# The decline of job security in the 1990s: Displacement, anxiety, and their effect on wage growth

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## Introduction and summary

The news media frequently suggest that American workers have suffered a significant decline in job security during the 1990s. Of course, planned and actual employment reductions at major corporations such as AT&T, IBM, and General Motors have been important stories. But articles such as those in the 1996 *New York Times* book *The Downsizing of America* go beyond reporting individual cases of layoffs to suggest that there has been a fundamental change in the employment relationship. According to such articles, workers in general have suffered a loss of job security, and long-term employment relationships are a thing of the past. Moreover, such articles claim that this decreased job security has left workers feeling more anxious about their futures.

The perception of declining job security is shared by many policymakers and other analysts, who believe worker anxiety to be a major reason wage inflation in the 1990s has remained modest in the face of historically low levels of unemployment. Perhaps most famously, Federal Reserve Board Chairman Alan Greenspan testified to Congress in February 1997 that "atypical restraint on compensation increases has been evident for a few years now and appears to be mainly the consequence of greater worker insecurity."[1] Former U.S. Labor Secretary Robert Reich recently made much the same point when he wrote, "Wages are stuck because people are afraid to ask for a raise. They are afraid they may lose their job."[2]

Labor economists, however, have often been skeptical of claims of widespread declines in job stability and security. They note that media accounts are long on anecdotes and short on evidence based on nationally representative survey data. Moreover, the most carefully executed studies using scientifically designed survey data collected through the early 1990s often reached conclusions quite at odds with media reports. For instance, Diebold, Neumark, and Polsky (1997) concluded that "aggregate job retention rates have remained stable." Similarly, Farber (1998) found that "there has been no systematic change in the overall distribution of job duration over the last two decades."

More recently, however, researchers have begun to analyze survey data from the mid-1990s, and conclusions somewhat more in line with media reports are emerging. For instance, Neumark, Polsky, and Hansen (1997) reported that "there is some evidence that job stability declined modestly in the first half of the 1990s. Moreover, the relatively small aggregate changes mask rather sharp declines in stability for workers with more than a few years of tenure." Similarly, Farber (1997b) concluded that "after controlling for demographic characteristics, the fraction of workers reporting more than ten and more than 20 years of tenure fell substantially after 1993 to its lowest level since 1979." Thus job stability—the tendency of workers and employers to form long-term bonds—seems to be declining somewhat. Moreover, evidence of significant change is especially apparent in more direct measures of worker security, such as Farber's (1997a) tabulations of the number of workers reporting involuntary job loss. The extent of changes in job tenure, turnover, and displacement reported in these more recent studies is much too modest to justify

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the most sensationalistic news reports. Nevertheless, some decline in job security, especially for workers who have attained significant seniority, now seems reasonably clear.

In this article, we review some of the findings of this research on job stability and job security. We then present some new tabulations of rates of job loss for high seniority workers based on the Bureau of Labor Statistics' (BLS) *Displaced Worker Surveys* (DWS). Next, we look directly at workers' own perceptions of their job security using data from the National Opinion Research Center's *General Social Survey* (GSS). Finally, we attempt to relate our measures of displacement and worker anxiety to wage growth by examining time-series data for the nine U.S. census divisions.

Our tabulations of annual displacement rates from the DWS focus on workers with five or more years of tenure. We find that among such workers, job loss due to "shift or position abolished," which among the surveys' possible reasons for job loss comes closest to capturing the notion of "downsizing," increased quite dramatically from annual rates of two tenths or three tenths of a percent throughout the 1980s to a range of six tenths or seven tenths of a percent in the mid-1990s. Determining the trend in displacement more generally is complicated by changes in the DWS. However, our preferred estimates suggest that overall displacement rates were higher in 1995, the most recent year for which we have data, than at any time since the data began in 1979. We estimate a 1995 displacement rate of about 3.4 percent for workers with five or more years of tenure. By comparison, the rate for 1982, which was in the middle of a severe recession, was only about 2.5 percent. We consider this a substantial increase in the risk of displacement for high-seniority workers.

We also find that displacement has become somewhat more "democratic" in the 1990s. Previously, high-seniority workers who were highly educated, were in white-collar jobs, or were employed in the service producing industries were relatively immune to displacement. More recently, however, displacement rates for these groups have risen especially fast, while those for some groups who had high rates of displacement in the 1980s, such as those with at most a high school education, those in blue-collar occupations, and those working in manufacturing, rose less or even fell relative to their peaks in the early 1980s. As a result of this increased democratization of displacement, many more workers may now consider themselves at risk for job loss.

The GSS data suggest that workers' own perceptions of their job security have also declined. The fraction of workers not responding "very unlikely" to the question, "How likely is it that you will lose your job in the next year?," rose from about 31 percent in 1989 to about 40 percent in 1996, the most recent year for which data are available. The 1996 figure approximately matches the highest reading since this question began to be asked in 1977. The 1996 reading is especially remarkable given that unemployment was generally below 6 percent, while in 1982, when such a level of anxiety was previously reached, unemployment was nearly 10 percent. One should not, however, exaggerate the extent to which workers' anxiety over job loss has increased. The main change has been an increase in the number of workers responding that it is "not too likely" rather than "very unlikely" that they will lose their jobs. The percentages of workers responding that it is "fairly likely" or "very likely" have risen more modestly.

With a few exceptions, the groups of workers who have experienced the largest increases in displacement rates have also had the largest increases in reported probabilities of job loss. For instance, an increase in perceived likelihood of job loss has been especially great among white-collar workers. Perceived job security has actually increased for blue-collar workers. Another interesting finding concerns the relationship between workers' perceptions of their job security and the use of computers in their industry. In the early 1980s, workers in industries with greater computer usage felt more secure on average than other workers. By the mid-1990s, however, the relationship had reversed, with workers in industries with greater computer usage feeling less secure.

Finally, we attempted to judge to what extent our findings of an increase in displacement rates and workers' perceptions of their chances of job loss are related to changes in aggregate wages. Standard short-run Phillips curve analyses such as Gordon (1997) have tended to predict higher levels of wage inflation than have actually occurred in the last two or three years. Though significant forecast errors are nothing new for such models,[3] the importance of the question to policymakers has led to a great deal of speculation as to why wage inflation has remained subdued. Our findings and those of other researchers, which suggest that job security has declined in recent years, add some plausibility to the case

that worker anxiety has played a role in restraining wage inflation.

However, one could easily point to other recent changes in labor markets or elsewhere in the economy that might be affecting wage growth.[4] Why should one particular change—that towards reduced job security—be considered the key factor? To make a more convincing case for the importance of worker insecurity, one would want to observe that in the past, when displacement or anxiety was high relative to unemployment, wage inflation had also been subdued. Moreover, to obtain any sense of the quantitative importance of worker insecurity in restraining wage inflation, one needs to examine historical evidence. Unfortunately, with annual measures of displacement and insecurity that go back only to the late 1970s, there is not much hope of extracting such information from the U.S. time-series data.

Our strategy is to look cross-sectionally as well as over time and ask whether census regions that have had higher displacement rates or worker perceptions of insecurity have tended to have lower wage growth. Such a strategy parallels the "wage curve" analyses of Blanchflower and Oswald (1994), although, as in Blanchard and Katz (1997), we employ the traditional Phillips curve specification in which the change in wages, rather than their level, is related to unemployment and measures of job security.

We pool separate data from the nine census regions to estimate the effect of displacement rates or perceptions of insecurity on forecasts of wage inflation. We find that, holding constant the unemployment rate, higher values of both displacement and worker insecurity are associated with lower wage growth. However, even with the additional source of data variation that comes from pooling the separate census divisions, our estimates of the magnitude of these effects are imprecise. Indeed, if we allow for the possibility that there may be some permanent, unmeasured characteristics of regions that are associated with different levels of wage growth, we cannot reject the hypothesis that the true effect of job security on wages is zero and that estimates the size of those we obtain could have arisen by chance. Nonetheless, our best estimates suggest that increases in displacement rates and workers' own anxiety about their job security could be responsible for restraining wage growth by about three tenths to seven tenths of a percentage point per year during the mid-1990s. Such an effect would explain all or most of the puzzle of lower than expected wage inflation.

## Previous research on turnover and displacement

The *New York Times* describes its book on downsizing as putting "a human face on a historic predicament that is as ubiquitous as it is painful."[5] A large body of research demonstrates that job loss is painful, at least for workers who have attained significant tenure.[6] For example, Jacobson, LaLonde, and Sullivan (1993c) found that even six years after job loss, earnings losses among a sample of Pennsylvania workers displaced in the early 1980s were still equal to about 25 percent of their predisplacement earnings levels. What has been somewhat less clear to researchers is whether job loss has become any more ubiquitous in recent years.

It is helpful to divide the relevant research into two parts—that on job stability and that on job security. By stability we mean the tendency for workers and firms to develop long-term relationships. Research on job stability questions the many media accounts claiming that such long-term employment relationships have gone the way of buggy whips. By security we mean workers' ability to remain in employment relationships as long as their own performance is satisfactory. Research on job security asks whether there has been an increase in involuntary job loss due to reasons beyond workers' control. Job stability depends on workers' own choices, in addition to the factors that influence job security. For instance, if a group of workers increase their commitment to the labor force or to their particular employers, then their job stability may rise even if they are increasingly subject to threats of displacement. As we shall see, research suggests a larger 1990s decline in job security than in job stability.

In our view, trends in job security are much more relevant to the discussion of whether special factors might be restraining wage inflation than are trends in job stability. Indeed, if declines in job stability are less dramatic than declines in job security, it must largely be because workers are less likely to leave jobs voluntarily, and a decreased tendency to quit jobs may itself signal worker insecurity. Nevertheless, we begin with a short account of research on job stability.

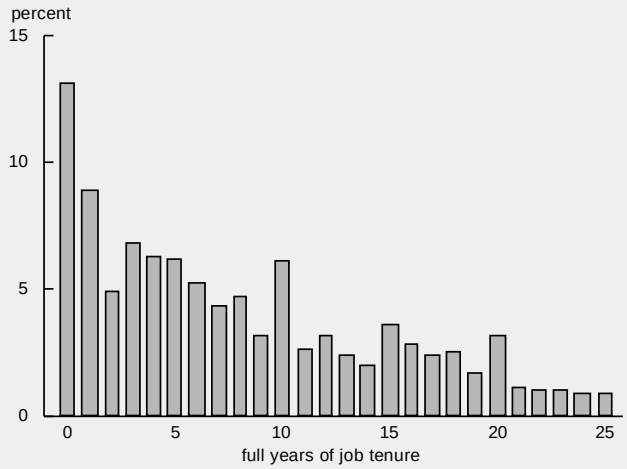
The starting point for much of the research on job stability is the distribution of job tenure.

Most of what is known about this distribution derives from a series of supplements to the *Current Population Survey* (CPS).[7] As an illustration, figure 1 displays the distribution of job tenure for employed men between the ages of 35 and 44. These data were collected from the most recent *Mobility Supplement* to the CPS, which was conducted in February 1996. The figure shows that the most common tenure levels are the shortest—for example, less than one year and between one and two years—with a roughly monotonic decline in the number of workers with successively longer tenure. Nevertheless, there are many workers with substantial levels of job tenure. For men in the 35 to 44 age group, the median tenure is about 6.1 years. Moreover, about 33 percent of such workers have been in their current jobs at least ten years and about 22 percent have been in their current jobs at least 20 years.

Figure 2 shows how median job tenure has changed over time for men and women in three age groups, 25 to 34, 35 to 44, and 45 to 54. These data are derived from CPS *Mobility Supplements* conducted in January of 1963, 1966, 1968, 1973, 1978, 1983, 1987, and 1991 and February of 1986.[8] Not surprisingly, older workers typically have longer job tenures than younger workers. Also, men typically have longer tenures than women, who are more likely to have interrupted their careers for family reasons. However, our primary interest is in the aggregate trends in these data.

**FIGURE 1**

**Distribution of job tenure for men aged 35–44, February 1996**




Source: Authors' calculations based on data from U.S. Department of Commerce, Bureau of the Census, *Current Population Survey*, *Mobility Supplement*, February 1996.

