WORKING PAPER #416
PRINCETON UNIVERSITY
INDUSTRIAL RELATIONS SECTION
MAY. 1999

The High-pressure U.S. Labor Market of the 1990s

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Revised: May 1999
First Draft: March 1999

We are grateful to Bill Brainard, Olivier Blanchard, Gary Burtless, Bill Dickens, Henry Farber, Jeff Kling, Bob Gordon, George Perry, Jim Stock, Stephen Wandner and Mark Watson for helpful comments. We thank Cindy Ambler, David Autor, Jared Bernstein, Hoyt Bleakley, Jeff Kling, Michael Lettau, Robert Shimer, Aaron Siskind, and Mark Watson for providing data. This paper was prepared for the Spring 1999 Brookings Panel Meeting.

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ABSTRACT

This paper examines the impact of selected labor market changes on the decline in the unemployment rate in the 1990s. The first section provides an overview of aggregate unemployment trends, inflation, and price and wage Phillips curves. The second section examines the effect of demographic changes on the unemployment rate. The third section examines the impact of the 150 percent increase in the number of men in jail or prison since 1985 on the unemployment rate. The fourth section examines the impact of evolving labor market intermediaries (namely worker profiling by the Unemployment Insurance system and the growth of the temporary help industry) on the unemployment rate. The fifth section explores whether worker bargaining power has become weaker, allowing for low unemployment and only modest wage pressure, because of worker job anxiety, the decline in union membership, or increased competitive pressures. The final section examines the impact of the tightest labor market in a generation on poverty. Our main findings are that changes in the age structure of the labor force, the growth of the male prison population, and, more speculatively, the rise of the temporary help sector, are the main labor market forces behind the low unemployment rate in the late 1990s.

The recent performance of the U.S. economy has been nothing short of extraordinary. In 1998 both inflation and unemployment reached their lowest levels since 1965 and 1969, respectively. Although estimates of the NAIRU -- the level of unemployment consistent with stable inflation -- are imprecise, the unemployment rate has been below the 5-percent lower-bound of Staiger, Stock and Watson's 95 percent confidence interval for the NAIRU for the last 20 consecutive months, and the rate of price inflation *declined* in 1997 and 1998.[1] What accounts for the unexpectedly strong performance of the U.S. economy?

It is unclear whether the unusual combination of low unemployment and low inflation in the 1990s is due to fortuitous developments originating in the labor market or to changes in product and financial markets. If labor market developments are responsible, they may represent lasting structural changes that could permanently lower the NAIRU, while developments outside the labor market are more likely to represent favorable transitory shocks that only temporarily allow low inflation and unemployment. Robert Gordon, for example, attributes the shift in the Phillips curve largely to favorable price shocks (e.g., computer and energy prices), changes in the measurement of inflation, and reduced growth in employer health care costs.[2] Others, including James Stock, point to the fact that the relationship between price inflation and capacity utilization, industrial production, and other measures of the business cycle have remained stable in the 1990s, which suggests that something in the labor market has changed to accommodate low unemployment and low inflation.[3] An understanding of the forces that have created the fortunate combination of low inflation and low

[1]See Staiger, Stock and Watson (1997a, table 1).

[2]Gordon (1998).

[3]Stock (1998).

unemployment is critical to predict whether these conditions will continue, and to devise the policies, if any, that could help sustain low unemployment. Moreover, the circumstances that have generated the lowest unemployment and inflation in a generation might also have altered the consequences of a tight labor market. In this paper we investigate the labor market causes and consequences of low U.S. unemployment in the 1990s. The goal of this paper can be thought of two ways: (1) we seek to explain why the unemployment rate at the peak of this ongoing expansion is 0.8 percentage points lower than at the peak of the last expansion, and 1.4 percentage points lower than at the peak of the expansion before the last one; (2) we seek to explain why the NAIRU has fallen by an estimated 1.2 percentage points since the mid 1980s. Even economists who question the utility of the NAIRU and Phillips curve should still find the first goal of interest.

We begin by providing an overview of trends in employment, unemployment, wage growth and price inflation in the next section. We first explore the stability of textbook macroeconometric relations between price inflation, wage inflation, and unemployment. A contribution of our analysis is that we use CPS data to examine the sensitivity of wage growth to unemployment for various education groups and wage deciles. Our overview of the macroeconomic evidence suggests that features of the labor market may have changed to allow for low unemployment and low inflation in recent years. The wage Phillips curve appears to have shifted since 1988. Additional evidence suggests that the Beveridge curve, which relates job vacancies to unemployment, has also shifted favorably. At a minimum, our review of the macro evidence suggests there is value in exploring labor market changes at a more disaggregate level.

Our approach is to explore the plausibility of various contending explanations for the decline in unemployment and restrained wage growth. We evaluate four main hypotheses concerning labor market changes that might partially or fully account for the contemporaneous decline in unemployment and wage pressure: (1) demographic trends which have led to a more mature and stable workforce; (2) the 1990s surge in the prison population may have reduced the measured unemployment rate because the institutional population is not counted as part of the labor force in official statistics, and individuals in prison historically have low rates of employment when they are not in prison; (3) improved efficiency of labor market matching, possibly due to the rise of the temporary help industry and the provision of job search assistance by the Unemployment Insurance system; and (4) the "weak backbone hypothesis," which holds that workers have been reluctant to press for wage gains in this recovery because they are anxious about their job prospects or because unions are weak. This type of exercise does not always lead to hypotheses that can be cleanly or directly tested. As a consequence, we cast a broad net, and try to gather strands of evidence where we can find them.

In the second section we evaluate the role of changing demographics. In particular, we focus on changes in the age and education structure of the work force. Changes in the age composition of the labor force, driven by the maturing of the baby boom, can account for an estimate 0.4 percentage point decline in the overall unemployment since the mid-1980s.[4] But naive compositional adjustments for increases in the educational attainment of the work force should have persistently reduced the NAIRU over the past several decades. We argue that adjustments to the unemployment rate for age structure changes are more plausible than adjustments for changes in the educational composition.

[4]Also see Horn and Heap (1999) and Shimer (1998).

The third section examines the role of the explosion of the prison population. Nearly 2 percent of the adult male population is currently incarcerated. The prison population has almost doubled in the last decade. Since convicted criminals typically have weak attachment to jobs prior to being arrested, it is possible that the labor market is not as tight as the low unemployment rate suggests. This explanation presumes that low measured unemployment in the 1990s is illusory: the unemployed have just been relabelled as the prison population. Our calculations suggest that the increase in the incarcerated population can account for roughly a 0.3 percentage point decline in the *male* unemployment rate since 1985.

In the fourth section we explore changes in labor market matching. First, we evaluate the effect of the new Worker Profiling and Reemployment Services (WPRS) program, which has been a major initiative to improve the efficiency of the Unemployment Insurance system in the 1990s. Most importantly, Job Search Assistance (JSA) has been much more widely used as a consequence of WPRS. Our analysis indicates that JSA, and Worker Profiling more generally, are unlikely to affect sufficiently large numbers of workers to significantly influence the aggregate unemployment rate. Second, and of more consequence, we examine the impact of the temporary help industry on unemployment and wage growth. Although the temporary help industry (help supply services) only employs 2.2 percent of the workforce, the industry has grown rapidly in recent years with its employment level doubling from 1992 to 1998.[5] A significant share of workers also flow through the temporary help industry. Estimates for the state of Washington indicate that 3.7 percent of

[5]Payroll employment in the help supply services industry (SIC 7363) increased from 1.41 million in 1992 to 2.82 million in 1998 according to the establishment survey data from the BLS Current Employment Statistics (CES) program.

workers flowed through temporary help jobs at some point during 1994, and 5 percent flowed through the industry between 1993:Q1 and 1994:Q4.[6] The availability of temporary help jobs may provide an alternative to short-term unemployment and job search for job seekers. We present some preliminary (and quite speculative) cross-state panel regressions suggesting that the availability of temporary help workers to firms may lessen wage pressures that ordinarily accompany tight labor markets, possibly by enabling firms to quickly fill vacancies without having to adjust their overall wage structure. Our results suggest the growth of the temporary help sector may account for as much of the decline in unemployment as demographic shifts.

The fifth section tests the weak backbone hypothesis. We first focus on the role of unions. Union membership has declined steadily within private sector industries for the last 30 years. And the PATCO strike appears to have been a major watershed in terms of union organizing and strike activity. It is possible that the threat of unionization is so low in many industries that the labor market has crossed a "tipping point" beyond which unions and the threat of unionization have very little influence on wage setting. We also explore more sociological based explanations for wage moderation. For example, 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 consequences of greater worker insecurity."[7] And in a February 1999 speech he elaborated: "The rapidity of change in our capital assets, the infrastructure with which all workers must interface day-

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[6]Segal and Sullivan (1997a). The employment share of the temporary help in industry in Washington in 1994 of 2.24 percent is quite similar to the national share in 1998.

[7]Greenspan (1997).

by-day, has clearly raised the level of anxiety and insecurity in the workforce."[8] Paul Krugman has emphasized a related argument for timid wage demands on the part of workers: "These days competition among firms is more intense (why? good question), and nobody wants to let costs get out of line."[9] We find the evidence for worker anxiety causing wage restraint murkier but less compelling. Most importantly, worker surveys do not reveal widespread insecurity, and the link between insecurity and wage growth across regions is tenuous at best. Wage growth in recent years has been weaker for sectors of the economy that have been exposed to more intense competition, such as goods-producing industries and unionized firms, but wage growth in these sectors has only been slightly below overall wage growth, especially once historical cyclical patterns are taken into account.

We next briefly explore some of the social and distributional consequences of tight U.S. labor markets since the mid-1990s. The prolonged macroeconomic expansion of the 1990s finally appears to be having a payoff in terms of real and relative wage growth for low-wage workers and improvements in family incomes for the disadvantaged even in the face of major social policy changes such as welfare reform. The conclusion summarizes our main findings and considers whether the factors we have identified are likely to be temporary or permanent.

Unemployment, Wage and Inflation Trends

The first seven columns of Table 1 report measures of the unemployment rate by sex and spell duration for each of the last 30 years. Although the unemployment rate historically has been

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[8] Quoted in Bureau of National Affairs (1999).

[9] Krugman (1999, p. 38).

higher for women than for men, beginning in the early 1980s the unemployment rate for men has exceeded or roughly equaled that for women. The female unemployment rate reached its lowest level since 1953 last year. Columns 3-6 show the number of ongoing spells of unemployment of various lengths as a percent of the labor force. Interestingly, the short-term unemployment rate is near an all-time low, while the long-term unemployment rate (defined as 26 weeks or longer) is slightly higher than it was in 1989 at the peak of the last recovery. Because the composition of unemployment has shifted toward longer-term spells, the average length of an on-going unemployment spell was 22 percent higher in 1998 than in 1989, and 34 percent higher than in 1979. These statistics suggest that factors that caused a decline in short-term joblessness hold the key to understanding why unemployment is lower now than it was at the peak of previous business cycles.

Column (7) reports a different measure of the unemployment rate: the "work experience unemployment rate." This variable measures the proportion of individuals in the labor force each year (for at least one week) who experienced any unemployment during the course of the year. Notably, in 1997 the work experience unemployment rate reached its lowest level since the BLS began this series in 1958. The low incidence of unemployment in this recovery is probably closely connected to the decline in short spells of unemployment. Thus, at a given unemployment rate, a lower fraction of the work force appears to flow through unemployment in the late 1990s than in the past.

Less than a decade ago, Juhn, Murphy, and Topel argued that the natural rate of unemployment increased in the 1980s because the demand for less skilled male workers declined, causing a rise in permanent joblessness.[10] The data in column (9) of Table 1 indicate a steady decline in the percent of men who are employed in the 1970s and early 1980s. Since the early 1980s, however, the male employment-population rate has held relatively steady, oscillating between 70 and 72 percent. This pattern is not only due to changes in the age structure: Joseph Quinn finds that the decline in employment rates came to a halt in the 1980s for older males as well.[11] By contrast, female employment rates have grown throughout this period, with slightly slower growth in the 1990s than in earlier decades. The combination of persistently rising female employment rates and stable male employment rates has caused the overall employment-to-population rate for the civilian noninstitutional population to reach an all-time high in each of the last 7 years.

Recent U.S. unemployment performance is even more impressive when compared to other major industrialized economies. Table 2 summarizes the unemployment experiences of the major industrialized countries and the OECD as a whole from 1950 to 1998. Major OECD nations shared a common pattern of rising unemployment from the 1960s to the 1980s, although the magnitudes of the increases vary widely across countries. The table highlights the strikingly distinctive aspects of the evolution of U.S. unemployment. The Unites States has moved from having consistently higher unemployment than the OECD as a whole in the 1960s, 1970s, and 1980s to having a substantially lower average rate in the 1990s. The United States is the only major OECD economy to have a lower average unemployment rate in the 1990s than in the 1970s: 5.8 percent in the 1990s versus 6.1 percent in the 1970s. Furthermore the United States and the United Kingdom are the only major countries to have lower unemployment today than in the early 1990s. Although decelerations

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[10]See Juhn, Murphy and Topel (1991).

[11]Quinn (1999).

in price inflation are present in many major economies, the United States stands out for combining lower inflation with lower unemployment in the 1990s.[12]

Table 3 displays recent price and wage inflation developments for the United States. The first three columns of Table 3 present the main indicators of price inflation for the United States: the Consumer Price Index (CPI-UX1), the Personal Consumption Expenditure deflator, and the Gross Domestic Product (GDP) implicit price deflator. As many commentators have pointed out, these measures of inflation have recently reached their lowest levels in decades.

Because labor compensation comprises more than two-thirds of the cost of producing the gross domestic product, wage growth that is unmatched by productivity growth (or by a decline in profit) tends to generate price inflation. The last three columns of Table 3 present three measures of nominal labor compensation growth. Additionally, Figure 1 displays five different wage series, each deflated by the CPI-UX1. Before analyzing their trends, we briefly describe the various wage series. The compensation per hour measure for the nonfarm business sector derived by BLS from the National Income and Product Accounts (NIPA) data is perhaps the most widely used measure of labor costs by macroeconomists and Wall Street economists. This series has the advantage of defining compensation broadly -- perhaps too broadly, since it includes some compensation of corporate owners and payments to retired workers. The Employment Cost Index (ECI) was designed by BLS to provide a gauge of inflationary pressure coming from the labor market.[13] The

[12]We restrict our analysis in this study to recent changes in U.S. labor market performance. Comparative studies of the role of macroeconomic shocks and labor market institutions in differences in the evolution of unemployment among OECD countries include Ball (1997) and Blanchard and Wolfers (1999).

[13]Although the ECI is widely considered the best measure of wage pressure, it is worth mentioning a few potential limitations of the index. First, the ECI is a fix-weight Laspeyres index, which may over state compensation pressures due to "substitution bias" as relative wages change, just as the CPI may overstate

ECI measures wage increases within a fixed set of establishments and jobs, much like the CPI measures price inflation. The table and figure display the ECI for private sector workers. The figure also displays an experimental compensation per hour measure that was derived from the ECI data; this measure uses current hours weights to calculate total compensation costs per hour in the private sector.[14] Labor economists tend to focus on Current Population Survey (CPS) wage data, which have the advantage of providing microdata but lack information on fringe benefit costs. The CPS data used here are from the May CPS for 1973-78, and the Outgoing Rotation Group files for 1979-98.[15] Finally, the average wage of production and nonsupervisory workers, which is estimated from the BLS's monthly Current Employment Statistics (CES) survey of establishments, is closely watched by the financial markets, and covers some 80 percent of the workforce.[16]

Nominal hourly compensation growth as measured by the NIPA or ECI compensation data has averaged about 1.0 percentage point less between from 1992 to 1998 than from 1983 to 1989. The wage and salary component of the ECI has grown about 0.75 points less in this recovery than in the previous one. Nonetheless, unlike price inflation, Table 3 indicates that nominal wage growth

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increases in the cost of living because of substitution bias. Second, and also parallel to issues raised by the CPI, there may be unmeasured changes in the quality of workers within industries and occupations. The secular increase in the education of the workforce which has occurred within occupations and jobs, for example, would be expected to cause the ECI to overstate the growth in labor costs. During expansions, however, upgrading of less qualified workers could possibly cause the ECI to understate wage pressures. And third, if technological change is skill biased, the ECI would provide a misleading measure of cost pressures because job categories that are in high demand will receive too little weight. An analogous problem would arise with the CPI if preferences change.

[14]The experimental compensation per hour from the ECI was derived along the lines described in Barkume and Lettau (1997). We thank Michael Lettau for providing these data.

[15]These data were supplied by Jared Bernstein of the Economic Policy Institute.

[16]Abraham, Spletzer and Stewart (1999).

clearly has increased in the last few years.

Worker well being depends more on real compensation than nominal compensation. Hence, we deflate the wage series by the CPI-UX1 in Figure 1. Two features of Figure 1 stand out. First, the wage series display divergent trends prior to 1996. Second, since 1996 all of the wage series have grown by 1-3 percent. From 1980 to 1996, the ECI and NIPA compensation per hour showed a steady upward trend, while the average wage from the CPS, average wage of production and nonsupervisory workers, and compensation per hour derived from the ECI data showed flat or declining trends. The disparate trends in the NIPA, CPS and CES wage series have been explored by Katharine Abraham, James Spletzer and Jay Stewart.[17] They conclude that different trends in hours account for half of the faster growth of the NIPA compensation measure than CPS wage measure between 1973 and 1997, and that different trends in total payroll account for the remaining discrepancy.

Part of the divergence in the wage series can be ascribed to fringe benefits, since the CPS and CES data exclude fringes. The wage component of the ECI has grown faster than the nonwage component in each of the last four years; only once in the preceding 15 years has the nonwage component grown less than the wage component. Employer health insurance costs have grown particularly slowly, declining by 14 cents per hour between 1994 and 1998, but other fringe benefits have also grown slowly or declined. For example, the hourly cost of providing workers' compensation insurance declined by 5 cents between 1994 and 1998.[18] Because wage and nonwage

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[17]Abraham, Spletzer and Stewart (1999).

