment duration in the United States.

These findings are of interest for several reasons. Rising unemployment duration (at a given unemployment rate) is likely to have adverse implications for social welfare, by increasing the burden of uninsurable labor-income risk on workers (Abraham and Shimer 2002). Moreover, explaining the source of rising unemployment duration may have important implications for aggregate dynamics and monetary policy. Rising duration can raise or lower the wage pressures associated with a given unemployment rate, thereby altering the natural or "non-accelerating inflation rate" of unemployment (NAIRU). In the concluding section of the paper, these implications are discussed in light of the paper's specific findings.

2. CPS Unemployment Data

The data used in this study are constructed from the complete set of monthly survey records from the U.S. Current Population Survey for the period January 1976 through December 2004 (data files obtained from Unicon Research Corporation). Observations were pulled for all individuals identified as unemployed in the survey. The analyses in this draft were restricted to individuals age 16 to 64, with alternative sample restrictions imposed as described below. All tabulations were weighted by the sample survey weights to render them representative of the broader U.S. population.

The key analysis variables are the individual duration of unemployment and reasons for unemployment, along with identifiers for selected individual characteristics (sex, age, and education). Unemployment duration is measured as the duration of in-progress spells at the time of the survey. This variable is used to form estimates of expected completed duration for an individual entering unemployment in a particular month, as described in the next section. The reasons for unemployment identified in the survey fall into five categories: job losers, for whom the survey distinguishes between those on temporary layoff (i.e., those expecting recall to the firm from which they were laid off) and permanent job losers (permanent layoffs, firings, or completion of temporary jobs); voluntary job leavers; re-entrants to the labor force; and new entrants to the labor force.

The U.S. Bureau of Labor Statistics (BLS) has published monthly figures on unemployment duration since 1948 and monthly figures on unemployment by reason since 1967. Although the analyses in this manuscript are limited by data availability (Unicon CPS files and unemployment by reason) to years beginning in 1976, this not a significant constraint, as past research (Valletta 1998) found that the upward trend in duration was most pronounced beginning in the mid- to late-1970s.

Figure 2 displays several measures of unemployment duration that are published by the BLS. No clear trend is evident in the average spell duration, as the peaks and troughs in this series have been quite similar over the last three U.S. business cycles. On the other hand, because the unemployment rate has trended downward over the sample period (look ahead to Figure 3 for this series), average duration has been increasing relative to the unemployment rate Moreover, average duration has remained high farther into the last two recoveries than it did in the aftermath of the more severe 1982-1983 recession (consistent with descriptions of the two recent episodes as “jobless recoveries”). The share of short-term unemployment in total unemployment (unemployed <5 weeks) exhibits a slight downward trend. The share of very long-term unemployment (at least 6 months) has been relatively flat. Like the average duration measure, however, the decline in the overall unemployment rate over the past three decades implies that long-term unemployment has become more significant at a given unemployment rate. More precise measurement of the cyclical and trend properties of unemployment duration can be obtained through the estimation of expected completed duration, as described in the next section.

An important issue for these data is the impact of major redesign of the basic monthly CPS survey beginning with the January 1994 survey. Using data from a parallel survey administered in 1992-1993, Polivka and Miller (1998) found that the new survey design generated a trend break in the measured duration of unemployment, increasing it relative to its measurement using the earlier survey design, and also altered the calculation of unemployment shares by reason. I use Polivka and Miller’s adjustment factors to yield consistent time-series estimates of unemployment duration and unemployment by reason for estimation purposes (see the Appendix for additional details).[1] These adjustments necessarily are imprecise and do not guarantee elimination of survey redesign effects on the estimated trends. However, sub-sample results discussed in section 4 indicate that main trends identified in this paper also are evident when the analysis is stratified using the pre- and post-redesign periods, which suggests little or no influence of the redesign on my key results.

3. Completed Duration Estimates from Grouped Data

The CPS data described in the previous subsection provide information on the length of existing unemployment spells up to the date of the survey. The average duration measure formed from these data (and published by the BLS) will not in general correspond to the expected duration of a completed spell for a new entrant to unemployment, particularly under changing labor market conditions such as rising unemployment (i.e., “nonsteady state” conditions). The general nonsteady-state approach to estimating expected completed duration using grouped duration data is a “synthetic cohort” approach (see Kaitz 1970, Perry 1972, Sider 1985, Baker 1992a).[2] This approach relies on the estimation of monthly continuation rates—i.e., the probabilities that an unemployment spell will continue from one month to the next. These rates in general will vary over the length of a spell due to individual heterogeneity or underlying duration

[1] Abraham and Shimer (2002) argue that estimated rates of long-term unemployment (15 weeks or more) were largely unaffected by the survey redesign. Rather than adjusting duration estimates, they restrict their analyses to the CPS “incoming rotation groups”, whose reported durations are unaffected by the introduction of dependent interviewing in the redesigned survey. Because I adjust the pre-1994 duration estimates upward based on Polivka and Miller’s (1998) results, my estimates of the upward trend in duration are conservative relative to Abraham and Shimer.

[2] This is a “synthetic cohort” approach in that with a rotating monthly sample such as the CPS, the estimate of unemployment continuation probabilities is formed by comparing different groups over time, rather than by following the same individuals through time.

dependence, and they also will vary from month to month as economic conditions change.