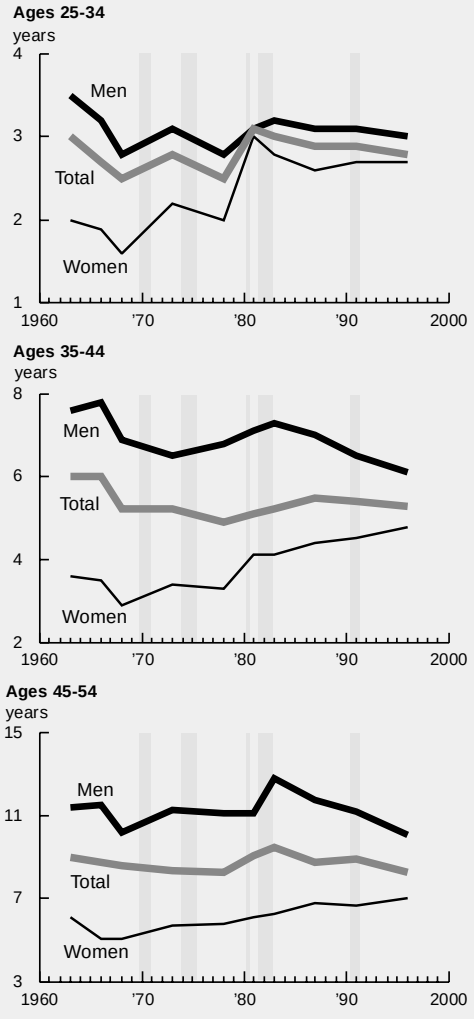
For men, especially those in the two highest age groups, median job tenures declined from 1991 to 1996, which is consistent with claims of decreased job stability. However, women's job tenure rose for all age groups. So, overall, there has been relatively little change in median job tenure during the 1990s. Moreover, the drop in male tenure for the two oldest groups seems to be mainly a continuation of a trend that was evident throughout the 1980s. Thus, it is difficult to conclude that job stability has suffered more than a modest decline in the 1990s.[9]

Farber (1997b) shows that similar, though somewhat more dramatic, changes took place during the 1990s at the high end of the tenure distribution. In particular, he shows that the percentages of workers reporting more than ten and 20 years of tenure declined significantly between 1991 and 1996. The proportion of workers aged 35 to 64 with ten or more years of tenure declined from 38.3 percent to 35.4 percent. For men the decline was more dramatic, from 44.3 percent to 40.0 percent, while among women, the decline was from 31.4 percent to 30.3 percent. Similar drops were reported for workers with different educational levels. However, the occupational groups that have historically had the highest long-term employment levels, such as managerial, professional and technical, and blue-collar workers, had the largest declines in the 1990s. Similarly, the declines were greatest in industries, such as transportation, communications, and public utilities, in which long-term employment had been most common. As a result, the frequency of long-term employment is now more similar across occupations and industries.

Farber's (1997b) results, as well as the trends in median job tenure shown in figure 2 suggest that job stability among men has declined modestly during the 1990s. For women, job stability has either declined very modestly or continued to rise, depending on whether it is measured by median tenure or the proportion of workers with high tenure levels. For both sexes, the changes appear too modest and gradual to support the sensationalistic media reports proclaiming the end of long-term employment relationships.

**FIGURE 2**

**Median tenure, ages 25–54**




Note: Shaded areas indicate recessions.

Source: Authors' calculations based on data from U.S. Department of Labor, "Employee tenure in the mid-1990s," January 30, 1997, and Farber (1998).

To reach a tenure of ten years, a worker must survive from the first year of a job into the second, from the second to the third, and so on for ten years. Thus, the distribution of job tenures and, in particular, the fraction of workers with ten or more years of tenure can be thought of as depending on a sequence of survival probabilities going back many years. This means that if the probability of remaining in a job for another year were to have suddenly dropped sometime in the early 1990s, it would take a number of years for this change to show up fully in the tenure distribution. In this case, results such as Farber's (1997a) and those shown in figure 2 might not reveal the full extent of change. For this reason, it is of interest to examine job survival or retention probabilities.[10]

Diebold, Neumark, and Polsky (1997) and Neumark, Polsky, and Hansen (1997) have carefully analyzed the trend in retention rates from the mid-1980s to the mid-1990s, using data on tenure distributions from CPS supplements for every four years from 1983 to 1995. The authors estimate four-year retention rates—the probability that a worker will remain in a job an additional four years—by dividing the number of workers with a given set of characteristics and a certain number of years of tenure in one survey by the number of workers with those characteristics but four fewer years of tenure in the survey four years earlier. Adjustments are made for a number of potential problems, including nonresponse to the survey, the tendency of workers to round their tenure to a multiple of five years, the differing levels of unemployment at the time of the surveys, and the special nature of the tenure data derived from the February 1995 CPS *Supplement on Contingent Work*.

Table 1 contains some representative results from Neumark, Polsky, and Hansen (1997).[11] Evidently, the trend over time in retention rates depends to a great extent on workers' initial level of tenure. For workers with less than two years of initial tenure, four-year retention probabilities are estimated to have increased from 32.9 percent for the 1983–87 period to 34.6 percent for the 1987–91 period to 39.1 percent for the 1991–95 period. However, for workers with two to less than nine years of tenure, rates first declined then rose slightly. The strongest evidence of a decline in job stability comes from the group of workers who initially had between nine and 15 years of tenure. The retention rate for these workers declined from 81.6 percent for 1987–91 to 74.8 percent for 1991–95. Retention rates also declined sharply between 1987–91 and 1991–95

**TABLE 1**

**Four-year job retention rate estimates**

| Initial tenure | 1983–87 | 1987–91 | 1991–95 |
|---|---|---|---|
| 0 to < 2 | 32.9% | 34.6% | 39.1% |
| 2 to < 9 | 58.6 | 54.8 | 56.4 |
| 9 to < 15 | 82.7 | 81.6 | 74.8 |
| 15 plus | 63.0 | 70.2 | 63.3 |
| Weighted average | 53.9 | 53.6 | 54.4 |

Source: Derived from Neumark, Polsky, and Hansen (1997).

for workers with 15 or more years of tenure, but returned to rates observed in the 1983–87 period. The weighted average rate was quite stable, falling just 0.3 percentage points from 1983–87 to 1987–91 and then increasing 0.8 percentage points from 1987–91 to 1991–95.

The results on retention probabilities are consistent with those on tenure levels in suggesting some modest declines in job stability for workers with several years of tenure. Several researchers have reported more dramatic declines in job stability during the 1980s and / or 1990s. For example, Boisjoly et al. (1994), Rose (1995), and Marcotte (1996) report evidence of declining job stability from the *Panel Study of Income Dynamics* (PSID) data. Similarly, Swinnerton and Wial (1995) reported significant declines in job retention rates in the 1991–95 period. However, in our view the combined results of Diebold et al. (1996), Swinnerton and Wial (1996), and Jaeger and Stevens (1997) show that the more dramatic declines reported in the literature were largely the result of researchers failing to take account of occasional changes in survey question wording. The most careful analyses of job stability trends imply that there have been at most modest declines in stability in the late 1980s and 1990s.[12]

Research suggests, we believe, larger declines in measures of job security. Farber (1997a) analyzes data from the seven *Displaced Worker Surveys* (DWS), CPS supplements that are described in the next section. He finds that "rates of job loss are up substantially relative to the standard of the last decade, particularly when some consideration is given to the state of the labor market." He also finds that displacement rates increased most for several groups, such as the more educated and those in white-collar occupations, that have traditionally had relatively low levels of displacement, which implies that displacement has become somewhat more democratic. Changes in the reasons workers give for their job loss also point to especially large increases in what the media might mean by "downsizing." Finally, Farber finds that the consequences of displacement in terms of time spent unemployed and reduced wage rates upon reemployment appear to be mainly a function of the business cycle. The consequences of displacement were worse during the recessions of the early 1980s and 1990s, but there is little evidence of a secular increase in the seriousness of displacement.

Valetta (1997) also finds an increasing number of dismissals in data from the PSID, which is consistent with Farber's results (1997a). Moreover, Valetta's finding that the increase is concentrated among workers with higher levels of tenure is consistent with our finding below that the increase in displacement rates for workers with five or more years of tenure has been especially dramatic.

## Displacement trends for high-seniority workers

Below, we present new measures of the rate of job displacement for workers with five or more years of tenure. We had two main goals. First and most important, we wanted the measures to be comparable over time so that we could accurately judge whether displacement was increasing. Given changes in the underlying survey methodology, this is not completely straightforward and, despite our best efforts, it is possible that certain of our measures change over time for reasons that have nothing to do with actual changes in the rate of job displacement. Our second goal was to create annual time series, the highest frequency possible, so as to be better able to examine the relationship between displacement and wage inflation.

Our measures of displacement are based primarily on the Bureau of Labor Statistics' DWS. These surveys were conducted as supplements to the CPS in January of even years from 1984 to 1992 and in February 1994 and 1996.[13] For the purposes of the survey, displacement is defined as involuntary job loss not related to a worker's performance. Thus, displacement excludes quits and cases in which workers are discharged for poor performance.[14] The surveys are retrospective, asking individuals whether they have experienced job loss any time over the last five years in the case of the 1984 to 1992 surveys and over the last three years in the case of the 1994 and 1996 surveys. Thus, our earliest information on displacement is for 1979 and our latest is for 1995.

For workers who report that they were displaced in the relevant time period, the DWS asks for the specific reason for their displacement. The possible responses are:

- Plant or company closed down or moved,
- Insufficient work,
- Position or shift abolished,
- Seasonal job completed,
- Self-operated business failed, and
- Some other reason.

This list of reasons is less than ideal. For example, insufficient work might be the reason why one of the other events occurred. A plant may have closed because there was insufficient work to do. Position or shift abolishment is probably supposed to cover instances of "corporate downsizing," but it is possible that those in nine to five work environments will be confused by the reference to shifts. In any case, it lumps together instances of complex "re-engineering" exercises, which presumably reflect long-run organizational changes, with closings of shifts in factories, which are more likely to be associated with short-run declines in demand. The seasonal job and self-employment categories don't correspond to many people's conception of job displacement and, in fact, make up only a trivial fraction of the job loss that we consider. Finally, perhaps because of some of the ambiguities of the preceding categories, "other" is a common response. In fact, growth in the "other" category is responsible for a large percentage of the total growth of displacement of high-seniority workers.

The first difficulty we face in constructing a consistent measure of worker displacement is that the DWS only collects information, such as the year of displacement, the worker's tenure, and other characteristics of the lost job, for at most one incident of displacement over the relevant period. If workers were displaced twice or more in the same period, they are instructed to answer the additional questions for the lost job on which they had the highest tenure. This inevitably leads to some undercounting of incidents of displacement. Moreover, as Farber (1997a) notes, the change in the length of the period over which the DWS asks workers to report on displacement creates a problem of comparability over time, since the undercounting problem is more severe when the interval covered is five years.

Farber's approach to this problem is to examine only displacement that occurred in the last three years of the five-year periods covered by the 1984 to 1992 surveys. As he notes, these rates are still not comparable to rates computed from the three-year intervals of the 1994 and 1996 surveys, because some workers may lose a job in year one or two of the five-year period before the survey and then lose another job in year three, four, or five. If the workers had accumulated less tenure on the second lost job than they had on the first, they would be recorded as losing a job in the 1994 and 1996 surveys, but not in the last three years before the 1984 to 1992 surveys. Farber's solution is to use PSID data to quantify the frequency of job loss patterns and adjust rates in the DWS to offset them.[15]

Our approach is to restrict our analysis to incidents of job displacement in which the affected workers had five or more years of tenure. Obviously, it is not possible to lose two such jobs in one three- or five-year interval, so the number of such job loss incidents should be correctly tallied no matter whether the year is part of a three- or five-year interval in the DWS. Of course, we will miss *all* displacement incidents in which workers had less than five years of tenure. However, the consequences of job loss are not likely to be particularly great for workers with little tenure and, thus, our measure may capture the most important forms of job displacement.

The DWS gives us estimates of the number of workers with five or more years of tenure who are displaced in a particular year. To calculate a displacement rate, we need to divide this estimate of the number of high-tenure displaced workers by the number of high-tenure workers who were at risk in that year. We derive the latter figure as the product of the level of total employment and the fraction of total employment accounted for by workers with five or more years of tenure.

