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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

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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 Inghino (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 i th 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 than 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.

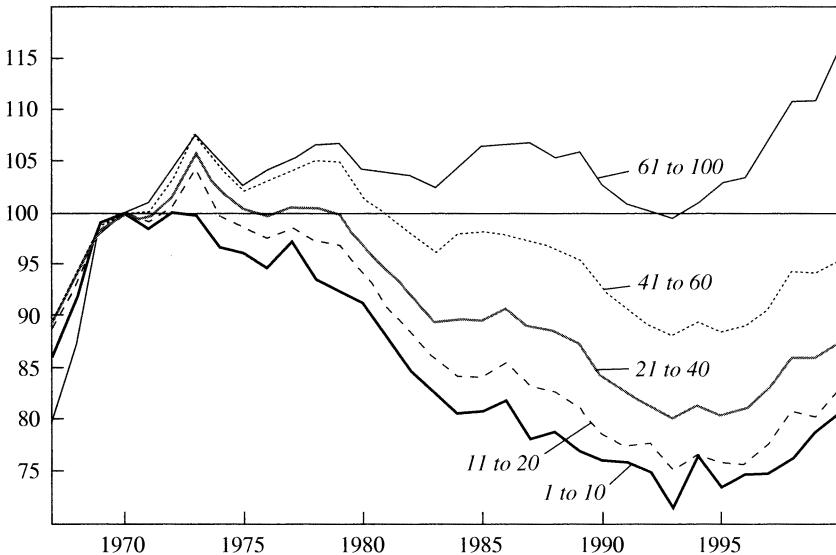
imputed wages are assigned probabilities of being in each of the categories.

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

Index, 1970 = 100



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

Period	Wage percentile group				
	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.

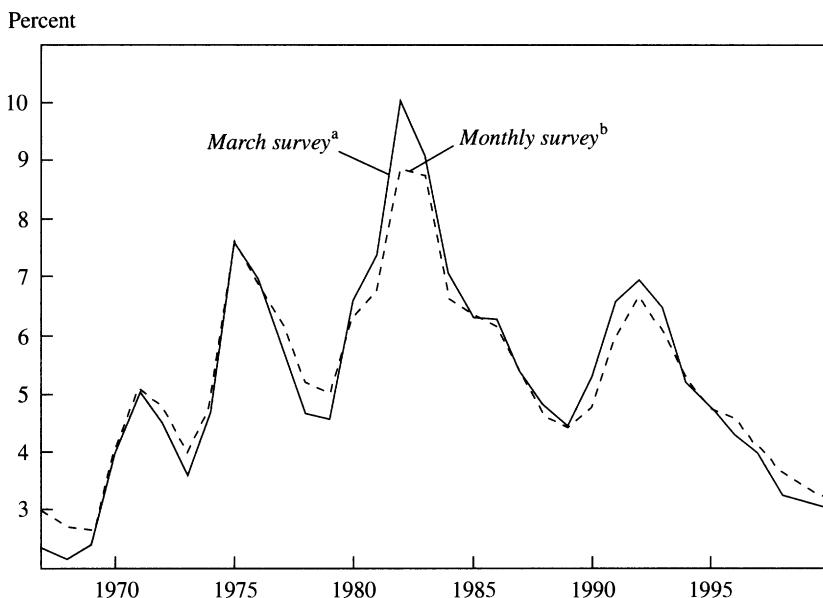
Even so, average wages of the least-skilled men were roughly 20 percent lower in 2000 than in the late 1960s (table 1), whereas those of men in the top 40 percent of the distribution increased by a roughly equivalent amount. As we showed in our 1991 Brookings Paper, wage declines were most prominent 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.