[18]Unpublished tables from the Bureau of Labor Statistics. These figures are from Employee Cost for Employee Compensation, and pertain to March of each year.

benefits are fungible, it is difficult to view the deceleration in benefit costs as a separate phenomenon from the wage trends; it is likely that wages would have grown more slowly had health insurance and other benefit costs not decelerated. Available evidence suggests that the slowdown in health insurance costs was not simply a result of a one-time switch to managed care. Krueger and Levy find that the slowdown in employer health care costs occurred because of a general slowdown in the growth of health insurance premiums, and because of a steady decline in employer-provided health care coverage.[19] It is unclear whether the quality of health care and extent of covered services declined as well.

Aggregate Price and Wage Phillips Curves

The coincidence of rather low and declining measured unemployment and price inflation in the United States from 1992 to 1998 is suggestive of a decline in the NAIRU relative to the 1970s and 1980s as well as of some favorable supply shocks over the past few years. Several recent econometric studies have found that the NAIRU declined by 0.7 to 1.5 percentage points between the mid 1980s and mid 1990s.[20] Much uncertainty remains concerning the magnitude, sources, and persistence of the decline in the NAIRU. And much debate remains concerning the extent to which the recent declines in price inflation and unemployment reflect transitory factors as opposed to structural changes in the labor market.[21]

We first summarize the macroeconomic patterns motivating a search for structural labor

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[19]Krueger and Levy (1997).

[20]See Stiglitz (1997), Staiger, Stock and Watson (1997a,b), and Gordon (1998).

[21]Gordon (1998) and Stock (1998).

market changes through the estimation of textbook (naive) price and wage Phillips curves. We then provide a more detailed analysis of wage growth and unemployment relations for sub-groups of the labor force.

We start with the simplest macreconometric model of the determination of the NAIRU. We specify a two equation system for price and wage determination of the following form:

(1) $\Delta p_t = \alpha_p + \Delta w_t + \epsilon_{pt}$

(2) $\Delta w_t = \alpha_w + \Delta p_{t-1} - \beta u_t + \epsilon_{wt}$

where Δp_t is the change in the logarithm of the price index in year t, Δw is the change in the logarithm of the nominal wage, and u_t is the unemployment rate. The first equation can be thought of as the first difference of a "price setting" or "demand wage" relation, and the second one as a "wage setting" or "supply wage" relation.[22] Textbook macroeconomic models imply the intercept in the price equation (α_p) will be $-q$, where q is expected productivity growth. In equation (2) lagged inflation is assumed to provide an adequate proxy for expected inflation. Supply shocks are not explicitly accounted for and are subsumed in the error terms. Substituting equation (2) into (1) yields the "expectations-augmented" Phillips curve:

(3) $\Delta p_t = \alpha + \Delta p_{t-1} - \beta u_t + \epsilon_t$

where $\alpha = \alpha_p + \alpha_w$ and $\epsilon_t = \epsilon_{pt} + \epsilon_{wt}$. Notice that Δp_{t-1} could be subtracted from each side of equation (3), yielding the "accelerationist" Phillips curve.

The NAIRU (u*) is the unemployment rate at which inflation is stable in the absence of shocks: $u^* = \alpha/\beta$. Thus the wage and price Phillips curves in equations (2) and (3) can be re-written

[22] This presentation follows Blanchard and Katz (1997), to which the reader is referred for a more elaborate discussion of these two equations.

as:

(4) $\Delta p_t = \Delta p_{t-1} - \beta(u_t - u^*) + \epsilon_t$

(5) $\Delta w_t = -\alpha_p + \Delta p_{t-1} - \beta(u_t - u^*) + \epsilon_{wt}$.

Equations (4) and (5) imply that price inflation tends to accelerate and expected real wages tend to grow faster than productivity when unemployment is below u*.

Figure 2 presents a scatter diagram of the "accelerationist" price Phillips curve, with the change in the PCE inflation rate on the vertical axis and the overall unemployment rate on the horizontal axis. The figure also shows the ordinary least squares (OLS) line fit to the observations using the years 1973-88. As Gordon and others have concluded from more sophisticated analyses, the large negative outliers in 1996-98 suggest a change in the Phillips curve relationship.[23]

To more formally test for a shift in the Phillips curve in the last decade, Panel A of Table 4 presents OLS regressions of year-over-year changes in price inflation on a lagged dependent variable, the overall unemployment rate, and a dummy variable that equals one in years after 1988. The lagged dependent variable is constrained to have a coefficient of one, but if it is unconstrained the coefficient is still very close to one. Additionally, some specifications include an interaction between the post-1988 dummy and the unemployment rate. The left part of the table presents estimates using the CPI-UX1 to measure price inflation, and the right part uses the PCE deflator. Columns 1 and 2 are estimated for the period 1973-98, while column 3 is estimated for the period 1962-98. The results are quite similar regardless of whether inflation is measured by the CPI or PCE. When the shorter period is used, the intercept of the Phillips curve is found to have shifted in

[23]Gordon (1998).

during the past decade, and the sensitivity of inflation to the unemployment rate is found to have become much weaker. The results for the sample in column (3), however, which include years prior to the productivity slowdown, show less evidence of a shift in the Phillips curve.[24]

Figure 3 provides a scatter diagram of a wage Phillips curve, using the percentage change in the wage and salary component of the ECI minus the lagged CPI inflation rate as the measure of wage growth and the overall unemployment rate. The figure also displays the OLS line fit through the points in the 1976-88 period. All of the observations for 1989-98 are below the predicted line, although the observation for 1998 is very close to the line. By fitting a Time-Varying Phillips curve to quarterly ECI data, Gordon finds that the latest observations (for wage and salary) for the first half of 1998 fall right on the line, which he interprets as evidence that the wage Phillips curve has been stable. Such a conclusion appears to be somewhat sensitive to the precise choice of specification of the wage Phllips curve and period examined.

To examine whether the relationship between wage growth, inflation and unemployment has changed in the last decade, Panel B of Table 4 reports a series of regressions based on equation (2) in which the dependent variable is the year-over-year change in the natural log of nominal hourly compensation, and the independent variables include the lagged growth of the CPI, the unemployment rate, and a dummy variable indicating years after 1988. We measure compensation by the NIPA compensation per hour, total ECI, wage and salary component of the ECI, and the average hourly wage from the CPS. Because the interaction between unemployment and the post-

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[24]If we include an interaction between the post-1988 dummy and the unemployment rate in the column (3) model, it is statistically insignificant, and thepost-1988 dummy and the interaction are jointly insignificant in both the CPI and PCE models.

1988 dummy is only significant for the NIPA data, we omit this variable from the other models. The regressions are estimated over various periods for which data are available. We estimate the equation constraining the coefficient on lagged (CPI) inflation to equal one, as in equation (2).

Similar to the price Phillips curve, results for all of the wage series in Table 4 indicate a shift in the wage growth-unemployment relationship. In the last decade, wage growth has been lower than one would predict based on the historical relationship between unemployment and wage growth. Moreover, the fact that the equation for the wage and salary component of the ECI appears to have shifted in at least as much as the equation for the total ECI shifted in, suggests that special factors due to slower growth in fringe benefits are not responsible for the post-1988 inward shift of the wage Phillips curve (compare columns 4 and 6). The pick up in wage growth over the last few years evidently is not sufficient to overturn the intercept-shift in the wage growth equation over the last decade as a whole. We do not want to push these regressions too far, however. We readily acknowledge that the 1989-98 period for the shift was chosen arbitrarily, and the results are sensitive to the time period we chose and not very precisely estimated.[25] But these results suggest that the something may have caused a change in the wage-setting relationship in the last decade, facilitating less wage-push inflation despite low unemployment. If nothing else, these results suggest it is worth probing what might have caused the wage growth-unemployment relationship to shift.

A final word on statistical measurement changes is required. It is well known that the BLS made several adjustments to the CPI in the mid 1990s that likely reduced the inflation rate. Gordon, for example, estimates that an approximately 0.2 percentage point decline in estimates of the NAIRU

[25]A grid search over possible years for the intercept shift typically finds that a post-1987 dummy maximizes the R-square of the wage growth equations.

from 1988 to 1998 may be due to changes in the measurement of the CPI.[26] It is also the case, however, that BLS redesigned the CPS questionnaire in 1994, which may have affected the measured *unemployment rate*. Polivka and Miller find that the redesign of the CPS may have raised the aggregate unemployment rate by 0.2 percentage points, with the effect being larger for women.[27] If the CPS revision increased the measured unemployment compared to what it would have been with the old questionnaire, then the NAIRU would have fallen by even more than 0.7 to 1.5 points since the mid-1980s.

Wage Trends and Phillips Curves for Sub-Groups of Workers

Wage growth has not been uniform for different groups in the labor market. It is well known, for example, that average wages grew more for those with a college education than those with a high school education or less in the 1980s and early 1990s.[28] Similarly, real wages have declined for low-wage deciles since the 1970s, and increased for those in high-wage deciles. Figure 4 illustrates real log wage growth relative to 1979 for workers at the 10th, 50th and 90th percentile each year, as well as the value of the minimum wage.[29] The minimum wage fell by 31 percent in real terms

[26]Gordon (1998, Table 6). The BLS is currently devising a consistent CPI series that adjusts the historical data to be comparable to the current data, which will be useful for future analyses.

[27]Polivka and Miller (1995).

[28]See, for example, Levy and Murnane (1992) and Katz and Autor (1999). One difficulty in comparing wages across education groups is that the average "quality" of the groups may change over time; for example, the quality of education could change. Scores on the NAEP exam for 17 year olds have remained relatively stable or increased since the early 1970s; see National Center for Education Statistics (1997).

[29]The CPI-UX1 is used to deflate the wage series. The wage data are from the May 1973-78 CPS and the Outgoing Rotation Group files for 1979-98. The data were provided by Jared Bernstein.

between 1979 and 1989. The wage at the 10th percentile of the distribution fell by 16 percent in this period, and has rebounded by 6.6 percent since 1989, with most of the increase occurring in 1997-98. Based on an analysis of regional variation in wages, David Lee attributes much of the 1980s decline in the relative earnings of workers at the bottom of the wage distribution to the declining relative (and real) value of the minimum wage.[30] The median worker saw a real decline of 2 percent in his or her earnings between 1979 and 1989, and a steeper decline in the mid 1990s, until real wages recovered during 1997-98. Finally, the worker at the 90th percentile of the distribution experienced a 4 percent gain in real earnings from 1979 to 1989, and another 5 percent gain between 1989 and 1998. One way in which the 1990s recovery differs from the 1980s recovery is that real wage growth was more widespread throughout the distribution seven to eight years into the 1990s recovery.

Table 5 presents estimates of wage Phillips curves using the average hourly wage of workers at different levels of education, derived from CPS data from 1973 to 1997. Panel A of Table 5 uses the overall unemployment rate to predict wage growth, whereas Panel B uses the unemployment rate specific to each education group; that is, in panel B the unemployment rate of high school dropouts is used to predict high school dropouts' wage growth, the unemployment rate of high school graduates is used to predict high school graduates' wage growth, and so on.[31]

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[30]Lee (1999).

[31]We experimented with including a variable measuring the change in the log minimum wage in the regressions in Table 5, but found that this variable had an insignificant and small effect. In addition, the other coefficients were unaffected by the inclusion of this variable. The education-specific unemployment rates are from U.S. Department of Labor (1997, Table 56) for 1974-96, and from <http://stats.bls.gov> for 1997. The education-specific unemployment rates pertain to March. The 1997 data are for those 25 and over, whereas the 1973-96 data are for those age 25 to 64. To make the 1997 data comparable, we subtracted off the differential unemployment rate that arose because of the difference in the ages of the samples in 1996.

The results in the top panel of Table 5 indicate that wage growth is more responsive to the overall unemployment rate for those with a lower level of education than for those with a higher level of education. This finding is in keeping with a large literature that finds that skill-upgrading is more common during periods of low unemployment, and that wage differentials tend to be compressed during periods of low unemployment.[32] Interestingly, when the education-group-specific unemployment rate is used in the regression in Panel B, the pattern is reversed: higher educated groups experience stronger wage growth when *their* unemployment rate declines by a percentage point compared to lower educated groups. Because the unemployment rate is much more variable over the business cycle for low-educated workers, this finding is not surprising: A tight labor market means that the labor market is especially tight for low-skill workers. Furthermore the much higher average unemployment rate for dropouts implies a smaller wage growth elasticity with respect to the group's unemployment rate for drop outs than for other education groups.

The results in Table 5A further suggest that the post-1988 intercept shift of the wage Phillips curve was primarily brought about by a shift for lower educated workers.[33] The Phillips curve for college graduates indicates no shift. We can calculate the unemployment rate required to generate positive expected real wage growth for each education group using the estimates in Table 5. Interestingly, the point estimates in the top panel imply that real wage growth arrived when unemployment was below 6.4 to 6.6 percent for all education groups prior to 1989. In the 1989-98

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[32]Okun (1973).

[33]The results in Panel B, however, indicate little shift in the Phillips curve forany of the education groups when group-specific unemployment rates are used, and when lagged inflation is constrained to have a coefficient of 1.0. But because the earlier aggregate results in Table 4 are based on the overall unemployment rate, the Panel A results are probably most relevant for understanding the underlying trends that influenced the aggregate Phillips curve.

period, the unemployment rate associated with zero expected wage growth is estimated to have declined to 4.6 percent for the less than high school group, and to 5.3 percent for the high school and some college groups. The estimates in Panel B suggest that the group-specific unemployment rates at which one would expect positive real wage growth are rather stable over time for all of the education groups.

Table 6 presents several additional estimates of wage Phillips curves, each using wage growth of workers occupying the 10th, 30th, 50th, 70th or 90th percentiles of the wage distribution as the dependent variable. In addition to the unemployment rate, these models also include the change in the logarithm of the nominal minimum wage as an explanatory variable.[34] Again, lagged CPI inflation is constrained to have a unit coefficient. These results also indicate that wage growth is more responsive to the overall unemployment rate for the least-paid groups of workers. According to the model with the unconstrained inflation rate, a one percentage point increase in the unemployment rate is associated with an increase in the 10th percentile wage of 1.5 percent, and increase in the 90th percentile wage of 0.4 percent. The bottom of the table reports the point estimates for the implied unemployment rate associated with zero expected real compensation growth (URZERCG) for each decile. Interestingly, in the pre-1989 period, the URZERCG tends to rise with the wage level as expected in a period of rising wage inequality and sharp labor shifts against less-skilled workers. In the 1989-98 period, however, the implied URZERCG is roughly constant for each of the wage deciles indicating a much more egalitarian impact of tight labor

[34]If the minimum wage increased in the middle of a year, we calculated the average wage in place during the course of the year. That is, we weighted the minimum wage by the number of months that it was in effect during the year. The results were qualitatively similar if we excluded this variable.

markets across the wage distribution in the 1990s. But a lower unemployment rate appears to have been necessary to generate positive real wage growth in the past decade relative to the 1974 to 1988 period.

Beveridge Curves

The Beveridge curve, or the relationship between vacancies and unemployment, can provide additional clues about the nature of possible structural changes in the labor market.[35] Labor market innovations that reduce the equilibrium unemployment rate by improving the efficiency of matching in the labor market and/or increasing search effort by the unemployed are likely to generate an inward shift in the Beveridge curve. Demographic shifts reducing the share of high turnover (high unemployment inflow) young workers in the labor force should also be associated with a favorable inward movement in the Beveridge curve. In contrast, wage restraint driven by pure reductions in worker bargaining power arising from a decline in union membership, increased worker psychological "insecurity," or increased international competition should shift in the wage Phillips curve relation but not systematically shift the Beveridge curve. Increased rates of economic turbulence from rising globalization and more rapid technological change (e.g., the new economy) could probably be re-interpreted as an increased rate of job reallocation and be expected to shift out the Beveridge curve. Thus the major alternative hypotheses for wage restraint and low unemployment make different predictions about change in the unemployment-vacancy relationship in the 1990s.

[35]See Blanchard and Diamond (1989) for a derivation of the theoretical underpinnings of the Beveridge curve and an assessment of its usefulness in identifying the sources of changes in unemployment.

The lack of a consistent national job vacancy series for the United States creates difficulties for assessing shifts in the U.S. Beveridge curve. Researchers are forced to rely on the Conference Board's help wanted index, based on newspaper help wanted advertising, as a proxy for the job vacancy rate. Katharine Abraham has shown that cyclical movements in the normalized help wanted index (the ratio of the help wanted index to total nonfarm payroll employment) tends to closely track cyclical movements in direct job vacancy measures in periods and locations for which both series are available.[36] But Abraham also finds that secular movements in the normalized help wanted index are likely to have deviated from those in the "true" underlying job vacancy rate because of changes in the newspaper industry and changes in employer recruiting practices (especially equal employment opportunity pressure that increased the use of help wanted advertising for a given level of job vacancies in the 1970s). We use a proxy for the job vacancy rate that is based on the normalized help wanted index and incorporate Katharine Abraham's adjustments through 1985. Because of the lack of much systematic analysis of changes in the use of help wanted advertising since 1985, we naively assume no change in the relation between the normalized help wanted index and the vacancy rate since 1985.[37]

Figure 5 provides a scatter of the U.S. unemployment-vacancy relationship from 1960 to

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[36]Abraham (1987).

[37]We are grateful to Hoyt Bleakley for providing us with the data on job vacancy proxies and on the Conference Board help wanted index. Bleakley and Fuhrer (1997) provide documentation for this job vacancy proxy and also provide a more detailed analysis of recent changes in the U.S. Beveridge curve and the efficiency of job matching. Our job vacancy measure is a re-scaled version of the normalized help wanted index including Abraham's adjustments through 1985. The variable is scaled to match earlier estimates of actual job vacancy rates following Blanchard and Diamond (1989). Our post-1985 vacancy proxy differs from the measure used in by Bleakley and Fuhrer since we do not assume a continued trend inward shift in the job vacancy rate relative to the normalized help wanted index after 1985.