Our estimated displacement rate is $r_t^5 = \frac{d_t^5}{n_t f_t^5}$,

where $d_t^5$ is the number of workers with five or more years of tenure displaced in year $t$, $n_t$ is total employment, and $f_t^5$ is the fraction of employment accounted for by workers with five or more years of tenure.

As noted above, we derive estimates of $d_t^5$ from the DWS. To estimate $n_t$, we use the CPS outgoing rotation files. The outgoing rotations are those CPS members who are in the fourth and eight months of their eight-month participation, about 25 percent in a given month. Pooling the outgoing rotations for all 12 months of the year yields a large data set that can be used to estimate employment levels quite precisely. To estimate $f_t^5$, we use the CPS tenure supplements described earlier. As noted, these were conducted in 1981, 1983, 1987, 1991, and 1996. To compute displacement rates for 1979 and 1980, we use the value of $f_t^5$ from 1981. For other years in which there was no supplement, we interpolate linearly from the preceding and succeeding tenure supplements. Because the fraction of workers

with five years of tenure changes very slowly relative to the number of displaced workers, this interpolation causes no problems.

There is another problem with the 1994 and 1996 DWS. The follow-up questions on the details of the displacement episode are not asked if workers do not give one of the first three standard reasons for displacement. This is unfortunate because, as we have already noted, a nontrivial and growing number of workers report "other" as their reason for displacement. To ignore workers not responding with one of the three standard reasons would, we feel, significantly skew our results.[16] However, for workers giving a nonstandard reason, we do not know whether they had five years of tenure or in what year they lost their job.

To deal with this problem, we estimated statistical models to gauge the percentage of displaced workers giving nonstandard reasons who had five years of tenure and, of those, the percentage who were displaced in each of the three years covered by the surveys. The details of our procedure are contained in box 1. The idea is to use the displaced workers giving nonstandard reasons in 1992 to determine which worker characteristics reported in the basic CPS were associated with having five years of tenure and then to use those characteristics to estimate the percentage of displaced workers reporting nonstandard reasons in the 1994 and 1996 surveys who had five years of tenure. Similarly, we used the workers reporting standard reasons and five years of tenure in the 1994 and 1996 surveys to determine the characteristics associated with being displaced in each of the three years covered by the surveys and then used those characteristics to predict which year workers reporting nonstandard reasons were displaced.

Table 2 shows our basic results for overall displacement rates. The rows of the table correspond to the years in which displacement occurred. The first five columns of the table correspond to the number of years after the displacement year that the displacement rate was measured. For example, the only information on the 1979 displacement rate comes from the 1984 DWS, which was conducted with a lag of five years. The estimated rate, 0.96 percent, is thus shown in the column headed five-year lag. For the majority of years, we have multiple measures of the displacement rate. For example, for 1985 we have estimates from the 1986, 1988, and 1990 DWS. These rates, shown in the columns for one-, three-, and five-year lags, are estimated to be 2.29 percent, 1.79 percent, and 1.43 percent, respectively.

**TABLE 2**

**Percent displaced among workers with five or more years tenure: Alternative measurement lags**

| Year | One-year lag | Two-year lag | Three-year lag | Four-year lag | Five-year lag | Overall estimate |
|---|---|---|---|---|---|---|
| 1979 | | | | | 0.96 | 1.46 |
| 1980 | | | | 1.23 | | 1.69 |
| 1981 | | | 1.59 | | 1.23 | 1.92 |
| 1982 | | 2.27 | | 1.84 | | 2.52 |
| 1983 | 2.41 | | 1.66 | | 1.52 | 2.25 |
| 1984 | | 1.75 | | 1.22 | | 1.80 |
| 1985 | 2.29 | | 1.79 | | 1.43 | 2.22 |
| 1986 | | 1.99 | | 1.56 | | 2.18 |
| 1987 | 1.83 | | 1.34 | | 1.34 | 1.84 |
| 1988 | | 1.44 | | 1.18 | | 1.61 |
| 1989 | 1.71 | | 1.62 | | | 1.85 |
| 1990 | | 1.90 | | | | 2.11 |
| 1991 | 2.76 | | 2.35 | | | 2.83 |
| 1992 | | 2.40 | | | | 2.67 |
| 1993 | 2.92 | | 2.31 | | | 2.89 |
| 1994 | | 2.21 | | | | 2.46 |
| 1995 | 3.44 | | | | | 3.44 |

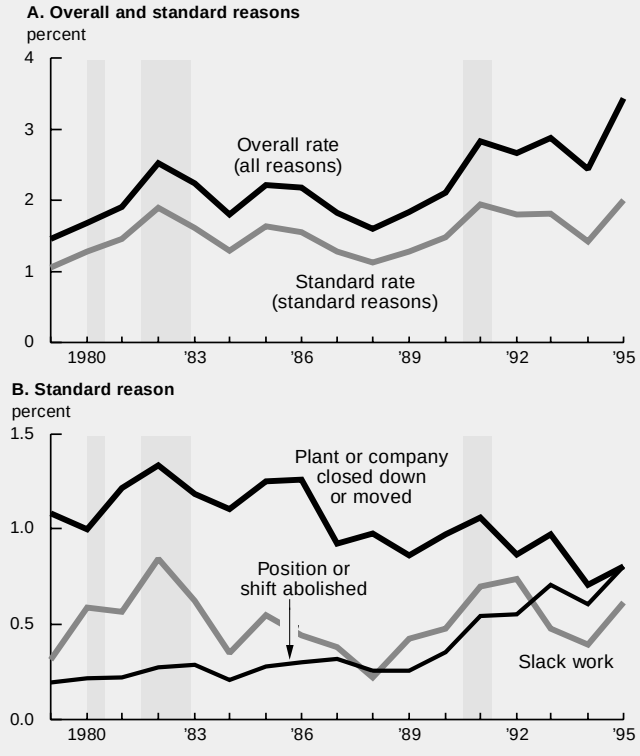
Source: Authors' calculations from data of the U.S. Department of Labor, Bureau of Labor Statistics.

The results for 1985 illustrate the final difficulty we face in constructing an annual measure of job displacement for workers with five or more years of tenure. That is, displacement rate estimates tend to drop as the time since the survey increases. Workers seem to forget incidents of displacement as time passes, a phenomenon noted previously by Topel (1990) and others. As a result, it is inappropriate to simply average the various measures to arrive at an overall displacement rate for that year. For instance, if we were to directly compare the single estimate for 1979 with the single estimate for 1995, we would be comparing a rate measured with a five-year lag with a rate measured with a one-year lag. Thus, the comparison would reflect not only differences in actual displacement rates between the years, but also the tendency of rates measured with a greater lag to be lower.

Table 2 reveals that estimated displacement rates tend to drop on average by about 11 percent for each additional year that the survey lags the year of displacement. Our solution to this problem, which is described in detail in box 1, is essentially to adjust rates based on lags greater than one year upward by about 11 percent for each additional year that the survey lags the year of displacement. Our final estimates of the annual displacement rates, which are shown in the last column of table 2, are averages of all the adjusted rates for the year in question. For instance, the estimated 11 percent annual decline suggests that if the rate for 1979 had been measured in 1980, it would have been 1.46 percent, rather than 0.96 percent. Thus, our final estimate for 1979 is 1.46 percent. In a year with multiple measurements, the measures are adjusted by different amounts, depending on how long after the year of displacement the survey was taken. For instance, the estimate for 1985 obtained with a one-year lag is left at 2.29 percent, but the rate obtained with a three-year lag is adjusted up from 1.79 percent to 2.18 percent to reflect the additional two years since the survey; the rate with a five-year lag is adjusted from 1.43 percent to 2.17 percent to reflect the additional four years since the survey. The adjusted rates of 2.29 percent, 2.18 percent, and 2.17 percent are then combined to obtain the final estimate of 2.22 percent.[17]

**FIGURE 3**

**Displacement rates, workers with 5 years tenure**




Note: Shaded areas indicate recessions.
Source: Authors' calculations based on data from U.S. Department of Labor, Bureau of Labor Statistics, *Displaced Worker Survey*, 1984–96.

The final results are plotted over time as the black line in figure 3, panel A. The overall displacement rate for workers with five years of tenure rose during the recessions of the early 1980s from 1.5 percent in 1979 to a peak of 2.5 percent in 1982. It then declined during the economic expansion that followed to a low of about 1.6 percent in 1988. Then, in the 1990s, it rose rather dramatically. It is not surprising that the rate should have risen during the recession of 1990–91, but the 1991 rate, at over 2.8 percent, was 0.3 percentage points higher than in 1982, even though by most measures the 1982 recession was much more severe. More noteworthy is the failure of the displacement rate to decline during the expansion of the mid-1990s. Indeed, in 1995 the rate shot up to 3.4 percent, its highest ever reading. The high overall displacement rates that we estimate for the mid-1990s are consistent with the view that job security declined significantly for workers with five or more years of tenure.

**BOX 1**

## Constructing an annual index of displacement

To estimate the fraction of workers reporting nonstandard reasons for displacement in the 1994 and 1996 DWS who had five or more years of tenure, we estimated a logistic regression model using the sample of such workers in the 1992 DWS.[1] The dependent variable in this model was an indicator for having five years of tenure and the independent variable consisted of dummy variables for the nine census regions, sex, ten-year age categories, race, marital status, education less than high school, high school graduate, some college, and college degree, as well as part-time status, one-digit occupation, and one-digit industry of the person's job as reported in the main CPS. We then used the estimates of the parameters of this model, along with the equivalent characteristics for workers reporting nonstandard reasons for displacement in 1994 and 1996 to form an estimate of the probability that such workers had five or more years of tenure at the time of their job loss.

We estimated the fraction of such workers that were displaced in each of the three possible years covered by the 1994 and 1996 surveys by estimating a multinomial logistic regression model on the sample of 1994 or 1996 displaced workers reporting standard reasons for displacement. In this model, the dependent variables were indicators for the year of displacement and the independent variables were the same as in the model above. The parameter estimates were then used to estimate the probability that workers were displaced in each of the three years covered by the 1994 and 1996 surveys.

In computing displacement rates based on the 1994 and 1996 surveys, we then counted all workers reporting nonstandard reasons for displacement as displaced in all three possible years. However, we multiplied the weights for such individuals by the estimated probabilities of having five or more years of tenure and of being displaced in the year in question. This procedure should provide estimates of displacement rates among those with five years of tenure that are consistent over time if 1) the relationship between the probability of five-year tenure and the independent variables remains constant from 1992 to 1994 and 1996, and 2) the distribution of year of displacement conditional on the independent variables is the same for workers displaced due to standard and nonstandard reasons.

The final task in computing annual displacement rates is to combine rates measured for the same year by different surveys into a single overall rate. We did this by estimating the following simple statistical model:

$$\log r_{st} = \alpha_t + \gamma(s - t - 1) + \varepsilon_{st},$$

where $r_{st}$ is the displacement rate for year $t$ measured by the survey in year $s$, and $\varepsilon_{st}$ is an error term assumed to have constant variance and to be uncorrelated across observations. The parameter $\gamma$ measures the rate at which estimates of displacement rates decline as time between displacement and the survey increases. Its estimate corresponds to an approximately 11 percent rate of decline. The overall rate is captured by the year of displacement effects, $\alpha_t$. Specifically, the estimate of the rate corresponding to a one-year lag between displacement and the survey is $\exp(\alpha_t)$. These are the estimates shown in the final column of table 2 and plotted in figure 3.

In order to compute estimates for separate demographic groups, we expanded the above model to

$$\log r_{dst} = \alpha_{dt} + \gamma(s - t - 1) + \varepsilon_{dst},$$

where $r_{dst}$ is the rate for demographic group $d$ in year $t$ as measured by the DWS of year $s$. The demographic specific rates are then $\exp(\alpha_{dt})$. We also computed estimates of displacement rates adjusted for changes in the age and sex distribution. These were based on models of the form

$$\log r_{dstk} = \alpha_{dt} + \beta_k + \gamma(s - t - 1) + \varepsilon_{dstk},$$

where $r_{dstk}$ is the rate for the age and sex group $k$. The presence of the $\beta_k$ controls for changes in the age and sex distribution that might affect estimates of overall rates. However, the adjusted rates were similar enough to the unadjusted rates that we only report the latter.

