Changes in the Structure and Duration of U.S. Unemployment, 1967–1998

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The unemployment rate is determined by the incidence and duration of unemployment spells. In this paper, I examine the time-series properties of unemployment incidence by reason and the duration of a typical unemployment spell. In line with earlier research, I find strong countercyclicality in unemployment durations, which is relatively uniform across the different reasons for entry into unemployment. However, I also uncover an upward trend in duration that is entirely attributable to rising incidence and duration of permanent job loss. These changes in the structure and duration of unemployment have various policy implications.

In 1998, the U.S. economy produced its lowest unemployment rates since the late 1960s. Despite this, the duration of unemployment spells has remained long compared to typical durations during previous expansions. This is consistent with the long-run secular trend toward rising unemployment durations noted by a number of analysts (e.g., Murphy and Topel 1987, Juhn, Murphy, and Topel 1991, and Sider 1985). Moreover, the structure of unemployment by reason during the 1990s expansion has remained heavily weighted towards permanent job loss rather than voluntary job search and labor force entry decisions. In this paper, I examine the time-series properties of unemployment duration and unemployment incidence by reason, and I investigate the links between them.

Although the unemployment rate is our primary labor market indicator, the underlying distribution of unemployment durations has played a prominent role in the macroeconomics and labor economics literatures since the 1970s. Whether unemployment spells are best characterized as long or short provides information about whether unemployment primarily is voluntary or instead reflects persistent insufficiency of aggregate demand. This information in turn provides insights into what, if any, macroeconomic policies might be appropriate to combat unemployment. The time-series properties of unemployment durations also are relevant to macroeconomic and labor market policies. For example, countercyclical variation in unemployment durations has implications for the behavioral effects and financing of unemployment insurance payments. Moreover, underlying secular trends in unemployment duration may imply long-run changes in the degree of labor market slack associated with a given unemployment rate.

A number of authors (notably Perry 1972, Sider 1985, and Baker 1992a) have investigated the cyclical component of time-series variation in unemployment duration. In this paper, I update earlier results regarding the cyclical sensitivity of unemployment durations, and I extend the analysis by examining secular trends in expected duration and incidence in more detail than did past work. A key component of this formulation involves linking reasons for the incidence of unemployment with changes in duration. In particular, the rising incidence of permanent job loss may be linked to a secular trend toward rising durations, because permanent job losers on average endure substantially longer spells of unemployment than do individuals unemployed for other reasons.

In Section I, I discuss previous work on the time-series properties of unemployment duration and incidence. I describe the data that I use in Section II; these data come from the U.S. Bureau of Labor Statistics' (BLS) monthly household survey. My sample period begins in 1967 and extends to May 1998. Because these data contain information on spells in progress, their use requires estimation of expected completed duration, as described in Section II. Section III presents estimation results; these include basic tabulations for the measures of unemployment incidence and expected duration and regression models that estimate their cyclical and secular properties. I summarize the results in the concluding section, where I also discuss some implications for macroeconomic performance and policy.

I. Unemployment Duration and Incidence

The unemployment rate reflects both unemployment incidence and duration. Although the unemployment rate by itself is our key indicator of labor market conditions, the underlying distribution of unemployment spell durations provides important additional information. In the 1970s and 1980s, the characterization of unemployment durations as "long" or "short" was the subject of substantial debate. The short view, as described most comprehensively by Feldstein (1973), emphasized the dynamic nature of unemployment. Proponents of the short view focused on job turnover and unemployment flows and argued that the pool of unemployed on average is characterized by a large number of individuals who experience relatively short spells of unemployment (i.e., a month or two). This view generally is consistent with voluntary search activity by unemployed individuals or implicit agreements between workers and firms regarding the use of temporary layoffs. In contrast, advocates of the long view, such as Clark and Summers (1979) and Akerlof and Main (1980, 1981), argued that the pool of unemployed typically is dominated by a small number of individuals who experience relatively long spells of unemployment, and who are best described as "involuntarily" unemployed.

These alternative views played an important role in the debate over appropriate macroeconomic and manpower policies aimed at combatting unemployment. In general, the short view is consistent with a less activist policy, given its implications concerning the voluntary nature of unemployment, widespread sharing of the burden of unemployment, and the implied efficiency of the associated employment and unemployment flows. In contrast, the long view of unemployment argues for more activist economic policy, under the assumption that persistent lengthy unemployment spells reflect a shortage of available jobs rather than an equilibrium matching process with frictions.[1]

A related issue is variation in unemployment durations over the business cycle. Changes in the duration distribution of unemployment spells over the business cycle provide information about the degree of demand insufficiency during cyclical downturns and its implications for the matching process between workers and firms. Moreover, changes in the incidence and duration of unemployment over the business cycle may have implications for optimal unemployment insurance policies. For example, Katz and Meyer (1990) documented the importance of temporary layoffs during the early 1980s, and they discuss the distortionary effects of unemployment insurance (UI) benefits on firms' and workers' behavior regarding temporary layoffs.