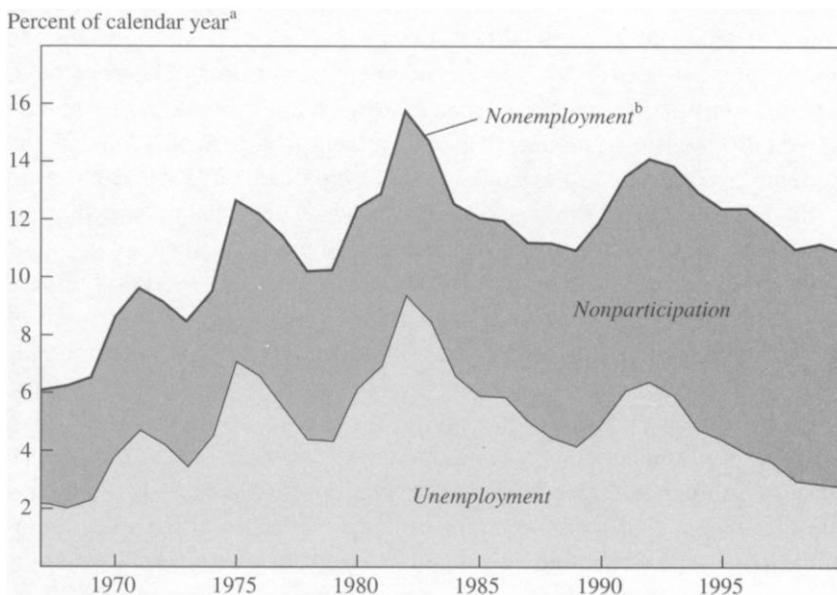
Figure 2 also shows annual unemployment rates for our sample of prime-aged males (calculated from the March CPS data as weeks unemployed divided by weeks in the labor force). Although the two series should not be identical because of differences in the underlying populations (our sample consists only 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

Period	Phase of business cycle	Labor force status (percent of calendar year ^a)		Nonemployed ^b	Change in unemployment (percentage points)	Change in nonparticipation (percentage points)
		Unemployed	Nonparticipating			
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

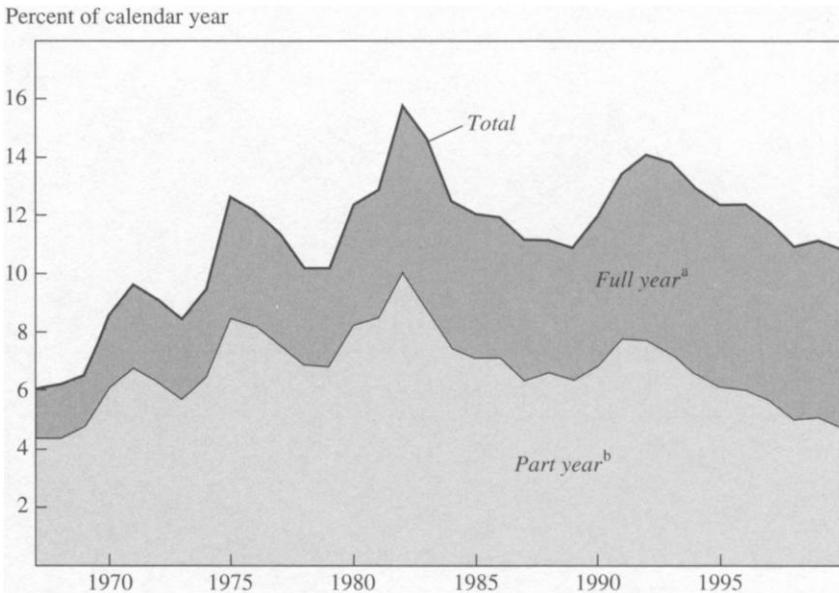
<i>Period</i>	<i>Phase of business cycle</i>	<i>Part-year^b</i>	<i>Full-year^c</i>	<i>Total</i>
1967–69	Peak	4.5	1.8	6.3
1971–72	Trough	6.5	2.9	9.4
1972–73	Peak	6.0	2.8	8.8
1975–76	Trough	8.4	4.1	12.4
1978–79	Peak	6.8	3.8	10.2
1982–83	Trough	9.4	5.8	15.2
1988–89	Peak	6.5	4.6	11.0
1991–92	Trough	7.7	6.0	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.