My application of the synthetic cohort approach to obtain nonparametric estimates of expected completed duration from grouped duration data follows M. Baker (1992a); see G. Baker and Trivedi (1985) for a more general overview. We begin with continuation probabilities, defined as the conditional probability that individuals whose unemployment spell has lasted *(j-1)* months at time *(t-1)* will remain unemployed into the next period:

$$f_j(t) = \frac{n(j,t)}{n(j-1,t-1)} \tag{1}$$

where *n(.)* represents the sampled number of individuals unemployed for a given number of months at the time of a particular monthly survey. In a rotating sample survey such as the CPS, the sample used to calculate the numerator and denominator differs, but under the assumption that each monthly sample represents the target U.S. population (as the CPS is constructed), this expression provides an estimate of the continuation probability for a fixed representative cohort.

The product of the continuation probabilities represent the empirical survivor function, or the proportion of individuals entering unemployment at time *(t-j)* who remain unemployed at time *t*:

$$G_j(t) = f_0(t)\, f_1(t)\, f_2(t)\, f_3(t) ... f_j(t) \tag{2}$$

In this expression, $f_0(t)$ is the continuation probability for the entering cohort, which is defined identically as one. Assuming that the duration intervals are not all identical (e.g., not all one month), the expected completed duration in a particular month *t*, *D(t)*, is estimated as:

$$D(t) = 1 + \sum_{j=1}^{m} G_j(T_j) * (T_j - T_{j-1}) \tag{3}$$

where the *T's* represent duration intervals (measured in units of the monthly sampling window) and T_m is the maximum duration measured or used.

Empirical implementation requires setting the width and number of duration intervals used for estimation. I follow Baker (1992a) in using 6 unequally spaced duration intervals and corresponding continuation probabilities:

$f_1(t)$: 5-8 weeks in month t to <5 weeks in (t-1)
$f_2(t)$: 9-12 weeks in month t to 5-8 weeks in (t-1)
$f_3(t)$: 13-16 weeks in month t to 9-12 weeks in (t-1)
$f_4(t)$: 27-39 weeks in month t to 13-26 weeks in (t-3)
$f_5(t)$: 53-78 weeks in month t to 27-52 weeks in (t-6)
$f_6(t)$: 100+ weeks in month t to 53-99 weeks in (t-12)

Note the variation in duration intervals for $f_4(t)$-$f_6(t)$, which must be incorporated into the duration estimate based on equation (3). Then the expected completed duration is formed as:

$$D(t) = 1 + f_1 + f_2 f_1 + f_3 f_2 f_1 + 3 f_4 f_3 f_2 f_1 + 6 f_5 f_4 f_3 f_2 f_1 + 12 f_6 f_5 f_4 f_3 f_2 f_1 \tag{4}$$

where the time identifier (t) has been suppressed on the right-hand side of (4) for simplicity. D(t) is defined as the expected duration of unemployment (in months) for a cohort that enters unemployment at *t* and faces current economic conditions throughout the unemployment spells of cohort members.[3] In the empirical work, I estimate expected completed duration for the full sample and for various groups (demographic groups and by reason for unemployment); estimation by group proceeds by first restricting the unemployment sample to the specified group, than estimating expected completed duration as described above.[4]

One additional estimation issue is "digit preference"—the tendency for respondents to report durations as integer multiples of one month or half-years (i.e., multiples of 4 or 26). Following previous analysts, I adjust for digit preference by allocating a fixed share (50 percent) of bunched observations to the next monthly interval (see the Appendix for additional details, along with a discussion of changing top-coding for the CPS duration variable). Baker (1992b) reports that although estimates of expected completed duration are sensitive to the allocation rule, cyclical elasticity regression results are not.

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[3] To relax the assumption that current economic conditions continue throughout cohort members' spells, Corak and Heisz (1996) propose and estimate a forward-looking nonsteady-state estimator, which reflects the evolution of continuation probabilities into the future for individuals entering unemployment in the current month. They find that their estimator has desirable properties relative to the standard backward-looking nonsteady-state estimator.

[4] The large sample size afforded by use of the complete set of monthly CPS records offers an advantage relative to the outgoing rotation group sample used by Baker (1992a), by enabling more reliable estimation of group-specific unemployment durations.

4. Results

Tabulations of Unemployment by Reason and Expected Duration

This sub-section displays and discusses the basic patterns in unemployment by reason and estimated expected duration (yearly averages of monthly values for each series).

Figure 3 shows the unemployment rate and the shares of layoffs and permanent job loss in total unemployment incidence, each expressed in percentage points, for the period 1977-2004.[5] Job losers on average account for about 42 percent of the newly unemployed, with permanent job losses outnumbering layoffs. The two series exhibit only limited cyclicality. Layoff incidence increased in 1994 and 1995 and has remained high; this is somewhat surprising, given declining shares of manufacturing employment and that sector's relative emphasis on temporary layoffs. Moreover, an upward trend is evident for permanent job loss.

Figure 4 displays unemployment incidence shares for voluntary job leavers (quits) and labor force entrants. Job leaving constitutes a relatively small share of unemployment incidence—12.5 percent on average—but the series exhibits a pronounced procyclical pattern. Re-entrant unemployment is frequent, although it exhibits a pronounced downward trend, while new entrants have constituted a small and relatively consistent share of unemployment incidence.

Figure 5 displays estimates of expected completed duration of unemployment, for all unemployed and by job loss category.[6] Each of these expected duration series exhibits

[5] Incidence shares are measured as the share of a group in the pool of individuals identified as unemployed for less than five weeks.
[6] Following past practice (e.g., Sider 1985), I multiplied estimates of expected duration in months by 4.3 to obtain expected duration in weeks for the charts.

noticeable counter-cyclicality. Permanent job losers on average endure long spells of unemployment; during the period 1976-2004, the expected duration of unemployment was 17 weeks for permanent job losers and 12 weeks for all unemployed individuals.