1998. The figure illustrates an outward shift in Beveridge curve in the 1970s. Katharine Abraham's analysis of the period through 1985 suggests that both demographic changes (an increased share of younger workers in the labor market) and increased regional dispersion in labor market performance played a role in this outward Beveridge curve shift.[38]

Figure 5 also suggests a large inward shift in the Beveridge curve from the mid-1980s to the 1990s that has more than reversed the earlier outward shift of the 1970s. This recent substantial inward movement in the Beveridge curve is potentially supportive of hypotheses emphasizing structural labor market changes that have increased the efficiency of job matching, demographic shifts towards older and more stable workers, and (possibly less plausibly) reductions in job reallocation intensity. An important caveat in drawing such a conclusion is the possibility of changing hiring practices that lead to less reliance on help wanted advertising for a given level of true job vacancies. For example, the growth of the temporary help industry and of internet job listings could both improve the efficiency of job matching and lead to a reduction in the number of newspaper help wanted ads placed for a given level of job vacancies. To the extent this latter effect is present, the use of data from the help wanted index as a proxy for job vacancies will tend to overstate the true inward shift in the Beveridge curve.

Demographic Change and the NAIRU

A venerable macroeconomic tradition is to examine the extent to which changes in the age and sex composition of the labor force can explain secular movements in the unemployment rate.

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[38]Abraham (1987).

The much higher unemployment rates for teenagers and young adults than for prime age adults makes it plausible that changes in the age structure of the work force can substantially affect the unemployment rate. Seminal studies by George Perry and by Robert Gordon provide strong evidence that changes in the age and sex composition of the work force (the labor market entry of the baby boom cohorts and rapid expansion of female labor force participation) contributed to an increase in the NAIRU in the 1960s and 1970s.[39] The convergence in male and female unemployment rates since the early 1980s indicates that the direct effect of sex-composition changes on the unemployment rate is unlikely to be important over the past two decades. But recent studies by Robert Shimer and by Robert Horn and Phillip Heap suggest that age-structure changes driven by the maturing of the baby boom cohorts can account for a substantial part of lower unemployment in the 1990s than in the 1970s and 1980s.[40] In this section, we reassess the role of age structure changes and also explore the possible consequences for the NAIRU of continuing secular increases in the educational attainment of the adult work force.

The potential importance of age-structure changes for the evolution of the aggregate unemployment rate is highlighted by the large differences in unemployment rates by age group and by the dramatic rise and then fall in the labor force share of young workers over the past four decades. Figure 6 illustrates the rise in the importance of youth and young adults (those aged 16 to 24 years) in the labor force from 1960 to 1998. The labor force share of those under 25 years of age increased from 16.6 percent in 1960 to 24.5 percent in 1978 and then declined to 15.8 percent in

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[39]Perry (1970); Gordon (1982).

[40]Shimer (1998); Horn and Heap (1999). In contrast, Gordon (1997) argues that increases in the labor force share of young workers help explain the rise in the NAIRU in the 1970s, but that dedines in the youth share failed to lower the NAIRU in the 1980s.

1997.[41]

Trends in unemployment rates for seven age groups (16-19, 20- 24, 25-34, 35-44, 45-54, 55-64, and 65+) from the 1960s to the 1990s are summarized in Table 7. The average unemployment rates for teenagers (16-19) and young adults (20-24) for the entire 1960-98 period were 16.8 and 9.6 percent, respectively, as compared to average unemployment rates of 4.2 percent and 3.7 percent for those aged 35 to 44 and 45 to 54 years. The higher unemployment rates of young workers largely reflect higher inflow rates to unemployment (greater employment instability), not longer durations of unemployment. Thus, the aging of the work force is consistent with the substantial decline inflow rates into unemployment in the 1990s compared to the 1980s and 1970s.[42]

Table 7 indicates the potential role for age-composition changes in differences in decadal unemployment experiences. Although the average overall unemployment rate was higher in the 1970s than it has been so far in the 1990s (6.2 versus 5.9 percent), the average unemployment rates for five of the six age groups of the working age population (those aged 16 to 64 years) were lower in the 1970s. But age-group specific unemployment rates were higher for all six of these age groups on average in the 1980s than in the 1990s, and were also higher for all six groups in 1989 than in 1998. Thus the stronger unemployment performance in the expansion of the 1990s than the expansion of the 1980s does not appear to be only attributable to age-composition effects.

We use a simple shift-share decomposition analysis to assess the mechanical effect of age-

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[41]The Bureau of Labor Statistics (BLS) forecasts the current low share of young workers in the labor force will persist over the next decade. The most recent BLS projections predict only a modest rise inthe labor force share of those under 25 from 15.9 percent in 1998 to 16.4 percent in 2006. See Fullerton (1997, Table 7).

[42]See, for example, Bleakley and Fuhrer (1997).

structure changes on the evolution of unemployment from 1960 to 1998.[43] We divde the labor force into seven age groups and ask the question: What would have happened to unemployment if the age structure of the labor force had remained constant over the 1960 to 1998 period? Our initial assumption is that if the age shares had remained fixed from 1960 to 1998, the disaggregate age-specific unemployment rates would have evolved no differently than the actual observed paths. The actual overall unemployment rate at time *t* (U_t) equals the weighted average of the age-group specific rates (u_{jt}'s where j indexes age groups) using the actual time *t* labor force shares (ω_{jt}'s) as weights:

(6) $U_t = \Sigma_j \, \omega_{jt} u_{jt}$.

The hypothetical age-constant unemployment rate at time t (UFW_t) is simply given by the weighted average of the group specific unemployment rates in t using a fixed set of age-group weights for some basline time period (ω_j's):

(7) $UFW_t = \Sigma_j \, \omega_j \, u_{jt}$.

The age-adjustment to the unemployment rate in period t is then simply given by the difference between the actual and age-constant unemployment rates (U_t - UFW_t).

The time pattern of the implied age-adjustments to the unemployment rate is relatively insensitive to the choice of base year. Table 8 illustrates the potential mechanical effects of age-structure changes on the unemployment rate using average labor force shares on the overall unemployment rate using fixed age-group labor force weights for two choices of base periods: the average shares for the 1960 to 1998 period and the age group share in 1979, the midpoint of the period.[44] The estimates using the full-period average age-group shares imply age-structure changes

[43]Our approach closely follows that of Summers (1986) and Shimer (1998).

[44]We normalize the age adjustments to equal 0 in 1979 for both choices of base periods.

can account for a rise on the unemployment rate of 0.63 percentage points from 1960 to 1979 and then a decline of 0.69 percentage points from 1979 to 1998.

An alternative approach, following Robert Shimer, to examine the impact of changes in the age structure on the unemployment rate is to directly calculate a measure of "age-driven" unemployment.[45] We define the age-driven unemployment rate in year t, UA_t, as:

(8) $$UA_t = \Sigma_j \, \omega_{jt} \, u_{jo},$$

where u_{jo} is the group-specific unemployment rate for group j in a base period. Changes in UA_t are entirely driven by changes in the age structure (the ω_{jt}'s). The last column of Table 8 summarizes the evolution of the age-driven unemployment rate with u_{jo} set equal to the average unemployment rate for age group j over the entire 1960-98. The age-driven unemployment rate increased by 0.71 percentage points from 1960 to 1979 and has since declined by 0.63 percentage points.

Thus, alternative age-adjustments lead to similar answers of likely significant reductions in unemployment from an aging population over the past two decades. Furthermore, BLS projections of changes in labor force composition over the next decade predict little change in age-driven unemployment through 2006 as shown in the last row of Table 8.

How far do mechanical effects of age-structure changes go to explaining lower unemployment in the 1990s than in the 1980s or 1970s? The age-adjustments in Table 8 can account for essentially all of the 0.5 percentage point decline in the unemployment rate from the unemployment trough experienced in 1979 to that of 1989. But age composition effects account for only around a 0.2 percentage point decline in unemployment from 1989 to 1998, or about one

[45]Shimer (1998).

quarter of the 0.8 percentage point actual change.

Age-structure changes also do not appear to be large enough to fully explain existing estimates of the decline in the NAIRU since the mid-1980s. Staiger, Stock and Watson provide a point estimate (with much uncertrainty) of a 1.4 percentage point decline in the time-varying NAIRU (TV NAIRU) using the core PCE from 1984 to 1994, while age-adjustments explain a decline in unemployment of approximately 0.3 to 0.4 percentage points over the same period.[46] Mark Watson's updated estimates of the TV NAIRU using the GDP deflator as the price measure indicate a decline in the TV NAIRU from 1985 to 1998 of 1.2 percentage points as compared to a decline in age-driven unemployment of 0.4 percentage points over the same period.[47] Figure 7 plots Watson's point estimates of the TV NAIRU and our measure of age-driven unemployment from 1962 to 1998. The figure indicates that age-driven unemployment tracks the TV NAIRU reasonably well through the end of the 1980s, but the TV NAIRU has diverged downward relative to the age-driven unemployment rate in the 1990s.

We conclude that age-structure changes can explain a significant fraction -- perhaps one-third -- of the decline in the NAIRU since the mid-1980s. The consideration of further demographic adjustments for changes in the sex or age-sex composition of the workforce does not alter these quantitative conclusions for the past two decades.[48] Thus, existing estimates of time-varying

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[46]Staiger, Stock, and Watson (1997a, Table 1).

[47]We are grateful to Mark Watson for providing us with estimates of the TV NAIRU from 1962 to 1998. The TV NAIRU estimates follow the methodology of Stager, Stock, and Watson (1997a,b) with quarterly data; use the GDP deflator as the price measure; and include controls for standard supply shock measures (food and energy price shocks, exchange rate movements, and Nixon price control indicators).

[48]Military personnel requirements are another factor that may influence the demographic composition of the civilian labor force. Since military personnel are disproportionately young adults and not included in

NAIRU's suggest a further decline in the NAIRU since the mid-1980s of at least 0.3 to 1 percentage point that cannot be accounted for by mechanical demographic-composition effects.

A key assumption behind the age-adjustments to unemployment in Table 8 is that changes in the age-composition of the labor force don't affect age-group specific unemployment rates. The labor economics literature on the effects of relative cohort size on the labor market outcomes of young workers has generated a somewhat mixed set of conclusions.[49] Robert Shimer has recently explored the effect of relative cohort size on the differences in unemployment across age groups in the United States. Shimer finds that age-group specific unemployment rates tend to rise relatively when age-group specific labor force shares increase.[50] This pattern suggests that shift-share age-adjustments may understate the effects of changes in the age structure on the unemployment rate as the composition effects of age-structure changes are magnified by impacts of relative cohort size on the unemployment rates of young workers. Shimer finds much larger age-structure effects under the assumption that changes in age-structure do not affect the unemployment of prime- age workers. Shimer's modified age-adjustments can completely explain almost all of the decline in estimates of

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measures of the civilian labor force, the substantial reduction in military personnel on active duty since the end of the Vietnam War in the mid-1970s has tended to increase the share of young workers in the civilian labor force. Military personnel on active duty measured as a percentage of the civilian labor force declined from 3.7 percent in 1970 to 1.9 percent in 1980 to 1.6 percent in 1990 to 1.1 percent in 1997 (U.S. Bureau of the Census 1998a, Table 582; U.S. Department of Labor 1999, Table A-1; Kosters 1999, Table 3). To the extent that military personnel have civilian labor market prospects typical of others in their age group, our age-structure adjustments incorporate the impacts of changes in military requirements on the measured civilian unemployment rate. Military downsizing in the 1990s has probably modestly attenuated the reduction in age-driven unemployment.

[49]Important early work on cohort size and earnings includes Freeman (1979) and Welch (1979). Recent studies have tended to find somewhat ambiguous results concerning the effects of relative cohort size on the employment and earnings of young workers. See, for example, Blanchflower and Freeman (1996).

[50]Shimer (1998).

the NAIRU from the late 1970s to the early 1990s, but lower unemployment in the late 1990s than the late 1980s still remains unexplained after his preferred demographic adjustment.

A further issue raised by Lawrence Summers in an analysis of high unemployment in the mid-1980s is the extent to which one should also attempt to adjust the unemployment rate for changes in the educational attainment of the work force.[51] Summers found that the implied compositional effects of increasing education levels offset the "adverse" changes in age-sex composition in the 1960s-1970s, and the combined effects of an aging work force and rising education levels should have greatly reduced the NAIRU in the 1980s. Robert Shimer finds the educational upgrading can "explain" a 1 percentage point decline in the unemployment rate from 1979 to 1997.[52]

The case for adjustments in the unemployment rate for education composition changes appears much weaker to us than the case for adjustments for changes in age composition. It is clear in a cross-section that more-educated workers have substantially lower unemployment rates than less educated workers. But to the extent that increases in education improve the productivity of the work force, most models of the equilibrium unemployment rate predict equal proportional increases in actual wages and worker's reservation wages thereby leaving the equilibrium unemployment rate unchanged.[53] This pattern implies we should not expect changes in educational attainment to necessarily affect the unemployment rate. This view can be reconciled with the cross-section

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[51]Summers (1986).

[52]Shimer (1998).

[53]See, for example, Blanchard and Katz (1997). Topel (1998) makes a similar point. Shimer (1998) argues against education adjustment both on empirical grounds and on the basis of a signaling model of education.

differences in unemployment by education level by recognizing that reservation wages (which depend on the generosity of government transfers, black market and illegal earnings opportunities, and home production) are likely to be higher relative to market wages for less-educated workers. Thus, even as productivity improvements associated with rising education levels increase wages, unemployment benefits and other determinants of reservation wages tend to rise by a similar proportion, and the smaller gap between the value of unemployment and legitimate labor market opportunities for the less educated tends to be preserved. Unemployment rates have not perennially trended downward in response to rising productivity and increasing education levels.

We conclude that changes in the age-structure of the labor force associated with the labor market entry and then maturation of the baby boom significantly contributed to increases in unemployment from the late 1950s to the late 1970s and to a decline in the NAIRU from the late 1970s to the early 1990s. But the estimated decline in the NAIRU since the early 1990s and most of the decline in actual unemployment from 1989 to 1998 remain unexplained even after accounting for both the mechanical and broader effects of age structure changes. If one adds mechanical adjustments for increases in educational attainment, then falling unemployment is no longer a mystery -- the mystery then is why hasn't unemployment secularly declined throughout the century in all advanced economies. But we are somewhat skeptical of the legitimacy of such adjustments for changes in education composition.

Rising Incarceration Rates and Measured Unemployment

Another major demographic shift that could influence the unemployment rate involves the movement of population into prisons and jails. Figure 8 displays the adult prison and jail population relative to the adult civilian, noninstitutional population. In 1970, 2 in 1,000 adults were in prison or jail; by 1998, the number increased to 9 in 1,000. The proportion of the population in prison or jail has doubled since 1985. About 90 percent of those in prison or jail are men. To put the magnitude of this social problem in perspective, note that in June 1998, the number of adult men in prison or jail equaled 2.3 percent the number in the male labor force. The United States has a much higher incarceration rate than all other developed countries, and the rate has grown exponentially since the early 1970s. Although most of the economics literature on crime has focused on the effect of economic conditions on criminal activity, or the effect of having an arrest record on subsequent labor market activity, little attention has been devoted to the direct effect of the high rate of incarceration in the U.S. on unemployment.[54] Since incarcerated individuals are not counted in either the numerator or denominator of the official BLS unemployment rate, and since those in jail and prison tend to have a high unemployment rate prior to being arrested, the surge in the prison population in recent years could account for some of the decline in measured unemployment. To the extent the decline in the official unemployment rate simply reflects the removal from the measured population through incarceration of a large number of individuals with high unemployment propensities, the decline should be interpreted as a compositional change rather than a "true" improvement in labor market performance for any given group of individuals.

We provide an illustrative set of calculations to explore the likely magnitude of the impact of the surge in the prison and jail population since 1985 on the official male employment and unemployment rates. Table 9 lays out our calculations. Jeffrey Kling finds that about 35 percent

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[54]Freeman (1995) provides an excellent overview of the literature on the economics of criminal activity, and the effect of criminal activity on subsequent labor market activity.

of individuals who served 1-2 year sentences in California for federal crimes were employed prior to being arrested.[55] This figure is similar to the employment rate he finds for a "control" group that was convicted but not sentenced to prison time, two years after their case was filed. Consequently, we assume that 35 percent of those in prison or jail would be employed were they not in prison or jail. Column (4) of Table 9 provides an estimate of the male employment-to-population rate for a hypothetical situation in which all incarcerated individuals were added to the civilian, noninstitutional population, and 35 percent were employed.[56] The 1998 employment-to-population rate is predicted to be 70.6 percent had the prison population been released, compared to the BLS estimate of 70.9 percent for the noninstitutional population. Notice also that the male employment-to-population rate is still estimated to rise from 1985 to 1998 with the adjusted data, but by 0.4 percentage points less than in the official unadjusted data which exclude the prison population.

Column (5) reports the male unemployment rate in 1985 and 1998. One additional assumption is needed to calculate the effect the incarcerated population has on the unemployment rate: the fraction of incarcerated individuals who would participate in the labor force had they not been incarcerated. In columns (6)-(9) we re-compute the unemployment rate under various assumed values of the labor force participation rate that incarcerated individuals would have if they were not incarcerated, and the assumption that 35 percent of the incarcerated population would be gainfully employed. We believe it is likely that the labor force participation for this population would exceed

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[55]Kling (1999).

[56]See Western and Pettit (1998) for an interesting study that calculates employment-to-population rates by race and age under the assumption that the prison population does not work.

40 percent, and most likely be less than 70 percent.[57] One way to estimate the likely labor force participation rate the incarcerated population would have if they were not in prison is to examine the rates for non-institutional populations with similar characteristics to the incarcerated population. Using the 1989 ORG CPS file we find that about one-third of high school dropouts, and of all workers aged 18-34, who are not employed are nonetheless in the labor force and counted as unemployed. If we continue to assume an employment rate of 35 percent, this assumption about the labor force participation rate of the non-employed implies the labor force participation rate would be 57 percent and the unemployment rate would be 38 percent for the incarcerated population if they were not in jail or prison.