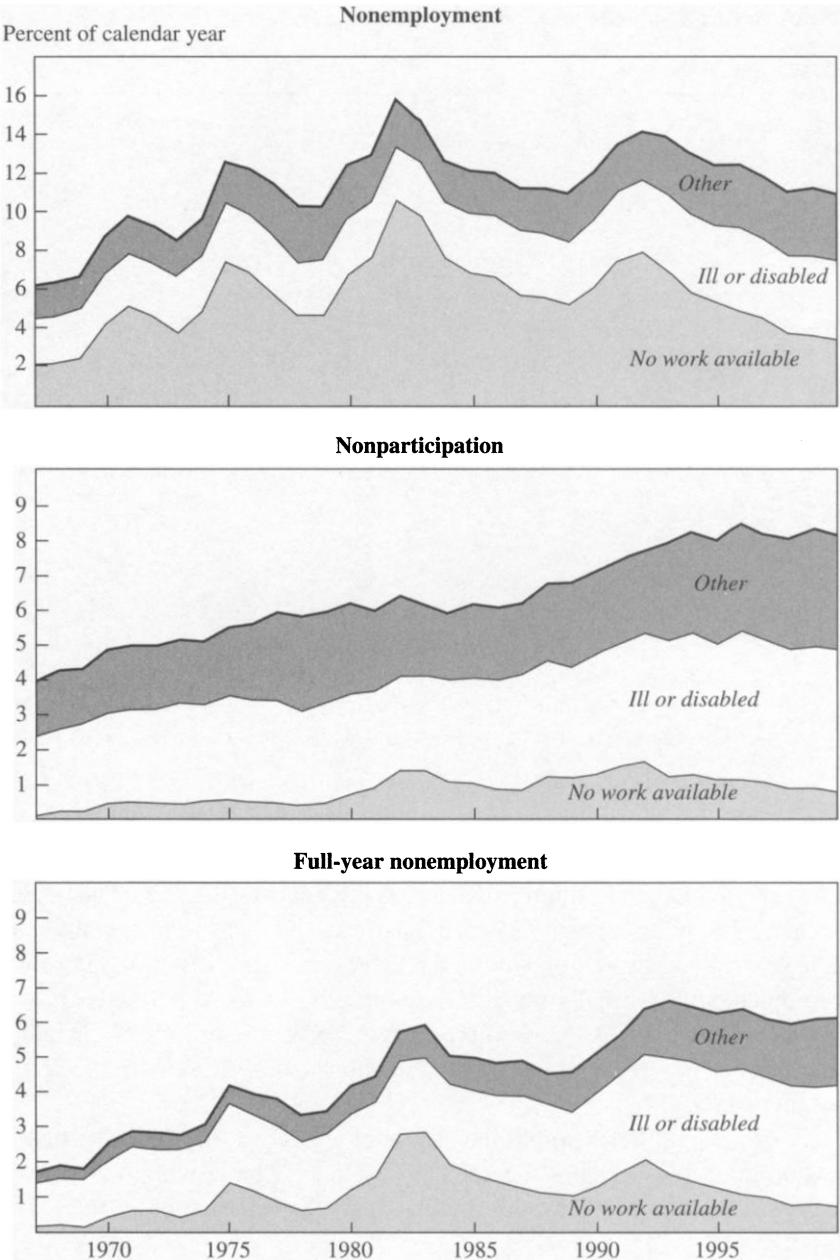
b. Fraction of males nonemployed for the entire year multiplied by the percent of the calendar year spent nonemployed (100 percent).

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*



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

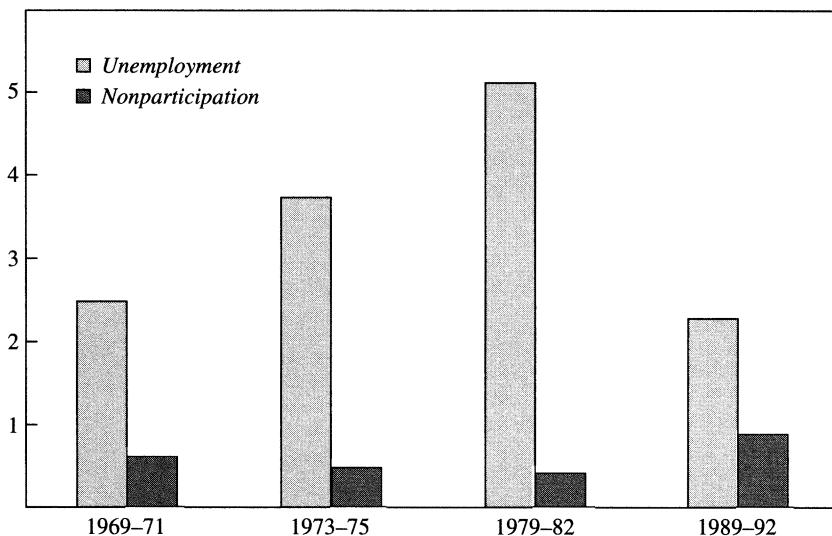
18. Autor and Duggan (forthcoming); Bound and Waidmann (2000).