Figure 6 shows expected completed duration for all unemployed, quits, and labor force entrants.[7] Individuals unemployed for each of these reasons on average experience durations around 10 weeks on average, slightly below the overall average, and all appear to exhibit counter-cyclicality.

The patterns in Figures 3-6 can be summarized as follows. An upward trend in permanent job loss is evident, with these individuals experiencing relatively long unemployment spells on average. In addition, unemployment durations by reason each exhibit substantial counter-cyclicality and have remained high since the 2001 recession (with the exception of unemployment durations for individuals on temporary layoff). These results suggest that the trend toward rising unemployment durations has been uniform by reason for unemployment, although it may also relate to the rising share and high durations for permanent job losers.

Regression Estimates of Cyclical and Trend Patterns

I use the regression framework of Sider (1985) and Baker (1992a) to estimate the degree of rising unemployment duration and assess its sources. In this framework, the monthly estimates of expected completed duration described above (corrected for the 1994 survey change) are regressed on the contemporaneous unemployment rate (not

[7] I merged the re-entrant and new entrant duration series together (weighted by relative incidence), because they are nearly identical over my sample frame (including an upward jump between 1993 and 1994, which does not appear due to survey redesign).

seasonally adjusted), three season dummies, and a linear time trend; unemployment duration and the unemployment rate are measured in natural logs.

The results for the full sample (all unemployed individuals age 16-64) are listed in Table 1. The first row shows that expected duration exhibits substantial cyclicality; the estimated elasticity with respect to the unemployment rate is about 0.7. Moreover, a strong upward trend is evident in the expected duration conditional on the unemployment rate (as suggested by Figure 1 earlier). This trend estimate implies that unemployment duration rose by about 35 percent between 1976 and 2004.[8]

The second through fourth rows of Table 1 list results for alternative time periods. The results in row two, which correspond to Baker's (1992a) sample period, largely replicate those earlier results, and they indicate that a small negative and statistically insignificant time trend. By contrast, the subsequent two rows for the longer sub-periods that eliminate any influence of the 1994 CPS survey redesign again reveal significant upward time trends in expected duration. This slope of this upward trend approximately doubled after 1993, consistent with the persistently high unemployment durations since 2001 that were depicted in Figure 5 and 6. The steeper time trend after 1994 constitutes evidence against an explanation for rising duration based on women's labor force attachment, because the convergence of women's and men's labor market attachment and experiences over the past few decades should largely eliminate the role of rising female attachment in the recent time period (Abraham and Shimer 2002).

More careful assessment of the sources of rising unemployment duration requires disaggregating the duration estimates and regressions by demographic group and reason for unemployment. Table 2 lists regression results for these groups, which demonstrate striking uniformity in the cyclical elasticity estimates and upward trend by sex, age, and education group. The upward trend is somewhat greater for women than for men, but the trend for men is large and statistically significant, which argues against an explanation for rising duration based on changing labor force attachment of women. On the other hand, the results in the bottom five rows of Table 2 show that the trend increase in expected duration has been most pronounced for labor market re-entrants and new entrants. The trend increase in expected duration is smaller but still pronounced for permanent job losers and voluntary job leavers (quits), but no upward trend in duration is evident for individuals on temporary layoff.

Understanding the contribution of duration by reason to the increase in overall duration requires an analysis of changing shares of unemployment incidence by reason. Table 3 lists regression results for linear probability models of unemployment incidence shares by reason, otherwise using the same specification as the duration regressions.[9] Permanent job losses and layoffs are strongly counter-cyclical; quits to unemployment and labor market re-entrance are pro-cyclical; and new labor market entrants exhibit a small and statistically insignificant counter-cyclical response to business cycle conditions. The results also reveal upward time trends in the incidence shares of permanent job loss (consistent with results using individual panel data in Valletta 1999) and temporary layoffs, which are largely offset by a significant downward trend in the

[8] The Durbin-Watson (DW) statistics indicate the presence of negative autocorrelation for most of the regression models, implying that the standard error estimates are conservative in most cases (i.e., the associated t tests mostly understate significance levels).

[9] Baker (1992a) reports no difference between results based on linear probability and logistic models in this setting, probably because the incidence variables all are sufficiently well-bounded away from zero and one.

incidence of re-entrant unemployment. The trend coefficients imply about a 54 percent increase in the incidence share of permanent job loss and a 50 percent decrease in the incidence share of re-entrants between 1977 and 2004 (relative to the sample mean incidence of these two unemployment types).

Decomposition Results

The regression results from the previous section suggest potentially important links between rising overall unemployment duration and changing incidence and duration by reason. I perform several decomposition analyses to investigate this link. I first apply a simplified variant of the decomposition that Baker (1992a) used to test the heterogeneity hypothesis of cyclical variability in unemployment durations, using unemployment by reason as my measure of heterogeneity. Expected unemployment duration across all groups can be expressed as a weighted average of expected duration by reason, with the weights equal to the shares of unemployment incidence by reason. This property enables decomposition of unemployment duration across all reasons for unemployment into two components:

D_{pc} ("probability constant") — expected completed duration holding expected duration for each reason at its sample average, but allowing the shares of unemployment incidence by reason to change

D_{sc} ("share constant") — expected total duration holding the shares of unemployment incidence by reason equal to their sample averages, but allowing expected duration by reason to change.