Kaitz (1970) and Perry (1972) made key early contributions to the analysis of the cyclical properties of unemployment duration. Kaitz used data for the years 1948–1969 and found greater cyclical variability in duration than did Perry for the years 1954–1971. These differences are explained by various differences in approach. Perry also found an upward time trend in duration during the period covered. Sider (1985) updated Kaitz's and Perry's work, emphasizing the importance of using nonsteady-state measures of unemployment duration when cyclical variability is of primary interest. Sider found evidence of substantial cyclical variability, along with an upward time trend for the period 1968–1982.

Michael Baker (1992a) updated these estimates to the late 1980s. He made a key contribution by noting that in contrast to analyses based on aggregate data, which typically uncovered countercyclicality in unemployment durations, some analyses that relied on individual data uncovered procyclicality. One explanation for this discrepancy is the influence of heterogeneity in unemployment incidence: if individuals or groups with long expected duration are more likely to enter unemployment during a downturn, aggregate data will display countercyclicality in unemployment durations even if expected unemployment duration for individuals displays no cyclical properties. Baker applied a decomposition technique that enabled a direct test of the heterogeneity hypothesis. The results did not support the heterogeneity interpretation. In particular, Baker found that pronounced countercyclical variation in duration during the 1980s was attributable to relatively uniform countercycli-

1. Of course, if the long-term unemployed simply lack the appropriate skills to acquire available jobs, aggregate demand management policies are likely to be ineffective.

cality across groups, rather than increasing incidence during downturns for groups with lengthy expected durations.

Although these papers also investigated the cyclical properties of unemployment incidence, none focused directly on the links between secular trends in duration and incidence. I focus on such links, particularly the link between secular increases in unemployment durations and the incidence of permanent job loss. Given that permanent job losers suffer longer spells of unemployment than do individuals unemployed for other reasons, it is likely that an increasing incidence of permanent job loss has contributed to the trend toward rising durations. As discussed in the conclusion, such a secular trend in aggregate labor market outcomes has potentially important implications for macroeconomic policy.

II. ESTIMATING UNEMPLOYMENT INCIDENCE AND EXPECTED COMPLETED DURATION

CPS Unemployment Data

I use monthly data on unemployment levels and rates, which are published by the BLS; the data are available beginning in 1948 and extending to the most recent survey month (May 1998 at the time this paper was written). The data reflect population-weighted counts from the Current Population Survey (CPS), the monthly household survey upon which official labor force statistics are based. The underlying sample is the civilian population aged 16 and over. The published BLS data include information on the total number unemployed, the number unemployed by reason, and the duration of spells in progress. I restrict my analysis to the period beginning in 1967, due primarily to relative consistency in the CPS survey since then. Some analyses are restricted to begin in 1976, because data on unemployment by reason only became available beginning in that year.

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., they expect 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. For total unemployment and unemployment by reason, the BLS data provide information on the monthly inflow into unemployment. As described in the next subsection, this is the key information used to form a steady-state estimate of expected completed duration.[2]

Use of the BLS unemployment data in a time-series framework requires that an adjustment be applied to the data beginning in January 1994, due to a significant redesign of the survey that became effective at that time; the labor force questions otherwise had been largely unchanged since 1967. As described in Cohany, Polivka, and Rothgeb (1994), results from a parallel survey administered in 1993 indicate that the new survey instrument produces lengthier unemployment durations and changes in unemployment shares by reason. The largest changes in shares are a substantial increase in re-entrants and corresponding decline in new entrants, due to removal of the requirement that to be classified as re-entrants respondents must have worked previously for at least two weeks in a full-time job.

I used three techniques to ensure that my results are not affected by the 1994 changes to the CPS survey. First, in each regression that uses both pre- and post-1993 data, I included a post-1993 dummy variable, so that the results are conditional on the intercept shift associated with the survey redesign; Perry (1972) used a similar approach to account for the 1967 survey redesign. Second, I imposed a direct adjustment to the post-1993 data, as implied by a comparison of the 1993 actual and parallel surveys. The adjustment is based on the percentage change in the total unemployment count and unemployment counts by reason. I also ran all regressions with the post-1993 period excluded. The results for these regressions were all very similar, which indicates that my conclusions are not affected by the 1994 CPS survey redesign.

Estimates of Expected Completed Duration

The CPS data described in the previous subsection provide information on the average length of existing unemployment spells up to the date of the survey. This "average interrupted duration" measure will not in general correspond to the expected duration of a completed spell for a new entrant to unemployment, particularly under changing labor

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2. Other studies of unemployment trends have used retrospective data from the March CPS Annual Demographic Supplement, which contains information on individuals' labor force experience during the entire preceding year (e.g., Murphy and Topel 1987). However, monthly data on spells in progress are available on a more timely basis and provide better variation for investigation of time-series properties such as cyclical sensitivity and secular trends. Moreover, as described by a variety of authors, unemployment data are plagued by response biases that are likely to be more severe in retrospective data (Akerlof and Yellen 1982, Levine 1993).

