Unemployment and wage growth: Recent cross-state evidence

Daniel Aaronson and Daniel Sullivan

Introduction and summary

The current economic expansion, now the longest on record, has delivered the lowest unemployment rates in 30 years. Yet nominal wage growth has remained relatively contained. This failure of wages to accelerate more rapidly suggests to some a shift, or even a complete breakdown, in the historical relationship between unemployment and wage growth. However, looking across the years, the relationship between unemployment and wage growth has always been relatively loose, implying that it might take many years to conclusively identify even a significant change in the link between unemployment and wages.

In this article, we look across the states for more timely evidence of a change in the relationship between unemployment and wage growth. We find, however, that even in recent years, there is a relatively robust, negative relationship between state unemployment rates, properly evaluated, and wage growth. In particular, states in which current unemployment rates are lower relative to their long-run averages tend to have faster wage growth than those in which unemployment is higher relative to average. We do find some evidence that the sensitivity of wage growth to unemployment may have decreased in recent years, but we consider that evidence to be somewhat weak.

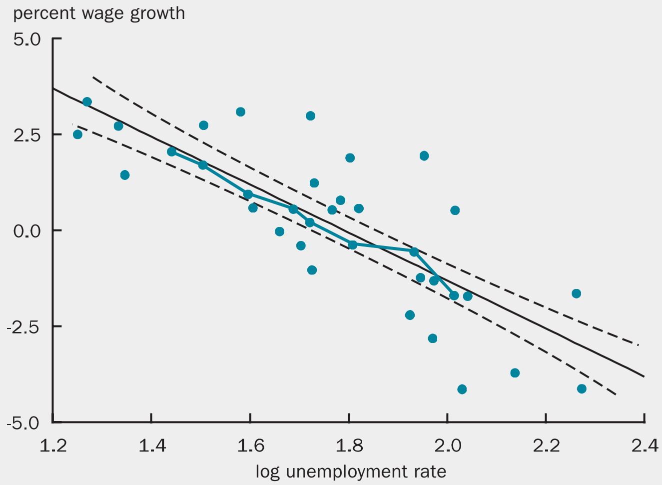
Before turning to the cross-state evidence, we briefly review some of the cross-year evidence that has led to speculation about a change in the relationship between unemployment and wage growth. That speculation has taken a number of forms, not all of which have been well reasoned. In particular, media analysts sometimes have characterized the lack of greater acceleration of nominal wages in the face of low unemployment as a failure of the "forces of supply and demand" in the labor market. But, the forces of supply and demand have direct implications not for *nominal* wage growth, but rather for *real*, or inflation-adjusted, wage growth.[1] Indeed, because nominal wage growth depends on the level of price inflation, which in turn depends on monetary policy, there is little reason to expect a long-run link between the level of nominal wage growth and unemployment. So it is not surprising that the statistical relationship between nominal wage growth and unemployment discovered by Phillips (1958) disappeared long ago.[2]

A more serious question is whether there has been a change in the relationship between unemployment and the growth of wages relative to expected inflation. A rough indication of the time-series evidence on this question can be gleaned from figures 1 to 3, which are scatter plots of annual data on the excess of wage growth over the previous year's price inflation versus the natural logarithm of the annual unemployment rate. In each case price inflation is measured by the change in the log of the annual Consumer Price Index. The three figures differ, however, in their measures of wage growth.[3] In figure 1 wage growth is the change in the log of the annual average of the Bureau of Labor Statistics' (BLS) Average Hourly Earnings (AHE) series. This closely followed monthly wage measure is limited to the wage and salary earnings of the approximately 80 percent of private industry workers who are classified as production or nonsupervisory workers. In figure 2 wage growth is derived from the hourly compensation measure from the BLS's productivity and cost data (Hourly Comp). This measure captures most wage and nonwage forms of compensation paid to all workers in the business sector and thus provides a superior measure of the compensation associated

Daniel Aaronson is an economist and Daniel Sullivan is a vice president and senior economist at the Federal Reserve Bank of Chicago. *The authors would like to thank Abigail Waggoner and Ken Housinger for research assistance and seminar participants at the Federal Reserve Bank of Chicago for helpful comments.*

FIGURE 1

Growth in average hourly earnings minus lagged CPI inflation versus unemployment



Note: Dashed lines indicate 90 percent confidence bands.

with an average hour of work. Finally, in figure 3 wage growth is given by the increase in the average value of the BLS's Employment Cost Index (ECI). This measure also reflects both wage and benefits costs for private employers and, in addition, adjusts for variation in the industrial and occupational mix of the labor force. Unfortunately, it only became available in 1983. So there are relatively few observations in figure 3.