Comparison of regressions using the constructed variables D_{pc} and D_{sc} with regressions using the unadjusted duration measure indicates the relative roles of changing duration by reason and changing shares by reason in the determination of the time-series properties of total unemployment duration. These results are listed in Panel A of Table 4. The first row repeats the results for the unadjusted duration measure (row 1 from Tables 1 and 2). The second row lists the results for the probability constant measure D_{pc} and the third row lists the results for the share constant measure D_{sc}. A comparison of the results in the final two rows indicates that virtually all of the cyclical variability in total unemployment duration is due to cyclical variability in expected duration by reason rather than variability in incidence by reason: the coefficient on the unemployment rate is very small in the D_{pc} equation, which holds expected duration by reason constant, and large in the D_{sc} equation, which holds incidence by reason constant. Essentially all of the upward time trend also is attributable to rising duration by reason (the time trend coefficient is very small in the D_{pc} equation).

The decomposition listed in Panel A of Table 4 groups all reasons for unemployment together. Recall, however, that the key changes over time have been in the incidence and duration of unemployment associated with permanent job loss and labor force entrance. The first row of Panel B lists results from an alternative decomposition that focuses on permanent job loss. The dependent variable used in the first row of Panel B is formed by holding the incidence and duration of permanent job loss constant at their respective sample averages. Comparison of these results with the results in the first row of Panel A reveals the effect on overall expected duration of rising incidence and duration of unemployment due to permanent job loss. The substantially smaller coefficients in row 1 of Panel B than in row 1 of Panel A indicates that rising duration and incidence of unemployment due to permanent job loss accounts for most of the cyclical and time-trend effects on total duration. In conjunction with the decomposition results from Panel A, which showed that changing incidence explains only a small portion of the time trend, the Panel B results indicate that rising duration associated with permanent job loss has played the dominant role in the trend toward rising duration of total unemployment.

Recall that Tables 2 and 3 also revealed significant changes in unemployment duration and incidence for re-entrants and new entrants to the labor force. The dependent variable in the final row of Panel B in Table 4 holds the incidence and duration of entrant unemployment constant. Comparison of these results with the total duration results from row 1 of Panel A reveals that labor force entrants account for a only a small portion of the cyclical variability in total expected duration. On the other hand, trends in the incidence and duration of labor force entrant unemployment account for a substantial share of the upward trend in overall unemployment duration, almost as much as the share accounted for by changes in permanent job loss.[10] This is consistent with the findings of Abraham and Shimer (2002), based on their analysis of labor force transition data. However, attributing the impact of changes in the incidence and duration for labor force entrants solely to women's labor force attachment seems unwarranted, given the findings presented above regarding rising duration for men and the role of permanent job loss.

[10] For the period 1976-1998, Valletta (1998) found no role for entrant unemployment in explaining rising overall duration.

5. Summary and Discussion

Using a synthetic cohort approach to estimate expected completed unemployment duration, this paper documents an upward trend in unemployment duration (conditional on the unemployment rate) over the period 1976-2004. Changes in CPS survey methodology made little difference for this trend, as the analyses adjust for the survey redesign. Indeed, it is striking to note that results from sub-period analyses—which eliminate any redesign effects that can be captured by a single level adjustment at the time of the redesign—show that the upward trend in relative duration accelerated after 1994. Continuation and acceleration of the upward duration trend in recent years constitutes informal evidence against an explanation for rising duration based on chaning labor force attachment for women, because the convergence of women's and men's labor market attachment and experiences over the past few decades should largely eliminate the role of rising female attachment in the recent time period.

Additional analyses tested explanations for rising duration based on changing labor force attachment by women (Abraham and Shimer 2002) and changes in the structure of unemployment that are consistent with declining job security (Valletta 1998, 1999). The results suggest an important role of rising duration for permanent job losers, consistent with the link to declining job security. On the other hand, the regression and decomposition evidence presented suggests that rising unemployment duration for labor force entrants also explains a substantial portion of the increase in overall duration. Thus, while it may be premature to attribute rising unemployment duration to rising labor force attachment by women, changes in entrant behavior for both sexes appear to be important. Future versions of this paper will attempt to pin down the roles of job loss and entrant unemployment more precisely, using more comprehensive decompositions of changing unemployment incidence and duration. Other planned methodological expansions include more careful treatment of the sampling properties of estimated completed duration and regression models of its determinants.

Pinning down the sources of rising unemployment duration may have important implications for aggregate dynamics and monetary policy. Rising duration can raise or lower the wage pressures associated with a given unemployment rate, thereby altering the natural or "non-accelerating inflation rate" of unemployment (NAIRU). Abraham and Shimer (2002) argued that the long-term unemployed may put less downward pressure on wages than do the short-term unemployed, through hysteresis effects on the employment prospects of the long-term unemployed. By contrast, Robert Solow (1970) argued that "People who have been unemployed a long time put more downward pressure on wages because they are more willing to undercut going wage rates in order to get a job." The theory and empirical findings in Campbell and Duca (2004) are consistent with Solow's view: they find that rising unemployment duration (and associated job insecurity) can account for a significant estimated reduction in the NAIRU during the 1990s. My findings regarding the important role of permanent job loss in the determination of rising duration reinforces Campbell and Duca's findings regarding the decline in the NAIRU.

Appendix: Data Adjustments

The text referred to several changes in CPS survey design and various data handling issues that affect measurement of unemployment by reason and duration over time.

Adjustments for the 1994 CPS Survey Redesign

The adjustment factors from Polivka and Miller (1998) were used to mitigate the influence of the 1994 CPS redesign. In particular, I applied the multiplicative adjustment factors from their Tables 2 and 3 to adjust the pre-1994 monthly estimates of unemployment shares by reason and by duration category.

