CHINHUI JUHN
University of Houston

KEVIN M. MURPHY
University of Chicago

ROBERT H. TOPEL
University of Chicago

Current Unemployment, Historically Contemplated

Eleven years ago, our Brookings Paper "Why Has the Natural Rate of Unemployment Increased over Time?" analyzed long-term changes in joblessness among American men.[1] We documented the dramatic rise between 1967 and 1989 in both unemployment and nonparticipation in the labor force among prime-aged males. Our main conclusion was that a steep and sustained decline in the demand for low-skilled workers had reduced the returns to work for this group, leading to high rates of unemployment, labor force withdrawal, and long spells of joblessness for less-skilled men. We found that time spent out of the labor force and time spent unemployed accounted in roughly equal measure for the long-term growth in joblessness. We concluded that structural factors, primarily the decline in the demand for low-skilled labor, had dramatically changed the prospects for a return to low rates of joblessness any time soon.

After that paper was published, things appeared to change. The 1990s opened with a brief recession that was followed by the longest sustained decline in unemployment in modern U.S. history. By the end of that expansion, the unemployment rate had reached its lowest level since the late 1960s, falling below 4 percent for the first time since 1969. Some macroeconomists argued that the so-called natural rate of unemployment

We thank the Brookings Panel participants for helpful comments. Earlier versions of this paper were presented at the National Bureau of Economic Research Labor Studies meeting and at the University of Missouri. We acknowledge financial support from the Center for the Study of the Economy and the State and the Milken Foundation.

1. Juhn, Murphy, and Topel (1991).

had permanently shifted to 5 percent or below.[2] Because we had emphasized changes in the structure of labor demand that had made a return to low rates of joblessness unlikely, these facts presented a challenge to our 1991 framework. Maybe we were just wrong—maybe the demand and supply framework of our previous work is inconsistent with rates of joblessness in the post-1990 period. If so, we would join a distinguished group of social scientists who have drawn attention to a significant empirical phenomenon only to watch that phenomenon disappear immediately thereafter.[3] As it turns out, however, the framework that we developed for thinking about pre-1990 patterns of joblessness also does fairly well in helping to understand jobless time in the post-1990 period.

In this paper we look in some detail at employment data from the 1990s, revisiting issues raised in our earlier work. Specifically, we ask:

—Have the trends we identified in our earlier paper—the concentration of nonemployment among the less skilled, the growth of nonparticipation in the labor force, and the increased duration of joblessness—been reversed with the fall in aggregate unemployment?

—Did the expansion of the 1990s really return the U.S. labor market to conditions of the late 1960s, as unemployment statistics seem to indicate?

—Does the economic framework of supply and demand we utilized a decade ago still help in understanding long-term developments in unemployment, nonemployment, and labor force participation?

Our answers are surprising. First, the basic trends toward longer spells of joblessness and rising nonemployment have continued in spite of the prolonged expansion of national output and the concomitant fall in unemployment rates. Long jobless spells and labor force withdrawal were more important in the 1990s than ever before. Second, the fall in unemployment to levels close to historical lows is very misleading. Broader

2. See, for example, Stiglitz (1997), Gordon (1998), and Staiger, Stock, and Watson (2001).

3. Malthus founded the club. His theory that the forces of endogenous population growth doomed the common people to perpetual poverty "explained" why incomes had failed to increase over the period his data covered. Publication of Malthus's theory was followed by two centuries of almost continuous progress. More recently, when the returns to a college education were at a record low in 1979, Richard Freeman (1976) offered a supply-based theory in *The Overeducated American,* only to see returns to a college education increase steadily over the next fifteen years, reaching a record high. To Freeman's credit, his model did predict a rebound, although not so large and sustained as the one that actually occurred.

measures of joblessness show that the labor market of the late 1990s was more like the relatively slack labor market of the late 1980s than like the booming labor market of the late 1960s. Finally, the basic forces of supply and demand identified in our previous paper continue to have explanatory power. The theory does a reasonably good job of explaining those trends that have continued, as well as those that have changed.

Recent data also provide considerable insight into what has happened in the labor market over the past decade. Over the 1990s, even as unemployment was falling, time spent out of the labor force was rising. In fact, the increase in time spent out of the labor force was so large that total joblessness—which combines the unemployed with those who have withdrawn from the labor force—was as high at the business-cycle peak in 2000 as it had been at the previous cyclical peak of 1989, even though the unemployment rate was roughly 2 percentage points lower. In terms of total joblessness, the often-praised boom of the 1990s really represented little in the way of employment progress for American males.

Although the growth in the amount of time American males spend out of the labor force continues a trend found in our earlier research, other features of the data changed in the 1990s. The real wages of less-skilled men, which had been falling steadily since the early 1970s, stabilized in the 1990s and even rebounded slightly in the second half of the decade. It appears that the thirty-year trend toward greater wage inequality has run its course, at least at the bottom of the wage distribution. The data on joblessness reflect the impact of the changing wage trends. The long-term divergence in employment rates between low-wage workers and those with higher wages, so pronounced in our earlier work, has stopped, and unemployment and wage gaps across skill groups have narrowed. The congruence between patterns of change in wages and in employment comports with our previous work, which stressed demand-driven wage changes as the dominant factor driving secular changes in employment rates.

We are not the first to study the decline in unemployment in the 1990s. Others have emphasized changes in the composition of the labor force as a source of this decline. Robert Shimer found that aging of the labor force is important in explaining the decline in unemployment, particularly compared with the late 1970s.[4] Lawrence Katz and Alan

4. Shimer (1998).

Krueger investigated to what extent the withdrawal of the incarcerated population from the labor force, among other factors, has led to a drop in the aggregate unemployment rate.[5] Other papers have explored the role of improvements in job search technology. For example, David Autor argues that temporary help agencies may have helped improve the efficiency with which job seekers are matched with employers, thus bringing about a decline in frictional unemployment.[6] The arrival of the Internet may have also reduced search costs, although its impact is less certain.[7]

We show in this paper that a sharp decline in the incidence of jobless spells accounts for the lower unemployment rates of the 1990s, but that at the same time durations of spells have remained high. This fact is inconsistent with a theory built on declining search costs, which would imply shorter unemployment spells.

A related line of research compares the divergence in employment outcomes between the United States and Europe. Although both the U.S. and the EU economies may have experienced similar patterns of labor demand during the 1970s and the 1980s, it is widely believed that more-flexible labor markets and wages kept American unemployment rates relatively low, while European rates rose. Along these lines, several papers emphasize the importance of interactions between macroeconomic shocks and labor market institutions.[8] These papers find that although neither macroeconomic variables (oil prices, real interest rates, total factor productivity, the labor share of income) nor labor market institution variables (unemployment benefits and duration, union coverage, collective bargaining, employment protection policies) alone can explain the differences between the United States and Europe, a model that allows for interaction effects fits the data well. But this shocks-plus-institutions framework is less successful in understanding recent changes in U.S. unemployment. For example, Giuseppe Bertola, Francine Blau, and Lawrence Kahn

5. Katz and Krueger (1999). They conclude that up to 0.4 percentage point of the rise in the male employment-to-population ratio from 1985 to 1998 could be due to the bias from ignoring the institutionalized population. For the sample of prime-aged males we study here, the bias could be larger, further underscoring our finding that labor market conditions did not improve much for prime-aged males in the 1990s.

6. Autor (2000a).

7. Autor (2000b); Kuhn and Skuterud (2000).

8. Bertola and Inchino (1995); Blanchard and Wolfers (2000); Ljungqvist and Sargent (1998); Bertola, Blau, and Kahn (2001).

reported that the model significantly underpredicts the decline in U.S. unemployment in the late 1990s.[9]

We revisit the evolution of joblessness in the United States, using thirty-four years (1967–2000) of microdata from the Current Population Surveys (CPS) conducted by the Bureau of the Census and the Bureau of Labor Statistics. Our main conclusions are the following:

—Falling unemployment rates over the 1990s greatly exaggerate the improvement in labor market conditions for prime-aged males. Rates of overall joblessness—which include time out of the labor force—remained roughly the same in the late 1990s as they had been in the late 1980s, even as unemployment rates fell. Rising labor force nonparticipation among prime-aged men largely offset declining unemployment, so that the employment-to-population ratio held constant.

—Trends toward longer durations of both unemployment and nonemployment continued in the 1990s, in spite of declining unemployment rates. The probability of entering unemployment (or nonemployment) fell dramatically during the 1990s. The decline in the incidence of jobless spells was so large that the likelihood of experiencing one reached its lowest level in the thirty-four years covered by our data. But there was no decline in the duration of unemployment spells—these were about 2.8 weeks longer in 1999–2000 than they had been a decade earlier—and the duration of *non*employment spells increased by over four months during the 1990s. Broadly speaking, all of the long-term growth in joblessness is the product of longer durations of jobless spells.

—Although nonemployment continues to be concentrated among less-skilled men, the trend toward rising joblessness among the least skilled reversed course somewhat in the 1990s. The largest declines in unemployment occurred among men in the lowest skill categories. Unemployment among men in the bottom 10 percent of the wage distribution fell by 4.6 percentage points between the cyclical peaks of 1989–90 and 1999–2000, while the decline in unemployment at the median of the wage distribution was about 1 percentage point. In contrast, over the longer term the growth in nonemployment is heavily weighted toward less-skilled men. Among men at the bottom of the wage distribution, the nonemployment rate increased by 13.5 percentage points between the late

9. Bertola, Blau, and Kahn (2001).

1960s and 2000, but by less than 1 percentage point for men with wages above the median of the distribution.

—The long-term decline in the real wages of less-skilled men stopped in the early 1990s and actually reversed itself slightly in the latter part of the decade. Although the wages of highly skilled men grew most rapidly of all during the 1990s—continuing past patterns of relative growth—inequality between men at the bottom of the wage distribution and men at the median contracted slightly over the decade. Overall, the trend toward greater wage inequality appears to have stopped for males in the bottom half of the wage distribution.

—Joblessness among less-skilled men has shown up increasingly as time spent out of the labor force rather than as time spent unemployed. Consistent with our earlier work, we believe that this continued trend toward labor force withdrawal reflects two factors: relatively low returns to work (real wages for the least skilled remain substantially lower than in the past) and increasingly attractive nonwork opportunities, such as collecting disability payments, which have shifted labor supply among the least skilled. We find that more than 40 percent of the growth in nonparticipation is associated with an increase in men claiming to be ill or disabled.

—Despite rising wages and rates of labor force participation for women, the high rate of joblessness among less-skilled men is not the outcome of improved labor market opportunities for their working wives. Nonemployment rates and rates of labor force withdrawal increased most among men who did not have a working wife. Looking across the male wage distribution, the proportion of men with a working wife actually fell among low-skilled men, whose wages and employment rates were falling, and rose among men in the top 40 percent of the wage distribution, where wages rose and employment rates were stable. We conclude that long-term changes in joblessness have been the result of adverse shifts in labor demand, perhaps coupled with policy-driven shifts in labor supply, among low-skilled men.

Data

Our data are drawn from the 1968–2001 Annual Demographic Files that supplement the March CPS. The CPS collects information monthly from a rotating, random sample of approximately 50,000 U.S. households. It forms the basis for published government statistics on earnings, employment, unemployment, and labor force participation, among other measures. Whereas published labor market statistics rely on questions about each survey respondent's employment status in the reference week of the survey (usually the third week of the month), we study retrospective information, collected each March, on labor market outcomes in the previous calendar year. Hence our data cover the thirty-four calendar years from 1967 through 2000.

In addition to personal and household characteristics for each respondent, the retrospective data in the March survey record the number of weeks during the previous year that the respondent worked, was unemployed, and was out of the labor force, as well as the respondent's number of unemployment spells. We measure time spent unemployed (U) as the percentage of the year spent in that state (for example, for the ith individual, U_i is the number of weeks unemployed divided by 52); time spent out of the labor force (O) and time spent nonemployed ($N = U + O$) are measured in analogous fashion. This differs from the usual method of measuring time in unemployment, which divides weeks unemployed by weeks in the labor force. Our method better summarizes the allocation of time across the three states, and it naturally aggregates across individuals.[10] Using methods described below, we use information on weeks worked, unemployed, and out of the labor force to calculate both the incidence and the duration of jobless spells.

The survey also records a respondent's annual earnings and usual weekly hours worked from all jobs as well as occupation, industry, and other characteristics for the longest job held during the previous year. We use the information on earnings, weeks worked, and hours worked to calculate average hourly wages and to assign individuals a percentile position in the overall wage distribution, as described below. This allows us to track changes in employment outcomes (U, O, and N) for persons in different parts of the wage distribution.

We focus our analysis on males because they were the focus of our earlier work and because labor force participation issues for women are significantly more complex. To avoid issues associated with early retire-

10. Since the denominator for U_i, N_i, and O_i is always 52, the corresponding jobless rate is simply the sample average of weeks in the state divided by 52.

ment, Social Security, and pensions, we focus on men who have one to thirty years of potential labor market experience. For high school graduates this cutoff yields men who are roughly nineteen to forty-nine years of age, with correspondingly higher age intervals for those with more schooling. We define years of labor market experience as the smaller of two numbers: age minus years of education minus seven, and age minus seventeen.[11] In addition, in order to avoid measurement problems for men who spent part of the year in school or in the military, we exclude those who report that they did not work part of the year because of school or military service.

The employment measures we study are based on CPS respondents' weeks worked, weeks unemployed, weeks out of the labor force, and number of unemployment spells during the previous year, as reported in the survey week. Using these data, we are able to identify the fraction of respondents who experienced some unemployment or time out of the labor force during the year, as well as the number who worked no weeks during the year. We refer to the latter event as full-year nonemployment.

Imputing Wages for Nonparticipants and Other Adjustments

We construct two samples for analysis. The "wage sample" contains non-self-employed men for whom valid observations are available on annual earnings, weeks worked, and usual weekly hours.[12] For men in the wage sample, we calculate an hourly wage as the ratio of annual earnings to the product of weeks worked and usual weekly hours. The "employment sample" includes the entire wage sample plus those men who lack valid wage data because they did not work. For men not included in the wage sample, we impute a statistical distribution of wages based on education, experience, and weeks worked. For each individual with recorded earnings, weeks, and hours we project the log hourly wage on a quartic function in potential experience, and we assign each individual a per-

11. We use age minus education minus seven rather the standard measure, age minus education minus six, because age is measured at the survey week and we wish to measure potential experience at the time of our wage and employment measures (which is the year before the survey).

12. For the early years (before the 1976 survey) we impute usual weekly hours from hours worked in the last week and individual characteristics, and we impute weeks worked and unemployed from the categorical data based on averages calculated for the 1976–80 period.

centile rank based on his position in the distribution of the residuals. For persons with zero weeks worked in the previous calendar year, we impute a wage distribution based on the observed distribution of wages for those who worked from one to thirteen weeks in that year.[13] The imputation assigns ten probability weights—each corresponding to the probability that the individual's wage would come from a given decile of the wage distribution—along with a mean wage for each decile.

Our 1991 paper sought to explain changes in jobless time by changes in wages across skill categories. As we showed then, the relationship between calculated wages and time worked during a year is contaminated by measurement error in the latter, and the relative importance of this type of measurement error declines with the number of weeks worked in the previous calendar year.[14] This builds in a negative relationship between labor supply (weeks worked) and calculated wages, particularly among men with high calculated wages. As in our 1991 paper, we use data on hourly wages for March respondents who were also in the outgoing rotation groups to calculate the wage adjustments that would equate the distributions of calculated retrospective wages from the previous calendar year and reported hourly wages from the survey week. We then apply these adjustments to each percentile of the wage distribution. The procedure effectively compresses the wage distribution in each year by an amount that we attribute to measurement error in calculated wages.