The retrospective data are most useful for joint analyses of unemployment and labor force nonparticipation, as in Juhn, Murphy, and Topel (1991). I focus on unemployment because the two states are behaviorally distinct (see Flinn and Heckman 1983) and unemployment is of independent interest as a macroeconomic variable.

market conditions. Expected completed spell duration depends on the probabilities of continuing in or exiting unemployment as the spell proceeds. Estimation of expected completed durations proceeds as follows.

If the labor market is in steady state—i.e., entry and continuation rates for unemployment spells are constant over time and over the length of a spell—then the total number of unemployed at a particular time can be expressed as the product of incidence and average duration:

$$U = f(0) \cdot D. \tag{1}$$

In (1), U is the number unemployed, D is average expected duration, and $f(0)$ denotes the number of new entrants to unemployment (incidence) in a particular month, which is assumed constant over time in steady state. Then the steady-state estimate of expected completed duration in months, D, is simply the total number of unemployed U divided by the new entrants to unemployment. Using the number of persons unemployed less than five weeks as a measure of the monthly inflow $f(0)$, I can compute this statistic for total unemployment beginning in 1967 and for unemployment by reason beginning in 1976. This is the simplest estimate of expected completed duration based on monthly household survey data.[3]

In contrast, other authors (Sider 1985, Baker 1992a, and Baker, Corak, and Heisz 1998) estimated expected completed duration based on an approach that does not impose steady-state assumptions, which they argue is of particular importance when the cyclical properties of unemployment duration are of interest. The general nonsteady-state approach to estimating expected completed duration using grouped duration data is a "synthetic cohort approach," developed by Kaitz (1970) and Perry (1972).[4] 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 dependence, and they also will vary from month to month as economic conditions change.[5]

Because Sider (1985) and Baker (1992a) started with individual data, they were able to perform relatively exact calculations of unemployment tallies in duration intervals corresponding essentially to the monthly sampling window. This enables relatively precise estimation of the monthly continuation probabilities needed to form a nonsteady-state estimate of expected completed duration.

The monthly BLS data that I use also provide information on unemployment tallies within duration intervals. However, these intervals do not correspond to the monthly sampling window.[6] Moreover, the BLS intervals do not provide adequate information regarding long spells of unemployment. Although such data have been used in the past to form monthly continuation probabilities (e.g., Kaitz 1970), this approach requires substantial data smoothing and reassignment.

I therefore focus on steady-state estimates of expected completed duration. As discussed in the results section, this estimator closely replicates the cyclical properties of Baker's (1992a) nonsteady-state estimator for a comparable sample period. This may seem surprising, given that the nonsteady-state estimator is specifically designed to account for cyclical variability. However, the lagged information used by the nonsteady-state estimator implies that it is not a measure of the expected duration for an individual entering unemployment in the current month.[7] In contrast, although the steady-state estimator requires only information from the current month, the unemployment level used in its denominator reflects much of the same lagged information on continuation rates as is used in the nonsteady-state estimator. Also, Baker (1992b) noted that estimates of expected duration are more sensitive to data smoothing and allocation rules than to steady-state assumptions.[8]

3. An alternative steady-state measure of unemployment duration is the "experience-weighted" measure discussed by Akerlof and Main (1981). It measures the length of the spell to which the average week of unemployment belongs, or the expected duration of in-progress spells. This measure is substantially larger in general than expected completed duration for an individual entering unemployment. Previous authors have not examined the time-series properties of this measure.

4. 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.

5. An alternative approach, which allows for changing continuation or escape rates over the length of a spell (duration dependence), is based on estimating the parameters of a specified distribution of escape rates using data on ongoing spells. Salant (1977) and others used a gamma density to characterize escape rates in such a model. In terms of its approach to changing continuation rates, this model lies between the pure steady-state approach described above and the nonsteady-state approach used by Sider (1985) and Baker (1992a).

6. The published BLS duration intervals are < 5 weeks, 5–10 weeks, 11–14 weeks, 15–26 weeks, 27–51 weeks, and > 51 weeks.

7. 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.

8. Although Sider (1985) found substantial differences in results based on nonsteady-state and steady-state estimators, he used the gamma density approach of Salant (1977) rather than the steady-state estimator that I use.