Digit preference and top-coding

To account for "digit preference"—the tendency for respondents to report durations as multiples of one month or half-years (i.e., multiples of 4 or 26)—I follow previous analysts by allocating a fixed share of bunched observations to the next monthly interval. In particular, I allocated 50 percent of respondents reporting the following durations of unemployment to the next weekly value: 4, 8, 12, 16, 26, 39, and 52 weeks. I also reset 50 percent of the responses of 99 weeks to 100 weeks (after imposition of the top code adjustment described in the next paragraph).

The CPS duration variable was top-coded at 99 weeks through 1993. For time-series consistency, I imposed this top-code on the post-1993 data as well. This constraint makes little difference for estimates of expected completed duration because: (i) the continuation probabilities are estimated by grouping the data for individuals with durations longer than 99 weeks (see Section 3 in the text); (ii) only a small number of observations (2-4%) are recorded as unemployed longer than 99 weeks after 1993, with durations recorded up to 118 weeks.

References

Abraham, Katharine, and Robert Shimer. 2002. "Changes in Unemployment Duration and Labor Force Attachment." In *The Roaring Nineties*, edited by Alan Krueger and Robert Solow. New York: Russell Sage Foundation.

Baker, Michael. 1992a. "Unemployment Duration: Compositional Effects and Cyclical Variability." *American Economic Review* 82(1): 313-321.

Baker, Michael. 1992b. "Digit Preference in CPS Unemployment Data." *Economics Letters* 39 (1, May): 117-121.

Baker, G.M., and P.K. Trivedi. 1985. "Estimation of Unemployment Duration from Grouped Data: A Comparative Study." *Journal of Labor Economics* 3(2): 153-174.

Campbell, Carl M., and John V. Duca. 2004. "The Impact of Evolving Labor Practices and Demographics on Inflation and Unemployment." Manuscript, Federal Reserve Bank of Dalla, June.

Corak, Miles, and Andrew Heisz. 1996. "Alternative Measures of the Average Duration of Unemployment." *The Review of Income and Wealth* 42(1, March), pp. 63-74.

Kaitz, Hyman B. 1970. "Analyzing the length of spells of unemployment." *Monthly Labor Review* 93 (Nov.): 11-20.

Perry, George L. 1972. "Unemployment Flows in the U.S. Labor Market." *Brookings Papers on Economic Activity* 3: 245-278.

Polivka, Anne, and Stephen Miller. 1998. "The CPS After the Redesign: Refocusing the Economic Lens." In *Labor Statistics Measurement Issues*, edited by John Haltiwanger, Marilyn Manser, and Robert Topel, pp. 249-289. Chicago: University of Chicago Press.

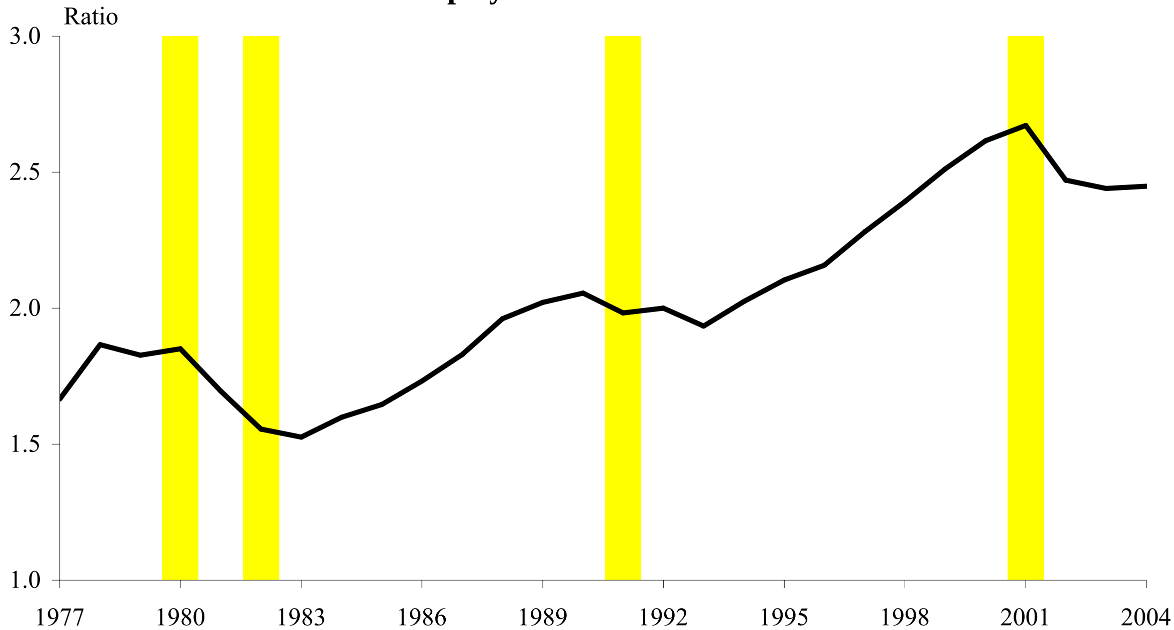
Sider, Hal. 1985. "Unemployment Duration and Incidence: 1968-82." *American Economic Review* 75(3, June): 461-472.

Solow, Robert. 1970. Comment on George L. Perry, "Changing Labor Markets and Inflation." *Brookings Papers on Economic Activity* 0(3): 1970, pp. 411-448.