The relationships depicted in figures 1–3 are analogous to the wage equations in some macroeconometric models.[4] They can be motivated by assumptions that 1) wages are set to exceed expected inflation by an amount that depends on the unemployment rate, and 2) expected inflation is equal to the level of inflation in the previous year. Of course, wage equations in actual macroeconometric models are considerably more elaborate than what is represented in the figures. In particular, they use quarterly rather than annual data and they allow for more complicated dynamics. They also include other variables, such as the level of productivity, that influence wage growth.[5] Nevertheless, figures 1–3 illustrate the basic nature of the time-series evidence on the relationship between wages and unemployment.

In at least the first two figures, there is a loose, but reasonably clear, negative correlation between unemployment and wage growth in excess of lagged inflation. The least squares regression lines shown in the figures all slope downward with elasticities that range from –0.044 for AHE to –0.055 for Hourly Comp to –0.013 for the ECI. The estimated standard errors of these estimates are 0.0095, 0.0090, and 0.0090.[6] Thus, if the relationships are stable over time, one can be reasonably confident that the true coefficients are different than zero for AHE and Hourly Comp. For the ECI, the evidence is less clear-cut, in part, perhaps, because the available sample is much shorter. Of course, in all three figures there is a sizable spread of values around the estimated line; the relationship between unemployment and wage growth is far from tight.

FIGURE 2

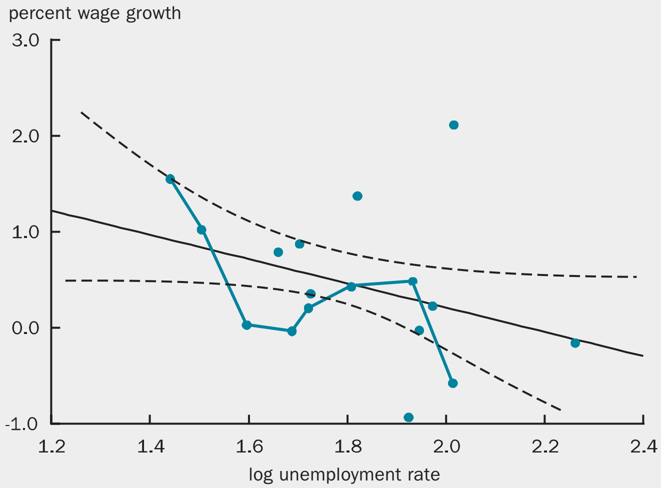
Growth in hourly compensation minus lagged CPI inflation versus unemployment

Note: Dashed lines indicate 90 percent confidence bands.

The data for the current expansion are highlighted in figures 1–3 by a line connecting the values from 1992 to 1999, when the unemployment rate was falling from 7.5 percent to 4.2 percent. Evidently, the extent of departure of recent data from historical patterns depends a good deal on the measure of wage growth. On the one hand, the recent AHE data shown in figure 1 have stayed remarkably close to the typical pattern. AHE growth from 1992 to 1999 did not differ from the estimated regression line by more than four-tenths

FIGURE 3

Growth of Employment Cost Index minus lagged CPI inflation versus unemployment



Note: Dashed lines indicate 90 percent confidence bands.

of a percentage point, while in some earlier years the deviation had been as much as 2 percentage points. On the other hand, the more comprehensive Hourly Comp data shown in figure 2 have departed fairly significantly from expectations over much of this expansion. In particular, the growth of Hourly Comp was a percentage point or more below expectations each year from 1993 to 1997. Though the data for the last two years have returned to the predicted line, the cumulative loss of wage growth over the expansion has been significant. Finally, the recent ECI data shown in figure 3 have also departed rather significantly from historical norms. As with the Hourly Comp data, ECI growth was significantly below expectations early in the expansion. But growth actually exceeded expectations late in the expansion, so the cumulative difference in wage growth is considerably less.

The differences in the performance of the three wage measures reflects the differing pattern of growth in wage and nonwage compensation over the sample periods as well as the coverage of the measures. Over most of the period covered in the graphs, nonwage compensation grew faster than wage compensation. For instance, according to data from the National Income and Product Accounts, the fraction of employee compensation paid in the form of wage and salary accruals fell from 92.4 percent in 1959 to 83.4 percent in 1980 to a minimum of 81.0 percent in 1994. Since 1994, however, the fraction of compensation paid in the form of wages and salaries has increased to 83.9 percent (in 1999), holding the growth of total compensation measures such as Hourly Comp and the ECI below that observed for AHE. In addition, over much of the period covered in the figures, wage growth has been more rapid for the more highly skilled, who are less likely to be classified as production and nonsupervisory workers and thus less likely to be covered in AHE.