[1]See, for example, Maddala (1983) for an explanation of the logistic regression model discussed below.

The colored line in figure 3, panel A shows the rate of displacement due to the first three standard reasons in the survey. Comparing the two lines, it is clear that a large part of the significant mid-1990s increase is due to an increase in the number of displaced workers giving "other" as their reason for displacement. However, even the colored line, which is not dependent on the imputations of tenure and year of displacement described in box 1, suggests that there has been some decline in security, especially given the level of unemployment. The rate of displacement for standard reasons is estimated to be higher in 1995 than it was in 1982, even though the unemployment rate was below 6 percent during most of 1995, while it was nearly 10 percent in 1982. Thus, even when limited to displacement for standard reasons, our results suggest a noticeable decline in job security.

Figure 3, panel B displays separate displacement rates for the three standard reasons. Evidently, the rate due to firms or plants closing or moving has declined somewhat in the 1990s, while the rate due to slack work has remained relatively high, given the state of the business cycle. However, the most notable feature of figure 3, panel B is the sharp increase beginning in 1990 of the displacement rate due to shifts or positions being abolished. This rate, which probably comes the closest to capturing corporate downsizing, was between 0.2 percent and 0.3 percent from 1979 to 1989, but rose to more than 0.8 percent in 1995. This two hundred or three hundred percent increase seems to represent a rather significant break from history.

Figure 4, panel A shows the overall displacement rate for men and women. For most of the period covered by our data, women were less subject to displacement than men, with the typical gap in rates being five tenths or six tenths of a percentage point. In the last three years, however, the gap has been much smaller, about one tenth of a percentage point. Thus, by our measure, women have suffered a larger decline in job security than men. This finding highlights the difference between the displacement rates estimated here and the trends in median tenure discussed earlier. Median tenure has generally been increasing for women relative to men. However, tenure levels are measures of stability, reflecting workers' own commitment to the labor force and individual employers in addition to forces beyond workers' control, such as displacement. In the comparison of male and female tenure levels, workers' own choices are likely the more important factor. For this reason, we would argue that displacement rates are the better measure of worker insecurity.

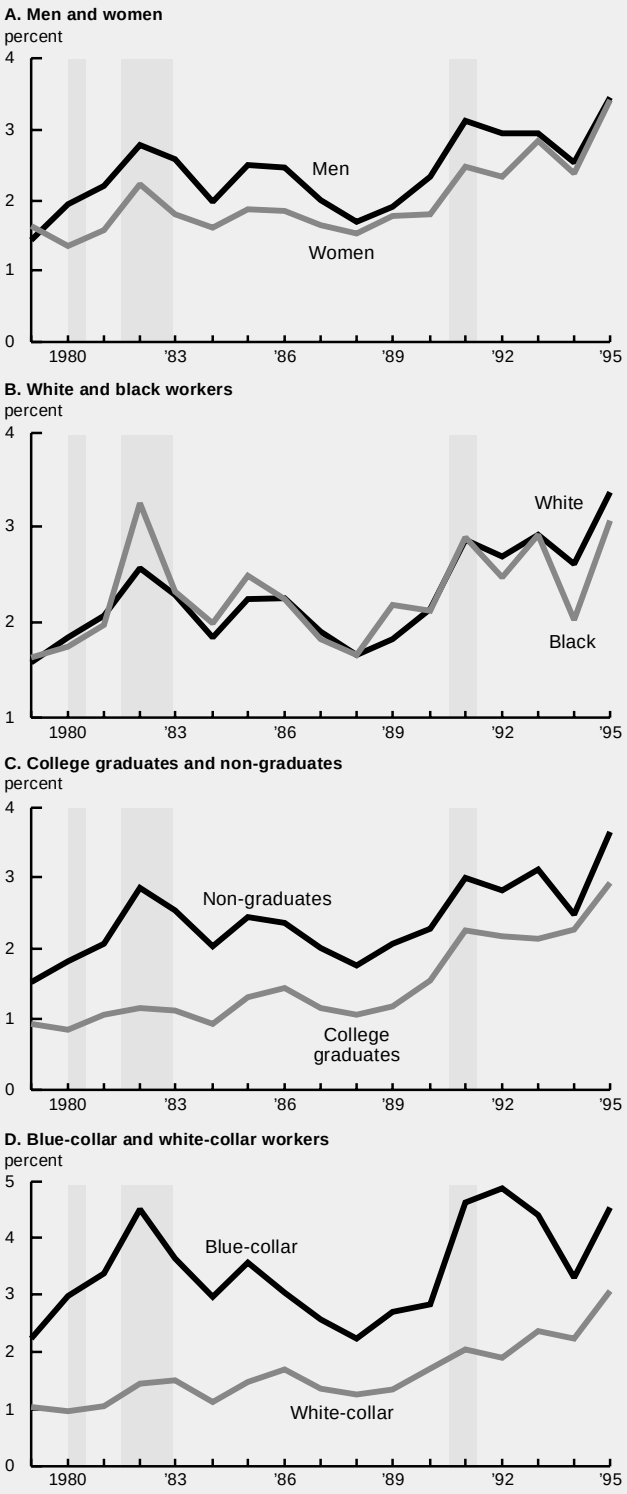
Figure 4, panel B displays displacement rates for white and black workers. Once we restrict the sample to workers with five or more years of tenure, the difference between the races is relatively minor. Still, there have been some changes over time. Early in the period covered, especially during the recession of the early 1980s, blacks had noticeably higher displacement rates. However, by the end of the period, whites had higher rates of displacement.

Figure 4, panel C shows the breakdown between those with a college degree and those without a college degree. Although displacement rates for college graduates remain much lower than those for workers without college degrees, the gap has narrowed considerably in the 1990s. Until 1990, displacement rates for college graduates never exceeded 1.3 percent and the gap between them and non-graduates was often a percentage point or more. In the 1990s, displacement rates for college graduates rose especially sharply, to levels of more than twice their previous peak. Thus, the gap in displacement rates between those with a college degree and those without has narrowed considerably, though rates for college graduates remain significantly lower.

Figure 4, panel D shows displacement rates for blue-collar and white-collar workers. Though displacement rates for high-tenure blue-collar workers remain about a percentage point higher than those for white-collar workers, the gap has clearly shrunk during the 1990s. For instance, even the recessions of the early 1980s had little effect on displacement rates for high-tenure white-collar workers, but since 1988, their rates of job loss have approximately doubled. By contrast, the recessions of the early 1980s caused a major increase in blue-collar displacement to levels only slightly lower than in recent years. Even more dramatic differences in displacement trends are observed between more narrowly defined occupations. For example, 1995 displacement rates for laborers are significantly lower than in 1982, while those for professional and technical workers are approximately three times higher.

**FIGURE 4**

**Displacement rates among workers with five or more years tenure, by demographic category**




Note: Shaded areas indicate recessions.

Source: Authors' calculations based on data from U.S. Department of Labor, Bureau of Labor Statistics, *Displaced Worker Survey*, 1984–96.

Figure 5 shows estimated displacement rates for workers in goods producing and service producing industries. Again, a large gap in rates in the 1980s has narrowed appreciably. Displacement rates for those in goods producing industries are still significantly lower than in 1982, but rates for those in the service producing industries are about two and half times greater. Even so, workers in goods producing industries remain significantly more at risk for displacement than those in service producing industries. More dramatic changes can be identified for certain industries. For instance, displacement rates for workers in the finance industries rose from about 0.5 percent to 1.0 percent in the 1980s to 2.8 percent in 1995.

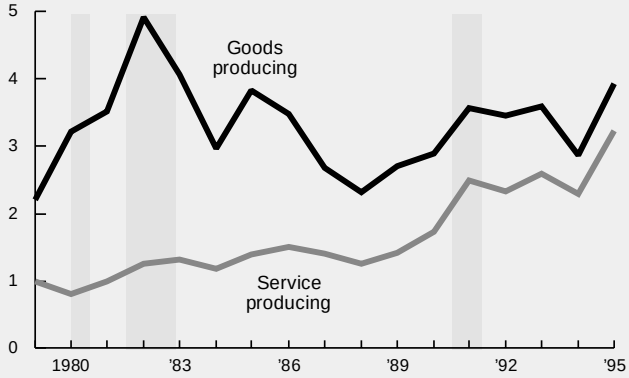
The results in figures 4 and 5 all point to the general increase in high-tenure displacement rates having been accompanied by a kind of democratization, in which those who had been relatively immune to job displacement have seen the fastest increase in displacement. Previously, those with a college education, in white-collar jobs, or in service producing industries might have considered themselves immune to job loss. Given the increase in displacement rates that we have estimated for these groups, this is probably no longer the case for many such workers. Thus, the number of workers who feel at risk may have increased even more than the increase in the displacement rate would suggest.

## Workers' perceptions of job security: The NORC-GSS

In a series of recent papers, Manski (1990, 1993) has observed that researchers know a great deal about the outcomes that individuals or groups experience but much less about the outcomes that they expect. This assertion is particularly relevant for job security research, which to date has focused on the measurement of displacement rates, tenure distributions, and other measures of actual employment outcomes. However, a primary issue in this literature concerns measuring perceptions of

**FIGURE 5**

**Displacement rate, by industry**




Note: Shaded areas indicate recessions.
Source: Authors' calculations based on data from U.S. Department of Labor, Bureau of Labor Statistics, *Displaced Worker Survey*, 1984–96.

the risk of future economic harm. Therefore, these measures are indirect, in the sense that expectations about risk, which are subjective in nature, must be inferred from individual or group realizations.[18]

The *General Social Survey* (GSS) data set allows us to address the perceptions question directly. Up to now, this data set has received some attention in the popular press but little among researchers studying job security.[19] The GSS is a nationally representative annual survey conducted by the National Opinion Research Center (NORC). The survey asks a series of demographic and employment questions, including, in most years since 1977, two questions about job security. Respondents are asked 1) "Thinking about the next 12 months, how likely is it that you will lose your job or be laid off—very likely, fairly likely, not too likely, or not at all likely?" and 2) "About how easy would it be for you to find a job with another employer with approximately the same income and fringe benefits that you now have? Would you say very easy, somewhat easy, or not easy at all?"

Data are available for 13 years between 1977 and 1996 covering roughly 10,000 individuals. Several years (1980, 1984, 1987) are missing because NORC did not ask the job security questions, and other years (1979, 1981, 1992) are missing because the GSS was not conducted. The sample includes all respondents who are currently employed, English-speaking, and aged 18 to 64. It is important to note that the sample makes no restriction on tenure because such information is not given in the GSS. Therefore, the job security perceptions sample is not strictly comparable to the displacement rate sample discussed earlier. This probably accounts for some different trends among subsamples of the population.

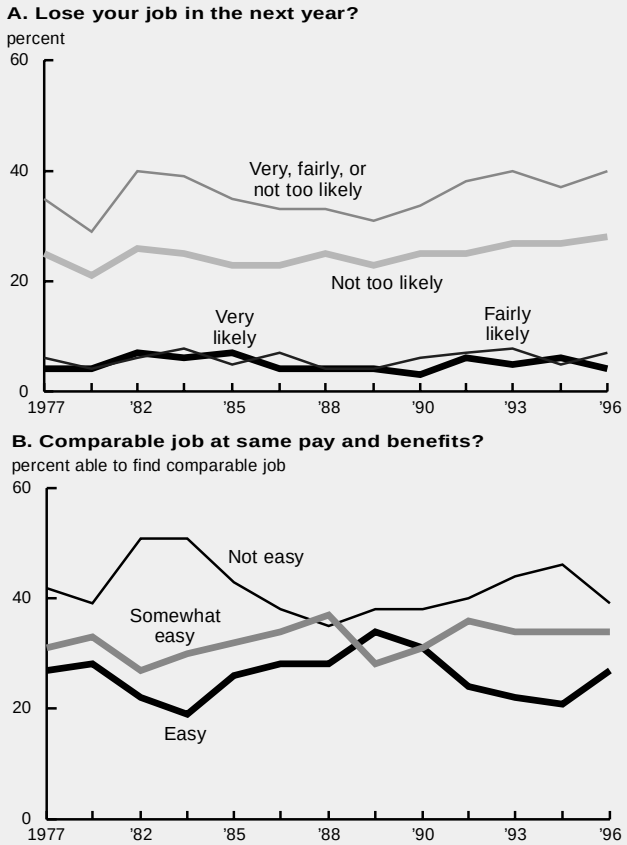
The GSS does have some important limitations.[20] Most noteworthy is that each GSS survey year consists of an independently drawn nationally representative sample of the population. Thus, unlike other national surveys such as the PSID, the GSS does not allow us to observe the same individuals across time. Surveys that follow individuals allow the use of panel data techniques to control for unmeasured individual-specific characteristics, such as ability or ambition, that change across the business cycle and are correlated with the other variables in the model. Such a survey format would allow us to investigate the future employment dynamics of workers and examine whether job anxiety predicts future job displacement or wage loss.