19. Bound and Waidmann (2000).

20. See Murphy and Topel (2001), for example.

Figure 6. Changes in Unemployment and Nonparticipation Entering Recessions

Percentage points



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).

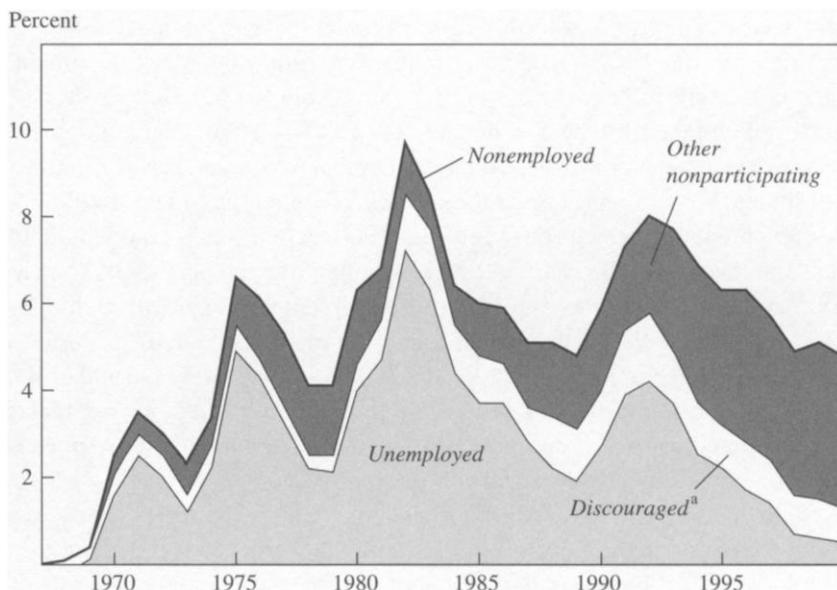
Table 4. Changes in Nonemployment, by Reported Reason for Nonemployment, 1967–2000^a

Percentage points

<i>Measure</i>	<i>Reason for nonemployment</i>			<i>Total</i>
	<i>No work available</i>	<i>Illness or disability</i>	<i>Other</i>	
<i>Change 1967–69 to 1999–2000</i>				
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
<i>Change 1988–89 to 1999–2000</i>				
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

Period	Number of weeks a year					
	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

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

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

Period	Number of weeks a year				
	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.

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

where

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

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

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

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 $\bar{U}(\tau)$ is the average unemployment rate over the year. With these definitions, $\bar{\lambda}_u(\tau)$ and $\bar{\lambda}'_u(\tau)$ 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

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

since spells are generated either by starting the year unemployed, $U_0(\tau)$, or by becoming unemployed during the year, $[1 - \bar{U}(\tau)]\bar{\lambda}_u(\tau)$. 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