A final estimation issue is "digit preference"—the tendency for measured durations to bunch at week values corresponding to integer multiples of one month and half-years (i.e., multiples of 4 or 26). Previously, analysts handled this problem by allocating a fixed percentage of bunched observations to the next monthly interval. For example, Sider (1985) and Baker (1992a) assigned 50 percent of the bunched observations to the next monthly interval. Baker (1992b) reports that although estimates of expected completed duration are very sensitive to the allocation rule, cyclical elasticity regression results are not. My estimator uses information only on the first monthly interval, and the weekly distribution within this interval is not identified. Given these considerations, I used a modified 50 percent allocation rule based on the grouped interval data. I assume a uniform distribution by weeks within the first monthly interval; the resulting allocation of 50 percent of the implied number at 4 weeks reduces the size of the entrant group by 12.5 percent. This makes my estimates of expected completed duration comparable to Sider's and Baker's; the regression estimates of the cyclical elasticity and time trend are unaffected.

III. RESULTS

Tabulations

Figures 1–4 show yearly average values of the unemployment rate, unemployment incidence by reason, and unemployment duration (total and by reason);[9] the unemployment rate is identified by the right-hand scale in each figure. Incidence is measured by the number of interrupted spells of less than five weeks duration during the sample month. The values for 1994–1998 in Figures 1–4 reflect adjustments intended to neutralize the impact of the 1994 survey redesign, as described in Section II.

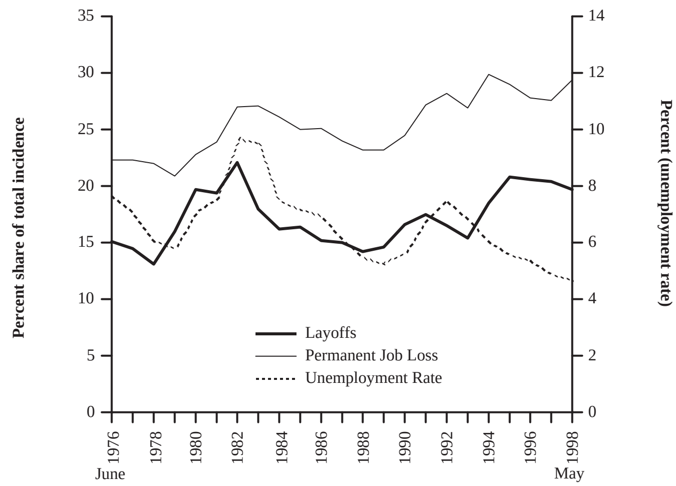
Figure 1 shows the unemployment rate and the shares of layoffs and permanent job loss in total unemployment incidence, each expressed in percentage points, for the period 1976–1998. Job losers on average account for about 43 percent of the newly unemployed during the period, with permanent job losses substantially outnumbering layoffs. Both series appear to be countercyclical, rising and falling with the unemployment rate. However, following the cyclical increase in the early 1990s, permanent job loss has remained very high throughout the decade, which is indicative of an upward trend. Moreover, layoff incidence increased sharply in 1994 and 1995 and has remained high. Overall, the rate of job loss in 1998 is above its sample period average, despite the low unemployment rate prevailing in 1998.

Figure 2 displays unemployment incidence shares for voluntary job leavers (quits) and labor force entrants. Job

9. The yearly average for 1998 is based on data for the first five months. Also, the incidence by reason figures first became available in June 1976, so the 1976 average is based on the last seven months of the year.

FIGURE 1

INCIDENCE OF INVOLUNTARY UNEMPLOYMENT (ANNUAL AVERAGES)



Note: Post-1993 figures corrected for 1994 survey change.

FIGURE 2

INCIDENCE OF VOLUNTARY UNEMPLOYMENT AND LABOR FORCE ENTRY (ANNUAL AVERAGES)

Note: Post-1993 figures corrected for 1994 survey change.

leaving constitutes a relatively small share of unemployment incidence—14 percent on average—but it exhibits a pronounced procyclical pattern. Re-entrant unemployment is frequent and appears to demonstrate moderate procyclicality, although its level has remained low in recent years. New entrant unemployment incidence exhibits limited cyclicality but an apparent downward trend.[10]

Figure 3 displays yearly average values for the steady-state estimate of expected completed duration of unemployment for all unemployed and by job loss category.[11] Each of these expected duration series exhibits noticeable countercyclicality. Permanent job losers on average endure long spells of unemployment; during the period 1976–1998, the expected duration of unemployment was 17 weeks for permanent job losers and 12 weeks for all unemployed.

Figure 4 shows expected completed duration for all unemployed, quits, and labor force entrants.[12] Individuals unemployed for each of these reasons experience durations around 10 weeks on average, slightly below the overall average, and all appear to exhibit countercyclicality.

Overall, the patterns in Figures 1–4 and underlying tabulations suggest that unemployment spells can be placed in long and short groups according to the reason for unemployment. Workers unemployed due to permanent job loss tend to suffer lengthy spells, with a high degree of countercyclicality in duration. Workers unemployed due to layoffs, voluntary mobility, or labor force entrance typically encounter relatively short spells. An apparent upward trend in unemployment duration may be linked to changes in incidence and duration of unemployment by reason, particularly the rise in permanent job loss. Implementation of more direct tests requires the regression approach described in the next subsection.