Armed with calculated wages for those in the wage sample and an imputed wage distribution for those without valid wage data, we group individuals into five "skill" categories based on their positions in the wage distribution. The percentile intervals are 1–10, 11–20, 21–40, 41–60, and 61–100. As described above, each individual's wage percentile is calculated based on his wages relative to those of men with the same level of experience in a given year. Individuals in the wage sample are assigned to one of the five categories based on their actual wage, whereas those with

13. Men with zero weeks worked resemble those with one to thirteen weeks worked in terms of years of schooling completed and in terms of living arrangements (living alone, with a spouse, or with other family). We also matched the outgoing rotation groups to the March survey, yielding data on current (March) hourly wages for those who worked during the survey week. Among individuals with zero weeks worked in the previous year but who worked in the survey week, average log wages are nearly identical to those of men who worked one to thirteen weeks in the previous year. See Juhn, Murphy, and Topel (1991) for further details.

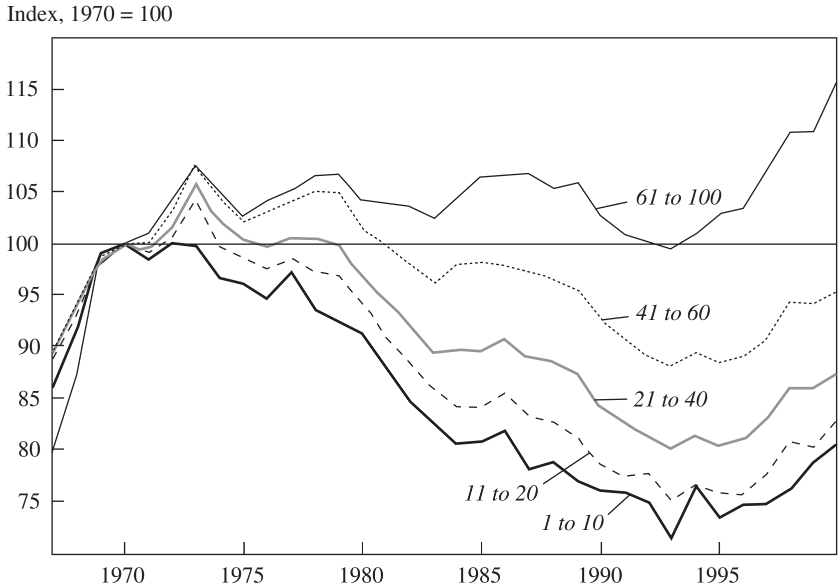
14. See Juhn (1992) for a more complete description.

imputed wages are assigned probabilities of being in each of the categories.

Changes in Wages, 1967–2000

In light of the prominence we assigned to wage changes in our 1991 paper, it is worthwhile to review what has happened to both wage levels and the distribution of wages since then. Figure 1 and table 1 summarize the main features of the data. Figure 1 shows trends in real hourly wages by wage percentile group (skill category) since 1967; the data are indexed to equal 100 in 1970. For our purposes the most interesting aspect of these data is that wage inequality stopped increasing in the 1990s, especially at the lower end of the wage distribution, and that the real wages of all skill categories increased after 1993. For less-skilled workers, real wage growth in the 1990s represented a slight reversal of a twenty-year decline in the returns to work, which had fallen by nearly 30 log points after 1972.

Figure 1. Real Wages, by Wage Percentile Group, 1967–2000[a]



Source: Authors' calculations based on annual March Current Population Survey (CPS) data.

a. Reported hourly wage (in natural logarithms) is projected on a quartic function in potential experience. Men are assigned a percentile category based on their position in the residual distribution. Wages for nonworkers and self-employed workers are imputed.

Table 1. Changes in Real Wages, by Wage Percentile Group, 1967–2000[a]
Percent

	Wage percentile group				
Period	*1 to 10*	*11 to 20*	*21 to 40*	*41 to 60*	*61 to 100*
1967–69 to 1988–89	–24.8	–19.9	–12.7	–4.0	5.5
1988–89 to 1994–95	–3.9	–7.1	–8.3	–7.6	–3.6
1994–95 to 1999–2000	6.1	6.7	6.8	6.4	10.7
1988–89 to 1999–2000	2.2	–0.4	–1.5	–1.2	7.1
1967–69 to 1999–2000	–18.7	–13.2	–5.9	2.4	16.2

Source: Authors' calculations using annual March Current Population Survey (CPS) data.
a. Reported hourly wage (in natural logarithms) is projected on a quartic function in potential experience. Men are assigned a percentile category based on their position in the residual distribution. Wages for nonworkers and self-employed workers are imputed.

Even so, average wages of the least-skilled men were roughly 20 percent lower in 2000 than in the late 1960s (table 1), whereas those of men in the top 40 percent of the distribution increased by a roughly equivalent amount. As we showed in our 1991 Brookings Paper, wage declines were most prominent among those whose employment outcomes are most sensitive to wage changes—the least skilled—whereas rising wages are concentrated among those with less elastic labor supply.

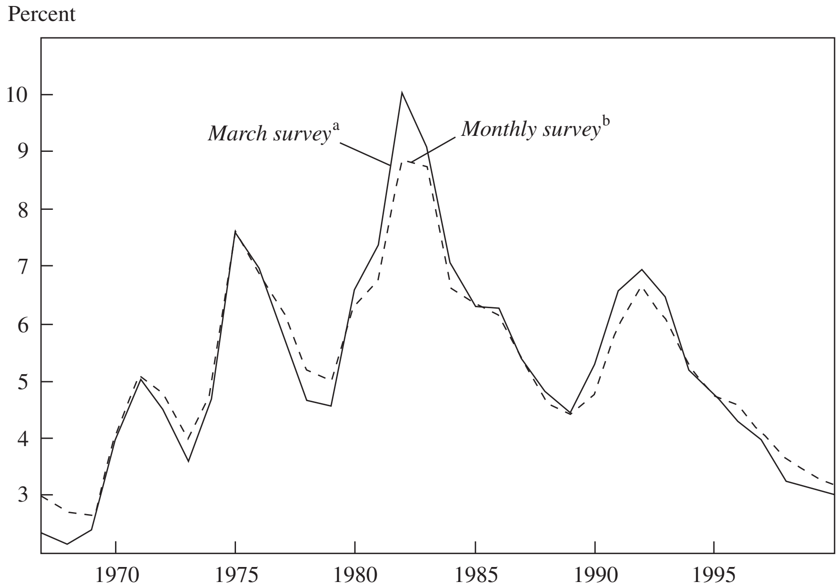
With this evidence as background, we turn to evidence on changes in joblessness, both in the aggregate and across the wage percentile groups defined above. We return to the implications of wage changes in the concluding section.

Unemployment, Nonparticipation, and Nonemployment

We begin by describing trends in unemployment and comparing our March CPS–based data with unemployment statistics from the monthly CPS published by the Bureau of Labor Statistics. Figure 2 shows that by 2000 the unemployment rate had reached its lowest level in thirty years, and unemployment rates in 1999–2000 were close to the extremely low rates seen during the late 1960s.[15] This is the culmination of a long downward trend in unemployment: in both the 1991–92 and the 2001–02 recessions (not shown), the peak unemployment rate was lower than the peak in the preceding recession, reversing a trend of rising peaks across the

15. The published rate has been adjusted downward by 0.86 percentage point to equate the means of the monthly and the March series over the sample.

Figure 2. Alternative Measures of the Unemployment Rate, 1967–2000



Source: Authors' calculations based on Bureau of Labor Statistics and March CPS data.
a. Subset including only prime-age males; calculated as the number of weeks unemployed divided by 52.
b. Each observation from the monthly survey is reduced by 0.86 percentage point to equate the means of the monthly and the March surveys over the sample.

1970–71, 1974–75, and 1982–83 recessions.[16] It appears from the figure that the U.S. economy has come full circle: unemployment rose for fifteen years (from 1968 to 1983) and then fell over the next seventeen years (from 1983 to 2000), with intervening cyclical swings. One might conclude from these data that the labor market conditions of the late 1960s and late 1990s were comparable.

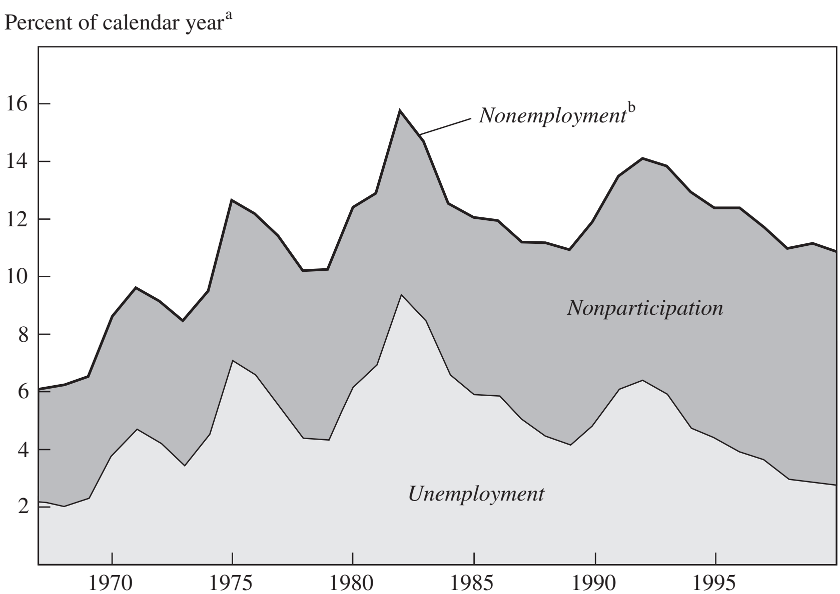
Figure 2 also shows annual unemployment rates for our sample of prime-aged males (calculated from the March CPS data as weeks unemployed divided by weeks in the labor force). Although the two series should not be identical because of differences in the underlying populations (our sample consists only of prime-aged males, whereas the overall unemployment data are from the full population of labor force participants aged sixteen and over), the two series are remarkably similar in terms of underlying trends and rankings of cyclical variations in unem-

16. The recession of 1980 did not fit this pattern, but as the figure shows, it did not represent much of a peak in terms of unemployment rates.

ployment. One would reach the same basic conclusions about unemployment trends from the published monthly series as from our calculations based on the March CPS data. From here forward we analyze the March CPS data exclusively.

A major finding of our 1991 paper was that the long-term growth in unemployment greatly understated the growth in joblessness. Recent data suggest that changes in unemployment shown in the aggregate and in the CPS statistics are even more misleading for the 1990s. This is illustrated in figure 3, which plots two series: the fraction of annual weeks spent unemployed and the fraction of annual weeks spent out of the labor force. The nonemployment rate is the sum of these fractions, so that the combined height of the two shaded regions represents the proportion of the year spent out of work. Figure 3 confirms that measured unemployment fell during the 1990s to levels comparable to those in the 1960s, but the conclusion in terms of overall jobless time is much different: the late 1990s were much like the 1980s, in that the decline in unemployment

Figure 3. Unemployment, Nonparticipation, and Nonemployment, 1967–2000



Source: Authors' calculations based on March CPS data.
a. Number of weeks a year in indicated state divided by 52.
b. Unemployment plus nonparticipation.

over the 1990s is not reflected in a lower overall rate of joblessness. This means that, on net, men who left unemployment did not find jobs but rather left the labor force, so that the employment-to-population ratio was unchanged from its level in the 1980s.

Table 2 summarizes the data in figure 3 by aggregating the data into nine time intervals corresponding roughly to peaks and troughs in the business cycle, as measured by aggregate unemployment rates. Unemployment shows a strong cyclical pattern as well as a long-run upward trend (whether measured peak to peak or trough to trough) until the recession of 1982–83. After 1982–83 unemployment rates fall or stay constant (again whether measured peak to peak or trough to trough). In contrast, the fraction of the year spent out of the labor force rises between every pair of intervals. In fact, whereas the unemployment rate in 1999–2000 is very close to its level in 1967–69, the nonemployment rate is 4.7 percentage points higher, and the fraction of the year spent out of the labor force is roughly double what it was in 1967–69. It is difficult to conclude from these data that employment conditions of the late 1960s and the late 1990s were "similar" in any meaningful sense.

Consider next the eleven-year interval between the business-cycle peaks of 1988–89 and 1999–2000. Over this peak-to-peak time span the unemployment rate fell by 1.3 percentage points—from 4.3 percent to 3.0 percent—but the percentage of men who were out of the labor force rose by exactly the same amount, from 6.7 percent to 8 percent. This left the nonemployment rate at the same level (11.0 percent) in 1999–2000 as in 1988–89, even though this period spans the longest sustained economic expansion, and the largest decline in unemployment, on record.

We next divide the growth in nonemployment along a second dimension. The percentage of weeks spent out of work is equal to the sum of two components: the fraction of men who did not work at all over the year (for whom the fraction of weeks spent out of work is 100 percent) and the fraction of weeks spent out of work for those who worked some positive amount (multiplied by the fraction of men who worked at least one week). In what follows we refer to these two components as "full-year nonemployment" and "part-year nonemployment," respectively. This decomposition allows us to examine how much of the growth in nonemployment is accounted for by men with very long stretches of joblessness—that is, spells that are so long that men do not work at all during a calendar year—and how much is due to men with "transitory" jobless spells.

Table 2. Unemployment, Nonparticipation, and Nonemployment during Business-Cycle Peaks and Troughs, 1967–2000

Units as indicated

		Labor force status (percent of calendar year[a])				
Period	*Phase of business cycle*	*Unemployed*	*Nonparticipating*	*Nonemployed[b]*	*Change in unemployment (percentage points)*	*Change in nonparticipation (percentage points)*
1967–69	Peak	2.2	4.1	6.3		
1971–72	Trough	4.5	4.9	9.4	2.3	0.8
1972–73	Peak	3.8	5.0	8.8	–0.7	0.1
1975–76	Trough	6.9	5.6	12.4	3.0	0.6
1978–79	Peak	4.3	5.9	10.2	–2.5	0.3
1982–83	Trough	9.0	6.3	15.2	4.6	0.4
1988–89	Peak	4.3	6.7	11.0	–4.7	0.5
1991–92	Trough	6.3	7.5	13.8	2.0	0.8
1999–2000	Peak	3.0	8.0	11.0	–3.3	0.5

Source: Authors' calculations based on March CPS data.
a. Number of weeks in indicated status divided by 52.
b. Unemployed plus nonparticipating; details may not sum to totals because of rounding.

The results, shown in table 3 and figure 4, are striking. The amount of joblessness accounted for by those working at least part of the year was only slightly higher in 1999–2000 than in 1967–69 (4.9 versus 4.5 percent). But the amount of joblessness accounted for by those who did not work at all over the year more than tripled, from 1.8 percent in the 1960s to 6.1 percent in 1999–2000. Moreover, whereas part-year nonemployment declined by 4.5 percentage points from its recessionary peak in 1982–83 to 1999–2000, full-year nonemployment increased slightly. This is particularly striking given that the intervening period is characterized by two of the longest economic expansions on record.