If the labor force participation rate of the incarcerated population would have been as low as 40 percent, then only 0.1 point of the 2.6 point fall in the male unemployment rate from 1985 to 1998 could be accounted for by the removal from the labor force statistics of an increasingly large incarcerated population. If the labor force participation rate of this group were 60 percent, which we consider a more plausible value, then a 0.3 percentage point contribution of the decline in the male unemployment rate since 1985 is possible. The low rate of incarceration for women suggests a 0.1 to 0.2 percentage point contribution of rising incarceration to decline in the overall unemployment rate since the mid-1980s. The effect is much more sizeable for sub-groups of less educated and minority men.[58] Of course, these calculations ignore the possible lasting negative effects of incarceration on the labor market prospects of individuals once they are released from

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[57]The labor force participation rate for the entire male civilian, noninstitutional population in February 1999 was 74 percent.

[58]See Western and Petit (1998) for an analysis of the impacts of incarceration on measured changes in black-white male employment rates.

prison. Such persistent effects would tend to raise measured unemployment and offset the mechanical reduction in measured unemployment from incarcerating more high-unemployment individuals.[59]

Worker Profiling, Contingent Jobs, Frictional Unemployment and Wage Pressure

On November 24, 1993 the U.S. Congress passed legislation requiring each state to implement a Worker Profile and Reemployment Services (WPRS) program for unemployed workers through its UI system.[60] Worker profiling involves using a statistical model (which varies across states) to identify individuals upon first receipt of UI benefits who are likely to exhaust their UI benefits and have difficulty finding a job, and then channel those workers to reemployment services, including job search workshops, counseling, job clubs and referrals to employers. (Reemployment services do not include training and educational services.) The program focuses on serving individuals who are predicted to suffer long-term unemployment, based on characteristics such as their recall status, first payment, industry or occupation, employment history, job tenure, education, and the local unemployment rate. Claimants referred to employment services are required to participate in those services as a condition of eligibility for UI. The WPRS initiative represents a break from the traditional approach of the UI program in the U.S., which primarily has been

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[59]The possible increasing magnitude of criminal records on aggregate labor market measures can be partially gauged by rise in the number of adults on probation or parole from 2.3 million (1.3 percent of the adult civilian noninstitutional population) in 1985 to 3.9 million (1.9 percent) in 1998 (U.S. Department of Justice, Bureau of Justice Statistics, <http://www.ojp.usdoj.gov/bjs.).

[60]The November 1993 legislation was preceded by legislation passed on March 4, 1993 that encouraged states to voluntarily establish a worker profiling system. Wandner, Messenger and Schwartz (1999) provide an overview and evaluation of the WPRS system, from which this section draws heavily.

concerned with providing temporary cash compensation to eligible unemployed workers while they search for a job. By implementing WPRS and related One Stop Career Centers, the UI service has embarked on a course to play a more active role in reducing unemployment.

All states phased in WPRS systems between 1994 and 1996. States that phased the program in early (e.g. by the end of 1994) included: Delaware, Hawaii Maryland, Missouri, New Jersey, New Mexico, Oregon, and West Virginia. Late adopters included Arkansas, South Dakota and North Dakota. In 1997, essentially all UI recipients were profiled, 30 percent were placed in the selection pool for services because they were deemed likely to exhaust their UI benefits, and 35 percent of those in the selection pool were referred to some type of service.[61] The intensity of assistance varies considerably across the states. Wandner, Messenger and Schwartz estimate that one third of the states provide only minimal reemployment services -- five hours or less, on average -- to WPRS participants.

A major part of the motivation for enacting WPRS was that several studies have found that job search assistance programs are effective at reducing unemployment spell duration. Bruce Meyer provides a summary of the effects of JSA in five states (South Carolina, New Jersey, Washington, Nevada and Wisconsin) that have randomly selected eligible claimants to receive various forms of JSA, and compared their performance to a randomly selected control group.[62] He finds: (1) JSA participants found a new job more quickly -- their average duration of UI benefits was reduced by about 0.5 to 4 weeks compared to the control group, with most estimates falling near the bottom of this range; (2) the reduction in UI benefits and increased tax revenue from faster reemployment made

[61]Wandner, Messenger and Schwartz (1999, Figure 2), and authors' calculations from ETA 9048 data.

[62]Meyer (1995).

the JSA programs cost effective for the government; (3) on average, JSA participants found jobs that paid about as much as the jobs found by the control group. A recent study by Orley Ashenfelter, David Ashmore and Olivier Deschenes suggests that the instructional component of JSA is essential for it to be effective; stricter enforcement and verification of worker search behavior alone do not appear to reduce unemployment spells.[63] The main activity of WPRS has been to provide various forms of JSA to dislocated workers. In 1994, only 10,773 workers reported for at least one type of reemployment service under WPRS, and 9,990 completed at least one service. In 1998, fully 999,208 reported for at least one type of service, and 747,904 completed a service.[64] Evidence in Wandner, Messenger and Schwartz suggests that most of the JSA services provided under WPRS are a net addition to the total amount of JSA that UI claimants receive.[65]

The following back-of-the-envelope calculation suggests that, even at its 1998 scale, WPRS is unlikely to significantly influence the aggregate unemployment rate. Suppose that one million additional dislocated workers received some type of reemployment assistance in 1998 because of WPRS. Using Meyer's range of estimates of the effect of JSA on UI spells, this would be expected to reduce the total number of weeks of unemployment in the U.S. economy by 0.5 to 4.0 million weeks. In 1998, 6.21 million workers were unemployed during the average CPS survey week, producing a total of 322.9 million weeks of unemployment. Thus, the total share of unemployed

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[63]Ashenfelter, Ashmore and Deschenes (1998).

[64]These figures were calculated by the authors for the 50 U.S. states and D.C. from ETA 9048 data. The reemployment services include: orientation, assessment, counseling, job placement services and referrals to employers, job search workshops, job clubs, education and testing, and a small self-employment program. In principle, these figures are based on unduplicated counts of claimants, although it is likely that some states double counted claimants who received multiple services.

[65]Wandner, Messenger and Schwartz (1999, p. 9).

weeks reduced by WPRS relative to total weeks of unemployment would be .15 to 1.24 percent. These estimates imply that the absence of WPRS would have increased the unemployment rate from its actual level of 4.5 percent in 1998 up to a range from 4.51 to 4.56 percent -- increments so small that they would be quite difficult to tell from sampling error in the unemployment rate. Moreover, these calculations probably overstate the effect of WPRS on aggregate unemployment because: (1) the average service provided under WPRS is probably less intensive than the average JSA treatment studied in the literature; (2) the literature may overstate the effect of JSA participation on unemployment duration because non-participants may incur longer unemployment spells if participants find jobs sooner; and (3) WPRS may have increased the net number of claimants receiving reemployment services by less than one million. On the other hand, if WPRS leads to more stable job matches, it could have a larger effect than this back-of-the-envelope calculation suggests.

As a final check on the effect of the WPRS system on unemployment, we exploit the interstate variability in the timing of the implementation of the state programs. Specifically, using state-level data for the years 1994-98, we estimate the following equation:

(9) $u_{jt} = \beta_0 + \beta_1 \text{PROFILE}_{jt} + \nu_j + \gamma_t + \epsilon_{jt}$

where u_{jt} is the unemployment rate in state j and year t as estimated by BLS, PROFILE_{jt} is a dummy variable that equals 1 if WPRS was in effect in state j during year t, ν_j is an unrestricted state effect, and γ_t unrestricted year effect. The results are shown in column (3) of Table 10. Consistent with the back-of-the-envelope calculation described above, the regression estimate indicates a trivial effect of WPRS on the aggregate unemployment rate once year and state fixed effects are held constant.[66] For comparison, we report results without year effects (column 1) and without state effects (column 2). These models highlight the importance of controlling for year and state effects. First, because profiling was implemented gradually in the second half of the 1990s, when unemployment was falling, failure to control for time effects induces a spurious negative correlation between the profiling dummy variable and unemployment. Second, because the states that implemented WPRS early tended to be high unemployment states, failure to control for state fixed effects induces a positive bias between unemployment and profiling.

Although much research suggests that JSA helps to reduce the duration of unemployment spells, and is cost effective for the government, results in this section suggest that programs to provide JSA more broadly such as WPRS are unlikely to have much effect on the aggregate unemployment rate. This does not imply that improving the reemployment system is not a worthy goal, but it does highlight the fact that even cost-effective micro policy interventions of modest scale are unlikely to have much affect on aggregate outcomes, such as unemployment.[67]

Temporary Help Agency Workers, Wage Pressure and Frictional Unemployment

A more promising explanation for the possible improvements in the efficiency of job matching and increased labor market competition in the 1990s is the rapid growth of private sector employment intermediaries (especially temporary help agencies). Payroll employment in the temporary help services industry increased from under 0.5 percent of U.S. employment in the early

[66]This conclusion is quite robust to extending the sample period (e.g., back to 1990) to allow longer time series in each state to the underlying state fixed effects.

[67]Heckman (1994) makes this point in the context of job training programs.

1980s to 1.1 percent in 1989 and to just over 2.2 percent in 1998. Employment growth in the temporary help services industry accounted for 8.2 percent of net nonfarm payroll employment growth in the economic expansion of 1992 to 1998, as opposed to 4.1 percent in the comparable 1983-89 period.[68] Recent work by David Autor indicates that temporary help agencies are playing an increasingly important labor market role in screening employees and providing some forms of computer training.[69] The possible greater ease for firms of locating qualified and screened employees through intermediaries may lower hiring costs, reduce labor market bottlenecks, facilitate better employment matches, and put greater pressure on the wage-setting of incumbent workers.

The scale of operations of temporary help agencies, employee leasing firms, and private sector employment intermediaries appears to have increased to a level of possible significance for the operation of the aggregate labor market. For example, approximately 3.1 percent of employed workers in the February 1997 CPS supplement on contingent work indicated that they were on-call workers or employee of a temporary help agency or contract firm. Sharon Cohany reports that 60 percent of the 1.3 million self-reported employees of temporary help agencies in the February 1997 CPS were temporary workers for economic reasons.[70] If half of these "involuntary" temporary workers would have continued to be searching for work through unemployment in the absence of the expanded temporary help industry, the official unemployment rate would have been increased

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[68]These tabulations use data from the BLS Current Employment Statistics program. Household survey data from the CPS indicates a lower share of the work force employed in the personnel services industry. See Polivka (1996) for a discussion of these discrepancies.

[69]Autor (1999a). Also see Segal and Sullivan (1997b) and Autor (1999b) for useful analyses of the growth of the temporary services work force.

[70]Cohany (1998, Exhibits 1 and 9).

in 1997 by around 0.2 to 0.3 percentage points.

Beyond such possible direct effects of shifting workers from job search through unemployment to job tryouts through temporary jobs, the growth of labor market intermediaries may facilitate wage restraint by increasing the ability of firms to locate substitute workers. The increased ability to establish "contingent" work arrangements may also allow employers to raise wages for "marginal" workers working for temporary help agencies or other intermediaries without creating internal equity problems that in the past may have necessitated increasing the wages of incumbent employees to prevent morale problems.

We next present a preliminary and highly speculative initial attempt to examine whether increased access to contingent employment options, as proxied by size of the temporary help industry, may play a role in wage restraint (and thereby possibly affect the NAIRU). We take advantage of differences across U.S. states in the relative scale of operations of the temporary help industry. In particular, we ask whether states with a better developed temporary help industry at the start of the 1990s -- measured by the average state employment share in the temporary help industry from 1985 to 1989 -- experienced greater wage restraint in the 1990s. The share of total employment in the temporary help industry from 1985 to 1989 using data from County Business Patterns averaged 0.9 percent, and ranged from less than 0.3 percent in states like North Dakota and Idaho to more than 1.2 percent in California, Florida, and Delaware.[71] There appears to be substantial persistence over the past two decades in the relative importance of the temporary help

[71]We are grateful to David Autor for providing us with state data on employment in the temporary help industry from both County Business Patterns and the CPS. The correlation of state measures of the share of employment in the temporary help industry in County Business Patterns and the personnel supply services industry in the CPS ORG files averaged over the 1985-89 period is 0.85. We focus on the more precisely estimated measures from County Business Patterns.

industry across states.

Our approach is to estimate wage Phillips curves using state panel data on (composition-adjusted) wages from the CPS ORG files and state unemployment rates from the Bureau of Labor Statistics for the 1980 to 1998 period.[72] We examine whether states with a greater initial presence of temporary help at the start of the decade -- the average temporary help industry employment share over 1985-89 (THSP) -- experienced lower than expected wage growth at given measured unemployment rates in the 1990s. We control for pre-existing state differences in NAIRUs through state fixed effects and common macroeconomic factors through a full set of year dummies. Our basic estimating equation is of the form:

(10) $\Delta w_{jt} = \alpha_j - \beta u_{jt} + \delta\,(THSP_j{*}d90) + d_t + \epsilon_{jt}$

where Δw_{jt} is the change in the (composition-adjusted) mean log wage for state j from period t-1 to t; u_{jt} is the state unemployment rate; d90 is an indicator variable equal to 1 after 1989 and zero before; α_j and d_t represent full sets of state and year dummies.[73] The hypothesis of greater wage restraint in the 1990s from a larger initial presence of the temporary help industry at the start of the decade implies δ less than 0.

Table 11 presents some simple regressions of the form of equation (10) to examine the

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[72]We adjusted the wage date by first estimating micro regressions each year of the log wage on education, experience, sex, and race, and then calculated the average residuals for each state using the fitted regressions.

[73]We create composition-adjusted wages by running a cross-section log hourly wage regression using the CPS ORG samples in each year with standard set of control variables (education, age, race, and sex indicator variables and interactions). The adjusted wage for each individual is the sum of the mean national log hourly wage and the individuals residual from the composition adjustment regression. The state adjusted wage is the mean adjusted wage in the state using the CPS sampling weights. The wage samples are limited to wage and salary employees. Our findings are quite similar for unadjusted (raw) mean state log hourly wages. We are grateful to David Autor for assistance in preparing the adjusted wage data.

possible effects of greater temporary help and contingent work options of overall wage growth. We include specifications with both the level and log of the state unemployment rate used as a cyclical indicator, and with and without allowing the effect of unemployment on wage growth to change in the 1990s. We consistently find modestly lower wage growth in the 1990s conditional on unemployment and pre-existing state wage growth patterns (state fixed effects) for states with a greater share of temporary help employment at the start of the decade. The estimates imply a standard deviation increase (a 0.25 percentage point increase) in the share of the temporary help industry in the late 1980s has been associated with almost 0.2 percent a year lower wage growth.

The regressions in Table 11 are suggestive of some potential role for increased labor market competition from the growth of labor market intermediaries in possibly preventing bottlenecks and restraining wage growth in tight labor markets in the 1990s. The rapid expansion of the temporary help industry also coincides with the inward shift in the Beveridge curve since the late 1980s displayed in Figure 5 suggesting a possible impact on improved labor market matching.

To derive a rough estimate of the effect of the growth in the temporary help sector in the 1990s on the NAIRU, we first calculated the implied intercept shift in the wage Phillips curve based on the regression in Table 11 and the expanded presence of the temporary help industry in the 1990s, and then converted this intercept shift into a decline in the NAIRU based on the estimated slope of the wage Phillips curve. Specifically, we multiplied the estimated effect of temporary help employment on wages (-.656) in the regression in first column of Table 11 by the growth in this sector from 1989 to 1998 (1.1 percentage points), and then multiplied this figure by the inverse of the slope of the aggregate wage Phillips curve (1/.93) based on the CPS wage data in panel B of Table 4. However, because we measure the 1990s presence of the temporary help industry in the regressions in Table 11 using the industry employment share from 1985-89, and the scale of this variable doubled in the 1990s, we divided the resulting estimate by 2. This approach yields an estimate of a decline in the NAIRU over the past decade of 0.39 percentage points from the impact of the temporary help industry (and other improvements in labor market intermediation correlated with the prevalence of this sector). Thus, the impact of improvement in labor market matching and competition from labor market intermediaries may be as large as the impact of demographic changes on the NAIRU since the 1980s.

Union Power, Worker Insecurity and the Wage Structure

Private sector union membership has declined steadily since reaching a peak in the mid 1950s, with the sharpest decline occurring in the 1970s and 1980s. In 1973, 24.6 percent of private sector, non-agricultural workers in the United States belonged to a labor union or employee association similar to a union; by 1998 the private sector union rate fell to 9.6 percent.[74] Farber and Krueger find that only one-quarter of the decline in union membership between 1977 and 1991 occurred because of shifts in employment from high-union industries and occupations to low-union industries and occupations, and from demographic changes in this period.[75] Much evidence suggests that unions raise wages for their members above what they would be in the absence of unions.[76] The union wage gap is larger for workers with relatively low earnings potential (based on education and

[74]In contrast, union membership levels in the public sector increased from 23 percent in 1973 to 37.5 percent in 1998. See Hirsch and Macpherson (1999, tables 1c and 1f).

[75]Farber and Krueger (1993). Linneman, Wachter and Carter (1990) also find that industry shifts account for very little of the decline in union membership 1973-86.

[76]See, e.g., Lewis (1986).

experience) than for workers with higher earnings potential.[77] Unions are also likely to raise wages for nonunion members, as employers raise compensation to discourage workers from unionizing.[78] Farber and Krueger find that only one third of nonunion members desired union representation in the mid 1980s and early 1990s. Employers in most industries today face probably face little threat of unionization. The decline in union membership has been rather steady and persistent. This pattern may make it seem unlikely that weakened unions could have a discrete effect on wage setting practices. However, it is possible that the union movement passed a tipping point, where its support fell so low that employers felt virtually no threat effect of unions. A majority of workers in a bargaining unit must vote for a union in order for the unit to be unionized, so passing the 50 percent threshold is key. It should also be noted that union recognition elections fell discretely following the failed PATCO strike in August 1981.[79] Figure 9 displays the fraction of working time lost due to strike activity in the U.S. each year since 1948. Strike activity fell sharply in the 1970s and early 1980s, and has not recovered. In 1998 there were only 34 strikes involving 1,000 or more workers. The percent of 0.2 percent of work time lost to strikes in 1998 was only slightly more than 1997's record low.