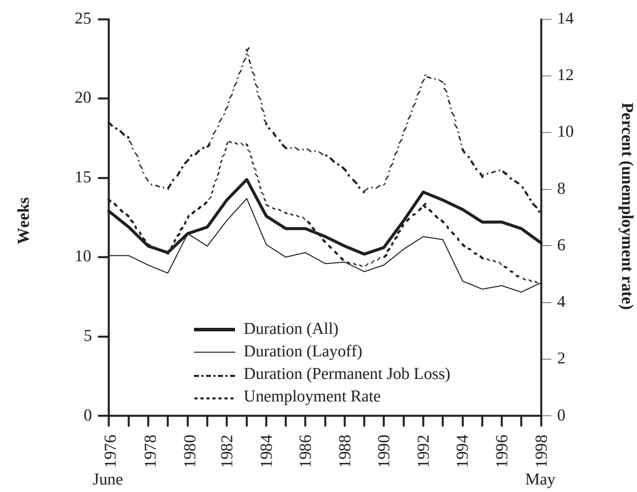
10. In Figure 2, the re-entrant series moves downward sharply between 1993 and 1994 (as does the duration of re-entrant unemployment in Figure 4). These movements may reflect the 1994 survey redesign and my data adjustment intended to overcome it. However, as noted in the text, the regression results reported in the next subsection are robust with respect to the survey redesign.

11. I multiplied estimates of expected duration in months by 4.3 to obtain expected duration in weeks.

12. 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).

FIGURE 3

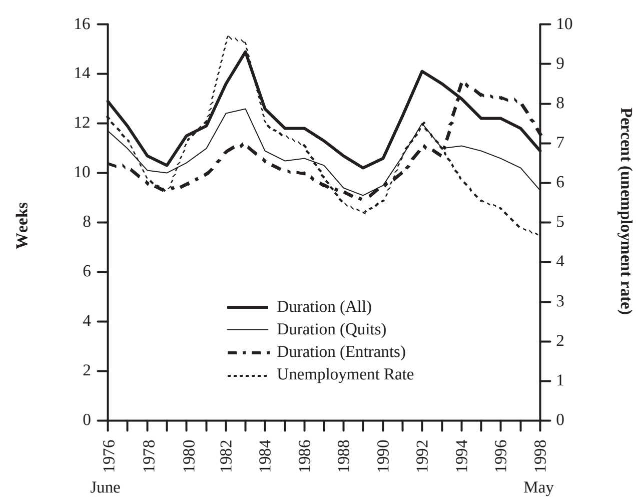
Expected Duration of Unemployment, Total Unemployment, and Job Losers (annual averages)



Note: Post-1993 figures corrected for 1994 survey change.

FIGURE 4

Expected Duration of Unemployment, Total Unemployment, Quits, and Labor Force Entrants (annual averages)



Note: Post-1993 figures corrected for 1994 survey change.

Regression Results

Following Perry (1972), Sider (1985), and Baker (1992a), I estimate time-series regression equations to explain the variation in monthly measures of unemployment duration and incidence. My basic approach is to regress these measures on the current unemployment rate (not seasonally adjusted), season dummies, a linear time trend, and a dummy variable that accounts for the CPS survey changes effective after 1993; unemployment duration and the unemployment rate are measured in natural logs.[13] I also report autocorrelation tests for all specifications; the tests are based on Durbin-Watson statistics for models without lagged dependent variables and Durbin's modified test (Durbin 1970) for models that include a lagged dependent variable.

Table 1 presents regression results for the steady-state unemployment duration models. Panel A lists results for total unemployment. The first row lists results for the full sample period of 1967–1998. The results indicate an elasticity of expected duration with respect to the unemployment rate of 0.45 and a small but statistically significant upward time trend. The Durbin-Watson statistic indicates that positive autocorrelation is present in this regression.

In the second row, I included a single period lag of the dependent variable in the regression. This specification is a partial adjustment model, in which the coefficients on the independent variables represent partial or short-run effects; the full long-run effects are obtained by adjusting the coefficients by a factor that is proportional to the coefficient on the lagged dependent variable.[14] Use of this specification in row 2 eliminates autocorrelation as a potential problem. Moreover, the adjusted coefficients indicate a full business cycle elasticity of 0.443 and full time trend effect of 0.026, which are identical to the estimates from the first row.[15] This time trend effect implies an increase in expected duration of about 10 percent between 1967 and 1998.