Taken together, the evidence in figures 1–3 for a significant recent shift in the relationship between unemployment and expected real wage growth appears to us to be relatively weak. As we have noted, when one focuses on the more comprehensive Hourly Comp measure, the departures from expectations over this expansion have at times been relatively great. But, such departures are far from unprecedented. In earlier years, the data have strayed further from expectations only to return to the basic pattern of low unemployment being associated with higher growth of wages relative to lagged inflation. Of course, the evidence in figures 1–3 also does not rule out a significant shift in the relationship between unemployment and inflation. Unfortunately, given the looseness of the historical relationship, it would take many years to confidently identify even a relatively large change in the relationship.

Some shift in the relationship between unemployment and wage growth would not be terribly surprising. Among the many changes in the labor market in recent years, the general drop in the level of job security, the aging of the work force, its higher levels of education, the growth of temporary services employment, the use of fax machines and the Internet in job search, and even the increase in the prison population could each be changing the relationship between unemployment and wage growth.[7]

Moreover, the theoretical basis for the relationships depicted in the figures is somewhat loose, which at least suggests the possibility of instability. The assumption that expectations of inflation are equal to last year's level of inflation is clearly ad hoc. Moreover, though a relationship between expected real wage growth and unemployment can be motivated by economic theory, such theory doesn't necessarily imply a special place for the standard civilian unemployment rate.

Indeed, in the simplest model of a competitive labor market, unemployment is not a well-defined concept because there is no distinction between workers being unemployed and out of the labor force. Rather, in that model wages adjust to clear the market, and workers for whom the equilibrium wage is below the alternative value of their time simply choose not to work. The competitive model would replace the relationship in figures 1–3 with a standard, aggregate labor supply curve. This is analogous to the relationship in figures 1–3, but with employment, rather than unemployment, as the variable predicting wage growth. Of course, (deviations from trend) fluctuations in these variables are highly correlated, so unemployment may predict expected real wage growth reasonably well even if employment is the theoretically preferable measure.

Economic theorists have gone beyond the simple competitive framework to formulate models in which unemployment is involuntary and in which the unemployment rate is related to wages. One class of such models explicitly recognizes the importance of the labor market search, the complex process by which workers desiring jobs and firms desiring workers are matched to each other. In such models, some workers and firms are left unmatched and thus unemployed or with vacancies. Moreover, in search models with wage bargaining, workers have greater bargaining power when the unemployment rate is low, since turning down a job offer with a low wage is more palatable when the unemployment rate is low.[8] This generates a link between unemployment and wages.

Another class of models in which unemployment can be involuntary and in which the unemployment rate is connected to wages incorporates what are known as efficiency wage considerations. In such models, involuntary unemployment arises because firms rationally choose to pay wages above market clearing levels in order to induce effort or reduce turnover.[9] For instance, when it is difficult to monitor workers' effort, firms may want to ensure that workers truly fear being discharged after having been found to exert insufficient effort. This will be the case if wages are high enough that workers prefer working to being unemployed. In such models, wages cannot fall enough to clear the labor market because if they did so, workers would have insufficient incentive to put forth appropriate effort. The connection of wages to unemployment emerges because when unemployment is low, discharged workers will face less time out of a job. Thus, wages need to be further above the value of workers' nonmarket uses of time to induce the same level of effort.

Even in search and efficiency wage models, the standard unemployment rate may not be the variable most directly related to wages.[10] Rather, in both classes of models, the exit rate, the rate at which workers leave unemployment, is a more direct measure of the cost to workers of becoming or staying unemployed than the unemployment rate itself, which also depends on the rate of entry into unemployment. Of course, since the exit rate and the overall unemployment rate are highly correlated, the latter may predict wages reasonably well even if the former is the variable that is truly linked to expected wage growth.

Even if one accepts the use of an unemployment rate as the measure of labor market conditions, there is still the question of which unemployment rate to use. The standard measure imposes requirements that nonemployed workers be available for work and have made an effort to find work in the last month. However, some out-of-the-labor-force workers, for example, those who say they want a job, are relatively similar to the unemployed and may exert an influence on wage growth. Conversely, some of those who are unemployed, such as those who have been unemployed for long periods, may be more similar to the out-of-the-labor-force pool.[11] Ultimately, which measure best captures the labor market forces influencing wages is an empirical question, the answer to which could be changing over time.