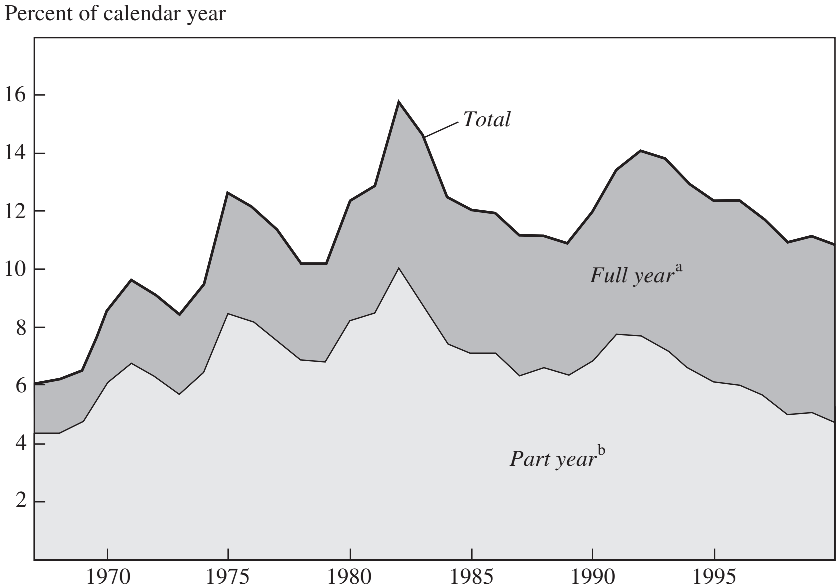
What explains this trend toward long-term joblessness? One possibility is that those men with the least favorable labor market prospects have simply dropped out of the labor market: the so-called discouraged worker effect. Figure 5 addresses this possibility by disaggregating nonemployment and nonparticipation, respectively, by reported reason. We distinguish among three main groups: those who reported that they could not find work, those who reported that they were ill or disabled, and a residual category we label "other." Over the period covered by our data, the figure shows only a small increase in the proportion of men who reported that they could not find work. Rising shares of the "ill or disabled" and "other" categories account for the largest changes of both nonemployment and nonparticipation. The larger impact of the "ill or disabled" category is on

Table 3. Part-Year and Full-Year Nonemployment, 1967–2000[a]
Percent of calendar year

Period	*Phase of business cycle*	*Part-year*[b]	*Full-year*[c]	*Total*
1967–69	Peak	4.5	1.8	6.3
1971–72	Trough	6.5	2.9	9.4
1972–73	Peak	6.0	2.8	8.8
1975–76	Trough	8.4	4.1	12.4
1978–79	Peak	6.8	3.8	10.2
1982–83	Trough	9.4	5.8	15.2
1988–89	Peak	6.5	4.6	11.0
1991–92	Trough	7.7	6.0	13.8
1999–2000	Peak	4.9	6.1	11.0

Source: Authors' calculations based on March CPS data.
a. Details may not sum to totals because of rounding.
b. Fraction of males nonemployed for part of the year multiplied by the average percent of the calendar year spent nonemployed for this group.
c. Fraction of males nonemployed for the entire year multiplied by the percent of the calendar year spent nonemployed (100 percent).

Figure 4. Part-Year and Full-Year Nonemployment, 1967–2000



Source: Authors' calculations based on March CPS data.

a. Fraction of males nonemployed for part of the year multiplied by the average percent of the calendar year spent nonemployed for this group.

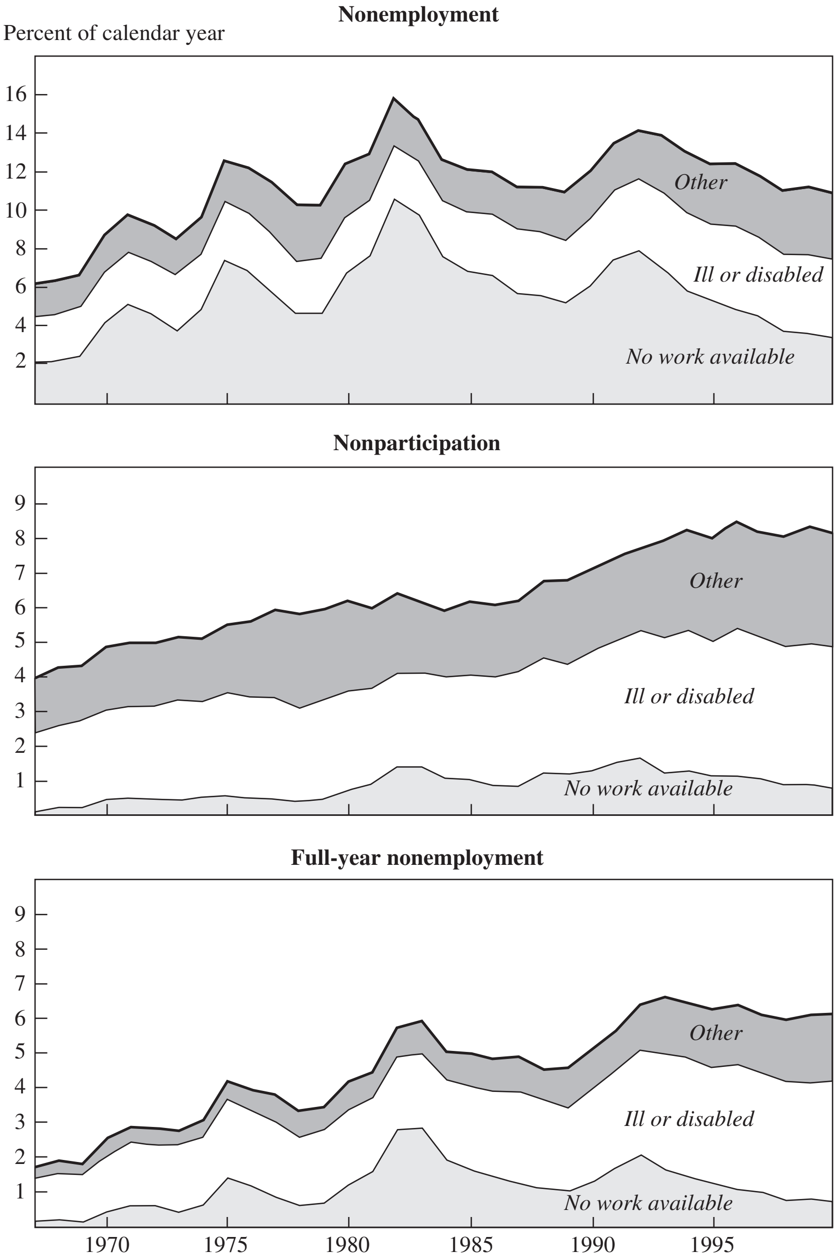
b. Fraction of males nonemployed for the entire year multiplied by the percent of the calendar year spent nonemployed (100 percent).

the nonparticipation rate (middle panel of figure 5): men in this category account for about 42 percent (0.8 out of 1.9 percentage points) of the increase in nonparticipation between 1982–83 and 1999–2000. The rest is "other." The bottom panel of figure 5 narrows the focus to men who were full-year nonworkers, for whom the effect of rising disability is more prominent still. Virtually none of the long-term increase in full-year joblessness is accounted for by discouraged workers: the "other" category and persons reporting joblessness for health reasons account for the secular increase.

A large literature examines the impact of changes in the disability benefits program on the labor market participation of male workers.[17] These papers document a substantial growth in the disability rolls in the early

17. Parsons (1980); Bound (1989); Bound and Waidmann (1992); Autor and Duggan (forthcoming).

Figure 5. Nonemployment and Nonparticipation, by Reported Reason, 1967–2000[a]



Source: Authors' calculations based on March CPS data.
a. Excludes individuals who report being students or retired.

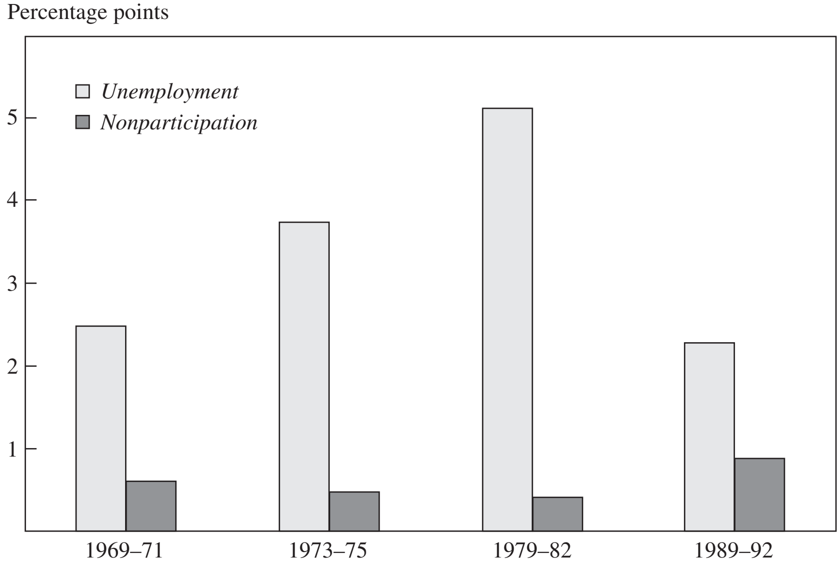
1970s, linked to the sharp decline in participation among older males. Because real wages were rising for the most part over this period, the earlier episode is consistent with a reduction in labor supply in response to the improving nonmarket alternative represented by disability payments. During the early 1980s, however, legislative and administrative changes tightened eligibility standards; these tighter standards led to reductions in new awards and terminated benefits for a substantial fraction of beneficiaries.

After 1984, eligibility criteria were substantially liberalized, and this led to increased receipt of disability payments.[18] Examining aggregate time series as well as cross-state variation, John Bound and Timothy Waidmann concluded that virtually all of the increase in nonemployment among those reporting that they were ill or disabled in the CPS could be explained by increased receipt of disability benefits.[19] Autor and Mark Duggan concluded that liberalization of eligibility for disability insurance interacts with adverse shifts in labor demand, as otherwise employable men opt for subsidized nonparticipation over unemployment or low-wage work. Figure 6 offers indirect supportive evidence on this point, comparing the changes in unemployment and nonparticipation between peaks and troughs of different business cycles. The figure shows that increased nonparticipation accounted for a much larger fraction of rising nonemployment in 1989–92 than in earlier recessions. The smallest contribution of nonparticipation was in the recession of 1992, when eligibility rules were tightened. The increase in nonemployment among the ill or disabled accounted for nearly 16 percent of the total change in nonemployment between 1989 and 1992 (not shown), a much higher proportion than in previous recessions. Nonemployment of the ill or disabled actually fell during the 1982 recession, an observation that also likely reflects the tightening of eligibility rules during this period.

Table 4 decomposes secular changes in nonemployment between 1967–69 and 1999–2000, as well as over the 1990s. In the 1990s the data indicate that roughly half (0.8 percentage point) of the 1.5-percentage-point increase in nonparticipation reflects a shift in labor supply caused by improving nonmarket alternatives to working. There is no reason to believe that the health of American men deteriorated over the decade (and much reason to believe that it improved).[20] Yet nonparticipation caused

18. Autor and Duggan (forthcoming); Bound and Waidmann (2000).
19. Bound and Waidmann (2000).
20. See Murphy and Topel (2001), for example.

Figure 6. Changes in Unemployment and Nonparticipation Entering Recessions



Source: Authors' calculations based on March CPS data.

by self-reported health reasons increased by 0.8 percentage point over the decade. Unlike in the early 1970s, when real wages were rising rapidly as nonparticipation increased, real wages remained low and were falling over the first half of the 1990s. This fact makes it more difficult to parcel out the component due to shifting labor supply. Yet with the increase in real wages in the latter half of the decade, continuing growth of nonparticipation indicates a shift in labor supply. In a manner analogous to interpretations of the European unemployment experience, the data indicate that the interaction of disability benefits and labor market shocks may be of key importance in understanding rising rates of labor force withdrawal.[21]

Figure 7 summarizes our previous results, showing long-term changes in three alternative measures of joblessness since the late 1960s. The unemployment rate shows the most dramatic improvement of the three measures in the 1990s, nearly returning to 1960s levels. By this common

21. Autor and Duggan (forthcoming).

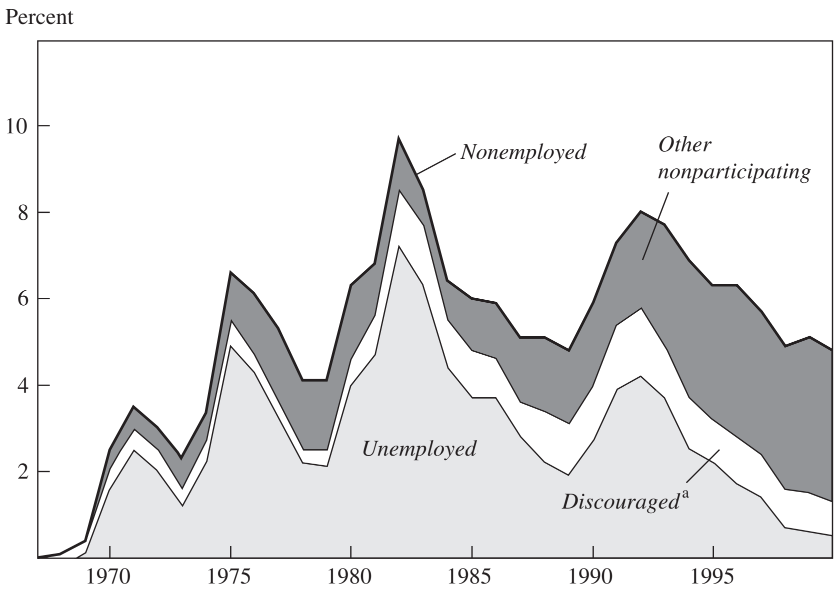
Table 4. Changes in Nonemployment, by Reported Reason for Nonemployment, 1967–2000[a]

Percentage points

Measure	*Reason for nonemployment*: *No work available*	*Illness or disability*	*Other*	*Total*
Change 1967–69 to 1999–2000				
Nonemployment	1.2	1.7	1.8	4.7
Out of labor force	0.6	1.7	1.7	4.0
Unemployment	0.6	0.0	0.0	0.6
Change 1988–89 to 1999–2000				
Nonemployment	–1.9	0.8	1.1	0.0
Out of labor force	–0.4	0.8	1.0	1.5
Unemployment	–1.5	0.0	0.0	–1.5

Source: Authors' calculations based on March CPS data.
a. Details may not sum to totals because of rounding.

Figure 7. Changes in Unemployed, Discouraged, and Nonemployed Workers Since 1967, 1967–2000



Source: Authors' calculations based on March CPS data.
a. Discouraged workers are those who report not being employed because they were unable to find a job; they are not included in the labor force.

measure of labor market performance, events have come full circle, and one might argue that the natural rate of joblessness has returned to previous low levels. Adding nonparticipants who are discouraged workers changes the conclusion slightly, although the figure also demonstrates that there has been no reduction in discouraged workers since the 1980s. Adding in other nonparticipants to give total nonemployment changes the interpretation substantially. By this measure there was no improvement in overall joblessness from the late 1980s to 2000, despite falling unemployment rates. In this sense, changes in unemployment provide a misleading picture of changes in employment opportunities and the likelihood of finding work.