In rows 3 and 4 of Panel A, I present results for periods beginning in 1976. Both the estimated business cycle elasticity and the time trend are larger for this sample period than for the full sample. The transformed coefficients reveal a long-run business cycle elasticity of 0.57 and a full time trend effect of 0.064. This time trend effect implies nearly a 17 percent increase in the expected duration of unemployment between 1976 and 1998. A separate regression (not reported) verified the absence of a time trend for the period 1967–1980, which explains the substantially larger time trend estimated for 1976–1998 than for the full sample period. In row 4, I estimate the model on data for the years 1976–1993, in order to ensure that the results for the 1976–1998 period are not influenced by inadequate adjustment for the 1994 survey redesign. The results for this restricted sample period are virtually identical to those for the full period from row 3.

The estimated business cycle elasticities in rows 1–4 of Panel A are smaller than Baker's (1992a) estimate of 0.62 for the period 1980–1988, although they are close for the samples that begin in 1976. It is unclear whether this is due to differences across subperiods in the data or to different properties of my steady-state duration estimator and Baker's nonsteady-state measure. Row 5 presents results using my duration estimator and Baker's sample period. The full long-run business cycle elasticity implied by the estimated coefficient is 0.58, which is very close to Baker's estimate. Although full validation of the steady-state approach would require direct comparison of the two approaches for various sample periods, it appears that the steady-state estimator captures cyclical variation reasonably well. Row 5 of Panel B also indicates an upward trend in expected duration during the 1980s, although the shorter sample period reduces the statistical significance of the trend estimate compared to the preceding regressions.

Panel B of Table 1 presents results for expected unemployment duration by reason. The results indicate significant countercyclicality in expected unemployment durations for all reasons, as found by Baker (1992a). The largest cyclical effect is associated with job loss: the transformed coefficients on the unemployment rate in rows 1–3 indicate a business cycle elasticity of about 0.7 in each case. Smaller cyclical effects are evident for labor market entrants; the transformed coefficients imply an elasticity around 0.3.

The time trend effects by reason exhibit less uniformity than the cyclical effects by reason. A substantial upward time trend is evident for unemployment durations experienced by permanent job losers and labor market re-entrants; these time trends imply secular increases of approximately 24 percent and 11 percent in the expected unemployment durations of permanent job losers and labor market re-entrants,

13. This essentially replicates Baker's (1992a) specification. Sider (1985) used deviations from the trend in the index of industrial production (IIP) as his cyclical measure, and he included month dummies in the regression. My results are virtually identical when I replace my season dummies with month dummies. However, I do not report results using the IIP, because my preliminary regressions indicated that its relationship with labor market conditions has changed over time.

14. Specifically, if β is the coefficient on an independent variable and λ is the coefficient on the lagged dependent variable, the full effect of changes in the independent variable is $\beta/(1 - \lambda)$.

15. I estimated the sampling distributions of the transformed coefficients using the standard normal bootstrap approach (Efron and Tibshirani 1993). The resulting estimates of statistical significance are similar to those implied by the unadjusted coefficients and standard errors, so I do not list the sampling statistics for the transformed coefficients.

TABLE 1

UNEMPLOYMENT DURATION REGRESSIONS

DEPENDENT VARIABLE	ln (unemployment rate)	TIME TREND (x 100)	LAGGED DEPENDENT VARIABLE	AUTOCORRELATION TEST[a]
PANEL A: TOTAL UNEMPLOYMENT				
ln(duration), 1967–98	0.446**	0.026**	—	DW = 1.64
	(0.021)	(0.006)	—	
ln(duration), 1967–98	0.330**	0.019**	0.255**	–0.069
	(0.033)	(0.006)	(0.057)	(0.105)
ln(duration), 1976–98	0.450**	0.051**	0.206**	–0.255*
	(0.053)	(0.011)	(0.073)	(0.126)
ln(duration), 1976–93	0.429**	0.050**	0.259**	–0.082
	(0.055)	(0.010)	(0.078)	(0.137)
ln(duration), 1980–88	0.474**	0.070*	0.189	0.041
	(0.083)	(0.032)	(0.112)	(0.196)
PANEL B: UNEMPLOYMENT BY REASON, 1976–1998				
ln(duration), TOTAL JOB LOSERS	0.382**	0.050**	0.462**	–0.244*
	(0.049)	(0.011)	(0.057)	(0.091)
ln(duration), temporary layoffs	0.376**	0.022	0.469**	0.166
	(0.058)	(0.012)	(0.068)	(0.106)
ln(duration), permanent job losers	0.418**	0.055**	0.408**	–0.518**
	(0.053)	(0.012)	(0.057)	(0.093)
ln(duration), JOB LEAVERS	0.339**	0.003	0.227**	–0.380**
	(0.049)	(0.011)	(0.068)	(0.141)
ln(duration), RE-ENTRANTS	0.284**	0.041**	0.021	–0.478*
	(0.045)	(0.012)	(0.073)	(0.204)
ln(duration), NEW ENTRANTS	0.251**	–0.014	0.187*	–0.466*
	(0.064)	(0.018)	(0.075)	(0.198)

NOTE: Standard errors are in parentheses. Other explanatory variables included in all regressions are three season dummies and a CPS survey redesign dummy for regressions that include years beyond 1993 (post-1993 = 1). The number of observations is 376 for 1967–98, 263 for 1976–98, 210 for 1976–93, and 108 for 1980–88. The 1976 data begin in June.