The easiest way to see how perceptions of job security have changed over time is to graphically examine the responses to the GSS questions. Figure 6, panels A and B show the distribution of responses to the two questions from 1977 to 1996. Each line represents a separate response except the highest line in panel A, which is the sum of the very, fairly, and not too likely responses. Between 30 percent and 40 percent of workers feel some degree of insecurity about losing their job in the next year, although only 10 percent of respondents feel very or fairly sure that job loss will occur. Between 35 percent and 50 percent of workers respond that it would not be easy to find a comparable job.

As with the displacement rates, the responses are fairly cyclical through the early 1990s. Using the job loss likelihood question, job security declined during recessions in the early 1980s and 1990s and increased during the expansion of the 1980s. But since 1991, the percentage of workers who answer that they are not at all likely to lose their job has fallen, despite the strong and widely felt expansion of the economy. Amazingly, in 1996, the fraction of workers who answered that they had some concern about their job's future

**FIGURE 6**

**Potential for job loss and comparable reemployment**




Source: Authors' calculations based on data from the National Opinion Research Center, *General Social Survey*, 1997.

was equal to the percentage that answered this way during the severe 1982–83 recession. However, most of this is due to an increase in the percentage of workers who answer they are not too likely to lose their job. Therefore, while there has been a noticeable shift in worker anxiety during this expansion, most of the change is due to workers acknowledging some, albeit a slight, likelihood of losing their job over the next year.

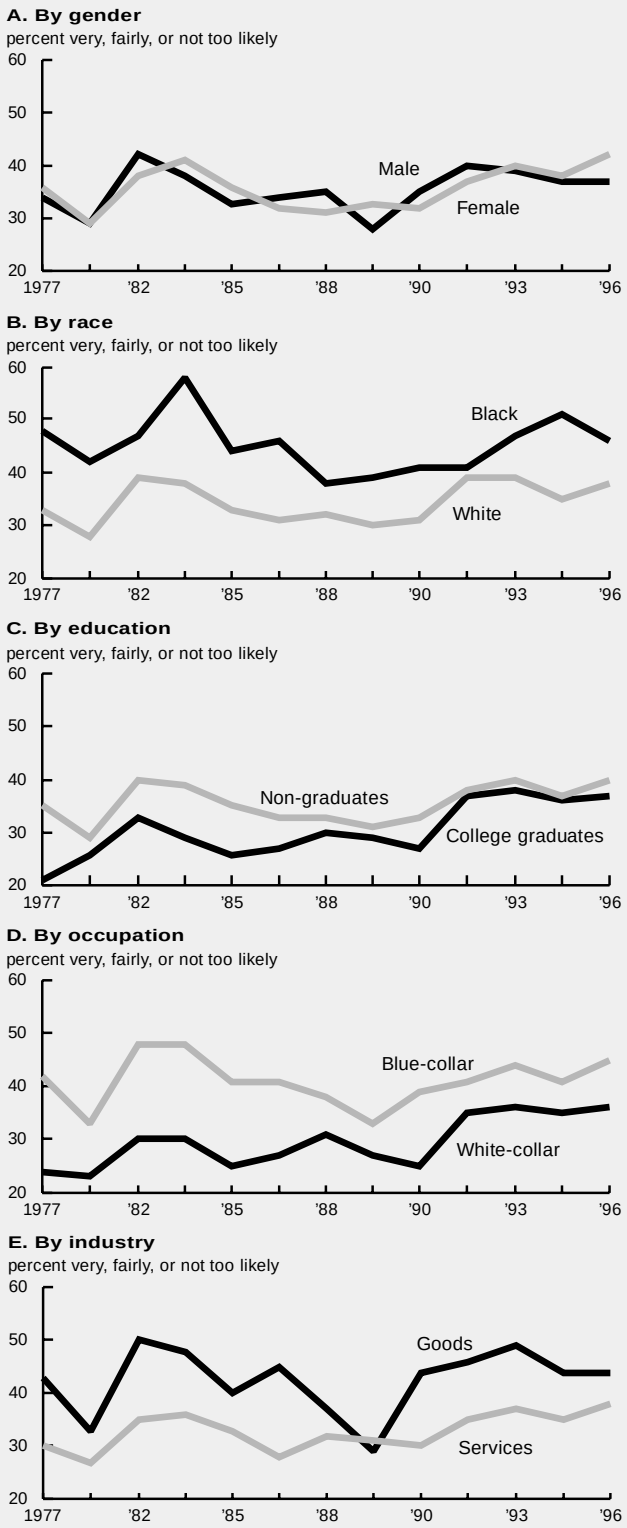
The job comparability question also tends to be cyclical, but showed signs of breaking this trend during the initial phase of the 1990s expansion. Beginning in 1988, the percentage of workers who answered that it would not be easy to find a comparable job at the same pay and benefits monotonically increased, peaking at almost 46 percent in the 1994 survey. However, in the 1996 survey, the "not easy to find a comparable job" response declined and the percentage answering it was easy to find a comparable job increased. Therefore, through 1996, workers seemed somewhat less concerned about their chances of finding a comparable job, but somewhat more concerned about the likelihood of losing their current job.[21]

Figures 7 and 8 show the trends in these two series by gender, race, education, industry, and occupation. The two primary differences between the GSS results and the displacement rate results are exhibited in figure 7, panels A and B. Panel A shows that there is no male–female gap in perceptions of job security throughout the sample period, while the displacement rates showed a large male–female gap that narrowed over time. Figure 7, panel B displays a large black–white gap in worker anxiety that has narrowed somewhat over time, while figure 4 showed no significant difference in displacement rates by race, except during the 1982 recession. We believe that a substantial portion of these differences may be due to the different tenure restrictions in the samples. That is, while the results on displaced workers come from a sample of workers with five years of tenure, we cannot make comparable restrictions on the GSS sample. The importance of this restriction is evident in other research. For example, Fairlie and Kletzer (1997) use the DWS to estimate displacement rate gaps between black and white workers but make no sample restrictions based on tenure. They find a 30 percent gap in displacement rates between the races from 1982 and 1991. Likewise, between 1982 and 1991, the black–white gap in the GSS job loss data is 29 percent.

On the other hand, figure 7, panels C and D look quite similar to the displacement rate results, pointing again to a democratization of job insecurity. White-collar and college-educated workers were relatively immune to job anxiety during the 1970s and 1980s, but have experienced substantial increases in job insecurity during the 1990s. The change has been large enough to basically eliminate the gap in job insecurity between college graduates and non-graduates. Blue-collar workers still feel less secure than white-collar workers, but the gap is less than half what it was in the 1970s and early 1980s.

**FIGURE 7**

**Likelihood of job loss, by demographic category**




Source: Authors' calculations based on data from the National Opinion Research Center, *General Social Survey*, 1997.

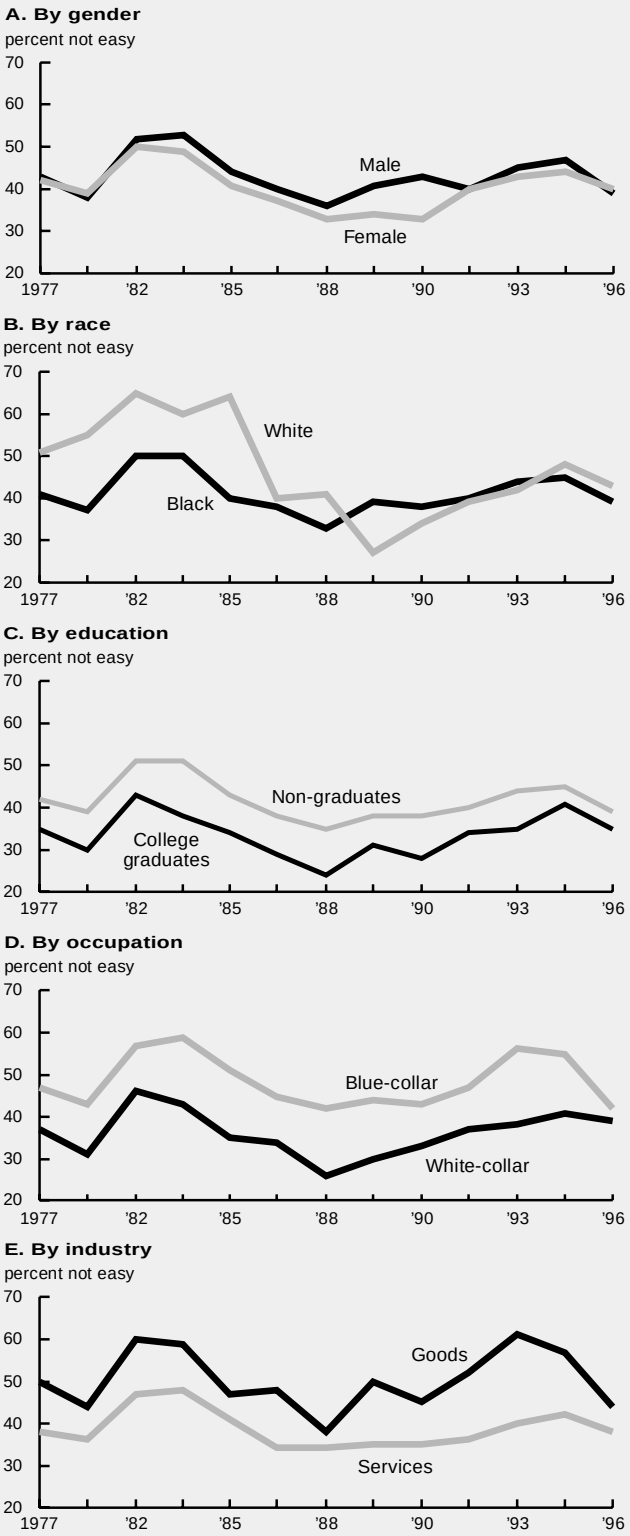
As shown in figure 7, panel E, job security has declined during the 1990s in the service sector but has remained relatively flat (other than a temporary drop in 1993) in the goods sector. Most of the decline in the service industry arises from the services, finance, insurance, and real estate (FIRE), and government sectors. Analogous to the displacement rate findings, perceptions of job security have dropped substantially in FIRE, with roughly 50 percent fewer workers saying that they are not at all likely to lose their job in the next year. The lack of movement among goods producing sectors hides some variance between specific industries. In particular, job insecurity (measured by the probability of losing your job) in manufacturing has doubled since 1989, surpassing the level of anxiety witnessed in 1982. In 1996, job insecurity in the manufacturing sector was substantially higher than in all other major industries. The goods sector has not increased because agriculture and construction workers have experienced corresponding declines in job insecurity over the past few years.

Figure 8, panel A shows the percentage of male and female workers who believe it is not easy to find a comparable job with the same pay and benefits. This graph shows a small but persistent male–female gap that is eliminated in 1996. Using this job security measure, most of the 1990s increase in anxiety appears to be due to female workers. Figure 8, panel B shows that a large black–white gap during the 1970s and early 1980s had disappeared by the end of the 1980s.