The Incidence and Duration of Jobless Spells

The data on full-year nonemployment suggest that the concentration of unemployment and nonemployment increased dramatically over the period covered by our data. Table 5 provides further evidence, showing, for various periods, the distributions of spells of joblessness during a calendar year. The trend toward long-term joblessness is unmistakable. For example, in the 1960s, when the nonparticipation rate was 6.3 percent, men who were jobless for the entire year accounted for 28.8 percent of nonemployment. But by the end of the 1990s—when nonemployment reached 11 percent—full-year nonemployment accounted for over half of all joblessness. A similar pattern holds for unemployment (table 6). Although unemployment rates in 1999–2000 were roughly comparable to those in the 1960s, the share of unemployment due to short spells (one to thirteen weeks) fell by one-third, from 30 percent to 20 percent. Individuals with more than six months of unemployment accounted for about a quarter of all unemployment in the 1960s, but 46 percent by the end of the 1990s. These shifts toward long-term joblessness mean that particular *rates* of unemployment and nonparticipation have much different meanings today than in past decades.

To examine the increased importance of long spells more closely, we use information in the CPS to estimate both the incidence and the duration of jobless spells. Focusing first on unemployment, we note that the rate of unemployment can be decomposed into the product of two components: the probability of an individual entering unemployment (the entry rate),

Table 5. Distribution of Nonemployment, 1967–2000[a]

Percent of nonemployed

	Number of weeks a year					
Period	*2 or fewer*	*3 to 12*	*13 to 25*	*26 to 38*	*39 to 51*	*52*
1967–69	2.4	17.2	18.9	18.0	14.8	28.8
1971–72	1.5	12.9	18.8	21.2	15.9	29.7
1972–73	1.6	12.4	17.2	21.7	15.3	31.9
1975–76	1.1	9.9	16.6	22.5	17.2	32.7
1978–79	1.5	13.1	17.7	20.2	14.4	33.2
1982–83	0.8	8.1	13.9	20.6	18.3	38.3
1988–89	1.0	10.3	13.2	19.0	15.2	41.3
1991–92	0.8	8.8	12.7	19.1	14.9	43.8
1999–2000	0.7	8.0	9.7	13.9	12.2	55.5

Source: Authors' calculations based on March CPS data.
a. Details may not sum to 100 because of rounding.

and the average duration of an unemployment spell. Denote the instantaneous transition rates from employment (e) and out of the labor force (o) to unemployment (u) at date t by $\lambda_{eu}(t)$ and $\lambda_{ou}(t)$, respectively, and the corresponding rates at which individuals leave unemployment by $\lambda_{ue}(t)$ and $\lambda_{uo}(t)$. Then the rate of change in the unemployment rate is

$$(1) \qquad \frac{du(t)}{dt} = e(t)\lambda_{eu}(t) + o(t)\lambda_{ou}(t) - u(t)[\lambda_{ue}(t) + \lambda_{uo}(t)].$$

The steady-state fraction of weeks spent unemployed, $[du(t)/dt = 0]$, corresponding to the entry and exit rates at any given point in time satisfies

Table 6. Distribution of Unemployment, 1967–2000[a]

Percent of unemployed

	Number of weeks a year				
Period	*13 or fewer*	*14 to 26*	*27 to 39*	*40 to 49*	*50 to 52*
1967–69	30.3	42.3	14.9	8.9	3.6
1971–72	20.4	40.6	21.5	10.9	6.6
1972–73	20.6	39.9	21.0	10.4	8.2
1975–76	17.8	31.2	22.8	14.5	13.7
1978–79	27.0	34.2	19.0	11.8	8.0
1982–83	13.7	28.9	21.7	17.9	17.7
1988–89	22.6	33.5	18.5	14.8	10.5
1991–92	16.9	31.7	20.6	16.4	14.3
1999–2000	20.4	33.4	18.2	14.6	13.4

Source: Authors' calculations based on March CPS data.
a. Details may not sum to 100 because of rounding.

(2) $$u^*(t) = [1 - u^*(t)]\frac{\lambda_u(t)}{\lambda'_u(t)},$$

where

(3) $$\lambda_u(t) = \frac{e^*(t)}{1-u^*(t)}\lambda_{eu}(t) + \frac{o^*(t)}{1-u^*(t)}\lambda_{ou}(t), \text{ and}$$
$$\lambda'_u(t) = \lambda_{ue}(t) + \lambda_{uo}(t).$$

Here λ_u is the rate at which individuals enter unemployment, being a share-weighted average of entry rates for persons who are employed and those who are out of the labor force. Similarly, $\lambda'_u(t)$ is the rate at which individuals leave unemployment by becoming employed or by leaving the labor force. Since $1/\lambda'_u(t)$ is the average duration corresponding to the contemporaneous rate of exit from unemployment, and $[1 - u^*(t)]\lambda_u(t)$ is the expected number of spells of unemployment per year at the current entry rate, equation 2 has a natural interpretation in terms of entry and duration. Growth in the steady-state fraction of the year spent unemployed can be decomposed into growth in the probability of becoming unemployed (entry) and the average duration of unemployment spells.

To implement this framework empirically, we use two identities that correspond to equation 1 integrated over the year. The change in the unemployment rate from the beginning to the end of year τ is

(4) $$U_1(\tau) - U_0(\tau) = [1 - \overline{U}(\tau)]\overline{\lambda}_u(\tau) - \overline{U}(\tau)\overline{\lambda}'_u(\tau),$$

where $U_1(\tau)$ is the unemployment rate (measured as a fraction of the population) at the end of year τ, $U_0(\tau)$ is the corresponding rate at the start of the year, and $\overline{U}(\tau)$ is the average unemployment rate over the year. With these definitions, $\overline{\lambda}_u(\tau)$ and $\overline{\lambda}'_u(\tau)$ are weighted averages of the instantaneous transition probabilities to and from unemployment.[22] The expected number of spells of unemployment over the year is then

(5) $$S(\tau) = U_0(\tau) + [1 - \overline{U}(\tau)]\overline{\lambda}_u(\tau),$$

since spells are generated either by starting the year unemployed, $U_0(\tau)$, or by becoming unemployed during the year, $[1-\overline{U}(\tau)]\overline{\lambda}_u(\tau)$. To estimate the entry and exit parameters, we use the data from the CPS together with

22. The weights in these weighted averages are $(1 - u)(\theta)$ and $u(\theta)$, respectively, where θ indexes weeks over the year.

monthly data on aggregate rates to interpolate the starting and ending numbers for each year. Solving equations 4 and 5 gives our estimating equations for unemployment transitions as

$$\bar{\lambda}_u(\tau) = \frac{S(\tau) - U_0(\tau)}{1 - \bar{U}(\tau)} \tag{6}$$

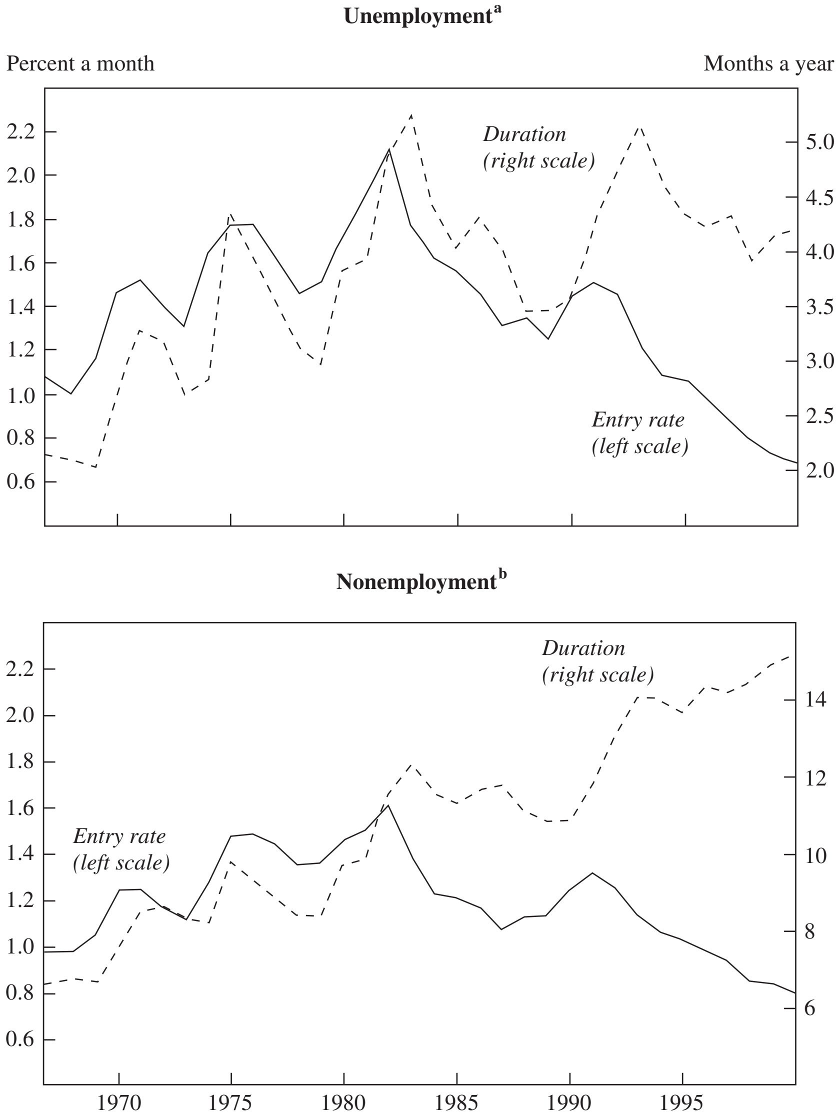
and

$$\bar{\lambda}'_u(\tau) = \frac{S(\tau) - U_1(\tau)}{\bar{U}(\tau)}. \tag{7}$$

The resulting estimates are shown in the top panel of figure 8 and in the first two data columns of table 7. For unemployment, the key finding is that an increase in durations accounts for the entire growth in unemployment over the 1967–2000 period. The entry rate into unemployment was actually lower in 1999–2000 (0.7 percent a month) than it was in 1967–69 (1.1 percent a month), whereas durations of unemployment spells doubled from 2.1 to 4.2 months. Notice from equation 3 that a declining incidence of unemployment spells can be caused either by a decline in the rate at which individuals lose their jobs, $\lambda_{eu}(t)$, or by a decline in the rate at which nonparticipants start to look for work, $\lambda_{ou}(t)$. These contributions are not separately identified, although it is likely that the $o \rightarrow u$ transition has declined substantially as nonparticipation has become a permanent labor force state for larger numbers of men. In any case, we cannot conclude from table 6 and figure 8 that the $e \rightarrow u$ transition has declined, that is, that jobs have become more stable.

According to figure 8, until the recession of 1991–92, cyclical fluctuations in unemployment were driven by changes in both the incidence and the duration of spells, with roughly equal weights on each component. But rising incidence played a minor role in the recession of 1991–92, while durations soared. Indeed, unemployment durations in 1993 were virtually the same as in the recession year 1983, which were the highest in all the years of our data, while the entry rate was about 25 percent lower. The ensuing decline in unemployment over the remainder of the decade is driven almost entirely by reduced probabilities of becoming unemployed; durations of unemployment remained high. From these data it appears that the main characteristic of the 1990s is that the historic correspondence between the incidence and the duration of unemployment spells

Figure 8. Estimated Entry Rates and Durations for Unemployment and Nonemployment, 1967–2000



Source: Authors' calculations based on March CPS data.
a. Estimated from the number of spells using equations 5, 6, and 7.
b. Based on incidence of nonemployment; see text for details.

Table 7. Estimated Entry Rates and Durations of Unemployment and Nonemployment, 1967–2000

Units as indicated

		Unemployment[a]		*Nonemployment*[b]	
Period	*Phase of business cycle*	*Entry rate (percent a month)*	*Duration (months)*	*Entry rate (percent a month)*	*Duration (months)*
1967–69	Peak	1.1	2.1	1.0	6.7
1971–72	Trough	1.5	3.2	1.2	8.6
1972–73	Peak	1.4	2.9	1.1	8.5
1975–76	Trough	1.8	4.1	1.5	9.6
1978–79	Peak	1.5	3.1	1.4	8.4
1982–83	Trough	1.9	5.1	1.5	12.0
1988–89	Peak	1.3	3.5	1.1	11.0
1991–92	Trough	1.5	4.5	1.3	12.4
1999–2000	Peak	0.7	4.2	0.8	15.1

Source: Authors' calculations based on March CPS data.

a. Estimated from the number of spells using equations 5, 6, and 7.

b. Based on incidence of nonemployment; see text for details.

came to an abrupt end. With fewer but longer spells, the population distribution of unemployment is much more concentrated than in earlier years.

The last two columns of table 7 and the bottom panel of figure 8 show corresponding calculations for the incidence and duration of nonemployment spells. In the case of nonemployment, the CPS does not provide information on the number of spells in a calendar year—separate spells of nonparticipation are not recorded—and so we use data on the incidence of nonemployment over the year (that is, the fraction of men with positive weeks of nonemployment) to infer the entry rate.[23] These calculations show that the contrast between entry and duration is even more extreme than in the case of unemployment. As with unemployment, the rate of entry to nonemployment is actually lower in 1999–2000 than it was in the 1960s, but durations show a steady upward trend over the thirty-four

23. The fraction of individuals who experience zero nonemployment (that is, who are employed for the full year) is given by $F(\tau) = E_0(\tau)\exp[-12\lambda^*(\tau)]$, where $\lambda^*(\tau)$ is the average monthly nonemployment hazard over the year for individuals who have not yet entered nonemployment, and $E_0(\tau)$ is the employment rate at the start of the year. In general, $\lambda^*_n(\tau) < \bar{\lambda}(\tau)$, where $\bar{\lambda}(\tau)$ is the average rate of transition to nonemployment for the population of employed people. This will cause our estimates of entry and exit rates to be biased downward. We attempted to assess the magnitude and variability in this bias with similar calculations for unemployment, where the number of spells is recorded. In that case the bias varied little over time, lending some confidence that this method should not be too far off.

years covered by our data, with no sign of slower growth in the 1990s. By the end of the decade the average duration of nonemployment spells was over fifteen months, which is more than double the length of spells in the late 1960s. The average duration of spells rose by over four months from the late 1980s to 1999–2000, reflecting the increasing proportion of men who have simply quit the labor force.