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

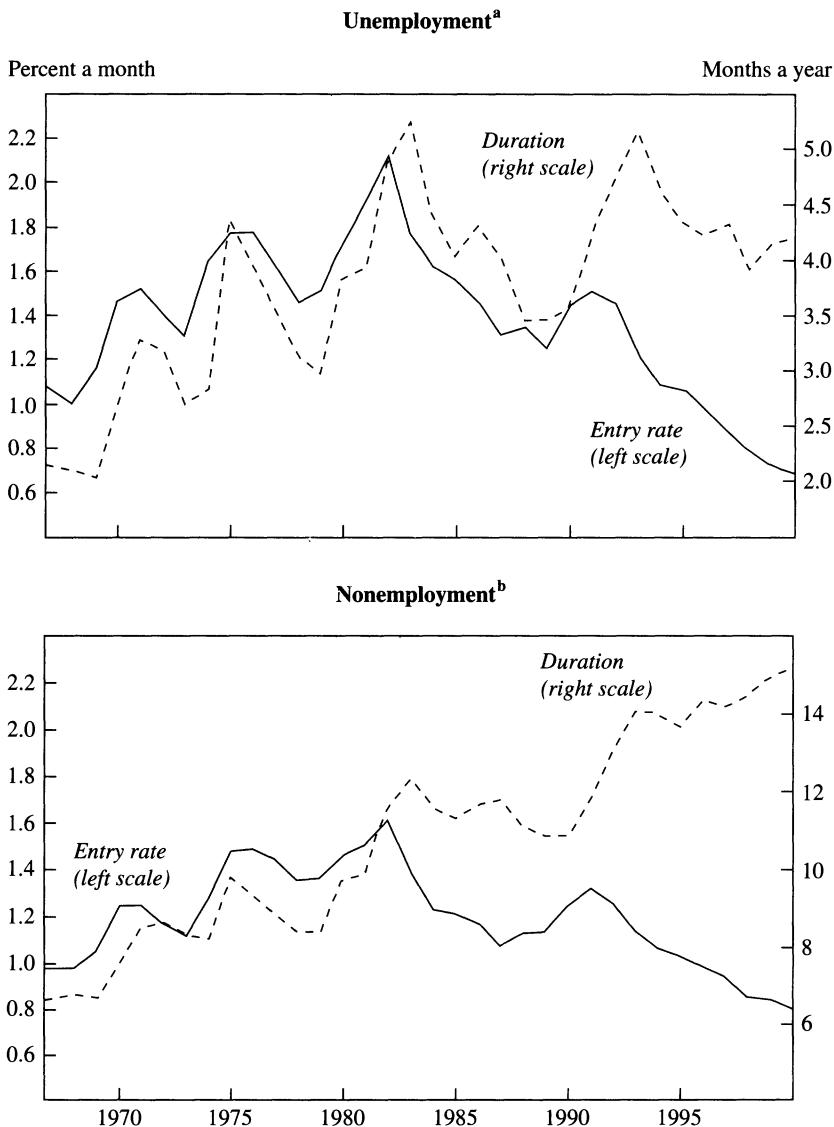
and

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

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

Period	Phase of business cycle	Unemployment ^a		Nonemployment ^b	
		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.