David Card attributes about 10 to 20 percent of the rise in wage inequality among men between the mid 1970s and early 1990s to the decline of unions, because union membership fell

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[77]Card (1998).

[78]See Dickens (1986) for a model of wage setting in response to the threat of collective action.

[79]Unpublished tabulations of NLRB elections by Henry Farber.

most for groups of low-wage men, and because the union wage premium is largest for these groups.[80] If workers have become more timid in their wage demands in the 1980s and 1990s, the low level of private sector unionization is a prime suspect for why they might be more docile: they lack the representation to press for wage gains. Some evidence suggests that the rents workers receive in the union sector may be eroding. Table 12 reports the percentage growth in the ECI for private sector nonunion and union workers in selected periods. According to these data, compensation growth was slower for union members in the 1980s and 1990s than for nonunion members.[81] It is likely that the faster growth of wages in the nonunion sector in this period is at least partially related to the rise in skill premiums more generally, since private sector nonunion members tend to have higher education, on average. Interestingly, in this recovery and in the previous one, wage growth was notably slower for union members in the latter stages of the recovery than it was for nonunion members. Compensation growth appears particularly sluggish in the unionized manufacturing sector in 1994-98 (see column 3), probably reflecting trade pressures resulting from the Mexican currency crisis, and more recent Asian currency crisis. But it is worth mentioning that these large trade shocks have not coincided with a rise in overall wage inequality (see Figure 4). Indeed, the wage structure narrowed at a time when the U.S. exchange rate and trade balance shifted most dramatically.

[80]Card (1998). Card also finds that the decline in union membership has had relatively little effect on wage dispersion among women.

[81]For a detailed study of the union wage premium in the 1973-86 period see Linneman, Wachter and Carter (1990). Using CPS data, they find that the conventionally estimated union premium was relatively stable in this period. Hirsch and Macpherson (1999, table 2a) find that the private-sector union wage gap fell by 4 percentage points between 1989 and 1998, after controlling for education, experience, demographics, industry, and occupation.

Evidence on Worker Insecurity

Such diverse observers as Robert Reich and Alan Greenspan have argued that wage growth has been sluggish because of worker insecurity. While evidence on worker insecurity as a cause of subdued wage growth is likely to be as inconclusive as evidence on sociological causes of downward wage rigidity and unemployment, it is worth considering whether worker anxiety about job prospects has caused wage demands to moderate. The labor market has visibly changed. The proportion of workers who use a computer increased from 25 percent in 1984 to 50 percent in the mid 1990s.[82] Workers are also concerned about international trade, regardless of whether such concerns are justified: In a 1996 survey, over two thirds of the public reported that an important reason the U.S. economy is not doing better is because "companies are sending jobs overseas."[83] But technology is always evolving. And the economy has flourished the last few years despite the boom in imports. What is the evidence that high levels of worker insecurity are influencing labor market behavior?

First, national data do show a slight decline in job tenure and increase in displacement rates in the mid 1990s.[84] Farber, for example, finds 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."[85] In other work he finds that the rate of worker

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[82]Autor, Katz, and Krueger (1998, Table IV).

[83]Blendon et al. (1997). By contrast, only 6 percent of members of the American Economic Association agreed that companies sending jobs overseas is a reason the U.S. economy is not doing better.

[84]See Aronson and Sullivan (1997) for a comprehensive review of the literature.

[85]Farber (1997).

displacement, especially among mid-level occupations, was higher in 1993-97 than in 1983-87. But it is probably the case that the magnitude of the rise in job instability in the 1990s is modest compared to the public attention the issue received in the mid 1990s.

Second, worker surveys display some tendency for higher than expected job insecurity in the mid 1990s, although the post-1996 data suggest that worker self-reported job insecurity has returned to relatively low, business-cycle-peak levels. For example, Schmidt and Thompson analyze three surveys (Gallup, General Social Survey (GSS), and U.S. Department of Labor survey) and finds "evidence of a growth in workers' concerns about job security since 1977; however, the most recent data (from 1996 and 1997) indicate that workers are no more worried about job security than they were during earlier economic recoveries."[86] For example, in the June 1997 Gallup poll 10 percent of workers said they were very likely or fairly likely to lose their job or be laid off in the next 12 months, compared to 12 percent in October 1979.[87]

Also, recent survey data from the Institute for Social Research (ISR), which tracks families' financial security as part of its Consumer confidence measure, finds a sharp increase in the net fraction of families who think they are currently better off financially than a year earlier in 1997 and 1998.[88] Indeed, the latest data for 1998 reached the highest level since 1965. In the 1992-96 period, the net fraction that felt better off financially was below the corresponding figures for the first four years of the 1980s business cycle upswing. In view of the low inflation the past two years, these

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[86]Schmidt and Thompson (1997). Also see Aronson and Sullivan (1998).

[87]Schmidt and Thompson (1997).

[88]See <http://athena.sca.isr.umich.edu/scripts/mine/Tables/table4.asp>. Siskind and Krueger (1998) find that the net fraction of families that have actually experienced a rise in real income is a good predictor of the ISR variable.

data suggest that workers are not suffering from money illusion, which would be the mechanism causing workers to reduce their labor supply in the original Phillips curve model. On the whole, trends in self-reported worker security suggest that insecurity may have contributed to wage restraint in the mid 1990s, but the return of these survey measures to their previous business-cycle-peak levels suggests that worker insecurity was not abnormally high 1996-98. Perhaps coincidentally, nominal wage growth also rebounded in those years (see Table 3).

A final issue concerns the link between worker job insecurity and wage growth. Aronson and Sullivan evaluate the effect on wage growth of self-reported worker insecurity from the GSS and job displacement rates by estimating regional-level wage Phillips curves, augment to include these additional explanatory variables. Their results are rather mixed. When they use annual earnings as their dependent variable, they find that these two measures are negatively related to earnings growth, although only the displacement rate has a statistically significant coefficient. However, this relationship may only stem from hours worked. A more relevant outcome measure for understanding wage pressure which they examine is the hourly wage rate. Their results for hourly wage growth are less supportive of the view that measured job insecurity has an important effect on wage demands: both the GSS insecurity index and the displacement rate have statistically insignificant, though negative, effects.

Competitive Pressure and Rent Erosion

A related explanation -- and one that might cause feelings of job insecurity -- is that the inability of businesses to raise prices in the face of heightened competition (e.g., resulting from the steady deregulation of U.S. industries, share-holder pressure, increased international trade, and exchange rate shocks) has caused employers to seek ways to restrain factor costs.[89] If wages were above competitive market levels in some sectors, then reducing worker rents would be one way to cut factor costs. This may have caused slow wage growth, and contributed to the remarkably slow growth of intermediate goods prices in recent years. The search for more efficient production practices spurred on by heightened competition may also account for why productivity growth has been stronger the last few years than during the end of the previous recovery. While profit maximizing employers always have an incentive to minimize their costs, this explanation presumes that firms are not always minimizing costs.

Indeed, a growing literature suggests that firms share some of their product market rents with workers, perhaps because managers have a preference to share profits with workers.[90] When profits are squeezed, pay tends to be squeezed as well. A number of recent studies have found that employee pay tends to fall prior to a plant closing, tends to fall when a firm's profits declines, tends to fall and become more dispersed following deregulation, and is related to "exogenous" changes in industry import and export prices.[91] This findings suggest that in many sectors workers are paid a premium over their alternative wage, which provides some scope for competitive pressures to induce firms to reduce wages.

The *increased competition* story has three potential empirical shortcomings, however. First,

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[89]This point was recently emphasized by Krugman (1999). Gordon (1996) argues that a re-allocation of rents from workers to managers accounts for the rising inequality in the 1980s.

[90]See, for example, Katz and Summers (1989).

[91]See Jacobson, LaLonde, and Sullivan (1993) on plant closings and wage growth; Blanchflower, Oswald, and Sanfey (1996) on wages and profitability; Rose (1987) and Card (1996) on the effects of deregulation on the labor market; and Revenga (1992) and Abowd and Lemieux (1993) on trade shocks and wage setting.

we would expect competition to have intensified most in the goods producing sector in years since 1994, as a result of the Mexican and Asian currency collapses. In addition, the goods producing sector is a high-wage sector that is widely thought to pay workers rents. Clearly, the traded goods sector has been more affected than the service sector by international competition the last few years. The compensation growth figures in Table 12 do indicate that since 1994 wage growth has been less in the goods-producing sector than in the service-producing sector, by 3.5 percent. However, during the corresponding years of the 1980s recovery, wage growth was 2.5 percent less in the goods sector than service sector. Thus, the weaker growth of wages in the goods producing sector does not seem particularly unusual.

Second, under some variants of the increased competition story one would expect labor's share to fall. For example, if rents have simply been redistributed from workers to firms, labor's share would fall. Evidence on a drop in labor's share is mixed. Poterba finds that the modest fall in labor's share between 1992 and 1996 was in line with past cyclical relations.[92] But alternative measures of labor's share, which are based on BLS ECI data instead of NIPA compensation data, suggest a larger fall in labor's share.[93] In any event, it is unclear how persuasive the evidence on labor's share can be, since the erosion of wage rents due to increased competition would not lead to a fall in labor's share if firms are continually on their demand curve and if the production function is Cobb-Douglas.

A third strand of evidence concerns the consequences of job loss. If competition has eroded

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[92]Poterba (1998).

[93]See Krueger (1999). Because some of the salary drawn by incorporated business owners and corporate officers is counted in labor's share in the BEA data, it is possible that computations of labor's share based on BEA compensation data miss transfers between workers and owners.

wages, then one might expect workers who lose their jobs because of plant closings and mass layoffs to suffer greater wage losses in recent years than in earlier periods. But work by Farber does not indicate that the displacement carries with it a more severe loss of earnings now than in the past, which suggests that labor market rents have not changed substantially.[94]

Social and Distributional Consequences of Tight Labor Markets in the 1990s

The tight labor markets of the past several years have occurred after a period of two decades of slow growth in family incomes, widening wage and income inequality, and perceptions of substantial crime problems. Although real wages were sluggish in the early 1990s, the prolonged macroeconomic expansion of the 1990s finally appears to be having a payoff in terms of significant real and relative wage growth for low-wage workers since 1996 (as illustrated in Figure 4). The 9 percent real hourly wage growth for the 10th percentile worker from 1996 to 1998 is striking in the face of substantial increases in competition in the low-wage labor market associated with welfare reform and large increases in the labor force participation of single women with children.[95] Improvements in earnings for low-wage workers with children are even more significant when one takes into account the large expansion in the generosity of the earned income tax credit (EITC) from 1993 to 1996.[96] Expanded employment opportunities and increased real wages have also meant a sizeable rise in the mean real incomes of disadvantaged families (e.g., those in the bottom quintile

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[94]Farber (1997).

[95]See, for example, Bartik (1998).

[96]See Liebman (1998).

of the family income distribution) since 1993.[97] The current macroeconomic expansion has also been associated with a sharp decline in the crime rate.[98]

Earlier work by David Cutler and Lawrence Katz showed that structural labor market shifts against less-skilled workers in the 1980s prevented the macroeconomic expansion of the 1980s from improving the economic position of the disadvantaged by as much as would be predicted from the experience of previous postwar expansions.[99] We extend their analysis to examine whether the same pattern persists into the 1990s.

Figure 10 displays the actual official poverty rate for persons from 1959 to 1997, and two predicted poverty rate series using the historical relationship between poverty and median (or mean) income over the 1959 to 1983 period. We form predicted poverty rates using Cutler and Katz's earlier regressions of the poverty rate on contemporaneous macroeconomic indicators over the 1959 to 1983 period.[100] The poverty rate has remained at a much higher level since 1983 than would be predicted using historical macroeconomic relations from 1959 to 1983. The actual poverty rate declined by 1.8 percentage points in the expansion of 1983 to 1990 as compared to predicted declines in the poverty rate of 3.3 and 4.4 percentage points using median and mean family income,

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[97]U.S. Bureau of the Census (1999, Table F3). See Okun (1973) for a seminal analysis of the social benefits of a high-pressure economy.

[98]Recent work by Gould, Weinberg, and Mustard (1998) using panel data on U.S. counties find strong negative effects of increases in wages for low-wage workers and reductions in unemployment on crime rates, especially property crime rates.

[99]Cutler and Katz (1991).

[100]We use the same regressions for the 1959 to 1983 period used by Cutler and Katz (1991, Table 1, first 2 rows). The poverty rate is regressed on ratio of the poverty line to median income (or ratio of the poverty line to mean income), the inflation rate measured by changes in the CPI-U, and the unemployment rate for men aged 25 to 54 years.

respectively. The actual decline in poverty of 1.8 percentage points from 1993 to 1997 is also below the predicted decline of 2.6 percentage points from either forecasting equation. But the income measure used in the official poverty rate fails to include the gains to low-income families from the large expansion in the EITC since 1993. Experimental poverty measures including income from the EITC suggest a further 0.8 percentage point decline in the poverty rate from 1993 to 1997.[101] Adjusting for the impact of changes in the EITC, we therefore find that macroeconomic performance since 1993 appears to have reduced poverty by as much as would have been predicted from the pre-1983 relations.

Thus, taking the EITC into account, tight labor markets since 1993 appear to be generating more widespread benefits for the disadvantaged than was the case in the 1980s expansion. Recent research by Richard Freeman and William Rodgers also finds that metropolitan labor markets with sustained low unemployment in the 1990s have generated large improvements in the employment and earnings of the group of workers who have fared the most poorly over the last couple decades: less-educated young men, especially African American men.[102] Tighter labor markets than those prevailing in the 1980s may be necessary to partially offset the strong secular relative demand shifts against less-skilled workers (documented in the wage inequality literature) and provide economic improvements for disadvantaged workers.[103] The structural labor market changes that we have examined (e.g., improved efficiency of job matches) may enable tight labor markets in the near

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[101]U.S. Bureau of the Census (1998b, Table C4). Adjustments for the EITC expansion of the mid-1980s would have a much smaller effect on changes in poverty from 1983 to 1990.

[102]Freeman and Rodgers (1999).

[103]See Katz and Autor (1999).

future without creating major labor market bottlenecks in the absence of adverse supply shocks. But the very recent improvements in the economic situation of low-wage workers and low-income families certainly have not restored them to their levels of two decades ago. Another key issue is whether the strong labor market gains for new entrants (especially those moving off welfare) and other disadvantaged workers of the past few years have improved labor market connections enough to cushion the effects of the next slowdown in macroeconomic performance in the face of major social policy changes reducing cash assistance for the non-employed.

Conclusions

We conclude by summarizing the contribution of different labor market changes to the (1) 0.8 percentage point decline in actual unemployment rate from 1989 to 1998, and (2) the approximate 0.7 to 1.5 percentage point decline in estimates of the NAIRU since the mid 1980s. Table 13 presents our "best" estimates of the contribution to the decline of unemployment since the mid 1980s of each of the labor market factors we have investigated. In some cases, we provide a range because we are particularly uncertain of the effect the factor has had on unemployment. These estimates are based on our subjective interpretation of the empirical evidence we have been able to garner in this paper. The evidence suggests to us that demographic shifts and the rise of labor market intermediaries (e.g., the temporary help industry) are the main labor market changes that have contributed to the decline in unemployment. The emphasis we place on demographics and labor market intermediaries is also consistent with our finding that the incidence of short-term unemployment spells has declined markedly, while the incidence of long-term unemployment spells exceeds that achieved in past business cycle peaks.

It is interesting to speculate as to whether the labor market changes we have investigated are likely to have a transitory or more lasting effect on the natural rate of unemployment. As noted earlier, population and labor force projections through 2006 imply that demographic shifts will exert very modest downward pressure on the unemployment rate -- leading to perhaps a 0.05 percentage point decline in the unemployment rate. There is certainly no evidence in the labor force projections that unemployment will rise early in the next millennium because of demographic shifts. Likewise, labor market shifts brought about by innovations in the temporary help industry are likely to represent lasting structural changes in the efficiency of the labor market. The future role of the incarcerated population is difficult to predict, however, because it largely depends on the course of sentencing guidelines and practices.[104]

In addition to the other caveats that we have noted throughout this paper, an additional caveat to bear in mind in considering our relatively rosy forecast that the labor market shifts we have studied are unlikely to have only a transitory effect on unemployment is that we have ignored labor market shifts that may raise equilibrium unemployment. To the extent that other structural shifts in the labor market have taken place that raise equilibrium unemployment, then the factors we have identified may well be offset, and the likelihood that the unusually low unemployment rate in the late 1990s is only a transitory phenomenon due to favorable price shocks or other factors would increase. Furthermore, future progress in lowering unemployment will likely require new approaches to reducing the incidence of long-term unemployment (and nonemployment) spells among less-skilled and disadvantaged individuals.

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[104]To appreciate the importance of sentence lengths, note that the prison population has continued to expand while the crime rate has declined in recent years.