Valletta, Robert G. 1998. "Changes in the Structure and Duration of U.S. Unemployment, 1967-1998." Federal Reserve Bank of San Francisco *Economic Review* 3: 29-40.

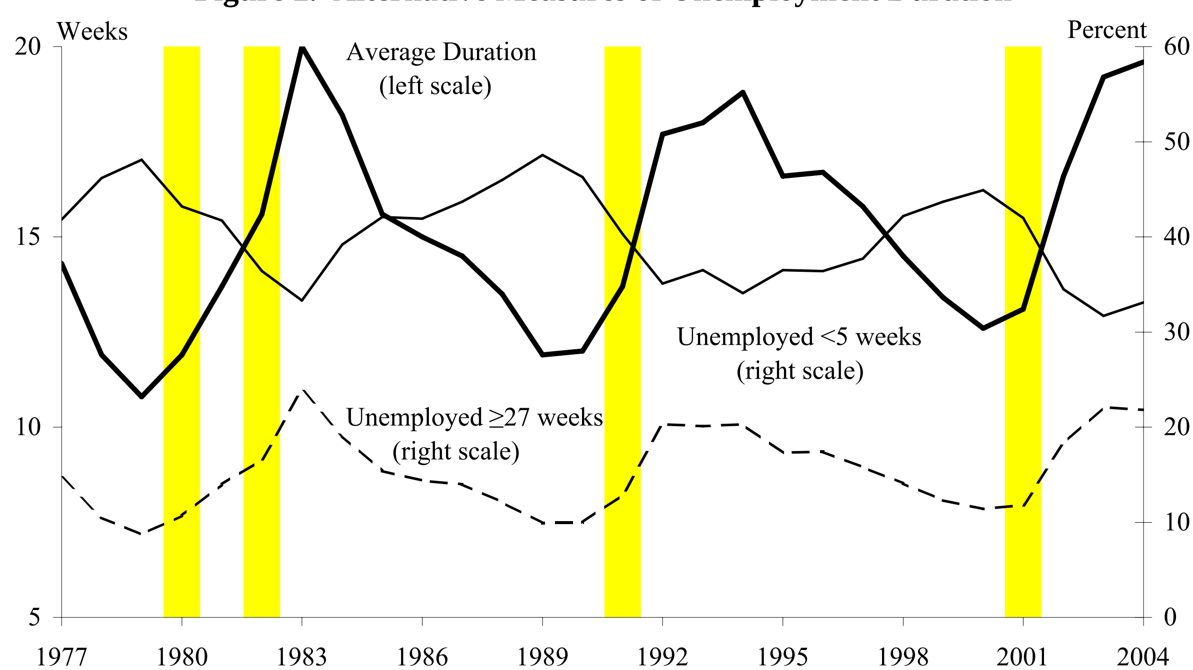
Valletta, Robert G. 1999. "Declining Job Security." *Journal of Labor Economics* 17(4, part 2): S170-S197.

Figure 1. Ratio of Expected Completed Unemployment Duration to the Unemployment Rate



Note: Author's calculations from U.S. BLS data (see text section 3 for duration methodology). Vertical bars denote recessions.

Figure 2. Alternative Measures of Unemployment Duration

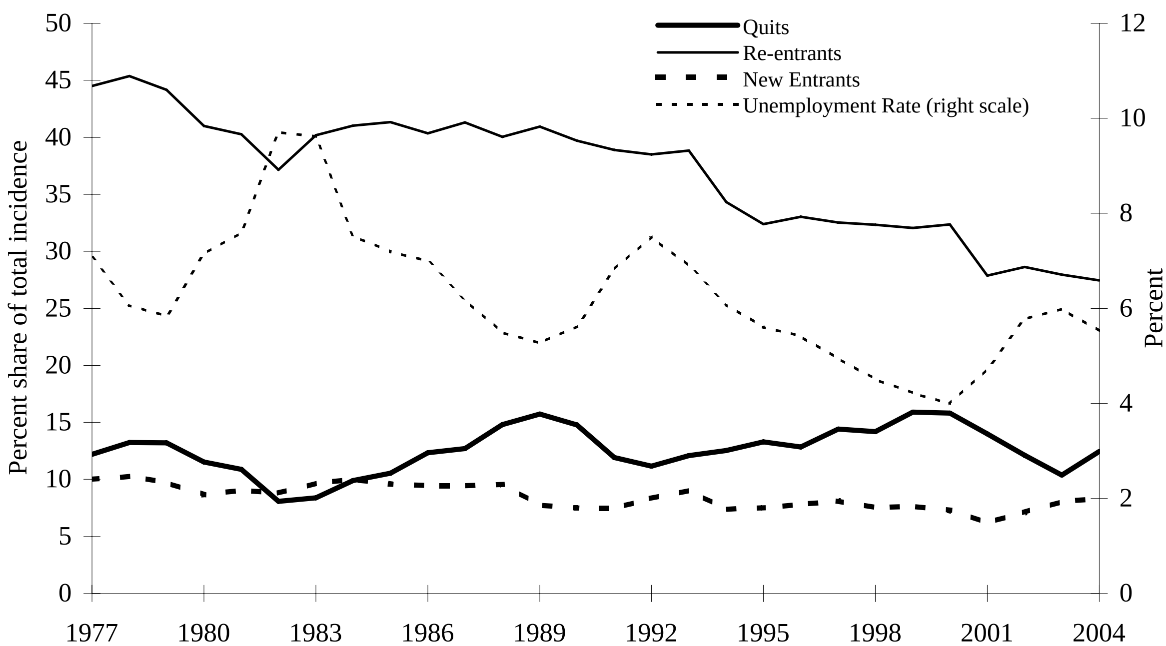


Note: Published data from the U.S. BLS, based on duration of in-progress spells. Series measured in percent are calculated as the share of all unemployed. Vertical bars denote recessions.