Figure 8, panels C and D again display the diminishing gap in worker anxiety between college-educated and non-college-educated workers and white- and blue-collar workers, respectively. Panel D shows that the historically large difference between white- and blue-collar employees all but vanished in 1996. White-collar workers are among the few groups that did not experience a sharp drop in anxiety

**FIGURE 8**

**Likelihood of finding comparable employment, by demographic category**




Source: Authors' calculations based on data from the National Opinion Research Center, *General Social Survey*, 1997.

about finding a comparable job in 1996, reflecting increased anxiety among professional workers. In 1996, 42 percent of professional workers responded that it was not easy to find a comparable job, up from 30 percent in 1989, matching the percentage that answered that way during the 1982 recession.

Finally, the 1996 drop in anxiety about finding a comparable job is mainly from the goods sector (figure 8, panel E). Anxiety about finding a comparable job for service sector employees peaked in 1994, but remains slightly above the levels seen during the last expansion. Nearly every group believed it would be easier to find a comparable job in 1996 than in 1994, exceptions being government employees and professional and sales workers.

## Controlling for population characteristics

The results presented thus far are based on raw data. However, other changes in the work force during the last 20 years, including shifts in the age and educational distribution of U.S. workers, may be confounding the time trends in worker anxiety. Is the trend in worker anxiety by industry, occupation, or education the same even after simultaneously controlling for multiple characteristics of the population? To find out, we estimate "ordered probit" regressions, an appropriate statistical technique for this problem because it accounts for the discrete and ordered nature of the job security questions. The details of the estimation procedure are described in box 2.

Table 3 reports the coefficients, standard errors, and marginal effects from a specification that uses the likelihood of losing your job as the dependent variable and industry, occupation, year, gender, race, age, marital status, education, and region dummies as controls.[22] The marginal effects measure the impact of a change in some variable, say whether the individual is a sales worker, on the probability

**BOX 2**

## Ordered probit regressions

The ordered probit model is based on a latent regression such as

$$1) \qquad y_i^* = \beta x_i + \varepsilon_i,$$

where $y_i^*$ is the unobserved job insecurity of person $i$, $x_i$ are demographic and other individual characteristics of person $i$ and $\varepsilon_i$ is a person-specific error term. The parameter $\beta$ is a vector of coefficients that measure the average impact of the demographic variables on the level of job security. While we do not observe $y_i^*$, we do observe the $k$ possible answers allowed by the survey, as represented by $y_i$:

$$\begin{aligned}
&y_i = 0 \quad \text{if} \quad y_i^* \le 0 \\
&y_i = 1 \quad \text{if} \quad 0 \le y_i^* < \mu_1 \\
&y_i = 2 \quad \text{if} \quad \mu_1 \le y_i^* < \mu_2 \\
&\vdots \\
&y_i = k \quad \text{if} \quad \mu_{k-1} \le y_i^*.
\end{aligned}$$

For example, in the GSS likelihood of losing your job question, $y_i = 0$ corresponds to answering "not at all likely to lose my job," while $y_i = 3$ corresponds to the "very likely to lose my job" answer. The $\mu_i$'s are unknown intercept parameters to be estimated in the model.

Assuming a normal distribution in the error term, we can calculate the probability of each of the $k$ answers as

$$2) \qquad \text{Prob}(y = j|x) = \begin{cases} \Phi(\mu_0 + \beta x) & \text{if } j = 0 \\ \Phi(\mu_j + \beta x) - \Phi(\mu_{j-1} + \beta x) & \text{if } 0 < j \le k-1, \\ 1 - \Phi(\mu_{k-1} + \beta x) & \text{if } j = k \end{cases}$$

where $\Phi$ is the standard normal cumulative distribution function. From equation 2, we can calculate the marginal effect (the impact of a change in the $x$ variable on the probability of event $y$ occurring) by

$$3) \qquad \begin{aligned}
&\frac{\partial \text{prob}(y=0)}{\partial x} = \phi(\mu_0 + \beta x)\beta \\
&\frac{\partial \text{prob}(y=j)}{\partial x} = (\phi(\mu_j + \beta x) - \phi(\mu_{j-1} + \beta x))\beta \\
&\frac{\partial \text{prob}(y=k)}{\partial x} = (1 - \phi(\mu_{k-1} + \beta x))\beta,
\end{aligned}$$

where $\phi$ is the standard normal density function and the $x$ variables are measured at their mean value. Equation 3 essentially calculates the effects of changes in the covariates on the cell probabilities.

It should be noted that many of the independent variables in our models are 0–1 indicators, such as whether the individual is a college graduate. In this case, the marginal effect is calculated as the difference between the cell probabilities when the event occurs (a college graduate) and when the event does not occur (not a college graduate):

$$\text{Prob}(y = j|x',1) - \text{Prob}(y = j|x',0),$$

where $x',1$ is the vector of covariates where the college graduate variable is set to 1 and $x',0$ is the vector of covariates where the college graduate variable is set to 0.

**TABLE 3**

## Likelihood of losing your current job in the next year: Ordered probit analysis

| | Coefficient | Standard error | Marginal effect on base case probability: Not at all likely | Not too likely | Fairly likely | Very likely |
|---|---|---|---|---|---|---|
| Base case probability[a] | | | 0.697 | 0.225 | 0.046 | 0.032 |
| Agriculture | 0.069 | 0.097 | 0.024 | –0.014 | –0.005 | –0.005 |
| Construction | –0.322 | 0.070* | –0.120 | 0.062 | 0.027 | 0.031 |
| Manufacturing | –0.234 | 0.054* | –0.086 | 0.046 | 0.019 | 0.021 |
| Transportation, communications, and utilities | –0.058 | 0.066 | –0.021 | 0.012 | 0.004 | 0.004 |
| Wholesale trade | –0.021 | 0.084 | –0.007 | 0.004 | 0.002 | 0.002 |
| Retail trade | –0.128 | 0.057* | –0.046 | 0.026 | 0.010 | 0.010 |
| Finance, insurance, and real estate | 0.060 | 0.069 | 0.021 | –0.012 | –0.004 | –0.004 |
| Services | –0.079 | 0.050 | –0.028 | 0.016 | 0.006 | 0.006 |
| Professional, technical | 0.148 | 0.047* | 0.049 | –0.030 | –0.010 | –0.009 |
| Managerial | 0.264 | 0.049* | 0.085 | –0.053 | –0.017 | –0.015 |
| Sales | 0.119 | 0.056* | 0.040 | –0.024 | –0.008 | –0.008 |
| Craftsman | –0.014 | 0.052 | –0.005 | 0.003 | 0.000 | 0.000 |
| Operative or laborer | –0.165 | 0.047* | –0.060 | 0.033 | 0.013 | 0.014 |
| Service worker | 0.126 | 0.048* | 0.042 | –0.026 | –0.009 | –0.008 |
| 1978 | 0.125 | 0.062* | 0.042 | –0.026 | –0.009 | –0.008 |
| 1982 | –0.194 | 0.061* | –0.071 | 0.039 | 0.016 | 0.017 |
| 1983 | –0.236 | 0.060* | –0.087 | 0.046 | 0.019 | 0.021 |
| 1985 | –0.089 | 0.060 | –0.032 | 0.018 | 0.007 | 0.007 |
| 1986 | –0.018 | 0.062 | –0.006 | 0.004 | 0.001 | 0.001 |
| 1988 | 0.034 | 0.069 | 0.012 | –0.007 | –0.002 | –0.002 |
| 1989 | 0.062 | 0.070 | 0.021 | –0.013 | –0.004 | –0.004 |
| 1990 | 0.022 | 0.070 | 0.008 | –0.004 | –0.002 | –0.002 |
| 1991 | –0.147 | 0.067* | –0.053 | 0.029 | 0.012 | 0.012 |
| 1993 | –0.184 | 0.066* | –0.067 | 0.037 | 0.015 | 0.016 |
| 1994 | –0.121 | 0.057* | –0.044 | 0.024 | 0.009 | 0.010 |
| 1996 | –0.151 | 0.056* | –0.055 | 0.030 | 0.012 | 0.013 |
| Female | –0.035 | 0.029 | –0.012 | 0.007 | 0.003 | 0.003 |
| Black | –0.265 | 0.039* | –0.098 | 0.052 | 0.022 | 0.024 |
| Other race | –0.095 | 0.070 | –0.034 | 0.019 | 0.007 | 0.007 |
| Age 18–24 | 0.031 | 0.043 | 0.011 | –0.006 | –0.002 | –0.002 |
| Age 44–65 | 0.130 | 0.029* | 0.044 | –0.027 | –0.009 | –0.008 |
| Never married | –0.111 | 0.034* | –0.040 | 0.022 | 0.009 | 0.009 |
| Divorced | –0.024 | 0.037 | –0.009 | 0.005 | 0.002 | 0.002 |
| Separated | –0.198 | 0.064* | –0.072 | 0.039 | 0.016 | 0.017 |
| Widowed | –0.181 | 0.076* | –0.066 | 0.036 | 0.014 | 0.015 |
| High school dropout | –0.124 | 0.038* | –0.044 | 0.025 | 0.010 | 0.010 |
| College graduate | 0.024 | 0.038 | 0.008 | –0.005 | –0.002 | –0.002 |
| Graduate school graduate | 0.055 | 0.056 | 0.019 | –0.011 | –0.004 | –0.004 |
| New England | 0.120 | 0.064 | 0.040 | –0.024 | –0.008 | –0.008 |
| Mid-Atlantic | 0.008 | 0.046 | 0.003 | –0.002 | –0.000 | –0.000 |
| East North Central | 0.093 | 0.044* | 0.032 | –0.019 | –0.006 | –0.006 |
| East South Central | 0.040 | 0.059 | 0.014 | –0.008 | –0.003 | –0.003 |
| South Atlantic | 0.045 | 0.044 | 0.016 | –0.009 | –0.003 | –0.003 |
| West North Central | 0.077 | 0.055 | 0.026 | –0.016 | –0.005 | –0.005 |
| West South Central | –0.032 | 0.052 | –0.011 | 0.006 | 0.002 | 0.002 |
| Mountain | –0.116 | 0.058* | –0.042 | 0.023 | 0.009 | 0.009 |
| Intercept 1[b] | 0.517 | 0.077* | | | | |
| Intercept 2[b] | 1.420 | 0.078 | | | | |
| Intercept 3[b] | 1.849 | 0.079* | | | | |
| Log likelihood | | –18,316 | | | | |
| Sample size | | 9,935 | | | | |

* = significant at the 5% level.

[a]Base case is a white, married male, aged 25 to 44, high school graduate, who worked a clerical government job in the Pacific region in 1977. Industry, occupation, region, and year dummies are relative to government (industry), clerical (occupation), Pacific (region), and 1977 (years).

[b]Each response has its own intercept. See box 2 for details. The three intercept terms are used to compute the marginal effects for the four categories of responses (final four columns).

Note: Dependent variable is the likelihood of losing your job in the next year. The possible answers are not at all likely, not too likely, fairly likely, and very likely.

Source: Authors' calculations based on data from the National Opinion Research Center, *General Social Survey*, various years.

of the individual responding to the job security questions in a particular way. The results are reported relative to a base case white, married, male, high school graduate, aged 25 to 44, who worked in a clerical job in the government in 1977. The first row shows that the probability of the base case person responding that he is not at all likely to lose his job is 69.7 percent. The third row, Construction, reveals that the probability of a not at all likely response from a clerical worker in the construction industry is 12.0 percentage points lower (or 57.7 percent) than for a clerical worker in the government in 1977.

Overall, the table shows that many of the characteristics that look significant in the univariate graphs, such as occupation and race, remain significant indicators, even after controlling for the demographic and employment variables. This is also true of ordered probit regressions where the comparable job question is the dependent variable. Table 3 gives further detail on the specific industry and occupational groups that traditionally experience higher levels of job anxiety, including workers in the construction and manufacturing sectors and operatives and laborers in all industries. While the probability of managerial workers responding that they are not at all likely to lose their job is 78.2 percent, the same probability for an operative or laborer is 63.7 percent. These industry and occupational differences are statistically significant.

We can test whether job security has changed over this expansion by calculating time trend effects within the ordered probit framework. The rows labeled 1978 to 1996 in table 3 show the results of such an exercise. As with the simple univariate graphs, perceptions of job security have been quite low since 1991 when measured by the likelihood of job loss. Controlling for demographic, industry, and occupation shifts cannot explain the recent high insecurity felt by workers. Job anxiety remains on the order of that seen during the last two recessions.