* Significant at the 5% level, two-tailed test.

** Significant at the 1% level, two-tailed test.

[a] The test statistics are the Durbin-Watson statistic or the t test on the listed regression coefficients for single-period lagged residuals.

respectively, over the period 1976–1998. The expected duration of temporary layoff unemployment also appears to have risen somewhat; the transformed coefficient is significant at the 5 percent level and implies an increase in expected duration of 11 percent between 1976 and 1998.[16]

Table 2 reports results for linear probability models of unemployment incidence by reason.[17] Layoffs and other job losses are strongly countercyclical; job leaving and labor market re-entrance is procyclical; and new labor

16. The presence of negative autocorrelation in these models implies that the standard error estimates are conservative (i.e., the associated t tests probably understate significance levels).

17. 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.

TABLE 2

REGRESSIONS OF UNEMPLOYMENT INCIDENCE (SHARES) BY REASON, 1976–1998

DEPENDENT VARIABLE	ln (unemployment rate)	TIME TREND (x 100)	LAGGED DEPENDENT VARIABLE	AUTOCORRELATION TEST[a]
% TOTAL JOB LOSERS	0.108**	0.019**	0.265**	0.137
	(0.021)	(0.005)	(0.080)	(0.116)
% temporary layoffs	0.041**	0.000	0.356**	0.222*
	(0.013)	(0.003)	(0.066)	(0.097)
% permanent job losers	0.069**	0.022**	0.014	–0.402*
	(0.010)	(0.003)	(0.073)	(0.162)
% JOB LEAVERS	–0.062**	0.000	0.454**	0.397**
	(0.011)	(0.002)	(0.073)	(0.125)
% RE-ENTRANTS	–0.037**	–0.014**	0.039	0.303*
	(0.010)	(0.003)	(0.068)	(0.153)
% NEW ENTRANTS	0.003	–0.010**	0.034	0.402**
	(0.010)	(0.003)	(0.075)	(0.104)

NOTE: Standard errors are in parentheses. Other explanatory variables included in all regressions are three season dummies and a CPS survey redesign dummy (post-1993 = 1). The number of observations is 263.
* Significant at the 5% level, two-tailed test.
** Significant at the 1% level, two-tailed test.
a The test statistics are the t tests on the listed regression coefficients for single-period lagged residuals.

market entrants do not respond to business cycle conditions. The transformed coefficients imply long-run elasticities with respect to the unemployment rate that are broadly comparable to those obtained by Baker (1992a), although my estimates are somewhat smaller in general. I also find a significant upward time trend in the incidence of permanent job loss, consistent with results using individual panel data in Valletta (1998). The transformed trend coefficient indicates nearly a 6 percentage point increase in the incidence of permanent job loss between 1976 and 1998; this equals nearly a 25% increase relative to the sample mean incidence of permanent job loss. The results also reveal downward trends in the incidence of labor market entrant unemployment; the coefficients indicate approximately 10 percent and 20 percent declines between 1976 and 1998 in the incidence of re-entrant and new entrant unemployment, respectively.[18]

Decomposition Analysis

Three key results arising from the analysis thus far are: (1) longer durations of unemployment for permanent job losers than for other groups; (2) secular increases in total unemployment duration and duration for job losers; (3) rising incidence of permanent job loss to unemployment. It seems likely that rising duration of total unemployment in (2) is due in large part to the change in the composition of unemployment implied by (1) and (3).

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.[19] Recall from equation (1) that expected unemployment duration in a particular month is the number unemployed divided by the number of new entrants to unemployment during the month. Then expected total unemployment duration equals a weighted average of expected

18. Positive autocorrelation is evident for all but the permanent job loser and total job loser series. However, estimation using the Newey and West (1987) approach to account for autocorrelation indicated that the impact on the estimated standard errors is minimal.

19. Baker also performed decompositions by region, industry, education, and demographic groups. These other decompositions provided even less evidence in favor of the heterogeneity hypothesis than did the decomposition by reason for unemployment.

duration by reason, with the weights equal to the shares of unemployment incidence by reason. This property enables decomposition of total unemployment duration into two components:

D_{pc} ("probability constant")—expected total 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 3. The first row repeats the results for the unadjusted duration measure (row 3 from Table 1, Panel A). 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 for D_{pc}, which holds expected duration by reason constant, and large for D_{sc}, which holds incidence by reason constant. Most of the upward time trend also is attributable to rising duration by reason. However, about 20 percent of the upward trend in total duration is due to changing incidence by reason.