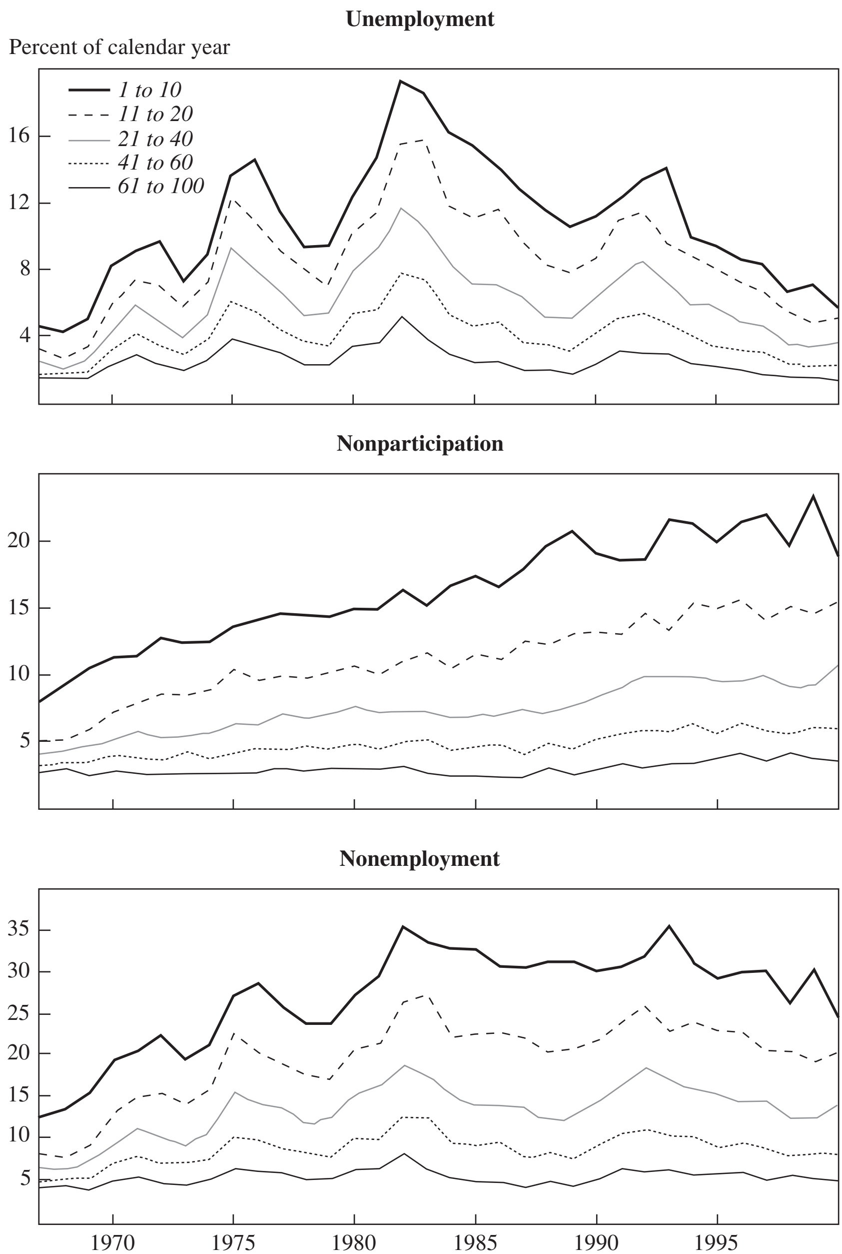
Table 7 and figure 8 paint a clear picture. Although the rates of entry into unemployment and nonemployment have returned to or even fallen below levels experienced during the late 1960s, the durations of jobless spells are more than twice as long at the end of the period. Indeed, jobless spells were longer in 1999–2000 than at any previous cyclical low of unemployment, and they exceeded the average duration of spells over the whole period of the data. It should be clear from these data that the employment patterns of the late 1990s resemble other periods of low unemployment—the late 1960s in particular—only in terms of the overall rate of unemployment and the rates at which individuals enter joblessness. The durations of spells are very different and very much longer. For the typical worker, the occurrence of a jobless spell is a far different event than it was in the past.

Unemployment, Nonemployment, and Wages

Our previous analysis of wage and employment data through 1989 found that the patterns of change in unemployment and nonemployment varied significantly across skill groups, as defined by percentile intervals of the overall wage distribution. Figure 9 and table 8 summarize our results based on wage percentile groupings for the period 1967–2000. Table 8 records changes in unemployment, nonparticipation, and total joblessness between 1967–69 and 1988–89 (the end of the data in our 1991 paper), between 1988–89 and 1999–2000, and over the full period of our data.

All components of nonemployment increased the most for low-wage workers, especially before 1989. Over the 1990s, nonparticipation continued to rise while unemployment rates declined sharply. Reversing previous trends, in the 1990s both unemployment and overall nonemployment fell the most for workers in the bottom 10 percent of the wage distribution. Other low-wage groups also experienced lower unemployment over

Figure 9. Unemployment, Nonparticipation, and Nonemployment, by Wage Percentile Group, 1967–2000



Source: Authors' calculations based on March CPS data.

Table 8. Changes in Unemployment, Nonparticipation, and Nonemployment, by Wage Percentile Group, 1967–2000
Percentage points

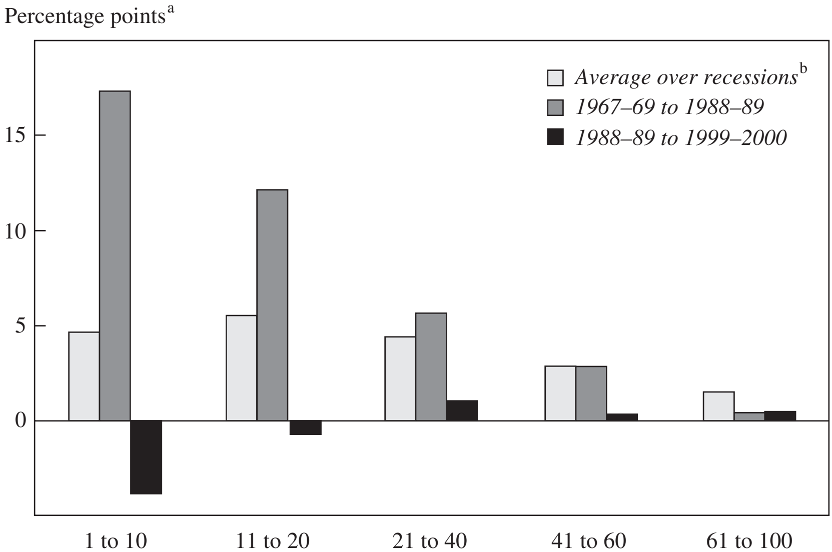
	Wage percentile group				
Period	*1 to 10*	*11 to 20*	*21 to 40*	*41 to 60*	*61 to 100*
Unemployment					
1967–69 to 1988–89	6.4	4.9	2.6	1.5	0.3
1988–89 to 1999–2000	–4.6	–3.1	–1.6	–1.0	–0.5
1967–69 to 1999–2000	1.8	1.8	1.1	0.5	–0.1
Nonparticipation					
1967–69 to 1988–89	10.9	7.2	2.9	1.3	0.0
1988–89 to 1999–2000	0.8	2.4	2.6	1.3	0.9
1967–69 to 1999–2000	11.7	9.5	5.5	2.6	0.9
Nonemployment					
1967–69 to 1988–89	17.3	12.1	5.6	2.8	0.3
1988–89 to 1999–2000	–3.8	–0.8	1.0	0.3	0.4
1967–69 to 1999–2000	13.5	11.3	6.6	3.0	0.8
Full-year nonemployment					
1967–69 to 1988–89	10.2	6.3	2.9	1.6	0.5
1988–89 to 1999–2000	1.7	2.7	2.4	1.3	0.9
1967–69 to 1999–2000	12.0	8.9	5.4	2.9	1.4

Source: Authors' calculations based on March CPS data.

the decade, although nonemployment was largely unchanged for workers above the bottom 10 percent, reflecting rising rates of labor force withdrawal. Even with the sharp decline in unemployment for low-wage workers in recent years, however, for the full period both unemployment and nonemployment increased most among the least skilled. Nonemployment rose by roughly 12 percentage points for the two lowest wage groups, but by less than 1 percentage point for men above the 60th percentile of the wage distribution. For workers near the median (percentiles 41–60), unemployment was essentially unchanged over the period as a whole, yet nonemployment increased by 3 percentage points.

Cyclical increases in joblessness are known to fall most heavily on the least skilled. Figure 10 compares the skill distributions of cyclical and secular changes in nonemployment. For each of the wage intervals shown in table 8, the figure shows the average cyclical change in jobless time going into and out of four recessions (1970–71, 1975–76, 1982–83, and

Figure 10. Cyclical and Secular Changes in Nonemployment, by Wage Percentile Group, 1967–2000



Source: Authors' calculations based on March CPS data.
a. Change in the percent of weeks a year nonemployed.
b. Average change entering (trough minus peak) and exiting (peak minus trough) four recessions (1970–71, 1975–76, 1982–83, and 1991–92).

1991–92),[24] the secular change between the cyclical peaks of 1967–69 and 1988–89 (the period covered in our 1991 paper), and the shorter secular change over the more recent 1988–89 to 1999–2000 period. Compared with business-cycle increases in nonemployment, the secular change in nonemployment from the late 1960s to the late 1980s was much more skewed toward low-skilled men, with virtually no impact on persons above the median of the wage distribution. The secular movement over the recent period is more nearly skill neutral, with the exception that nonemployment fell significantly for men in the first decile of the wage distribution. In this sense the 1990s represent a small reversal of declining

24. We measure this by the average of the change going into and out of each recessionary period. The periods are defined using the same year groupings used in the tables: 1967–69, 1970–71, 1972–73, 1975–76, 1978–79, 1982–83, 1988–89, 1991–92, and 1999–2000.

employment opportunities among the least skilled. We relate these changes to concomitant changes in the distribution of wages below, but first we take a brief detour to explore an alternative explanation for changing jobless rates among prime-aged men, namely, the labor market opportunities of their wives.

Did Working Wives Shift Male Labor Supply?

It remains our view that long-term changes in male joblessness were driven by changes in labor demand that disproportionately affected less-skilled workers. These adverse demand conditions continued into the 1990s, although somewhat mitigated, so that nonemployment continued to rise even as measured unemployment was falling. An alternative explanation—with far different welfare implications—is that increased wages and greater labor force participation of women have shifted men's labor supply: as the labor market opportunities of wives improved, husbands chose to work less and household utility rose. On this view, at least part of the long-term increase in nonemployment among men represents a welfare-improving change in household labor supply decisions.

Table 9 and figure 11 explore this issue. Table 9 records male earnings, the percentage of households with a working wife, average earnings of wives, and average household income for households in different percentiles of the male wage distribution. Among less-skilled men, where the largest increases in nonemployment occurred, the percentage of households with a working wife actually fell over time. For these men, declining marriage rates offset increased labor force participation of women, so that fewer low-wage men today reside with a working wife. For less-skilled men, average household income (which includes earnings of all household members) increased only slightly from 1972–73 (when the trend toward rising inequality began) to 1999–2000: average household income rose by 11 percent in each of the percentile intervals 1–10 and 21–40; in these groups, increases in the average earnings of working wives by 40 percent helped to offset declining male earnings. In contrast, the presence of a working wife is both higher and *rising* in households above the 60th percentile of the male wage distribution, where men's wages were rising and employment rates were stable. The share of these households in which a working wife was present increased by 9 percentage points after 1972–73 and by 12 percentage points from the 1960s, and

Table 9. Household Earnings Characteristics, by Wage Percentile Group, 1967–2000[a]
1996 dollars a year, except where noted otherwise

Item	*1967–69*	*1972–73*	*1978–79*	*1988–89*	*1999–2000*
1st to 10th percentile					
Male earnings	11,134	11,354	11,479	8,625	9,584
Percent of households with working wife	38	36	36	31	29
Wife's earnings[b]	3,050	3,392	3,766	3,975	4,770
Household income	22,965	26,366	28,140	25,952	29,194
21st to 40th percentile					
Male earnings	23,746	25,213	23,579	21,522	21,507
Percent of households with working wife	49	48	48	46	40
Wife's earnings	5,132	5,773	6,133	7,410	8,091
Household income	35,394	39,137	40,770	40,773	43,431
61st to 100th percentile					
Male earnings	44,213	49,375	47,147	51,449	59,843
Percent of households with working wife	41	45	50	55	54
Wife's earnings	4,786	5,918	7,219	11,537	14,993
Household income	54,456	61,722	61,861	71,937	88,507
All households					
Male earnings	31,230	34,121	32,686	33,193	36,789
Percent of households with working wife	44	45	48	49	45
Wife's earnings	4,713	5,548	6,388	8,895	10,854
Household income	42,263	47,538	48,560	52,821	61,541

Sources: Authors' calculations based on March CPS data and *Economic Report of the President,* 2001.

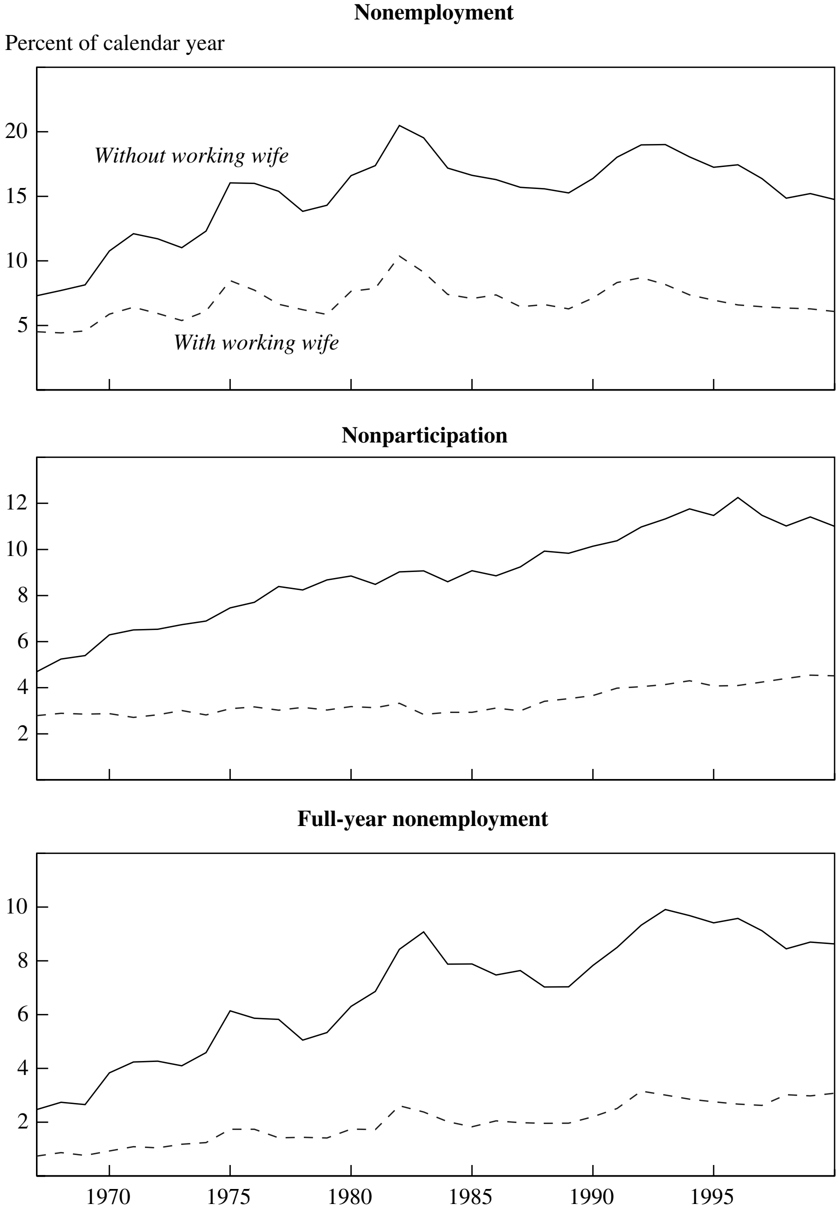
a. Earnings and household income are deflated by the price index for personal consumption expenditure. Male's and wife's earnings include income from wages, salary, and self-employment; household income includes all sources of earned and non-earned income of all household members.

b. Average of wife's earnings for all males in the sample; the observation is zero when there is no working wife.

the average earnings of these wives increased by 153 percent between 1972–73 and the end of the century.

It is difficult to square these facts with the view that long-term increases in labor force withdrawal among men have been driven by improved labor market opportunities for their wives. To settle the issue, figure 11 shows changes in unemployment, nonparticipation, and full-year nonemployment for men with and without working wives. The clear pattern is that rising unemployment and labor force withdrawal have been concentrated among men who do *not* have a working wife. The contrast in trends is particularly striking for nonparticipation and full-year nonemployment, where men without a working wife have steadily withdrawn

Figure 11. Male Nonemployment and Nonparticipation, with and without a Working Wife, 1967–2000[a]



Source: Authors' calculations based on March CPS data.

a. To qualify as a working wife, the wife must both live with her husband and have worked at least one week during the previous year.

from regular employment. From these data we conclude that a theory built on shifts in household labor supply will not go far in explaining changes in male joblessness.