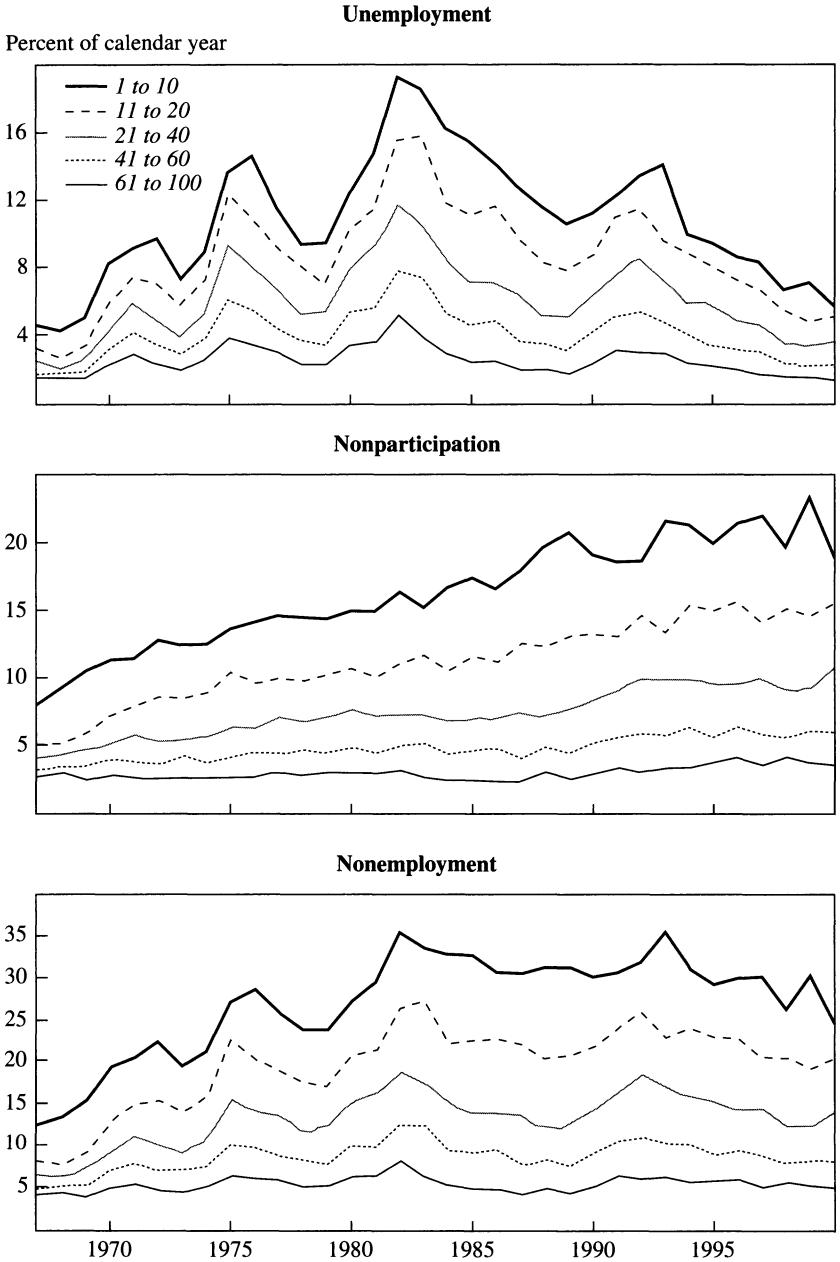
Table 7 and figure 8 paint a clear picture. Although the rates of entry into unemployment and nonemployment have returned to or even fallen below levels experienced during the late 1960s, the durations of jobless spells are more than twice as long at the end of the period. Indeed, jobless spells were 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

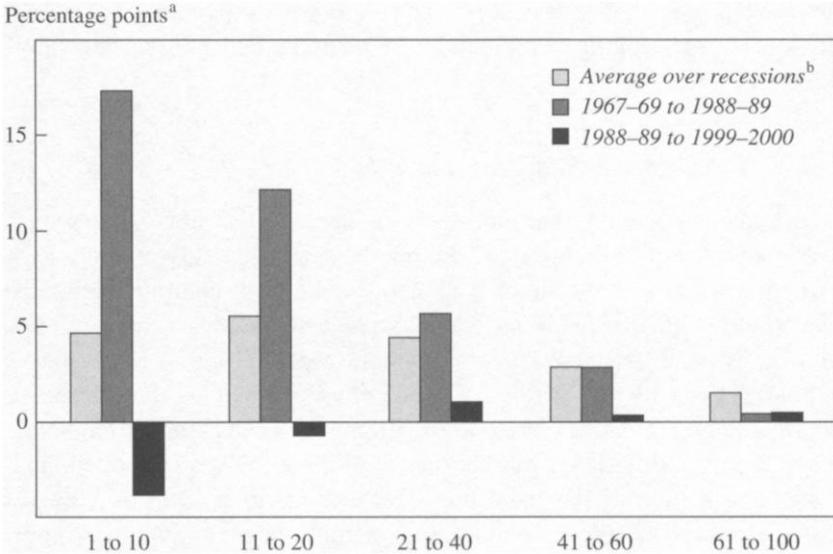
Period	Wage percentile group				
	1 to 10	11 to 20	21 to 40	41 to 60	61 to 100
<i>Unemployment</i>					
1967–69 to 1988–89	6.4	4.9	2.6	1.5	0.3
1988–89 to 1999–2000	–4.6	–3.1	–1.6	–1.0	–0.5
1967–69 to 1999–2000	1.8	1.8	1.1	0.5	–0.1
<i>Nonparticipation</i>					
1967–69 to 1988–89	10.9	7.2	2.9	1.3	0.0
1988–89 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
<i>Nonemployment</i>					
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
<i>Full-year nonemployment</i>					
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
1996 dollars a year, except where noted otherwise

<i>Item</i>	<i>1967–69</i>	<i>1972–73</i>	<i>1978–79</i>	<i>1988–89</i>	<i>1999–2000</i>
<i>1st to 10th percentile</i>					
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
<i>21st to 40th percentile</i>					
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
<i>61st to 100th percentile</i>					
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
<i>All households</i>					
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.