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Table 1: Unemployment and Employment Rates for Selected Groups (Percentage Points)

	Unemployment Rate							Employment/Population Rate		
				Duration (weeks)			Work Exp			
Year	All (1)	Men (2)	Women (3)	<5 (4)	5-25 (5)	>26 (6)	Unempl. (7)	All (8)	Men (9)	Women (10)
1968	3.6	2.9	4.8	2.0	1.4	0.2	12.4	57.5	77.8	39.6
1969	**3.5**	**2.8**	**4.7**	**2.0**	**1.3**	**0.2**	**12.5**	**58.0**	**77.6**	**40.7**
1970	4.9	4.4	5.9	2.6	2.1	0.3	15.3	57.4	76.2	40.8
1971	5.9	5.3	6.9	2.7	2.7	0.6	16.3	56.6	74.9	40.4
1972	5.6	5.0	6.6	2.6	2.4	0.7	15.5	57.0	75.0	41.0
1973	**4.9**	**4.2**	**6.0**	**2.5**	**2.0**	**0.4**	**14.3**	**57.8**	**75.5**	**42.0**
1974	5.6	4.9	6.7	2.8	2.4	0.4	17.9	57.8	74.9	42.6
1975	8.5	7.9	9.3	3.1	4.0	1.3	20.2	56.1	71.7	42.0
1976	7.7	7.1	8.6	3.0	3.3	1.4	19.1	56.8	72.0	43.2
1977	7.1	6.3	8.2	2.9	3.1	1.0	17.9	57.9	72.8	44.5
1978	6.1	5.3	7.2	2.8	2.6	0.6	15.9	59.3	73.8	46.4
1979	**5.8**	**5.1**	**6.8**	**2.8**	**2.5**	**0.5**	**15.8**	**59.9**	**73.8**	**47.5**
1980	7.1	6.9	7.4	3.1	3.3	0.8	18.1	59.2	72.0	47.7
1981	7.6	7.4	7.9	3.2	3.4	1.1	19.5	59.0	71.3	48.0
1982	9.7	9.9	9.4	3.5	4.6	1.6	22.0	57.8	69.0	47.7
1983	9.6	9.9	9.2	3.2	4.1	2.3	19.6	57.9	68.8	48.0
1984	7.5	7.4	7.6	3.0	3.1	1.4	17.4	59.5	70.7	49.5
1985	7.2	7.0	7.4	3.0	3.1	1.1	16.7	60.1	70.9	50.4
1986	7.0	6.9	7.1	2.9	3.1	1.0	16.2	60.7	71.0	51.4
1987	6.2	6.2	6.2	2.7	2.6	0.9	14.3	61.5	71.5	52.5
1988	5.5	5.5	5.6	2.5	2.3	0.7	12.9	62.3	72.0	53.4
1989	**5.3**	**5.2**	**5.4**	**2.6**	**2.2**	**0.5**	**12.9**	**63.0**	**72.5**	**54.3**
1990	5.6	5.7	5.5	2.6	2.4	0.6	14.7	62.8	72.0	54.3
1991	6.8	7.2	6.4	2.8	3.2	0.9	15.7	61.7	70.4	53.7
1992	7.5	7.9	7.0	2.6	3.3	1.5	15.8	61.5	69.8	53.8
1993	6.9	7.2	6.6	2.5	3.0	1.4	14.8	61.7	70.0	54.1
1994	6.1	6.2	6.0	2.1	2.8	1.2	13.5	62.5	70.4	55.3
1995	5.6	5.6	5.6	2.0	2.6	1.0	12.8	62.9	70.8	55.6
1996	5.4	5.4	5.4	2.0	2.5	0.9	11.7	63.2	70.9	56.0
1997	4.9	4.9	5.0	1.9	2.3	0.8	10.8	63.8	71.3	56.8
1998	**4.5**	**4.4**	**4.6**	**1.9**	**2.0**	**0.6**	**NA**	**64.1**	**71.6**	**57.1**

Notes: Data are for workers age 16 and older. Source: BLS web page (http://stats.bls.gov/). Work Exp. Unempl. is the percent of the labor force that was unemployed for at least one week during the year. Years in which the unemployment rate reached a trough are highlighted in bold.

Table 2: Unemployment Rates in Major Industrial Countries, 1950-1998[a]
Unemployment as a Percent of Total Labor Force

	1950s	1960s	1970s	1980s	1990s[b]	1992	1998[c]
United States	4.4	4.7	6.1	7.2	5.8	7.3	4.6
Canada	3.8	4.7	6.6	9.3	9.7	7.7	8.5
Japan	2.1	1.3	1.7	2.5	2.9	2.2	4.0
France	1.5	1.7	3.8	9.0	11.2	10.4	12.0
Germany[d]	4.9	0.6	1.9	5.7	6.6	4.6	7.5
Italy	7.2	3.8	4.7	7.5	10.5	8.3	12.2
United Kingdom	1.7	2.0	4.4	10.1	8.4	10.0	6.3
OECD Total	3.5	2.8	4.3	7.0	7.3	7.3	7.0

Notes:
[a]The reported numbers are OECD standardized unemployment rates.
[b]The average unemployment rate for 1990 to 1998.
[c] The average unemployment rate for the first 3 quarters of 1998.
[d]West Germany, data for the 1990s are from *Economic Report of the President* (1999, Table B-109).

Sources: Martin (1994, Table 1), Katz (1998, Table 1), and OECD (1999).

Table 3: Measures of Price Inflation and Nominal Compensation Growth, 1968-98

	Measures of Price Inflation			Measures of Compensation Growth		
Year	CPI-UX1	PCE	GDP	NIPA	Total ECI	Wage & Salary ECI
1968	3.9	4.0	4.4	9.3	NA	NA
1969	**4.5**	**4.1**	**4.7**	**7.5**	**NA**	**NA**
1970	4.8	4.7	5.3	6.7	NA	NA
1971	4.4	4.5	5.2	5.7	NA	NA
1972	3.0	3.5	4.3	7.1	NA	NA
1973	**6.3**	**5.4**	**5.6**	**8.9**	**NA**	**NA**
1974	10.0	10.1	9.0	11.0	NA	NA
1975	8.3	8.1	9.4	8.9	NA	NA
1976	5.7	5.7	5.8	9.2	NA	7.2
1977	6.4	6.6	6.4	7.6	NA	6.9
1978	6.8	7.3	7.3	9.4	NA	7.6
1979	**9.6**	**9.0**	**8.5**	**9.8**	**NA**	**8.7**
1980	11.2	10.9	9.3	11.0	9.6	9.1
1981	9.5	9.0	9.4	8.5	9.9	8.8
1982	6.1	5.8	6.3	6.8	6.5	6.3
1983	4.2	4.5	4.2	3.5	5.7	4.9
1984	4.3	3.8	3.8	4.6	4.9	4.2
1985	3.6	3.7	3.4	5.6	3.9	4.1
1986	1.9	2.8	2.6	4.7	3.2	3.2
1987	3.6	3.8	3.1	4.1	3.3	3.3
1988	4.1	4.2	3.7	4.2	4.8	4.1
1989	**4.8**	**4.9**	**4.2**	**2.9**	**4.8**	**4.1**
1990	5.4	5.1	4.4	6.0	4.6	4.0
1991	4.2	4.2	3.9	4.7	4.4	3.7
1992	3.0	3.3	2.8	4.5	3.5	2.6
1993	3.0	2.7	2.6	1.9	3.6	3.1
1994	2.6	2.4	2.4	1.9	3.1	2.8
1995	2.8	2.3	2.3	2.9	2.6	2.8
1996	3.0	2.0	1.9	3.8	3.1	3.4
1997	2.3	1.9	1.9	3.8	3.4	3.9
1998	**1.6**	**0.8**	**1.0**	**4.2**	**3.5**	**3.9**

Notes: NIPA compensation per hour is for the nonfarm business sector. The ECI data are for private industry. Years in which the unemployment rate reached a trough are highlighted in bold.

Table 4: Price and Wage Phillips Curves

Panel A: Price Phillips Curves

	CPI-UX1 Inflation			PCE Inflation		
	(1)	(2)	(3)	(4)	(5)	(6)
Constant	7.15	8.63	3.37	6.12	7.44	2.97
	(1.31)	(1.34)	(1.83)	(1.32)	(1.39)	(.78)
Lagged dependent variable	1.00*	1.00*	1.00*	1.00*	1.00*	1.00*
Unemployment rate	-1.00	-1.21	-0.53	-0.86	-1.05	-0.47
	(0.18)	(0.19)	(0.13)	(0.18)	(0.19)	(0.12)
Post-1988 dummy	-1.53	-7.35	-0.50	-1.71	-6.62	-0.56
	(0.49)	(2.48)	(0.44)	(0.50)	(2.57)	(0.41)
Post-1988 dummy X unemployment rate	--	0.95	--	--	0.85	--
		(0.40)			(0.41)	
P-value for joint F-test of post-1988 dummy and interaction	--	.002	--	--	.006	--
RMSE	1.10	1.00	1.18	1.11	1.04	1.11
Dubin-Watson statistic	1.70	2.23	1.62	2.07	2.52	1.80
Time period	1973-98	1973-98	1962-98	1973-98	1973-98	1962-98
N	26	26	37	26	26	37

Notes: CPI-UX1 data are from BLS, PCE data are from BEA. An asterisk indicates that the coefficient was constrained to equal 1. The mean [SD] for the dependent variable in columns 1-2 is 5.00 [2.54], in column 3 is 4.36 [2.45], in columns 4-5 is 4.85 [2.54], and in column 6 is 4.24 [2.43].

Table 4 (Continued)

Panel B: Wage Phillips Curves

Dependent Variable: Annual Percent Change in Hourly Compensation

	NIPA Compensation per Hour			ECI - Total Compensation	ECI - Wage and Salary		CPS Average Hourly Wage
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Constant	10.46 (1.47)	11.82 (1.57)	9.06 (0.93)	4.49 (1.02)	6.41 (1.04)	5.38 (1.10)	6.46 (1.32)
Lagged CPI-UX1 Price Inflation	1.00*	1.00*	1.00*	1.00*	1.00*	1.00*	1.00*
Unemployment Rate	-1.30 (0.20)	-1.49 (0.22)	-1.13 (0.15)	-0.60 (0.13)	-0.88 (0.14)	-0.78 (0.14)	-0.93 (0.18)
Post-1988 Dummy	-2.67 (0.56)	-8.02 (2.90)	-5.26 (2.77)	-.75 (0.37)	-1.25 (0.37)	-0.84 (0.40)	-0.96 (0.48)
Post-1988 Dummy x Unemployment Rate	--	0.87 (0.42)	0.52 (0.46)	--	--	--	--
P-Value for Joint F-Test of Post-1988 Dummy & Interaction	--	0.000	0.000	--	--	--	--
RMSE	1.24	1.17	1.26	.65	.74	.70	1.02
Durbin-Watson statistic	1.37	1.55	1.89	1.96	1.61	1.57	2.34
Time Period	1973-98	1973-98	1962-98	1980-98	1976-98	1980-98	1974-98
N	26	26	37	19	23	19	25

Notes: An asterisk indicates that the coefficient was constrained to equal 1. NIPA compensation per hour is from BLS, ECI total compensation and wage and salary data are from BLS, CPS average hourly wage data are from EPI. The mean [SD] of the dependent variable in columns 1-2 is 5.94 [2.82], in column 3 is 5.95 [2.54], in column 4 is 4.65 [2.05], in column 5 is 4.90 [2.11], in column 6 is 4.33 [1.83], and in column 7 is 5.01 [2.28].

Table 5: Wage Phillips Curves by Education Level, 1974-97

A: Regressions Using Overall Unemployment

	Less Than High School	High School	Some College	College +
	(1)	(2)	(3)	(4)
Constant	8.90 (2.26)	6.39 (1.40)	7.38 (1.88)	4.91 (2.21)
Lagged CPI Price Inflation	1.00*	1.00*	1.00*	1.00*
Overall Unemployment Rate	-1.39 (0.31)	-1.00 (0.19)	-1.12 (0.26)	-0.75 (0.30)
Post-1988 Dummy	-2.48 (0.80)	-1.14 (0.50)	-1.44 (0.66)	0.11 (0.78)
RMSE	1.68	1.04	1.40	1.65
Durbin-Watson statistic	2.19	1.96	2.35	2.30

B: Regressions Using Education-Specific Unemployment Rate

	Less Than	High School	Some College	College +
	(1)	(2)	(3)	(4)
Constant	4.64 (1.88)	2.94 (1.21)	5.29 (1.45)	2.08 (1.97)
Lagged CPI Price Inflation	1.00*	1.00*	1.00*	1.00*
Education-specific unemployment rate	-0.58 (0.18)	-0.60 (0.19)	-1.25 (0.29)	-1.07 (0.80)
Post-1988 Dummy	-0.13 (0.84)	-0.08 (0.55)	-0.35 (0.60)	1.16 (0.77)
RMSE	1.94	1.30	1.41	1.80
Durbin-Watson statistic	2.10	1.75	2.10	2.25

Notes to Table 5: Wage data are from EPI <http://www.epinet.org>, and were derived from May CPS files 1973-78, and ORG CPS files 1979-97. The education-specific unemployment rates are from U.S. Department of Labor (1997, Table 56) for 1974-96, and from <http://stats.bls.gov> for 1997. The education-specific unemployment rates pertain to March of each year. The 1997 data are for those 25 and over, whereas the 1973-96 data are for those age 25 to 64. To make the 1997 data comparable, we added the difference between the unemployment rates in the two sources in 1996 (when we have data from both series) for each education level. EPI only reports average wages for those with exactly a college degree and those with greater than a college degree separately. To derive the average wage of workers with a college and higher education, we calculated the weighted average of exactly college and higher than college wages, using .7 and .3 as weights, respectively. The mean [SD] of the dependent variable in columns 1-4, respectively, is 3.81 [2.98], 4.47 [2.24], 4.56 [2.66], and 5.16 [2.30].

Table 6: Wage Phillips Curves for Various Deciles of CPS Wage Distribution, 1974 - 98

Dependent Variable: Annual Change in Log Wage x 100

	10th Percentile	30th Percentile	Median	70th Percentile	90th Percentile
	(1)	(2)	(3)	(4)	(5)
Constant	9.74	9.44	5.89	7.47	7.19
	(3.27)	(1.70)	(1.79)	(1.92)	(1.82)
Lagged CPI-UX1 Inflation	1.00*	1.00*	1.00*	1.00*	1.00*
Unemployment Rate	-1.57	-1.40	-0.86	-1.02	-0.95
	(0.42)	(0.22)	(0.23)	(0.25)	(0.24)
Change Log Minimum Wage x 100	0.14	0.02	0.00	-0.04	0.00
	(0.08)	(0.04)	(0.05)	(0.05)	(0.05)
Post -1988 Dummy	-0.81	-1.66	-1.25	-1.65	-1.63
	(1.09)	(0.57)	(0.60)	(0.64)	(0.61)
RMSE	2.24	1.16	1.22	1.31	1.25
Durbin-Watson statistic	2.68	2.06	2.08	2.24	2.08
Implied URZERCG 1974 - 88	6.2	6.7	6.8	7.4	7.6
Implied URZERCG 1989 - 98	5.7	5.6	5.4	5.7	5.8

Notes: An asterisk indicates that the coefficient was constrained to equal 1. Sample size is 25 observations. Wage data were calculated from May CPS 1973-1978 and ORG CPS 1979-98 by Economic Policy Institute. Unemployment rate is from BLS. URZERCG is the unemployment rate associated with zero expected real compensation growth. The mean [SD] of the dependent variable for columns 1-5, respectively, is: 4.74 [3.47], 4.74 [2.38], 4.78 [2.32], 4.98 [2.43], and 5.36 [2.32].

Table 7: U.S. Unemployment Rates By Age Group, 1960-1998
Average Unemployment Rates in Percent for Each Period

Age Group	1960s	1970s	1980s	1990s[a]	1989	1998	1960-98
16-19	14.5	16.8	18.6	17.3	15.0	14.6	16.8
20-24	7.4	10.0	11.4	9.6	8.6	7.9	9.6
25-34	4.2	5.5	7.1	5.8	5.2	4.3	5.7
35-44	3.4	3.9	5.1	4.5	3.8	3.4	4.2
45-54	3.2	3.5	4.5	3.8	3.2	2.7	3.7
55-64	3.4	3.3	4.1	3.8	3.2	2.6	3.6
65 and over	3.5	3.9	3.1	3.5	2.6	3.2	3.5
Overall	4.8	6.2	7.3	5.9	5.3	4.5	6.0

Notes:
[a]The average unemployment rate for 1990 to 1998.

Source: U.S. Department of Labor, Bureau of Labor Statistics, <http://www.bls.gov>.

Table 8: The Age Composition of the Labor Force and Unemployment, 1960-2006
Precentage Points

Year	Unemployment Rate	Age Adjustment[a] 1960-98 LF Shares	Age Adjustment[a] 1979 LF Shares	Age-Driven Unemployment Rate[b]
1960	5.5	-0.63	-0.55	5.69
1963	5.7	-0.60	-0.58	5.74
1966	3.8	-0.35	-0.21	5.96
1969	3.5	-0.33	-0.21	6.04
1973	4.9	-0.29	-0.06	6.32
1976	7.7	0.06	-0.02	6.38
1979	5.8	0.00	0.00	6.40
1982	9.7	-0.02	-0.19	6.22
1984	7.5	-0.25	-0.29	6.12
1985	7.2	-0.30	-0.33	6.08
1989	5.3	-0.49	-0.44	5.90
1992	7.5	-0.65	-0.69	5.76
1995	5.6	-0.68	-0.67	5.72
1998	4.5	-0.69	-0.63	5.67
2006[c]	NA	NA	NA	5.62

[a]The age adjustment for each year is calculated by first creating an age-constant unemployment rate for each year that is a fixed-weighted average of the age-group specific unemployment rates for 7 age groups (16-19, 20-24, 25-34, 35-44, 45-54, 55-64, 65+) in each year using the average labor force shares in the base period as weights. The base periods are 1960-98 and 1979. The adjustment for labor force composition is then the difference between the actual overall unemployment rate and the age-constant unemployment rate. See equations (6) and (7) in the text. The age adjustments using 1960-98 weights are normalized to 0 in 1979.

[b]The age-driven unemployment rate (UA) tracks changes in unemployment predicted by changes in the age compositon of the labor force among our 7 age groups: $UA_t = \Sigma_j \omega_{jt} u_j$, where u_j is average group-specific unemployment rate for group j over the 1960-98 period and ω_{jt} is the labor force share of group j in year t.

[c]Age-group labor force shares for 2006 are based on BLS labor force projections.

Sources for Table 8: The data on unemployment rates and labor force shares by age group for 1960 to 1998 are from the U.S. Department of Labor, Bureau of Labor Statistics, <http://www.bls.gov>. The underlying data were kindly provided to us by Robert Shimer. The data on prjected labor force shares by age group for 2006 are from Fullerton (1997, Table 7).