Wage Changes, Labor Supply, and Nonemployment

So far our discussion has focused on changes in unemployment and nonemployment over time. Figure 1 and table 1 showed that many of the same patterns observed with regard to employment and unemployment hold for real wages. Inequality in real wages grew significantly from 1970 to 1990 across the full range of the wage distribution. Since 1990, inequality has continued to increase at the top of the wage distribution, but inequality has held steady or even narrowed slightly at the bottom: both low-wage and middle-wage workers experienced real wage increases starting around 1995. These increases in real hourly wages represent the first significant growth in real wages for low- to middle-wage males since the early 1970s. According to our earlier analysis, rising wages for these groups should lead to increased employment rates, especially among the least skilled, for whom we concluded that labor supply elasticities were largest.

At a general level, trends in nonemployment by wage percentile group (bottom panel of figure 9) and trends in real wages for these same groups (figure 1) reveal a similar pattern. In both cases low-wage workers fared far worse than their middle- and high-wage counterparts for much of the sample period, and in both cases the divergence stops in the 1980s (after roughly 1983 in the case of employment, and after roughly 1989 in the case of wages). Our earlier paper formalized this connection, arguing that declining rewards to work provoked labor supply responses from less-skilled workers, who chose to work less. To what extent does the demand-driven explanation we stressed in our earlier paper—that individuals respond to changing real wage opportunities—help us to understand the changes since 1989 in employment for men in different skill categories?

Table 10 presents estimated partial labor supply elasticities obtained from cross-sectional data for the years 1972–73 and 1988–89.[25] Our estimates correspond closely to those reported in our earlier paper. They

25. To obtain these elasticities, we fit a quadratic function using average wage and employment data by percentile category. We report the slope at each percentile.

Table 10. Estimated Labor Supply Elasticities and Changes in Nonemployment and Wages, 1972–2000[a]

Units as indicated

Wage percentile group	*Partial labor supply elasticity*[b]	*Change in real log hourly wage (percent)*	*Change in nonemployment (percentage points)*	
			Predicted	*Actual*
1972–73 to 1988–89				
1 to 10	0.287	–24.8	7.0	10.2
11 to 20	0.217	–22.3	4.7	5.7
21 to 40	0.170	–16.5	2.7	2.6
41 to 60	0.126	–9.2	1.1	0.8
61 to 100	0.048	–0.2	0.1	–0.1
Entire sample	n.a.	–9.9	2.0	2.2
1988–89 to 1999–2000				
1 to 10	0.287	2.3	–0.7	–3.8
11 to 20	0.217	–0.4	0.1	–0.8
21 to 40	0.170	–1.5	0.3	1.0
41 to 60	0.126	–1.2	0.2	0.3
61 to 100	0.048	7.0	0.1	0.4
Entire sample	n.a.	2.4	0.1	0.0

Source: Authors' calculations based on March CPS data.

a. Labor supply elasticities and observed changes in real wages are used to predict the change in nonemployment.

b. Elasticities in both panels are estimated using cross-sectional data on average wage and employment, by percentile, for 1972–73 and 1988–89.

show substantially higher elasticities at low wages: the employment rates of less-skilled workers are more responsive to wage changes. The top panel of the table illustrates the fact that the large wage declines (shown in the second data column), together with the estimated elasticities, can account for most of the rise in nonemployment from 1972–73 to 1988–89. The bottom panel of the table uses the same labor supply elasticities estimated from pre-1990 data, together with post-1989 wage changes, to predict changes in nonemployment during the 1990s. There is a reasonable correspondence between the predicted and the actual changes: we predict an improvement in employment for the lowest wage group and somewhat worsening conditions for the other groups below the median. Yet we also underpredict the improvement in employment for the least-skilled group. Although wages and employment are obviously linked in the long run, our interpretation of the results is that the labor supply model is less successful in predicting the dynamics of employment and wage changes over a relatively short period. Notice also that employment gains preceded the recovery in wages among the least skilled, which is inconsistent with a pure labor supply explanation of changing employment rates.

Conclusion

We have examined unemployment and nonemployment among prime-aged males in the United States using thirty-four years of data on labor market outcomes. Although recent unemployment rates have fallen to levels reminiscent of the 1960s, we find that rising nonparticipation rates have offset reductions in unemployment, leaving nonemployment rates unchanged. The rise in nonparticipation appears to be due to both an expansion of the disability benefits program—as previous research has argued—and continued low levels of real wages of less-skilled men during the 1990s.

Compared with earlier decades, the increase in nonparticipation in the 1990s is more evenly distributed across skill groups. Employment rates of the least skilled rose the most, even as their wages lagged behind other groups for much of the decade. This suggests that rising inequality, which characterized labor markets in the 1970s and 1980s, may have run its course.

Is the American labor market today fundamentally different from that of the 1960s? Despite the comparability of unemployment rates between the late 1960s and the late 1990s, the changing composition of nonemployment—from unemployment to nonparticipation, and from part-year to full-year nonemployment—suggests that the combination of low wages and the availability of nonwork alternatives has made out-of-work males increasingly less likely to enter new jobs. From this perspective, our assessment of the labor market for less-skilled men is rather grim.

Our earlier work concluded that the natural rate of unemployment, or of nonemployment, is not a constant toward which the economy gravitates over time. Rather, it varies with labor market conditions in a manner consistent with the original formulation of Edmund Phelps.[26] Over the long term, the natural rate of nonemployment has increased because

26. Phelps (1974).

changing patterns of labor demand have reduced the returns to work for less-skilled men. In this setting, the low unemployment rates of the latter half of the 1990s have a far different interpretation than comparable rates of the past. By the end of the 1990s, an important proportion of less-skilled men had withdrawn from the labor force for demand-related reasons. That they are not counted among those seeking work is not a sign of strength in current labor market conditions.

Comments and Discussion

Lawrence F. Katz: Chinhui Juhn, Kevin Murphy, and Robert Topel have produced an insightful and informative extension into the 1990s of their earlier important work on the evolution of joblessness among U.S. prime-aged males. Their earlier study documented a large increase in the nonemployment rate of prime-aged males from 1967 to 1989, concentrated among low-skilled (low-wage) workers and in long-term spells of joblessness.[1] Increases in the shares of men classified as unemployed and as out of the labor force contributed to this rise in the nonemployment rate. The authors concluded that a secular decline in the demand and labor market opportunities for low-skilled males was the driving force behind rising U.S. male nonemployment in the 1970s and 1980s.

In their new work, the authors find that some of the earlier trends continued into the 1990s and some did not. The share of prime-aged men not participating in the labor force continued to rise in the 1990s. The large reduction in unemployment in the 1990s for prime-aged men was completely offset by this increase in nonparticipation, so that the overall nonemployment rate for these men did not decline from the late-1980s business-cycle peak to the 1999–2000 peak. The rise in prime-aged male nonparticipation is concentrated in full-year nonemployment, and those listing illness or disability as the main reason for nonemployment account for a large share of the growth in male labor force nonparticipation (0.8 percentage point of a 1.5-percentage-point increase from 1988–89 to 1999–2000).

1. Juhn, Murphy, and Topel (1991).

On the other hand, the authors document that the rise in nonemployment among low-skilled men (those in the bottom 40 percent of the wage distribution) of the 1970s and 1980s stopped and may even have reversed itself in the 1990s. But nonemployment and nonparticipation continued to rise for men in the top half of the wage distribution. And the trend of large reductions in the real wages of low-skilled men stopped: these men saw real wage increases in the second half of the 1990s. Thus the 1990s expansion generated less inequality in labor market outcomes for prime-aged males than the experiences of the 1970s and 1980s.

The authors call their paper "Current Unemployment, Historically Contemplated." A wordier but more appropriate title would be "Almost-Current U.S. Prime-Aged Civilian Noninstitutional Male Unemployment, Contemplated through the Lens of the March Current Population Surveys." One reason is that the authors limit their analysis to prime-aged U.S. males (which they define as those with one to thirty years of potential experience) and focus on the information available in the March Current Population Survey (CPS) through calendar year 2000. Prime-aged males are indeed a key labor force group, but women and older workers (those with more than thirty years of potential experience) have become increasingly important labor force participants in recent years. In addition, their analysis is not fully "current," because it does not include the most recent recession. Furthermore, the use of the March CPS limits the analysis to the civilian noninstitutional population and thus fails to capture a major component of the rise in male nonemployment: the massive increase in U.S. incarceration rates over recent decades. The CPS does not include institutionalized groups (such as those held in state and federal prisons and in jails). Finally, the paper addresses only employment in the United States; some useful perspective on U.S. employment patterns could be gained through comparisons with other major developed economies.

The inclusion of other important labor force groups (women and older males) in the analysis, and comparison of the U.S. experience with that of other economies of the Organization for Economic Cooperation and Development (OECD), would generate a somewhat more positive overall assessment of U.S. employment performance over the full period studied by the authors, and especially the 1990s. U.S. female unemployment has declined and converged with male unemployment: there has been no rise in the unemployment rate for prime-aged females despite a more than

50 percent rise in the employment rate for women since 1969. And the long-term trend of a declining employment rate for near-elderly men (those fifty-five to sixty-four years old) ceased in the United States in the 1990s. Sustained real wage growth in the 1990s boom translated into greater reductions in measured poverty than in the 1980s and 1970s. Economic prosperity was more widely shared in the 1990s expansion than in the 1980s expansion.

Also, in comparative perspective, the U.S. labor market provided improving opportunities in the 1990s for prime-aged males relative to OECD Europe and to Japan. Over the 1990s the nonemployment rate for prime-aged males (those twenty-five to fifty-four years old) was stable in the United States (10.9 percent in 1990 and 11.0 percent in 2000), but it increased sharply in both OECD Europe (from 11.0 percent in 1990 to 13.6 percent in 2000) and in Japan (from 3.8 percent in 1990 to 6.5 percent in 2000). Over the same period, the nonemployment rate actually declined by 0.4 percentage point for U.S. men aged fifty-five to sixty-four, while it increased substantially (by 5.7 percentage points) in OECD Europe. And the employment rate remained much higher and the unemployment rate much lower for adult females in the United States than in OECD Europe.[2]

On the other hand, the expansion of the sample to include the incarcerated would undo the apparent improvement in employment that the authors find for low-skilled U.S. males in the 1990s. Such a more complete sample leads to an even more pessimistic set of conclusions concerning the labor market for low-skilled males at the end of the 1990s relative to that of the late 1960s, despite similar aggregate unemployment rates.

I will focus the remainder of my comments on three issues: the role of changing disability policies in rising male nonemployment rates; the impact of rising incarceration rates; and some puzzles related to the authors' supply-side analysis of the role of real wage movements in changes in prime-aged male employment rates by skill group.

The authors find that increases in the share of individuals reporting illness or disability as the main reason for nonemployment contribute substantially to the rise in male nonemployment over the last several decades. Their analysis of the March CPS data indicates that about one-

2. Data on nonemployment and unemployment rates by age and sex for the United States, OECD Europe, and Japan are from OECD (2001, table C, pp. 216–20).

third of the rise in prime-aged male nonemployment over the last three decades (1.7 percentage points of a 4.7-percentage-point increase from 1967–69 to 1999–2000) results from those indicating illness or disability as their main reason for nonemployment. And as already noted, this group accounts for more than half the rise in the rate of nonparticipation in the 1990s. The authors suggest that the growing attractiveness of disability benefits relative to work and job search could help explain this pattern.

I strongly agree with this conclusion. In fact, the share of the U.S. nonelderly adult population receiving disability benefits, either Social Security Disability Insurance (SSDI) or the means-tested Supplemental Security Income (SSI), has expanded substantially over much of the last forty years. The disability rolls grew rapidly in the 1960s and 1970s, especially SSDI for males from forty-five to sixty-four years of age; this was followed by a strong clampdown on disability recipiency in the early 1980s. Congressional legislation in 1984 ended the clampdown and relaxed the eligibility rules and screening criteria for SSDI, with a broader definition of disability (especially providing more flexibility on allowing claims of mental illness and pain), ending Continuing Disability Reviews for existing recipients, and providing applicants and their own medical providers greater opportunity to influence the decision process. The relaxation of eligibility requirements also applied to SSI. In the late 1980s and early 1990s, Congress mandated outreach efforts to inform potentially eligible low-income individuals of SSI benefits and to put greater weight on information from an SSI or SSDI applicant's own medical provider in award decisions.[3]

The result has been a resurgence in the growth of SSDI and SSI rolls since the mid-1980s. The share of nonelderly adults on SSDI increased from 1.8 percent in 1985 to 3.0 percent in 2000, with a similar increase for SSI, from 1.3 percent to 2.3 percent, over the same period.[4] Since about one-fourth of SSDI recipients also receive SSI,[5] the share of the nonelderly adult population receiving disability benefits (either SSDI or SSI) may be around 4.5 percent today, compared with 2.7 percent in

3. Autor and Duggan (forthcoming).

4. Data on the number of nonelderly adult SSDI and SSI recipients (those aged eighteen to sixty-four) are from the 2001 *Annual Statistical Supplement* to the *Social Security Bulletin,* tables 5.D3 and 7.A9. Data on the nonelderly adult population are from tabulations of the CPS provided by David Autor and Mark Duggan.

5. Bond, Burkhauser, and Nichols (2001).

1985. In fact, the rise in the disability rolls more than offset the more familiar decline in the welfare rolls (Aid to Families with Dependent Children and, after 1995, Temporary Assistance to Needy Families) of the 1990s in terms of the number of nonelderly adults supported by cash transfers. SSDI is now the largest income transfer program directed toward nonelderly adults, with cash transfers of approximately $50 billion in 2000.

The rise in the SSDI rolls is likely to be far more important for understanding rising labor force nonparticipation among prime-aged males than the rise in the SSI rolls. About 85 percent of SSDI applicants were employed for several years before applying, whereas only a small fraction (under 30 percent) of SSI applicants have been employed in the years before applying.[6]

The recent increase in the share of prime-aged men leaving the labor force to go on disability reflects the relaxation of screening criteria since the mid-1980s and the increasing generosity of SSDI benefits relative to work for low-skilled males. The progressive nature of SSDI benefits means that there has been a substantial increase in the replacement rates for low-wage males in the face of declining real wages for these workers since the 1970s. David Autor and Mark Duggan estimate that the cash income replacement rate for a male at the 10th (from the bottom) percentile of the wage distribution increased from 46 percent in 1979 to 54 percent in 1999 (and from 59 percent to 84 percent over the same period if one includes the value of Medicare benefits and in-kind employee benefits in the calculation).[7] They also present strong evidence that adverse labor demand shocks led to larger increases in disability applications and high rates of labor force withdrawal for less-educated males in the period since the liberalization of disability insurance benefits in the mid-1980s.

The growth in prime-aged adults receiving disability benefits seems to be a response to changes in screening and in economic incentives to enter disability programs and not due to a decline in true health status. The available evidence suggests improving rather than declining health over this period.[8] And Autor and Duggan find that the new flow of SSDI recipients increasingly comes from those whose disabilities are characterized

6. Bond, Burkhauser, and Nichols (2001).
7. Autor and Duggan (forthcoming).

by lower mortality rates and longer typical durations on the disability rolls (for example, musculoskeletal and mental disorders).[9]

How much of the rise in male nonemployment and nonparticipation documented by Juhn, Murphy, and Topel can be explained by a shift of workers onto the SSDI rolls? Table 1 of this comment shows that the SSDI recipiency rate for prime-aged males (here defined as those eighteen to fifty-four years old) increased from 1.0 percent in 1970 to 2.2 percent in 1999. Thus, possibly 1.2 percentage points of the 4-percentage-point rise in the share of (noninstitutional) prime-aged males out of the labor force since the late 1960s can be explained by the growth of SSDI. The growth in SSI might explain a little bit more. And SSDI growth is much larger for the low-skilled (less-educated) males who experienced the largest increases in nonemployment over the past three decades.