The decomposition listed in Panel A of Table 3 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. The first row of Panel B lists results from an alternative decomposition that focuses on permanent job loss. I formed the dependent variable used in the first row of Panel B 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 total expected du-

TABLE 3

Unemployment Duration Regressions, 1976–1998, Adjusted by Reason for Unemployment

Dependent Variable	ln (unemployment rate)	Time Trend (x 100)	Lagged Dependent Variable	Autocorrelation Test[a]
Panel A: Decomposition, Probability and Share Constant, All Reasons				
ln(D) (unadjusted)	0.450**	0.051**	0.206**	–0.255*
	(0.053)	(0.011)	(0.073)	(0.126)
ln(D_{pc}) (probability constant)	0.034**	0.013**	0.151*	–0.517**
	(0.005)	(0.002)	(0.069)	(0.166)
ln(D_{sc}) (share constant)	0.388**	0.037**	0.247**	–0.302*
	(0.047)	(0.010)	(0.070)	(0.122)
Panel B: Incidence and Duration by Reason Held Constant				
ln(D), perm. job loss constant	0.176**	–0.006	0.112	–0.086
	(0.023)	(0.005)	(0.075)	(0.155)
ln(D), entrant constant	0.384**	0.061**	0.184*	0.062
	(0.047)	(0.010)	(0.080)	(0.119)

Note: Standard errors are in parentheses. Other explanatory variables included in all regressions are three season dummies and a CPS survey redesign dummy (post-1993 = 1). The number of observations is 263.

* Significant at the 5% level, two-tailed test.

** Significant at the 1% level, two-tailed test.

[a] The test statistics are the t tests on the listed regression coefficients for single-period lagged residuals.

ration 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 effect and all of the time trend effect 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.

Finally, recall that Tables 1 and 2 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 3 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 only a small portion of the cyclical variability in total expected duration. Moreover, trends in the incidence and duration of labor force entrant unemployment reduced rather than increased total expected duration over time: the estimated time trend effect is larger when entrant unemployment duration and incidence are held constant (Panel B, row 2) than when they are allowed to vary (Panel A, row 1). This result occurs because the upward trend in expected duration for re-entrants (Table 1, Panel B) is offset by declining incidence for re-entrants (Table 2).

IV. CONCLUSIONS

Using steady-state measures of expected duration of unemployment, I find strong countercyclicality in unemployment durations during the period 1967–1998, which is relatively uniform across the different reasons for entry into unemployment. This result updates the previous literature into the 1990s. However, I also uncover an upward trend in duration that raised expected durations by approximately 17 percent between the years 1976 and 1998 (conditional on the unemployment rate as a measure of business cycle conditions). Like previous researchers, I also found substantial cyclical variability in unemployment incidence by reason for unemployment. I extend previous research by uncovering an upward trend in the incidence of permanent job loss and downward trends in the incidence of labor force entrant employment between 1976 and 1998. A decomposition of the increase in overall duration reveals that rising incidence and duration of unemployment due to permanent job loss can account for the full time trend effect during this period, with rising duration playing the dominant role. The incidence and duration of unemployment for other reasons made essentially no contribution to rising total unemployment duration.

Given the potential welfare losses associated with lengthy unemployment spells, the secular trend toward rising unemployment duration merits additional investigation. A useful research goal might be to identify the underlying economic forces that generate the link between rising permanent job loss and rising unemployment durations. In very recent work, Baumol and Wolff (1998) argue that technological progress can lengthen unemployment duration by increasing employment churning and skill mismatches. Their estimates suggest that several measures of technological change in the workplace are strongly associated with increases in the average interrupted spell duration between 1971 and 1994. However, drawing a reliable inference of a causal link between these two phenomena requires additional research. Other possible explanations of the trend toward rising duration include changing job search strategies by job losers and measurement issues related to movements between labor force states. Such hypotheses should be tested directly.

Depending on the underlying cause, the trend toward rising unemployment duration may have macroeconomic implications. Perry (1970) suggested that for a given unemployment rate, differences in the demographic structure of unemployment imply different degrees of wage pressure, since the value of output on the job varies across groups. In his comments on Perry's paper, Solow (1970) argued instead that the duration structure of unemployment may be more important than the demographic structure: "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" (p. 445). The model of Blanchard and Diamond (1994) formalizes this reasoning in a matching model of the labor market, and Duca (1996) presents evidence suggesting that longer unemployment durations help explain low wage and price inflation in the early 1990s. This interpretation of recent weak wage pressure is reinforced by the key role of permanent job loss in explaining lengthening durations: research on reemployment prospects of displaced workers indicates that they suffer larger wage losses upon reemployment than do workers who changed jobs for other reasons. Thus, the secular trend toward rising permanent job loss and rising durations may help to explain limited upward wage and price pressure in the current lengthy expansion.

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