Table 9: Likely Effects of Growth in Prison and Jail Population on Employment-Population Rate and Unemployment Rate for Men

	Number of Men in Jail or Prison	Incarcerated Relative to LF	BLS Male Emp./Pop. Rate	Adjusted Emp./ Pop. Rate	Civilian Noninstituional Unemployment Rate	Adjusted Unemployment Rate Assuming Labor Force Participation Rate of Incarcerated Would Be:			
						0.4	0.5	0.6	0.7
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1985	683173	0.011	0.709	0.706	0.070	0.070	0.071	0.072	0.073
1998	1716237	0.023	0.726	0.719	0.044	0.045	0.047	0.049	0.052
Change	1033064	0.012	0.017	0.013	-0.026	-0.025	-0.024	-0.023	-0.021

Notes: 1998 data are for June. Calculations assume that 35 percent of incarcerated men would be employed were they not incarcerated, based on Kling (1998). Sources: BLS; Sourcebook of Criminal Justice Statistics, 1997; Prison and Jail Inmates at Midyear 1998, Bureau of Justice Statistics (1999).

Table 10: Regression Estimates of the Effect of Worker Profiling on the Unemployment Rate; State-Level Analysis, 1994-98

Variable	(1)	(2)	(3)
Constant	5.676 (.086)	5.545 (.181)	5.683 (.069)
Profiling Dummy Variable (=1 if Profiling in Effect)	-.751 (.097)	1.015 (.380)	.008 (.163)
4 Year Dummies	No	Yes	Yes
50 State Dummies	Yes	No	Yes
Adj. R^2	.80	.12	.87

Notes: Dependent variable is BLS estimate of state unemployment rate. Mean of dependent variable is 5.07. Mean of profiling dummy variable is 0.81. Sample size is 255 state-by-year observations. Profiling dummy is derived from ETA 9048 data provided by Cindy Ambler of the U.S. Department of Labor.

Table 11: Wage Inflation, Unemployment, and the Temporary Help Industry
State Panel Data Analysis, 1980-98

	(1)	(2)	(3)	(4)
State Unemployment Rate (u_{jt})	-0.529 (0.038)	-0.544 (0.041)		
u_{jt}*d90		0.072 (0.069)		
Log State Unemp. Rate			-0.0379 (0.0026)	-0.0423 (0.0029)
log(u_{jt})*d90				0.0131 (0.0041)
$THSP_j$*d90	-0.656 (0.334)	-0.724 (0.339)	-0.579 (0.331)	-0.742 (0.333)
State Dummies	Yes	Yes	Yes	Yes
Year Dummies	Yes	Yes	Yes	Yes
Adj. R^2	0.69	0.69	0.69	0.69
Sample Size	950	950	950	950

The dependent variable is the change in the state (composition-adjusted) mean log hourly wage (Δw_{jt}). All regressions are weighted by state shares of total national employment in each year. Hourly wages are for wage and salary workers from the CPS ORG and are regression adjusted each year to account for differences in education, age, sex, and race. The (weighted) mean of the dependent variable is 0.043 with a standard deviation of 0.021. The (weighted) mean state unemployment rate is 0.0658 with a standard deviation of 0.020.

THSP = the mean employment share of temporary help industry employees from 1985 to 1989, based on establishment employment counts from County Business Patterns. The (weighted) mean of THSP is 0.0089 with a standard deviation of 0.0024.

d90 = 1 if year greater than or equal to 1990; 0 otherwise.

The data on state adjusted wages and temporary help employment from County Business Patterns were provided to us by David Autor. State unemployment rates are from the U.S. Department of Labor, Bureau of Labor Statistics.

Table 12: Nominal Growth in the ECI for Selected Sectors and Time Periods, Private Sector

Period	Nonunion (1)	Union (2)	Unionized Manufacturing (3)	Service Producing (4)	Goods Producing (5)
1979-89	74.1%	68.5%	NA	NA	NA
1985-89	*18.7*	*13.0*	*14.9*	*18.2*	*15.7*
1989-98	36.8	35.1	34.2	37.2	34.8
1994-98	*13.7*	*10.7*	*8.8*	*14.4*	*10.9*

Notes: Figures are nominal growth in ECI from fourth quarter of each year. Source: BLS web page: <ftp://ftp.bls.gov/pub/special.requests/ocwc/ect/echistry.txt>.

Table 13: Summary of Contribution of Labor Market Factors to Declining Unemployment Rate since Mid 1980s

Factor	Best Estimate of Effect	Comments
1. Demographic Shifts	0.4	Assumes age-specific unemployment rate unaffected by cohort size, and that education composition has no effect.
2. Changes in Labor Market Efficiency due to Developments in the Temporary Help Industry	0.0 - 0.4	Effect of temporary help on wage moderation is highly speculative.
3. Growth of Incarcerated Population	0.17	Estimated effect is 0.3 for men. We multiplied by male fraction of labor force to derive overall estimate.
4. Weak Backbone Hypothesis; Decline in Unions; Increased Competition	0.0 - 0.1	Difficult to assess.

Figure 1: Real Compensation Growth, 1980-98 (1988=100)

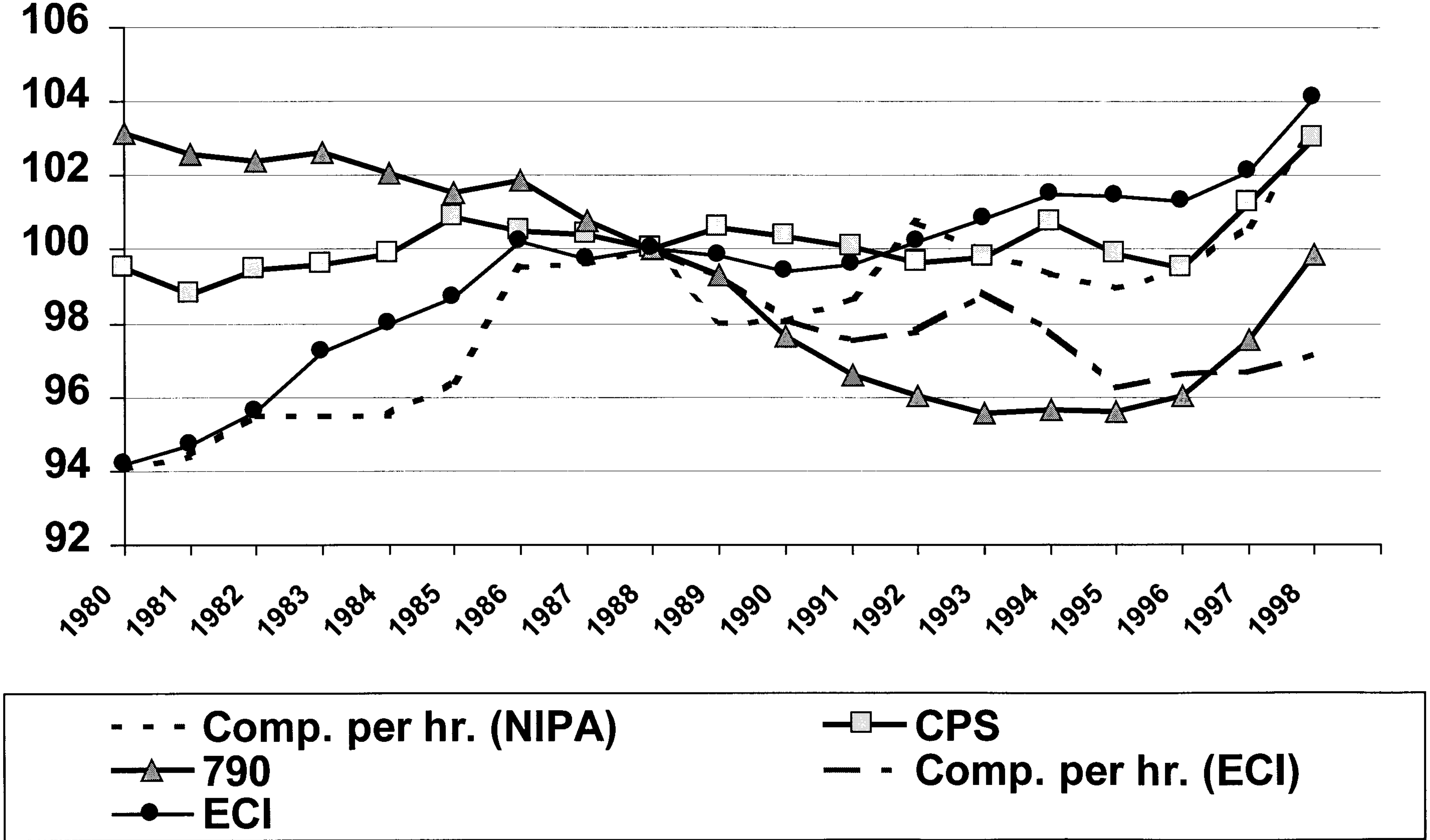
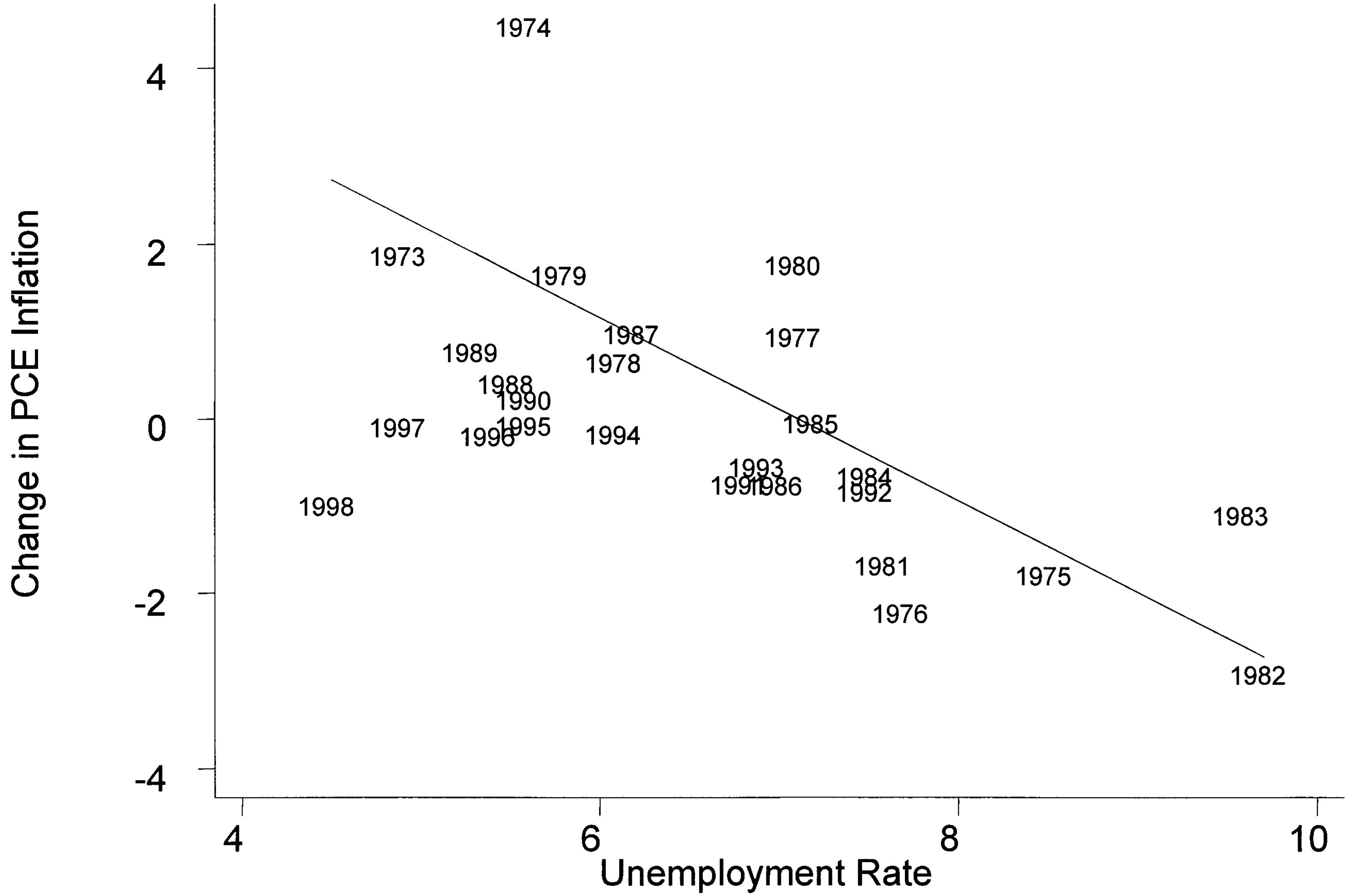


Figure 2: PCE Deflator Phillips Curve, 1973-98



Fitted line based on 1973-88.

Figure 3: Wage and Salary ECI Phillips Curve, 1976-98

Fitted line based on 1976-88.

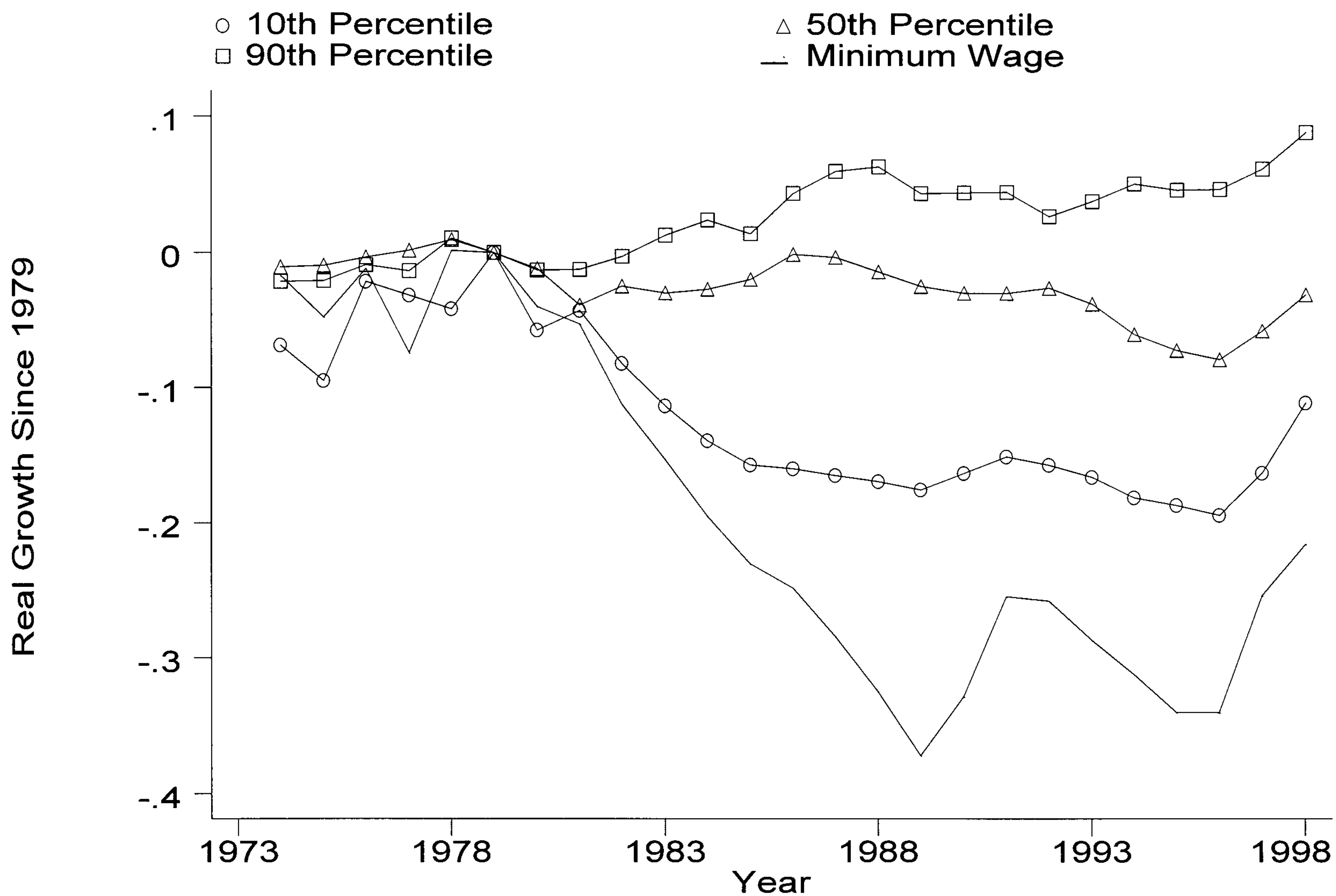


Figure 4: Log Real Wages and Minimum Wage, 1973-98 (1979=0)

Source: Authors' calculations using CPI-UX1 from BLS and wage data from EPI.