Figure 3. Incidence of Involuntary Unemployment

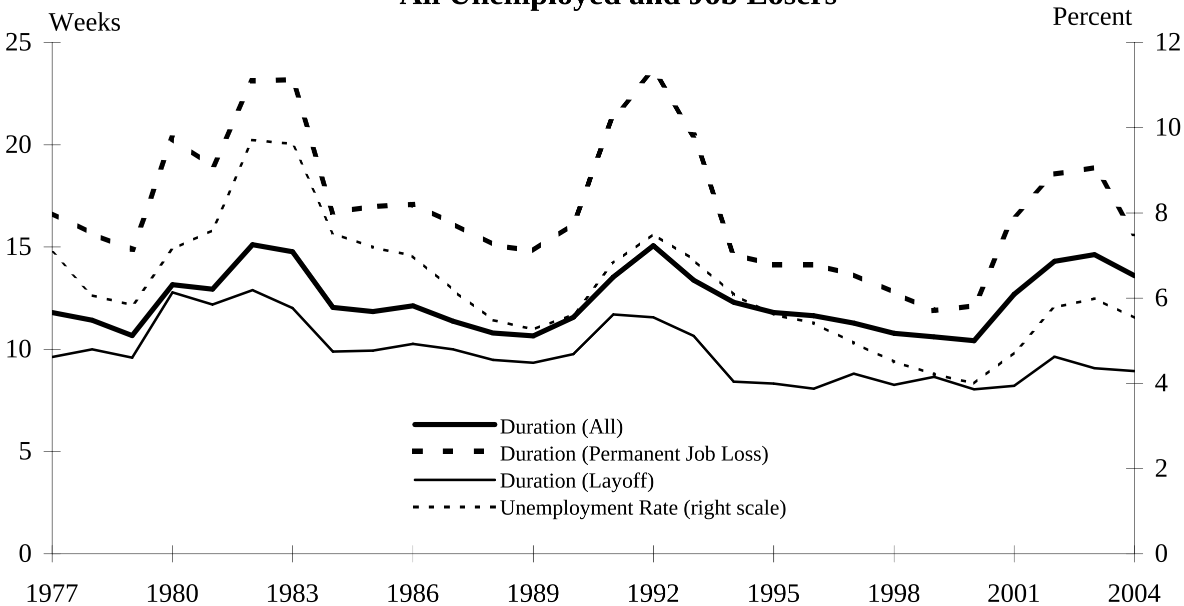
Note: Annual averages of monthly data (author's tabulations of CPS microdata). Pre-1994 figures are adjusted for the 1994 change in survey methodology.

Figure 4. Incidence of Voluntary Unemployment and Labor Force Entry



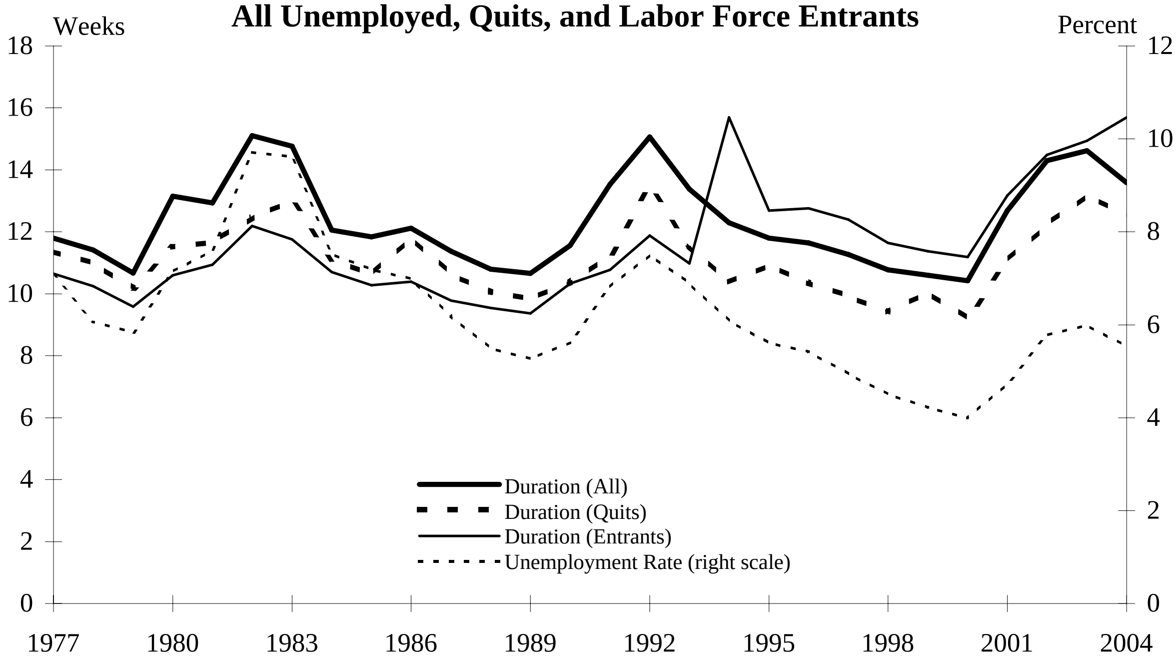
Note: Annual averages of monthly data (author's tabulations of CPS microdata). Pre-1994 figures are adjusted for the 1994 change in survey methodology.