Also, we stratified the sample by gender, race, education, occupation, and industry and ran separate ordered probit regressions for different categories of workers. The purpose of this exercise is to see whether the time trends reported in the graphs still exist after controlling for other demographic, industrial, and occupational structural shifts. By running separate regressions for each demographic group, we allow the parameters on other covariates to change across groups. This flexible specification allows, say, the effect of being married to exert a different influence on perceptions of job security for high school dropouts and college graduates. However, the main inferences from these results (not shown) do not change much. The recent trend in increased job anxiety arises primarily from better educated and white-collar workers. On the other hand, workers who are high school dropouts are more secure about their job in 1996 than at any other time since 1977, with the exception of 1989, the end of the 1980s expansion. Managerial and professional workers have witnessed increases in job insecurity, while there is no statistical trend apparent in other detailed occupations. Increased anxiety appears in manufacturing, services, and government, while construction workers, who have traditionally had a high probability of job loss because of the seasonal nature of the work, have seen an increase in job security during the 1990s.

Lastly, we looked at the perceptions of job loss among a few nonstandard groups of workers. Table 4 reports the coefficients, standard errors, and marginal effects from additional variables that are asked (sometimes periodically) in the GSS or are computed from other data sources. The first group of variables is other work characteristics, including union membership, the size of the employee's work site, whether the firm pays fringe benefits, whether the organization has gone through a merger or reorganization during the last five years, computer usage in the industry, and employment conditions in the industry and region. The second group of variables lists several hardships that individuals have recently experienced, including poverty, unemployment, work problems, financial problems, other hardships, and health problems. The letter at the beginning of each row indicates a separate regression that includes all of the demographic and employment variables reported in table 3. The sample sizes from each of these regressions are reported in column 1. Since some questions begin to be asked after 1977, all marginal effects are calculated relative to the same base case person in the first year that the question is included in the GSS.

The first row shows that union members are likely to be more insecure about their future job prospects than nonunion members even after controlling for compositional differences in occupation and industry between the groups. This finding may be confounded by the choice to join

a union (workers who are more insecure about their future employment are more likely to join unions) and therefore suffers from what econometricians call endogeneity bias. It is a bit surprising, because much of the research on unions suggests that union workers are less sensitive to business cycles because wages and employment are set in multiyear contracts. However, union wages have been growing slower than nonunion wages recently, suggesting that workers are concerned enough about job security that they are willing to trade off wage growth for more security. Furthermore, the decline in union membership over the last few decades could signal reduced bargaining power of union employees.

In a way, the union results are similar to some limited evidence on workers in organizations undergoing change. In 1991, the GSS included information on whether a respondent's firm has gone through a merger or a reorganization. Because we only have one year of data, the precision of the point estimates is low. Nevertheless, the magnitude of the marginal effects is consistent with stories about restructuring and downsizing leading to more insecurity during the 1990s. Unfortunately, because the sample is redrawn each year, we cannot test whether these workers have faced greater job loss frequencies in subsequent years.

In regressions that add the size of the employee's work site, the results suggest that those who work at smaller sites are less likely to be concerned about their job. However, the question asks the size of the work site not the size of the organization.

**TABLE 4**

**Effect of other variables on likelihood of losing your current job in the next year: Ordered probit analysis**

| Other variables | Sample size | Coefficient | Standard error | Marginal effect on base case probability: Not at all likely | Not too likely | Fairly likely | Very likely |
|---|---|---|---|---|---|---|---|
| **Work characteristics** | | | | | | | |
| a. Union member | 4,761 | –0.124 | 0.049* | –0.040 | 0.025 | 0.008 | 0.007 |
| b. Current organization merged | 550 | –0.209 | 0.132 | –0.080 | 0.039 | 0.019 | 0.022 |
| c. Current organization reorganized | 552 | –0.123 | 0.120 | –0.046 | 0.024 | 0.011 | 0.012 |
| d. No fringe benefits | 463 | –0.491 | 0.196* | –0.193 | 0.077 | 0.042 | 0.074 |
| Size of work site | | | | | | | |
| e. 1–9 employees | 3,079 | 0.158 | 0.079* | 0.060 | –0.032 | –0.013 | –0.015 |
| e. 10–49 employees | 3,079 | 0.107 | 0.078 | 0.041 | 0.021 | –0.009 | –0.011 |
| e. 100–499 employees | 3,079 | 0.076 | 0.080 | 0.029 | –0.015 | –0.006 | –0.008 |
| e. 500+ employees | 3,079 | 0.031 | 0.081 | 0.012 | –0.006 | –0.003 | –0.003 |
| f. Region unemployment rate[a] | 9,935 | –0.086 | 0.019* | –0.026 | 0.016 | 0.005 | 0.005 |
| g. Industry unemployment rate[a] | 9,935 | –0.029 | 0.020 | –0.010 | 0.006 | 0.002 | 0.002 |
| h. Industry computer use[b] | 9,529 | –0.193 | 0.078* | –0.070 | 0.039 | 0.015 | 0.015 |
| **Work and other problems** | | | | | | | |
| i. Below the poverty line | 7,620 | –0.352 | 0.051* | –0.132 | 0.066 | 0.031 | 0.035 |
| j. Unemployment spell in last 5 yrs. | 3,946 | –0.448 | 0.045* | –0.135 | 0.090 | 0.026 | 0.020 |
| k. Unemployment spell in last 10 yrs. | 5,752 | –0.403 | 0.035* | –0.135 | 0.083 | 0.029 | 0.023 |
| l. Problems at work | 281 | 0.004 | 0.184 | 0.001 | –0.000 | –0.000 | –0.000 |
| l. Financial problems | 281 | –0.124 | 0.216 | –0.043 | 0.028 | 0.010 | 0.006 |
| l. Other hardships | 281 | –0.535 | 0.219* | –0.201 | 0.110 | 0.052 | 0.039 |
| m. Not healthy | 5,292 | –0.229 | 0.047* | –0.087 | 0.043 | 0.020 | 0.025 |

* = significant at the 5% level.

[a]Industry and regional unemployment rates are calculated from the March 1977–96 *Current Population Survey*.

[b]Industry computer use is from Autor, Katz, and Krueger (1997). They calculate the share of computer users by three-digit Standard Industrial Classification codes from the October 1984, 1989, and 1993 *Current Population Survey*. The computer use data are linearly interpolated between 1984 and 1993, set at 1984 levels in year prior to 1984, and at 1993 levels in years post 1983.

Notes: The letter at the beginning of each row indicates a separate regression that includes all of the control variables listed in table 3. The base case is the same as table 3. If the variables listed in this table were not reported in 1977, the base case is the first year the question was asked. Except for the unemployment rates and the computer usage variable, the marginal effects are reported as the difference between the base case where the characteristic is not present (say, the respondent is not a union member) and where the characteristic is present (is a union member). The size of work site coefficients are relative to a company with 50–99 employees. See box 2. The base case probabilities, which are not reported, are available upon request.

Source: Authors' calculations based on data from the National Opinion Research Center, *General Social Survey*, 1997.

When we add industry and regional unemployment rates to the statistical model, we find, not surprisingly, that workers in regions and, to a much smaller extent, industries that are experiencing higher unemployment are less secure about their own prospects. (The unemployment rates are calculated from the March CPS.)

Finally, we add a variable that measures the share of computer users in each individual's three-digit SIC industry. The data are compiled by Autor, Katz, and Krueger (1997) using the October 1984, 1989, and 1993 CPS. As part of the *Education Supplements*, the three CPS surveys asked workers whether they used a computer at work, where a computer is defined as a desktop terminal or PC and not a hand-held data device or electronic cash register. We interpolate computer shares by industry between the 1984 and 1993 end dates and hold years before 1984 and after 1993 constant at the 1984 and 1993 levels. Surprisingly, the results suggest that workers in industries that are more computer intensive are less secure about their jobs, after controlling for demographics, time, industry, and occupation.[23] When the computer usage variable is interacted with the time dummies, it becomes apparent that this computer industry–job insecurity correlation is driven by the 1993 to 1996 period. Prior to the 1990s, there is a positive relationship between working in a computer-intensive industry and job security. Unfortunately, we have no data on whether the individual respondents are computer users.

The bottom of table 4 reports the parameters from the "problem" variables. Many of these coefficients are significant and negative, suggesting that job insecurity goes hand-in-hand with other work- and non-work-related problems. Furthermore, the large effects from previous unemployment suggest that these workers are more prone to insecurity than those who have not experienced a spell of unemployment in the last ten years. This result suggests that past job loss may be a reasonable indicator of future anxiety and is consistent with studies that use displacement rates as an indicator of job security. However, the results on past unemployment may be driven by unobserved characteristics, such as ability.

## Job security and wage growth

While a number of papers have measured recent trends in job security or stability, none that we are aware of attempt to link these trends to wage growth. Yet the allegedly slow rise in compensation during this tight labor market expansion is one of the driving forces behind public policy concerns about job security. Many analysts argue that workers have sacrificed wage growth for a more secure relationship with their current employers.

To investigate this question, we follow an estimation strategy pursued in Blanchflower and Oswald (1994), among others. In particular, we look cross-sectionally as well as over time and ask whether census regions that have had higher displacement rates or worker perceptions of insecurity have tended to have lower wage growth. The regressions that we use are similar to the original Phillips curve, which posits a negative relationship between the rate of wage change and the contemporaneous unemployment rate.[24]

We use three wage measures, annual, weekly, and hourly earnings, that are computed from the 1977 to 1996 March CPS. However, our preferred wage measure is hourly earnings, because it does not confound changes in wages with changes in hours worked. This is an important distinction because annual hours are highly correlated with the job security measures (as well as unemployment rates). Therefore, we have to be careful to distinguish a wage effect from a labor supply (hours) effect. This is probably less of a concern with the weekly measure.

We include controls for one of two job security measures (a security index calculated from the GSS and a displacement rate calculated from the DWS), the contemporaneous unemployment rate (calculated from the March CPS), and time- and location-specific indicator variables. The time- and location-specific variables account for unexplained characteristics of wages that are common across time and regions (essentially, they allow the intercept term to vary over time and region). For example, the time variables will account for changes in productivity growth and expectations of inflation that are common across regions within the U.S. Box 3 gives the technical details of our estimation procedure.

Table 5 highlights some of our findings. Panels A and B are from separate regressions. The security measure is the insecurity index in panel A, and the log displacement rate in panel B.

The wage effect is reasonably consistent across the two job security measures. The coefficients (since the variables are measured in logs, the coefficients are elasticities) suggest that, using the annual earnings measure, a 10 percent increase

**BOX 3**

## The impact of job security on wage growth

The wage-job security relationship is estimated from a regression of the form:

$$4) \qquad w_{rt} = \alpha y_{rt} + \lambda U_{rt} + \varphi w_{rt-1} + v_r + v_t + v_{rt},$$

where $w_{rt}$ is the log wage, $y_{rt}$ is a measure of the level of job security, and $U_{rt}$ is the log unemployment rate. The variables are aggregated into a market $r$ at time $t$. We aggregate individuals into the nine census regions since geographical labor markets smaller than regions are not available in the GSS. The wage and unemployment measures are computed from the 1977 to 1996 March *Current Population Surveys* (CPS). The annual earnings measure is a sum of all income earned in the previous year. The weekly earnings measure is calculated as annual earnings divided by the number of weeks worked in the previous year. The hourly earnings measure is calculated as annual earnings divided by the number of weeks worked in the previous year times the number of hours worked per week in the previous year.

There are several ways to estimate equation 4. Perhaps the simplest way is to average all individuals in market $r$ at time $t$ and use the cell means as the observation unit. This is essentially what was done for the displacement rates (see box 1). However, the standard errors will be biased downward because common unmeasured factors of individuals may be attributed to local employment conditions (Moulton, 1990). Instead, for the job loss likelihood index, we use a "two-step" procedure. In the first step, we estimate ordered probit regressions like equation 1 in box 2 but augment them to allow a calculation of a region-specific security index for each year. In particular, we regress the loss likelihood responses, $y_i^*$, on demographics ($X_{it}$), and year interacted with region dummies ($R_{it}v_t$):

$$y_{it}^* = \beta X_{it} + \delta R_{it} v_t + \varepsilon_{it}.$$

The vector of region–year dummy coefficients ($\delta$) is equivalent to the mean residuals by year and region and can be interpreted as indexes of job insecurity, after controlling for differences in education, gender, age, income, and marital status of workers in particular areas. We also run the final wage equations with job security indexes that are not demographically adjusted and find that this adjustment does not make a significant difference to the inferences.