Figure 5: The U.S. Vacancy-Unemployment Relationship, 1960-98

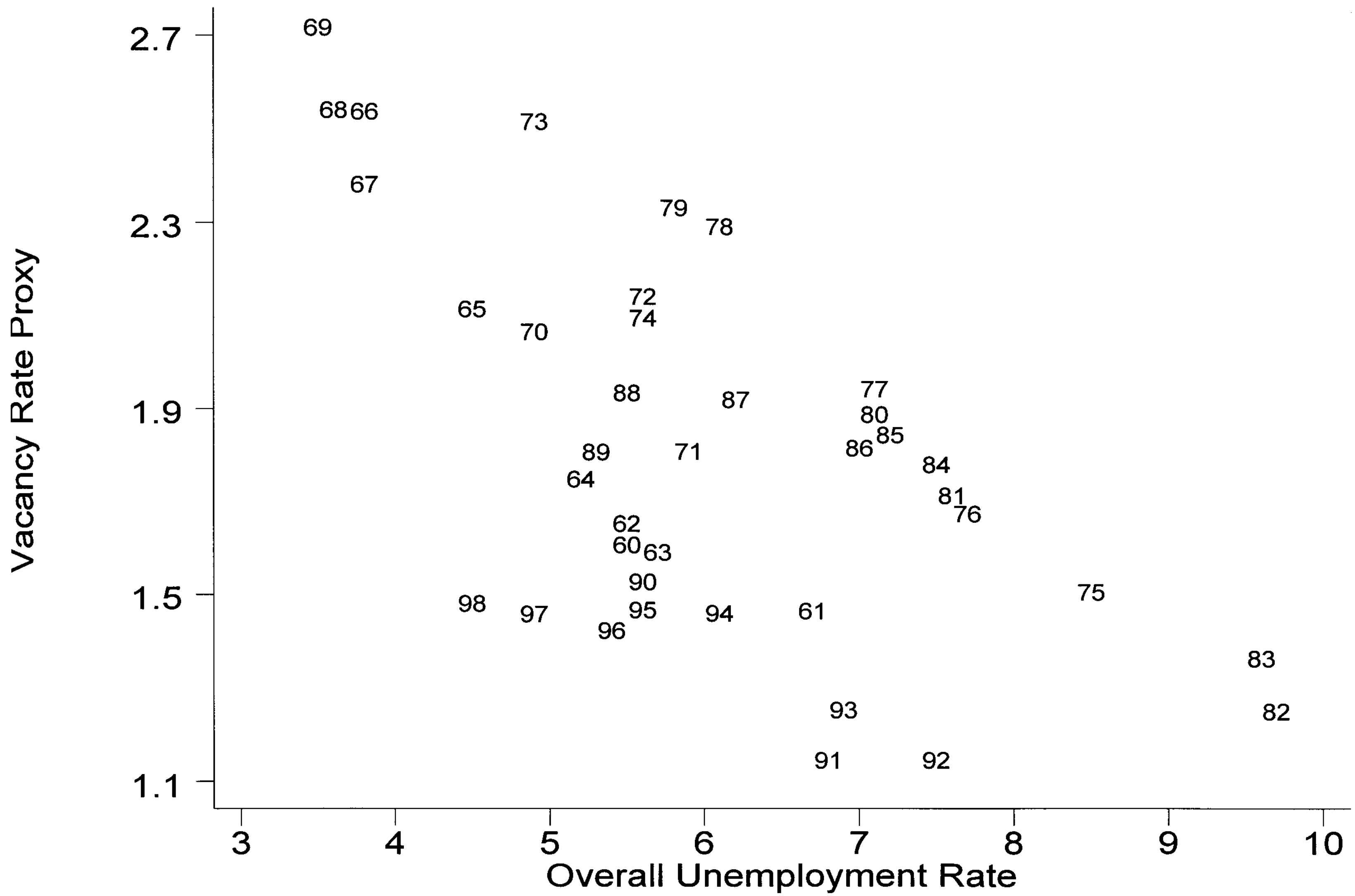


Figure 6: Labor Force Share of Young Workers, 1960-98

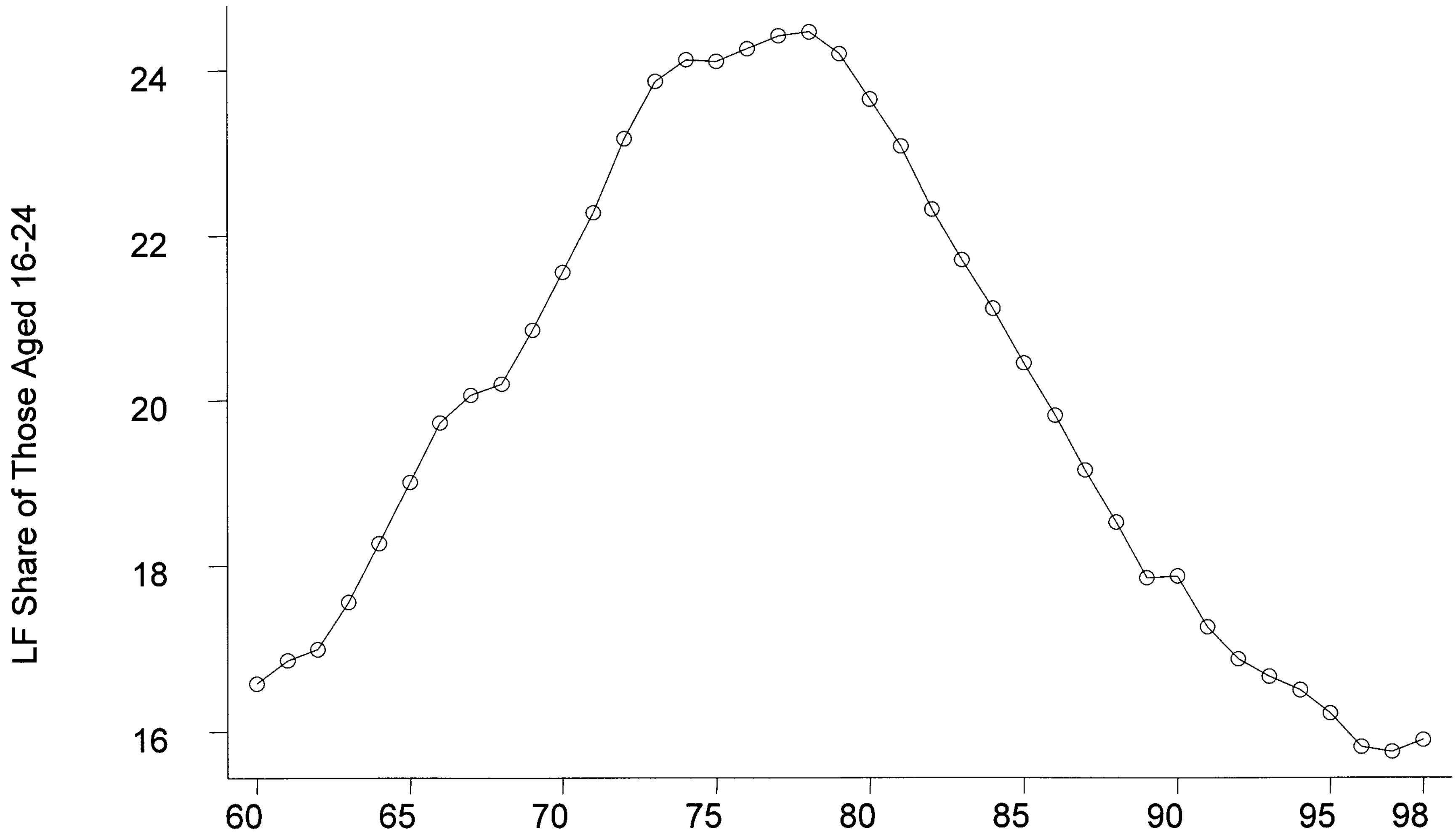
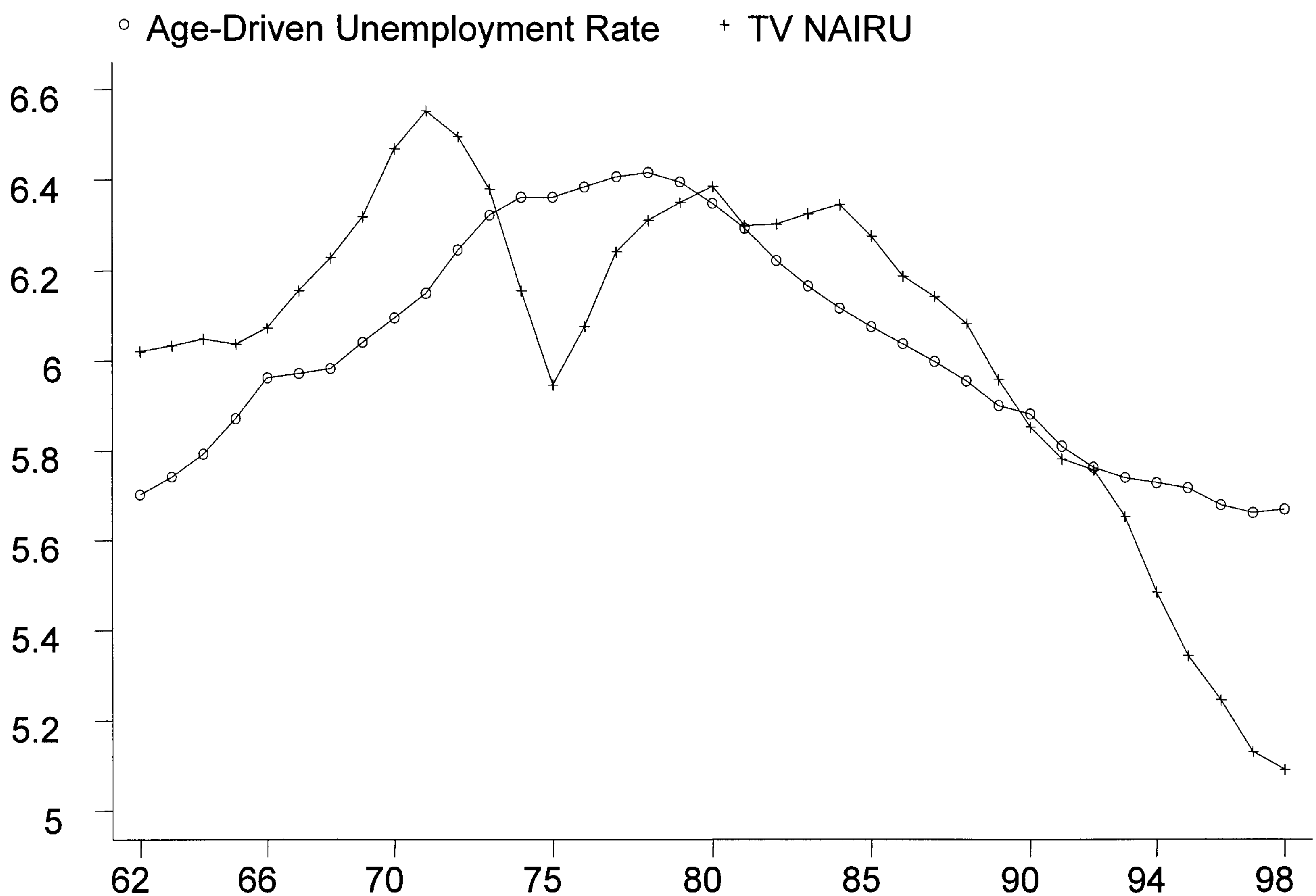
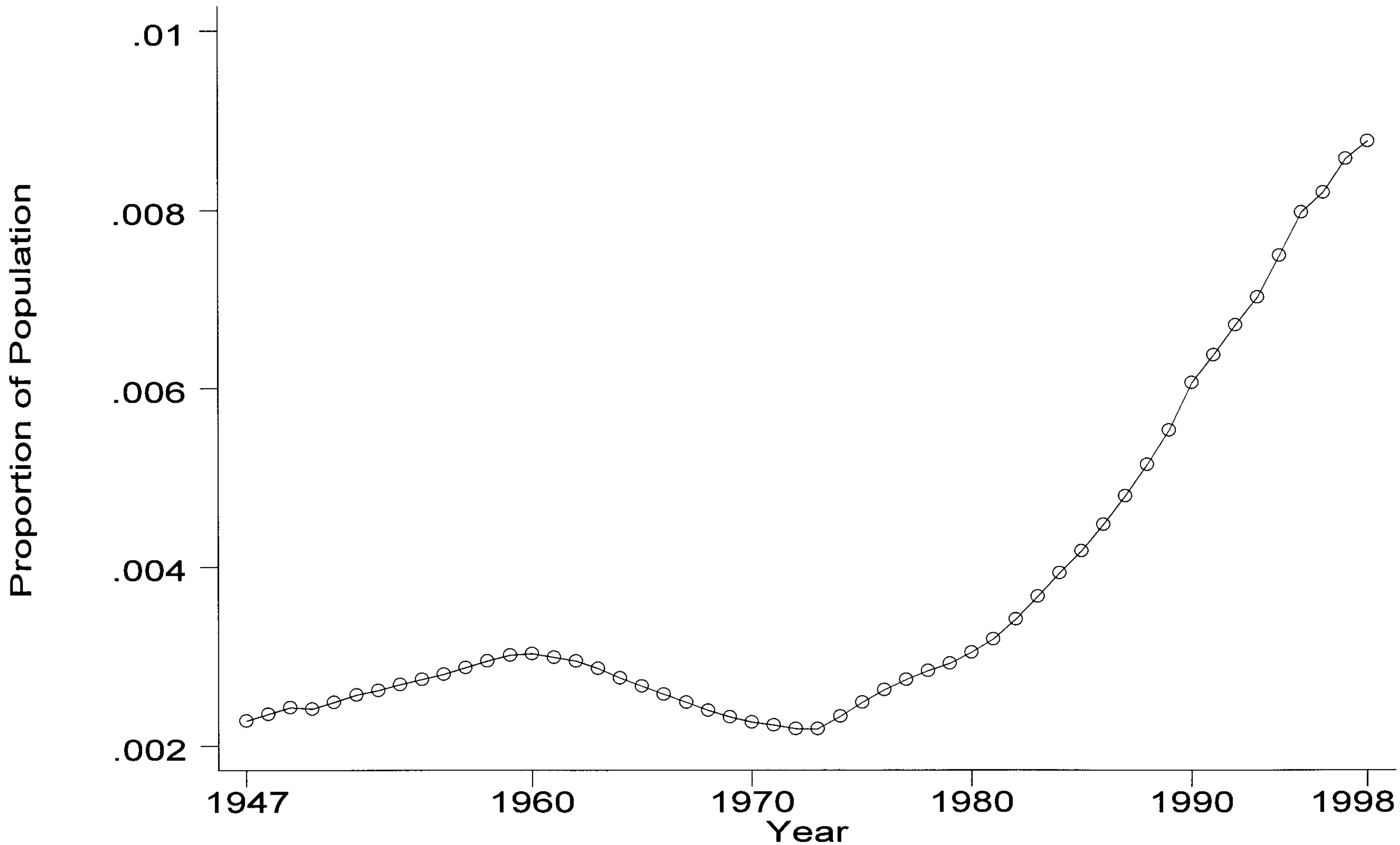


Figure 7: Age-Driven Unemployment Rate and Time-Varying NAIRU, 1962-98



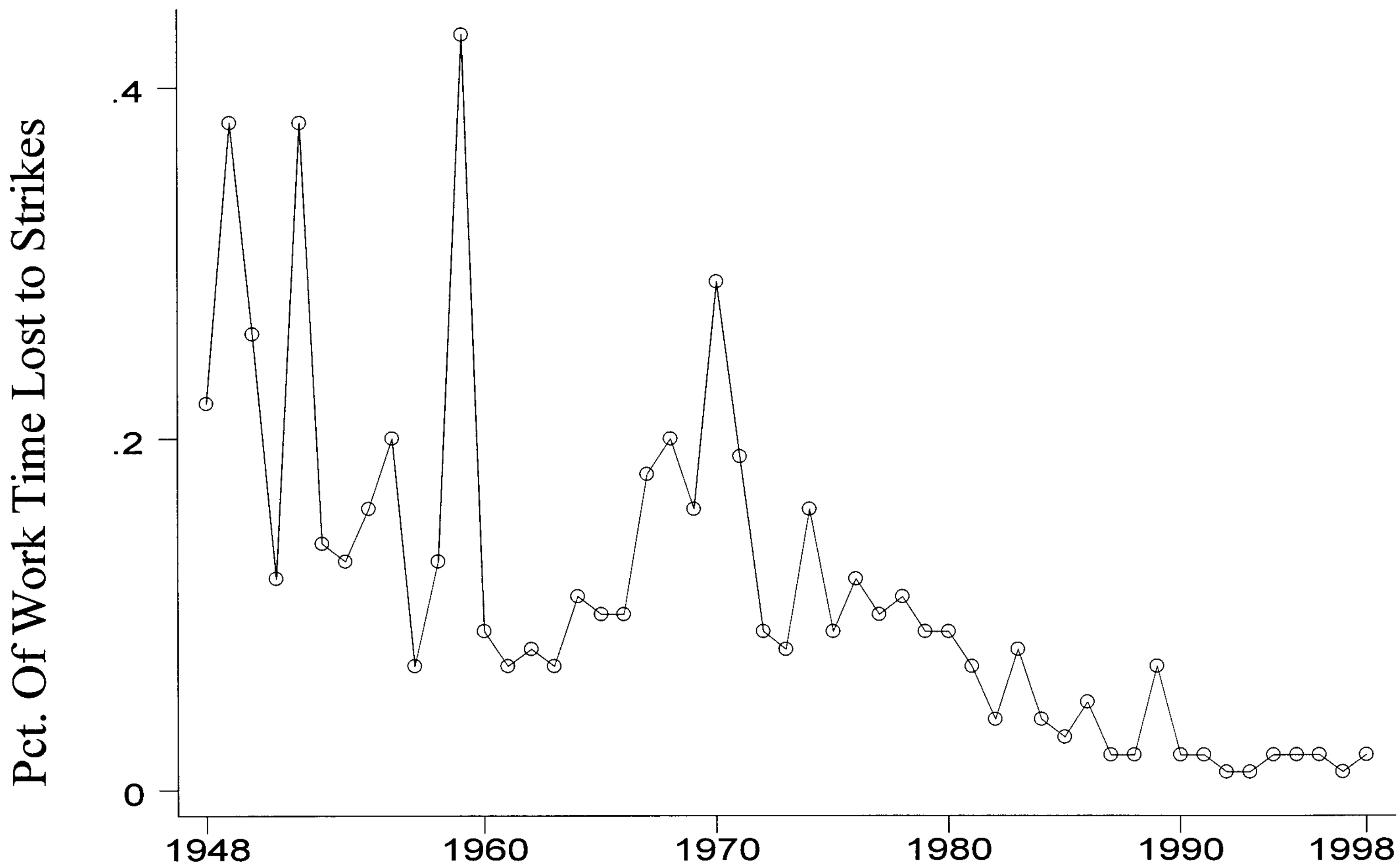
Sources: Age-driven unemployment rate is from Table 8. TV NAIRU was provided by Mark Watson.

Figure 8: Proportion of Population in Prison or Jail, 1947-98



Notes: Figure shows ratio of prison and jail population to civilian, noninstitutional population. Authors' calculations based on Freeman (1995), Bureau of Justice Statistics (1999), and BLS population data.

Figure 9: Percent of Estimated Working Time Lost to Strikes



Source: *Daily Labor Report*, February 11, 1999, p. D-2. Figures are for strikes involving 1,000 or more workers.

Figure 10: Actual and Predicted U.S. Poverty Rates, 1959-97