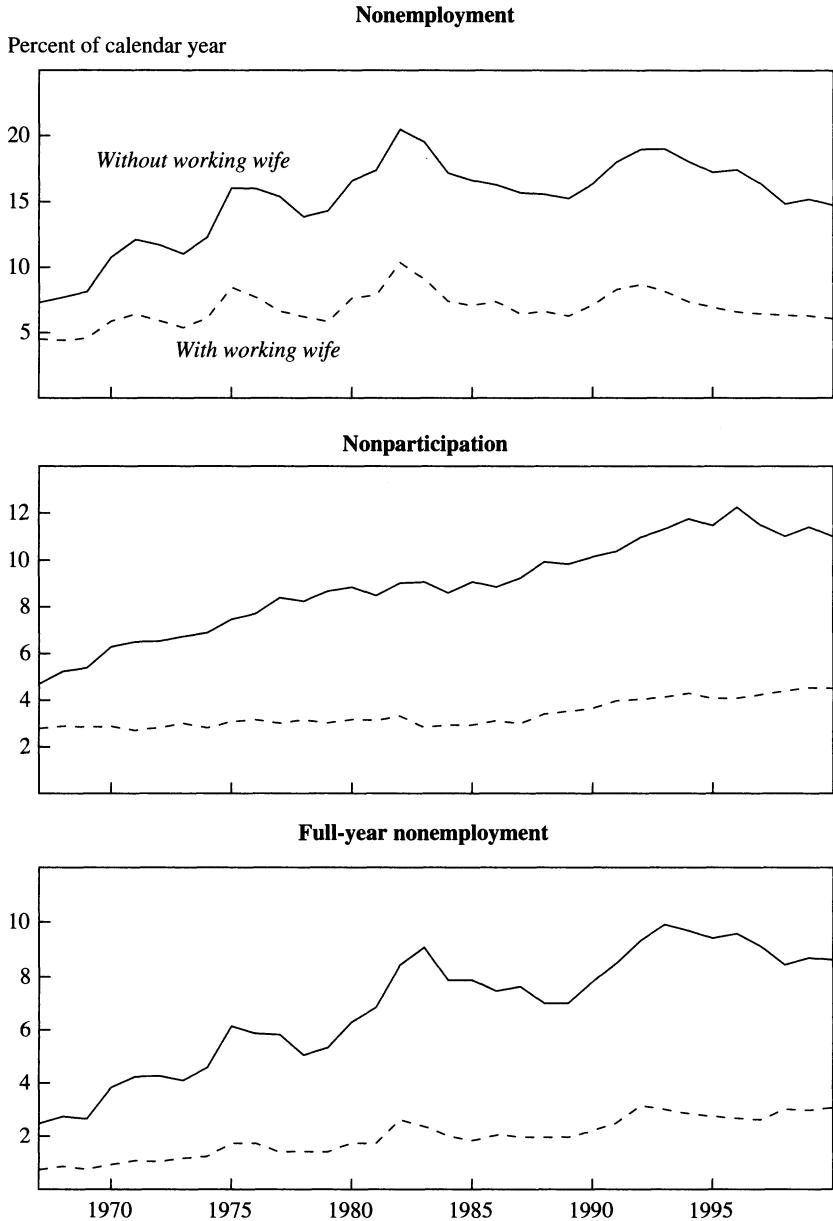
a. Earnings and household income are deflated by the price index for personal consumption expenditure. Male's and wife's earnings include income from wages, salary, and self-employment; household income includes all sources of earned and non-earned income of all household members.

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
<i>1972–73 to 1988–89</i>				
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
<i>1988–89 to 1999–2000</i>				
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.