Figure 5. Expected Duration of Unemployment: All Unemployed and Job Losers



Note: Annual averages of monthly data (author's tabulations of CPS microdata). Pre-1994 figures are adjusted for the 1994 change in survey methodology.

Figure 6. Expected Duration of Unemployment: All Unemployed, Quits, and Labor Force Entrants



Note: Annual averages of monthly data (author's tabulations of CPS microdata). Pre-1994 figures are adjusted for the 1994 change in survey methodology.

Table 1. Regression Estimates of Cyclical and Trend Effects on Ln(Expected Duration)
(Full sample, age 16 - 64)

	Coefficient estimates		DW	N
Years	ln(unemp. rate)	Time Trend (*100)		
1977 - 2004	0.694	0.104	1.774	336
	(0.028)	(0.006)		
1980 - 1988	0.598	-0.043	1.801	108
	(0.051)	(0.030)		
1977 - 1993	0.690	0.075	1.812	204
	(0.034)	(0.010)		
1994 - 2004	0.705	0.169	2.045	132
	(0.050)	(0.020)		

Note: Samples use monthly data for the indicated years, with expected duration estimated from microdata (see text). Pre-1994 data are adjusted for the 1994 change in survey methodology. Regressions also include 3 season dummies and a constant. Standard errors in parentheses. DW is the value of the Durbin-Watson statistic.

Table 2. Regression Estimates of Cyclical and Trend Effects on Ln(Expected Duration) by Group, 1977 - 2004

Group	Coefficient estimates		DW
	ln(unemp. rate)	Time Trend (*100)	
Aggregate	0.694	0.104	1.774
	(0.028)	(0.006)	
Men (all)	0.779	0.083	1.979
	(0.036)	(0.008)	
Men (age 25 - 54)	0.781	0.083	1.967
	(0.036)	(0.008)	
Men (HS Degree or less)	0.694	0.087	1.905
	(0.032)	(0.007)	
Men (> HS Degree)	0.824	0.116	1.881
	(0.045)	(0.010)	
Women (all)	0.559	0.120	1.611
	(0.028)	(0.006)	
Women (age 25 - 54)	0.554	0.119	1.606
	(0.028)	(0.006)	
Women (HS Degree or less)	0.651	0.110	1.850
	(0.029)	(0.007)	
Women (> HS Degree)	0.694	0.123	1.583
	(0.041)	(0.009)	
By reason for unemployment:			
Permanent Job Loss	0.934	0.071	1.405
	(0.044)	(0.010)	
Temporary Layoff	0.579	-0.013	2.102
	(0.057)	(0.013)	
Quit	0.559	0.072	2.146
	(0.046)	(0.011)	
Re-entrant	0.453	0.160	1.417
	(0.042)	(0.010)	
New Entrant	0.414	0.163	1.602
	(0.066)	(0.015)	

Note: Samples use monthly data for the indicated years, with expected duration estimated from microdata (see text). Pre-1994 data are adjusted for the 1994 change in survey methodology. Regressions also include 3 season dummies and a constant. Sample size is 336. Standard errors in parentheses. DW is the value of the Durbin-Watson statistic.

Table 3. Regression Estimates of Cyclical and Trend Effects on Unemployment Incidence (entrance shares) by Reason, 1977 - 2004 (Full sample, age 16 - 64)

Reason for Unemployment	Coefficient estimates: ln(unemp. rate)	Coefficient estimates: Time Trend (*100)	DW
Permanent Job Loss	0.071	0.040	1.655
	(0.007)	(0.002)	
Temporary Layoff	0.045	0.025	1.199
	(0.011)	(0.002)	
Quit	-0.096	-0.006	1.215
	(0.006)	(0.001)	
Re-entrant	-0.031	-0.055	1.482
	(0.011)	(0.003)	
New Entrant	0.010	-0.007	1.666
	(0.006)	(0.001)	

Note: Samples use monthly data for the indicated years, with unemployment entrance shares estimated from microdata (see text). Pre-1994 data are adjusted for the 1994 change in survey methodology. Regressions also include 3 season dummies and a constant. Sample size is 336. Standard errors in parentheses. DW is the value of the Durbin-Watson statistic.

Table 4. Unemployment Duration Regressions, 1977 - 2004, Adjusted by Reason for Unemployment

Panel A: Decomposition (Probability and Share Constant), All Reasons

	Coefficient estimates		DW
Dependent Variable	ln(unemp. rate)	Time Trend (*100)	
ln(D) (unadjusted)	0.694	0.104	1.774
	(0.028)	(0.006)	
$\ln(D_{pc})$ (probability constant)	0.026	0.010	1.646
	(0.004)	(0.001)	
$\ln(D_{sc})$ (share constant)	0.658	0.104	1.693
	(0.029)	(0.007)	

Panel B: Incidence and Duration by Reason Held Constant

	Coefficient estimates		DW
Dependent Variable	ln(unemp. rate)	Time Trend (*100)	
ln(D), perm. job loss constant	0.283	0.025	2.220
	(0.023)	(0.005)	
ln(D), entrant constant	0.474	0.035	1.311
	(0.028)	(0.007)	

Note: Samples use monthly data for the indicated years, with expected duration estimated from microdata (see text). Pre-1994 data are adjusted for the 1994 change in survey methodology. Regressions also include 3 season dummies and a constant. Sample size is 336. Standard errors in parentheses. DW is the value of the Durbin-Watson statistic.