As I have mentioned, the authors' estimate of prime-aged male nonemployment from the March CPS covers only the noninstitutional population. The rapidly growing number of incarcerated males disappears from the population covered by the CPS and thus is missing from both the numerator and the denominator of the authors' estimates of nonemployment and nonparticipation rates. This incarcerated group is heavily concentrated among the less educated and the less skilled. An expanded measure of nonemployment (and nonparticipation) that consistently includes the incarcerated as nonemployed (and out of the labor force) indicates a substantially larger rise in male nonemployment and nonparticipation rates in recent decades. And such an expanded measure implies that the increase in male nonemployment since 1970 is even more concentrated among the less skilled than indicated by the authors' tabulations from the March CPS.

Table 1 also shows that the share of U.S. prime-aged males who are incarcerated increased even more rapidly than the SSDI rolls from 1970 to 1999. It increased by 1.7 percentage points, from 0.7 percent in 1970 to 2.4 percent over that period, including a 0.9-percentage-point increase in the 1990s alone. Thus, when the incarcerated are included, the prime-aged male nonemployment rate, rather than being stable in the 1990s, actually increased by 0.8 percentage point. And the overall increase in the prime-aged male nonemployment rate from 1967–69 to 1999–2000 rises by

8. As documented, for example, by Cutler and Richardson (1997).
9. Autor and Duggan (forthcoming).

Table 1. Disability Insurance Recipiency Rates and Incarceration Rates for Males Aged 18 to 54, 1970–99

Percent

Year	*Social Security disability insurance recipiency rate*[a]	*Incarceration rate*[b]
1970	1.02	0.67
1980	1.45	0.83
1983	1.19	0.97
1989	1.51	1.47
1990	1.59	1.55
1999	2.24	2.35

Source: Author's calculations based on data from Bureau of the Census, *Statistical Abstract of the United States*, various years; *Social Security Bulletin, Statistical Supplement*, 2000; Bureau of Justice Statistics, *Sourcebook of Criminal Justice Statistics*, 2000.

a. Ratio of SSDI recipients to resident population.

b. Ratio of jail inmates plus state and federal prisoners to resident population. Number of incarcerated males aged 18 to 54 is estimated from data on the total number of adult males incarcerated. Based on the characteristics of prisoners in state and federal correctional institutions in 1991 and 1997, it is assumed that 3.5 percent of incarcerated adults are aged fifty-five or older. For 1970 and 1980, where separate estimates by sex are not available, it is assumed that 90 percent of inmates are male; this is similar to the share observed in the early 1980s.

30 percent, from 4.7 to 6.1 percentage points. (The corresponding increase in the nonparticipation rate is an even greater 35 percent, from 4.0 to 5.4 percentage points.)

It is less clear what the causal impact of changes in incarceration policies has been, but this mechanical measurement effect should be taken into account for an accurate reading of recent changes in the labor force status of prime-aged U.S. males. A disproportionate share of the incarcerated are likely to be nonemployed when out of custody.[10]

Surveys of prisoners clearly indicate that the growth in the incarcerated population is concentrated among less-educated, low-wage, and minority males. This suggests that the March CPS may particularly understate the growth of nonemployment for low-skilled males. If, under a possibly conservative assumption, 80 percent of incarcerated males would be in the low-skilled group (the bottom 40 percent of the wage distribution), then the nonemployment rate for low-skilled males increased by an additional 3.4 percentage points from 1970 to 1999. The inclusion of the incarcerated also implies a slight rise in the nonemployment rate for low-skilled males in the 1990s, in contrast to the authors' finding of a modest decrease.

These patterns suggest that a major issue for social policy in the coming decade is how the labor market will treat the rising number of low-

10. Kling (2002) estimates that only 35 percent worked in the year before incarceration.

skilled males with criminal records and under the continuing supervision of the criminal justice system. Nevertheless, the economic recovery of the 1990s was associated with sharp reductions in crime rates. Tight labor markets, with better earnings opportunities for low-skilled males in the legitimate economy relative to criminal opportunities, may play an important role in reducing crime and decreasing permanent reductions in human capital from increases in the share of potential workers stigmatized by criminal records. Much evidence shows a strong positive response of property crime rates to local unemployment rates and wages in low-wage sectors, and this response is observed across U.S. metropolitan and regional labor markets.[11]

Finally, wage inequality increased dramatically for U.S. males, and the real wages of low-skilled males declined substantially, from the early 1970s to the early 1990s. The authors emphasize the role of declining rewards to work in the 1970s and 1980s in generating labor supply responses from less-skilled males, who chose to work less and, in many cases, to drop out of the labor force. The authors point to a decline in the relative demand for less-skilled workers as the key factor behind these real wage declines. This pattern contrasts with stable real wages and a rather stable employment rate for high-wage males (those above the 60th wage percentile) over the same period. The authors also argue that an increase in real wages in the 1990s played a key role in the shift toward rising employment rates for less-skilled males. But I would like to point out some puzzles for the authors' simple framework in which stable labor supply curves combined with labor demand shifts drive the observed employment changes by skill group. In particular, as the authors point out, their framework does a reasonable job for long periods, but it has problems matching the actual dynamic patterns of employment and wage changes in the data.

First, a large part of the secular rise in nonemployment for low-skilled males occurred before the period of declining real wages. The nonemployment rate of low-skilled males increased substantially from its 1967–69 peak to its 1972–73 peak, despite substantial real wage growth in this period, and it does not seem to have accelerated with declining real wages over next two decades. In fact, the nonemployment of less-skilled

11. See, for example, Gould, Mustard, and Weinberg (2002).

men did not really rise from the early 1980s to the early 1990s, despite continuing noticeable declines in real wages. These figures suggest that cyclical factors affected employment changes beyond the response to real wage changes along a stable labor supply curve. And a permanent adverse labor supply shift (possibly from the expansion of disability programs, and possibly from changes in illegal labor market opportunities) appears to have occurred from 1967–69 to 1972–73.

Second, a key remaining question for the authors' approach is what caused an apparent slowdown in adverse demand shifts against less-skilled males in the 1990s relative to the previous two decades. Indicators of skill-biased technological change (such as computer investment) continued rising in the 1990s, and trade with less developed countries grew more rapidly in the 1990s than in the 1980s. It may be that tight labor markets and rapid productivity growth themselves improve the demand for less-skilled males, because optimistic firms are willing to take chances on workers whom they would not hire in a weaker labor market. The authors' framework needs a clearer and testable story about what drives the relative demand shifts that are doing the work in their story of changes in U.S. male nonemployment rates.

Robert Shimer: This is a provocative paper. The conventional wisdom is that, in the late 1990s, the U.S. unemployment rate fell to levels not seen in three decades because demand for labor was so strong. This paper points out that the nonemployment rate, the fraction of individuals who are either unemployed or out of the labor force in an average week, has behaved very differently in recent years, at least for prime-aged men. For example, between the business-cycle peak at the end of the 1980s and that at the end of the 1990s, the unemployment rate for prime-aged men fell from 4.3 percent to 3.0 percent. In contrast, as the authors' figure 3 shows, the nonemployment rate for this group remained constant at 11.0 percent across those cyclical peaks. Moreover, as their figure 8 reveals, these numbers mask an important increase in nonemployment durations, from eleven months on average at the end of the 1980s to about fifteen months a decade later.

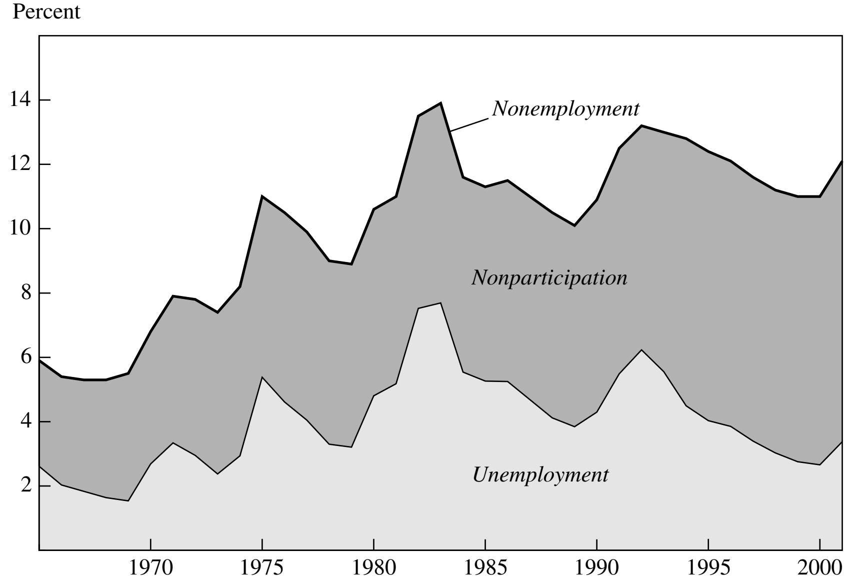
That nonemployment among prime-aged men is high should be uncontroversial, even if its causes and consequences are not. Indeed, there is no need to use microdata to uncover this fact. According to the standard Bureau of Labor Statistics time series, available from the agency's homepage and displayed in my figure 1, unemployment among twenty-five to fifty-four-year-old men reached a cyclical low in 1989 at 3.8 percent, but in the subsequent cycle it fell further, to 2.7 percent in 2000.[1] Over the same eleven-year period, however, the fraction who were nonemployed rose from 10.1 percent to 11.0 percent. By comparison, during the deepest post–World War II recession, in 1983, the nonemployment rate for this group peaked at an only modestly higher 13.9 percent. By this measure, then, the labor market for prime-aged men at the end of the century remained slack.

The increase in nonemployment duration, on the other hand, is hard to measure, and its occurrence should be more controversial. This does not mean that nonemployment duration is unimportant. Assuming labor insurance markets are incomplete, even a pure utilitarian cares about both the incidence and the duration of nonemployment spells. Workers can self-insure against nonemployment by building up a buffer stock of savings; however, that buffer stock will generally be too small to allow the individual to maintain his or her accustomed level of consumption during a very long spell of nonemployment. Put differently, the average worker's utility will be lower if 8.3 percent of the population is out of work for the entire year than if everyone is nonemployed for one month. Of course, many other social welfare functions would also imply that long nonemployment durations are intrinsically undesirable, for example because they yield a more unequal distribution of income. In addition, it is plausible that a very long spell of nonemployment makes it increasingly difficult for a worker to reenter employment, because that worker's basic labor market skills begin to atrophy. Such forces lead to hysteresis in nonemployment.

What, then, is the evidence that nonemployment duration increased from eleven to fifteen months during the 1990s? Unfortunately, the Current Population Survey does not ask individuals how long it has been since they were last employed. Instead, the authors infer the duration of

1. These numbers are based on Bureau of Labor Statistics time series LFU21003301 (the unemployment rate) and LFU1603301 (the employment-to-population ratio, or 100 percent minus the nonemployment rate) and refer to a slightly older group of men than do the authors. This probably explains why I find a slightly lower unemployment rate. To calculate the unemployment rate for men with one to thirty years of potential labor market experience, as the authors do, one must look at the microdata.

Figure 1. Unemployment, Nonparticipation, and Nonemployment among Men Aged 25 to 54, Using Monthly CPS Data, 1965–2001



Source: Bureau of Labor Statistics.

nonemployment indirectly from another question, which asks the number of weeks worked, including paid vacation and sick leave, during the previous year. Let E denote the fraction of the relevant population who report being employed during a typical week of the year and N the fraction who report being nonemployed. If we assume for simplicity that these fractions are constant during the year, then the fraction of the population who find a job, $N\lambda_{ne}$, must equal the fraction of the population who lose a job, $E\lambda_{en}$, where λ_{ne} and λ_{en} are the fraction of nonemployed workers who become employed and the fraction of employed workers who become nonemployed, respectively. Since $E + N = 1$, this gives us one important equation, $N\lambda_{ne} = (1 - N)\lambda_{en}$. A second equation comes from the definition of the fraction of the population who experience a spell of nonemployment, S. This is assumed to be equal to the fraction of people who begin the year nonemployed, N, plus the fraction who become nonemployed during the year, $E\lambda_{en}$. Thus $S = N + (1 - N)\lambda_{en}$. Combining these equations gives

$$(1) \qquad \frac{1}{\lambda_{ne}} = \frac{N}{S - N}.$$

The left-hand side of equation 1 is the inverse of the likelihood of a nonemployed worker finding employment, or, equivalently, the average duration of a nonemployment spell in years. This is a function of the fraction of nonemployed workers N and the fraction of workers who experience a spell of nonemployment S. Both of these quantities are then measured using the retrospective question in the March CPS. N is the fraction of the year that the average worker reports that he did not work, and S is the fraction of workers who report working less than fifty-two weeks during the year. Similarly, the same two equations imply that the entry rate into nonemployment, that is, the incidence of nonemployment, satisfies

$$(2) \qquad \lambda_{en} = \frac{S - N}{1 - N}.$$

Both the measured increase in nonemployment duration and the decrease in nonemployment incidence reflect a decline in $S - N$, the difference between the fraction of workers who experience a spell of nonemployment and the fraction of weeks spent nonemployed. This, in turn, is primarily due to the enormous increase in the incidence of full-year nonemployment (documented in the authors' figure 4).

The authors' finding, using this methodology, that nonemployment duration nearly doubled during the 1980s and 1990s is so striking that it seems worth attempting to verify it using an alternative methodology. This can be done by looking at data on gross worker flows between employment and nonemployment, constructed from matched monthly CPS files. The CPS sample is a rotating panel, although the public-use microdata do not contain unique individual identifiers. Still, there are well-established procedures for matching individual records across months—and these procedures have well-known shortcomings, which I will discuss shortly. Following the authors' approach, I focus my analysis on men with one to thirty years of potential labor market experience, which I define as age minus years of education minus six.[2] In each month

2. The authors define potential experience as age minus years of education minus *seven* because the questions in the March CPS refer to employment during the previous year. The

from 1976 to 2001, I calculate the fraction of employed men who leave employment (the incidence of nonemployment) and the fraction of nonemployed men who become employed (the inverse of the duration of nonemployment). I then aggregate these to get annual average data, which should be comparable to the numbers in the paper.