The wage variables are estimated from a log wage equation of the form:

$$\ln w_{irt} = \beta X_{irt} + \mu_{irt}.$$

Again, the $X$ matrix controls for education, marital status, and other standard human capital controls. We use the mean residual by region and year ($\bar{\mu}_{rt}$) as a measure of the wage adjusted for these demographics.

In the second step, we estimate ordinary least squares regressions of the mean residual from the first stage wage equation ($\bar{\mu}_{rt}$) on the contemporaneous unemployment rate ($U_{rt}$), the lagged dependent variable, the security index ($\delta_{rt}$), and region and year indicator variables (or fixed effects):

$$\bar{\mu}_{rt} = \alpha\delta_{rt} + \lambda U_{rt} + \varphi\bar{\mu}_{rt-1} + v_r + v_t + v_{rt}.$$

The sample size varies depending on the job security index that is used. The regressions that include the displacement rate are run on 17 years (17 years times nine regions equals 153 observations) and the regressions that include the security index are run on 13 years (117 observations).

in the job security measure results in a 0.2 percent decline in wage growth. This is statistically significant at conventional levels. However, the hourly wage coefficient implies that about half of this decrease is a wage effect and the other half is an hours-worked effect. Furthermore, the wage effect is imprecisely enough estimated that we cannot reject the hypothesis that the true effect is zero. However, if this effect is real, the impact on nominal wage growth during the 1990s has been fairly large. Referring to figure 3, displacement rates rose from around 2.0 percent in the 1980s

**TABLE 5**

## Relationship between wages and job security

| | Change in log annual earnings | | Change in log weekly earnings | | Change in log hourly earnings | |
|---|---|---|---|---|---|---|
| | No region controls | Region controls | No region controls | Region controls | No region controls | Region controls |
| **A. (sample size = 117)** | | | | | | |
| Insecurity index | –0.018 | –0.010 | –0.013 | –0.007 | –0.013 | –0.008 |
| | (0.012) | (0.013) | (0.012) | (0.013) | (0.010) | (0.011) |
| Log unemployment rate | –0.029 | –0.048 | –0.029 | –0.045 | –0.024 | –0.040 |
| | (0.012) | (0.015) | (0.012) | (0.015) | (0.010) | (0.013) |
| **B. (sample size = 153)** | | | | | | |
| Log displacement rate | –0.021 | –0.018 | –0.020 | –0.017 | –0.010 | –0.008 |
| | (0.008) | (0.008) | (0.008) | (0.008) | (0.007) | (0.007) |
| Log unemployment rate | –0.028 | –0.052 | –0.028 | –0.051 | –0.026 | –0.044 |
| | (0.009) | (0.011) | (0.009) | (0.012) | (0.008) | (0.011) |

Notes: All regressions include year controls. See text for explanation of variables and sample. The unit of observation is the nine census regions.

Sources: Authors' calculations based on data from the Bureau of Labor Statistics, *Current Population Survey*, 1996; the Bureau of Labor Statistics, *Displaced Worker Survey*, 1984–96; and the National Opinion Research Center, *General Social Survey*, 1997.

to 2.75 percent in the early 1990s to almost 3.5 percent in 1995. Using a job insecurity wage elasticity estimate of –0.01, this suggests that job insecurity lowered wage growth by 0.3 percentage points a year in the early 1990s and roughly 0.7 percentage points in 1995, relative to what would have happened if displacement rates had stayed at the 1980s level. The job anxiety index grew approximately 25 percent during the 1990s, suggesting a 0.3 percentage point decline in wages per year from the results in table 5, panel A. However, our estimates for hourly wages cannot reject the possibility that these effects arose purely by chance.

Our analysis is just a first step in estimating the impact of job security on wage inflation. There is much more work to be done on this question. First, we plan to explore micro-data-based techniques to solve technical problems associated with having two measures, such as wage growth and job security, that are jointly determined. Second, as it is currently measured, the security index encompasses a fair amount of noise or measurement error. This measurement error leads to a downward bias in the wage–security relationship. Finally, a key question is causation. Does high job security cause high wages or vice versa? This question could be examined by estimating vector autoregressive models, which allow a flexible relationship between wages, unemployment, and job security.

## Conclusion

Our review of the literature and our new results on displacement for high-tenure workers reveal a modest decline in job stability and a larger decline in job security, especially for workers with higher levels of job tenure. Apparently, some of the increases in displacement that have been observed in the 1990s have been offset by declines in quit rates. The higher displacement rates suggest that workers have more reason to be worried about their job security in the 1990s, and the lower quit rates suggest they may be less confident about their job prospects. Consistent with these findings, our tabulations of workers' evaluations of their chances of job loss reveal a noticeable increase in the proportion of workers who feel that they are at least at some risk of job loss.

When we relate variations in displacement rates and anxiety levels over time and across census divisions to the corresponding variations in wage growth, we find estimates of the effect of insecurity on wages that would be large enough to explain all or most of the puzzle of slow wage growth in the 1990s. Of course, these estimates are rather imprecise and may even have arisen by chance. Still, we believe that these results add to the case for worker insecurity having restrained wage growth and justify further research on the topic.

## NOTES

[1]Greenspan (1997).

[2]Reich (1997).

[3]See, for example, Staiger, Stock, and Watson (1997).

[4]The manner in which workers and employers are matched to each other has changed quite noticeably in the 1990s. The process may have been made more efficient by the rapid expansion of the temporary services industry. (See Segal and Sullivan, 1995, 1997). Also, Internet job postings may make interregional job search more efficient. Such developments may reduce the likelihood of bottlenecks and spot labor shortages that contribute to inflationary pressures. Or, in the language of the short-run Phillips curve, we can argue that they have reduced the natural rate of unemployment independently of any increase in worker anxiety.

[5]*New York Times* (1996).

[6]Notable contributions to this literature include Podursky and Swaim (1987), Kletzer (1989), Topel (1990), Ruhm (1991), and Jacobson et al. (1993a, 1993b, 1993c). Fallick (1996) and Kletzer (1997) provide recent surveys of this literature.

[7]The CPS is a monthly mini-census of about 45,000 households that is the source for such familiar statistics as the unemployment rate. When appropriately weighted to account for the scientifically designed sampling procedures employed by the BLS, the CPS yields nationally representative estimates.

[8]The figures are taken from U.S. Department of Labor (1997) and Farber (1998).

[9]Tenure data were also collected as part of the CPS *Pension and Benefits Supplements* of May 1979 and April 1993. The latter, which found higher median tenure for most age and sex groups than the *Mobility Supplement* of January 1991, supported the conclusion of Farber (1998) that job durations were relatively stable in the 1980s and 1990s. However, the tenure data from the *Pension and Benefits Supplements* are based on a slightly different question than those from the *Mobility Supplements*. The latter asks how long workers have been *continuously* employed by their current employers, while the former simply asks how long workers have been employed by their current employers. Omitting the condition that the employment be continuous could raise the tenure estimates. Suppose a worker was employed by a firm for five years, left for two years, and then returned for another five. In the *Mobility Supplements*, in which the question refers to continuous employment, the worker is likely to report a tenure of five years. However, in the *Pension and Benefits Supplements*, in which the question simply asks workers how long they have worked for their employers, the worker is likely to report a tenure of ten years. Thus, it is possible that some or all of the higher tenure reading in the 1993 survey was due to the omission of the word continuous in the key question rather than an actual increase in job stability.

Changes in question wording also complicate the interpretation of the trends in figure 2. Before 1983, the *Mobility Supplements* asked workers when they started working for their current employers. Tenure was then calculated based on workers' responses. Since 1983, workers have been asked how long they have continuously worked for their current employers, which yields the tenure information directly. Of course, if workers correctly answered all questions, it would make no difference whether tenure was solicited directly or calculated from the start date of their jobs. But workers do not always report accurately; figure 1 shows that workers have a tendency to report tenures that are multiples of five years. In the earlier *Mobility Supplements*, there was a tendency to report start years that were multiples of five, such as 1960, 1965, and so on. This change compromises the comparability of the data over time.

[10]Hall (1982) studied retention probabilities using data from a single cohort which, as shown by Ureta (1992), requires a stable rate of job beginnings, as well as a stable set of retention probabilities.

[11]Neumark, Polsky, and Hansen (1997) report several alternative estimates of retention probabilities. Those shown in table 1 are, we believe, their preferred estimates.

[12]The lack of a major decline in job stability is also consistent with the work of Stewart (1997), who analyzed the March CPS annual demographic files and found no increase in the rate of job change from the previous calendar year.

[13]See Hipple (1997).

[14]Of course, since this is a survey of individuals, there may be instances in which respondents misreport by saying, for example, that they were displaced when they were actually fired. Such mismeasurement is possible with any household survey.

[15]The PSID has a number of advantages for studying turnover and displacement. Unfortunately, the sample size is too small to estimate disaggregated rates. Thus, Farber computes a single set of adjustment factors that he applies to all workers.

[16]Tabulations, such as those in Hipple (1997), that do not count workers displaced for nonstandard reasons do not show an increase in the current period comparable to what we find below.

[17]Our procedure yields what we believe to be consistent comparisons across years. It is possible, however, that the overall level of the estimates could be off by some constant percentage. Suppose, for example, that the reason the displacement rates decline by 11 percent for each additional year that the survey lags the year of displacement is not that the rates measured with a one-year lag are correct and the other years reflect forgetting, but that the rates measured five years later are correct and earlier surveys reflect "spurious remembering." (In our opinion, this is a much less likely scenario but one that we obviously can't rule out.) Then our estimates will all be too high by about 52 percent (11 percent compounded for four years), but the pattern across years will be unaffected.

[18]Some observers (for example, Neumark and Polsky, 1997) argue that attitudinal questions about job security may not provide convincing evidence of actual job loss if perceptions are formed from misinformation. They point out that much of the reporting on job security relies on anecdotal evidence and, therefore, is not based on random sampling. It is always possible to find someone who is struggling, even in a booming economy. During the early 1990s, the recession hit journalists and editors, as well as other white-collar workers especially hard, perhaps resulting in more stories about displacement than were warranted. Since press reports may help form perceptions of the chance of job loss among readers, there is the danger that we might observe an increase in perceptions of job insecurity that has little to do with actual job loss.

[19]An exception is Schmidt and Thompson (1997). Several polling agencies, such as Gallup and Yankelovich, and survey organizations, such as the University of Michigan Survey Research Center (SRC), have also been soliciting perceptions of worker security over the last two decades. See Otoo (1997) for an analysis of the SRC data. Dominitz and Manski (1997) describe the new *Survey of Economic Expectations*, which asks respondents their level of concern about losing their job, losing part of their income, losing their health insurance, and being victimized by a burglary. However, this survey began in 1994 and therefore provides no information on longer-term trends in job security.

[20]See Dominitz and Manski (1997) for a criticism of the wording of the GSS questions.

[21]The Gallup poll also found that the number of workers answering "not at all likely" to lose their job decreased during the 1990s. On the other hand, Yankelovich, which asks whether losing your job worries you, found little change in response between 1992 and 1995. But both polls found some job security differences across the 1990s by education or income. Likewise, Otoo (1997) found significant increases in job anxiety between 1988 and 1995 using the SRC micro data.

[22]Additional results, including an analogous table for the comparable job question, are available from the authors upon request.

[23]With no such controls, the correlation between industry computer usage and job security is positive and highly significant.

[24]Blanchflower and Oswald estimate models that relate the *level* of wages to unemployment. But using wage levels implies that the coefficient on a lagged wage variable is less than one, whereas most of the literature (see Card, 1995, and Blanchard and Katz, 1997) has found that it is close to one. When we ran level wage equations, we also found the coefficient on the lagged dependent variable to be one, suggesting that the relationship is really between unemployment and security and changes in wages. The change specification also avoids technical problems associated with having a lagged dependent variable on the right-hand side with a regional fixed effect and serial correlation in the error term.

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