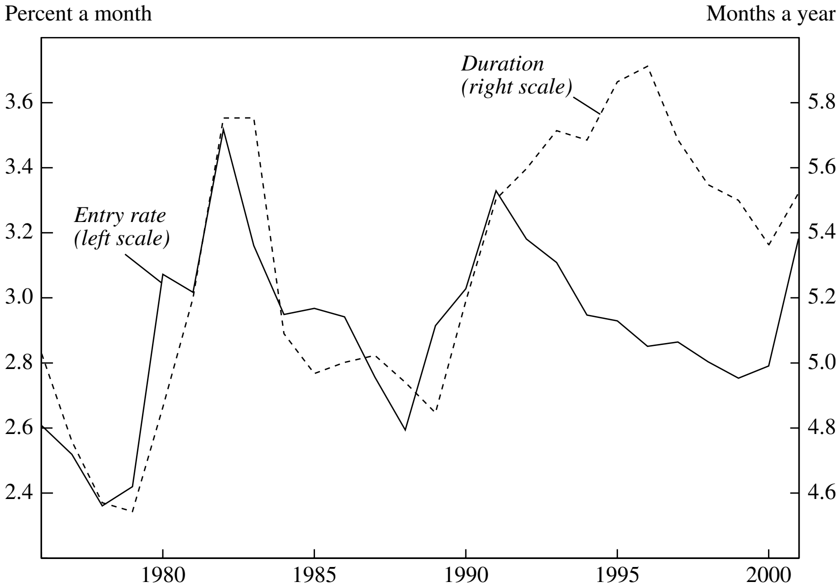
Figure 2 shows the results from this exercise. I find that nonemployment durations increased from 4.5 months in 1978 to 4.8 months in 1988 and then to 5.4 months in 1999, a cumulative increase of about 20 percent. This is much smaller than the increase in nonemployment duration that the authors report. The flip side of this is the incidence of nonemployment, which they find decreases from 1.4 percent a month to 0.8 percent during the 1980s and 1990s, whereas I find that the incidence actually increased slightly, from 2.4 percent to 2.8 percent. Thus there is a difference in both the level and the trend of nonemployment duration and incidence between the gross flows data and the retrospective data from the March CPS.

Measurement and classification error explains part of the reason why I find such a short duration and a high incidence of nonemployment. An individual who is mistakenly recorded as employed in one month will generate two spurious transitions, first from nonemployment to employment and then from employment back to nonemployment, and similarly for an individual who is mistakenly recorded as nonemployed. This shortcoming of the gross flows data has been analyzed at great length, notably by John Abowd and Arnold Zellner and by James Poterba and Lawrence Summers.[3] Abowd and Zellner used data from households that were interviewed twice to conclude that as many as 40 percent of reported transitions are spurious. This would, roughly speaking, reduce the incidence of nonemployment by 40 percent and raise the duration of nonemployment by a similar amount. Unfortunately, there is no way to tell whether these reporting errors have increased or decreased over time, because, to my knowledge, no one has updated Abowd and Zellner's study. Although it is conceivable that the redesign of the CPS instrument in 1994 reduced reporting errors—which would mask an increase in nonemployment duration—work that I have done with Katharine Abraham indicates that this is

basic monthly CPS questions refer to contemporaneous employment and nonemployment experiences, and so potential experience is one year greater.

3. Abowd and Zellner (1985); Poterba and Summers (1986).

Figure 2. Entry Rates and Durations of Nonemployment among Men Aged 25 to 54, Using Monthly CPS Data, 1976–2001



Source: Author's calculations based on CPS data.

not the case.[4] One can see this in my figure 2, where nothing special occurs in 1994.

The importance of measurement error can be addressed directly by matching CPS files across three consecutive months. An individual who is measured as employed in January and February is likely to have been in fact employed in both those months. This means that, by looking only at workers who are employed in both months, we avoid contaminating the employed population with misclassified nonemployed workers. The fraction of these workers who become nonemployed in March therefore reflects both genuine incidents of nonemployment and misclassification of workers who remain employed, but it does not reflect misclassification of nonemployed workers. Likewise, by looking at the probability of entering employment of workers who are nonemployed for two consecutive months, the sample does not include employed workers who were mis-

4. Abraham and Shimer (2001).

classified. I find that doing this nearly doubles the measured duration of nonemployment and halves the incidence. However, the trends—or, more precisely, the lack of trends—remain the same. There is no evidence in gross worker flow data that the nonemployment duration of prime-aged men has sharply increased during the last decade.

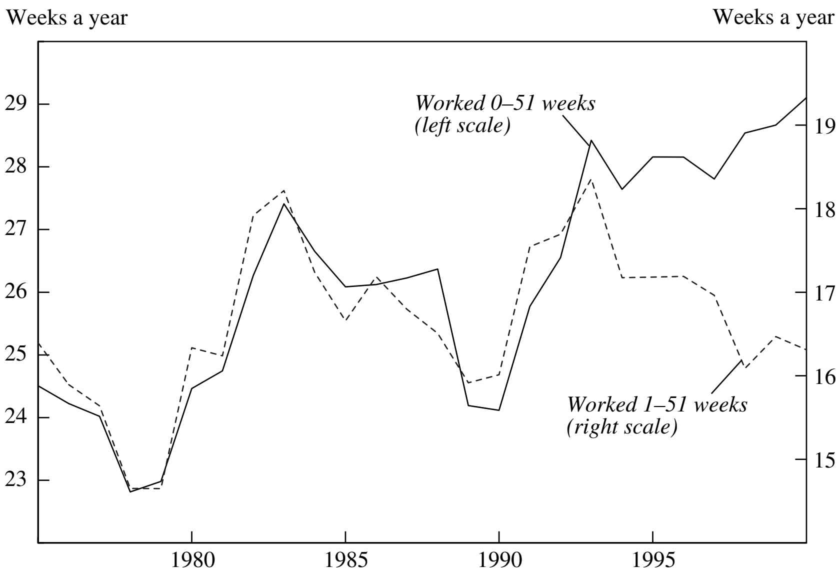
So the question is, Which numbers are correct? Has nonemployment duration sharply increased and nonemployment incidence fallen during the last decade, as the authors argue? Perhaps not surprisingly, I will make the case that the gross flows data are more reliable than the March CPS data.

One shortcoming of the March CPS data is that the measured number of spells of nonemployment is capped at one per person. If we lived in a world in which people were either employed or nonemployed but were otherwise identical, this would not be a big problem. In such a world, all employed individuals are equally likely to become nonemployed in a given month, and all nonemployed individuals are equally likely to become employed. In other words, there is a two-state Markov process for individual employment status. If this were the case, it would imply that very few people find a job and then lose it within the same year, because the relevant transition rates are extremely small. But, of course, we do not live in such a world. Newly employed workers are much more likely to become nonemployed than are workers with long tenure. This means that many workers are likely to experience multiple spells of nonemployment during a year, and this biases downward the authors' measure of the number of nonemployment spells S. Nevertheless, there is no evidence that this bias has changed much over time. For example, there is no secular shift in the likelihood of a newly employed worker losing his job, compared with that of a worker who has at least two months' tenure. Thus I think the explanation for the different results must lie elsewhere.

Another possibility is that there are time-varying biases in the way that people answer retrospective questions in the March CPS. Recall that the driving force behind the authors' finding was the increase in the fraction of nonemployment accounted for by full-year nonemployment. Perhaps in recent years people have been more likely to give extreme responses, that is, to say that they worked either zero or fifty-two weeks in the previous year, rather than recollect a short intervening spell of employment or nonemployment. This might be the case if respondents have become increasingly careless in the way they answer questions over time.

To see whether such an explanation is plausible, it is useful to look at the mean weeks of nonemployment among individuals who report between one and fifty-one weeks worked. If there has indeed been an increase in nonemployment duration, mean nonemployment duration should have increased for the part-year nonemployed, not only the fraction of workers experiencing full-year nonemployment. Figure 3 below compares the mean weeks of nonemployment for all prime-aged men who report less than fifty-two weeks worked with the mean weeks of nonemployment for prime-aged men who report one to fifty-one weeks worked. The two time series track each other from 1975 to 1993 before diverging. As a result, from 1989 to 2000 nonemployment in the full sample of prime-aged men increased by 4.9 weeks a year, while nonemployment in the sample excluding the extreme response increased by only 0.4 week a year. In other words, all of the increase in nonemployment duration is accounted for by an increase in full-year nonemployment. Although there are other possible explanations for this finding, it seems plausible that this

Figure 3. Average Number of Weeks Nonemployed among Men Aged 25 to 54, Using Monthly CPS Data, 1975–2000



Source: Author's calculations based on CPS data.

reflects at least in part a change in how individuals report their employment status retrospectively.

To reiterate, this is a provocative paper. Despite the strong labor market in the United States in the 1990s, the nonemployment rate for prime-aged men did not fall. Moreover, the paper argues that the constant rate of nonemployment masks a sharp increase in the duration of nonemployment and a sharp decrease in its incidence. If this is really the case, it is likely to have significant and adverse welfare implications. However, other data sources do not indicate that there was much of an increase in nonemployment duration or a decrease in nonemployment incidence for this group of men. Instead, the measured increase in nonemployment duration may be due to a change in the way people answer retrospective questions in the CPS. Nevertheless, the finding that the labor market for prime-aged men was no stronger in 2000 than it was a decade earlier is surprising indeed.

General discussion: Several participants discussed what to make of the rising trend toward nonemployment among men. Katharine Abraham reasoned that the low unemployment rates of the last half of the 1990s made it unlikely that any man who wanted a job could not find one. The low real wages of less-skilled workers and the easier access to disability insurance during the past decade both could be factors behind men's decision not to work. She warned, however, against the easy interpretation that the rising nonemployment of men is necessarily a bad thing, noting that few would jump to that interpretation if women were the group under consideration, and she commented that it would be of interest to know more about how nonemployed men were spending their time.

The impact on nonemployment of more comprehensive disability insurance generated further discussion. Robert Gordon argued that the growth in disability insurance may have contributed to the decline in the NAIRU during the 1990s just as increased incarceration of the unemployed had done according to Lawrence Katz and Alan Krueger's 1999 study in the *Brookings Papers*.[1] Robert Hall added that the open-ended nature of the current disability insurance program has undesirable side effects: an individual's duration on disability is correlated with a decreased likelihood of eventually reentering the work force and with a

1. Katz and Krueger (1999).

deterioration of work skills. William Nordhaus compared the coverage in the disability insurance program to the poverty line, suggesting that both seemed to be ratcheting up over time. For example, carpal tunnel syndrome would probably not have been considered a disability thirty years ago, but today it is. Erik Brynjolfsson interpreted the rise in disability rolls as partly the result of a policy decision, in which individuals who were previously considered fit for work are now categorized as disabled. He took this as a social value judgment that these people should no longer have to work, and he suggested that this was not necessarily a bad thing.

William Dickens observed an important difference between the findings in the present paper and those in the authors' paper of a decade ago: at the time of the earlier paper, unemployment and nonparticipation in the labor force were moving in the same direction, whereas now they seem to be moving in opposite directions. This suggests that the supply-demand model that is useful for explaining nonemployment is not good for explaining unemployment. Olivier Blanchard noted that a striking feature of the data was the decrease in the flow out of employment. He argued that this phenomenon deserved more interpretation than the paper had given it. Nordhaus suggested that a change in survey design in the early 1990s may have biased some of the entry and duration results.

References

Abowd, John M., and Arnold Zellner. 1985. "Estimating Gross Labor-Force Flows." *Journal of Business and Economic Statistics* 3(3): 254–83.

Abraham, Katharine G., and Robert Shimer. 2001. "Changes in Unemployment Duration and Labor-Force Attachment." In *The Roaring Nineties: Can Full Employment Be Sustained?* edited by Alan B. Krueger and Robert M. Solow. New York: Russell Sage Foundation.

Autor, David H. 2000a. "Outsourcing at Will: Unjust Dismissal Doctrine and the Growth of Temporary Help Employment." Working Paper 7557. Cambridge, Mass.: National Bureau of Economic Research (February).

———. 2000b. "Wiring the Labor Market." Working Paper 7959. Cambridge, Mass.: National Bureau of Economic Research (October).

Autor, David H., and Mark G. Duggan. 2001. "The Rise in Disability Recipiency and the Decline in Unemployment." Working Paper 8336. Cambridge, Mass.: National Bureau of Economic Research (June).

———. Forthcoming. "The Rise in Disability Rolls and the Decline in Unemployment." *Quarterly Journal of Economics*.

Bertola, Giuseppe, and Andrea Inchino. 1995. "Wage Inequality and Unemployment: United States vs. Europe." In *NBER Macroeconomics Annual*, edited by Ben S. Bernanke and Julio J. Rotemberg. MIT Press.

Bertola, Giuseppe, Francine D. Blau, and Lawrence M. Kahn. 2001. "Comparative Analysis of Labor Market Outcomes: Lessons for the US from International Long-Run Evidence." Working Paper 8526. Cambridge, Mass.: National Bureau of Economic Research (October).

Blanchard, Olivier, and Justin Wolfers. 2000. "The Role of Shocks and Institutions in the Rise of European Unemployment: The Aggregate Evidence." *Economic Journal* 110(462): C1–C33.

Bound, John. 1989. "The Health and Earnings of Rejected Disability Insurance Applicants." *American Economic Review* 79(3): 482–503.

Bound, John, and Timothy Waidmann. 1992. "Disability Transfers, Self-Reported Health, and the Labor Force Attachment of Older Men: Evidence from the Historical Record." *Quarterly Journal of Economics* 107(4): 1393–1419.

———. 2000. "Accounting for Recent Declines in Employment Rates among the Working-Aged Disabled." Working Paper 7975. Cambridge, Mass.: National Bureau of Economic Research (October).

Bound, John, Richard Burkhauser, and Austin Nichols. 2001. "Tracking the Household Income of SSDI and SSI Applicants." Michigan Retirement Research Center Working Paper 99-03. University of Michigan (May).

Cutler, David M., and Elizabeth Richardson. 1997. "Measuring the Health of the U.S. Population." *BPEA, Microeconomics*, 217–72.

Gordon, Robert J. 1998. "Foundations of the Goldilocks Economy: Supply Shocks and the Time-Varying NAIRU." *BPEA, 2:1998*, 297–333.

Gould, Eric D., David B. Mustard, and Bruce A. Weinberg. 2002. "Crime Rates and Local Labor Market Opportunities in the United States: 1979–1997." *Review of Economics and Statistics* 84(1): 45–61.

Juhn, Chinhui, Kevin M. Murphy, and Robert H. Topel. 1991. "Why Has the Natural Rate of Unemployment Increased over Time?" *BPEA, 1:1991,* 75–126.

Katz, Lawrence F., and Alan B. Krueger. 1999. "The High-Pressure U.S. Labor Market of the 1990s." *BPEA, 1:1999*, 1–65.

Kling, Jeffrey R. 2002. "The Effect of Prison Sentence Length on the Subsequent Employment and Earnings of Criminal Defendants." Woodrow Wilson School Economics Discussion Paper 208. Princeton University (April).

Kuhn, Peter, and Mikal Skuterud. 2000. "Job Search Methods: Internet versus Traditional." *Monthly Labor Review* 123(10): 3–11.

Ljungqvist, Lars, and Thomas J. Sargent. 1998. "The European Unemployment Dilemma." *Journal of Political Economy* 106(3): 514–50.

Organization for Economic Cooperation and Development. 2001. *OECD Employment Outlook*. Paris.

Parsons, Donald O. 1980. "The Decline in Male Labor Force Participation." *Journal of Political Economy* 88(1): 117–34.

Phelps, Edmund S. 1974. "Economic Policy and Unemployment in the 1960's." *Public Interest* 34(Winter): 30–46.

Poterba, James M., and Lawrence H. Summers. 1986. "Reporting Errors and Labor Market Dynamics." *Econometrica* 54(6): 1319–38.

Shimer, Robert. 1998. "Why Is the U.S. Unemployment Rate So Much Lower?" In *NBER Macroeconomics Annual*, edited by Ben S. Bernanke and Julio J. Rotemberg. MIT Press.

Staiger, Douglas, James H. Stock, and Mark W. Watson. 2001. "Prices, Wages and the U.S. NAIRU in the 1990s." Working Paper 8320. Cambridge, Mass.: National Bureau of Economic Research (June).

Stiglitz, Joseph. 1997. "Reflections on the Natural Rate Hypothesis." *Journal of Economic Perspectives* 11(1): 3–10.