In this article we look for evidence of such changes in the cross-state relationship between unemployment and wage growth. Previous work has demonstrated a relationship between unemployment and wage growth across states that is analogous to that in time-series data.[12] The basic assumption underlying this work is that inflation expectations are approximately the same for all states in a given year. Given that the U.S. has a single, national monetary policy, this is plausible, though clearly one could imagine deviations from this assumption. If inflation expectations are constant across states, differences in wage growth across states are unaffected by inflation expectations. Similarly, to the extent that other variables, such as productivity, that affect wage growth are constant across states in a given year, comparisons of states' wage growth rates are also unaffected by these variables.

A major advantage of the cross-state approach is the greatly increased number of degrees of freedom available from the wide variation in state unemployment rates. This makes it possible to estimate the response of wage growth to unemployment separately for relatively short periods. Thus, it may be possible to identify changes in that response that would take many years of time-series data to uncover.

Despite its attractions, the cross-state approach requires some care in its implementation. In particular, differences across states in unemployment rates persist for long periods, reflecting differences in factors such as demographics, industry composition, and generosity of social insurance that don't necessarily translate into differences in wage growth. The cross-state approach can allow for such persistent differences across states by employing multiple years of data. The empirical analysis then amounts to measuring the tightness of a state's labor market by its deviation from its own average unemployment rate over the entire sample period.

Deviations from mean unemployment rates reveal a different view of where labor markets are tight than the simple level of unemployment. For example, Wisconsin unemployment averaged 3.1 percent in 1999, six-tenths of a point less than in Michigan where unemployment averaged 3.7 percent. But, Michigan has historically had much higher unemployment than Wisconsin. For instance, over the 1980–99 period, Michigan's average unemployment rate was 8.4 percent, versus 5.7 percent in Wisconsin. Thus, Michigan in 1999 was 4.7 percentage points below its average, while Wisconsin was only 2.6 points below its average. Our empirical analysis finds that such unemployment-deviation measures are a better guide to labor market tightness than the standard unemployment rate.

That empirical work confirms the negative cross-state correlation between unemployment and wage growth found by previous researchers for the years 1980–99. We also find that the elasticity of wages with respect to unemployment has fallen over successive five-year intervals, a result that does not seem to be the result of a compositional shift toward college-educated workers. However, we regard this evidence of a weakened relationship between unemployment and wage growth as itself somewhat weak. In particular, when we estimate an elasticity for each year from 1980 to 1999, there is enough year-to-year variability that a downward trend in the magnitude is not obvious. Rather, the extent of change observed in the relationship depends on the necessarily arbitrary decision of where to draw the line between periods. Moreover, if one considers the response of wage growth to the level of unemployment, rather than its logarithm, there is very little evidence of a recent change in the sensitivity of wage growth to unemployment.

A recent study by Lehrman and Schmidt (1999) of the Urban Institute for the U.S. Department of Labor suggests that the level of unemployment across states is not related to wage growth. We believe those authors' results differ from ours for at least the following reasons: their measure of unemployment is not well matched in time to their measure of wage growth, their procedure does not allow for differences across states in other factors that affect wage growth, and their statistical procedure, which does not impose a linear relationship between wage growth and unemployment, has high variability with only 50 state observations. Thus, we agree with Zandi (2000), who concludes that the results of Lehrman and Schmidt (1999) prove little about the relationship between unemployment and wage growth.[13]

Our main results concern possible changes in the sensitivity of wage growth to unemployment. But we also briefly examine how the level of wage growth for particular levels of unemployment may have changed over time. We find that the levels of real wage growth associated with high, medium, and low unemployment rates have been reasonably constant in recent years. The real wage growth levels associated with typical values of unemployment were somewhat higher in the early 1980s, but since then have been relatively constant, with the wage growth associated with high unemployment rates actually rising somewhat in the late 1990s. Similarly, the unemployment rate associated with the average rate of real wage growth fell after the early 1980s, but has been relatively constant since then.

Because, as we noted, there is no compelling theoretical reason for the standard civilian unemployment rate to be the best measure of labor market conditions for predicting wage growth, we investigated a number of alternative measures of labor market tightness. These included the employment-to-population ratio, broader and narrower measures of unemployment, separate measures of short-term and long-term unemployment, and a measure of the exit rate from unemployment. Most of these measures predict wage growth about as well as the standard unemployment rate. Most also show the same decline in the magnitude of their elasticity with respect to wage growth that we observe over five-year intervals for the unemployment rate. The decline in the coefficients associated with the exit rate and short-term unemployment measures are, however, more severe. Such findings suggest that further work on improved measures of labor market tightness may be fruitful.

Finally, our results have implications for inflation forecasting, a task that plays an important role in the formulation of monetary policy. One of the most widely used approaches to such forecasting has been the short-run, or expectations-augmented, Phillips curve.[14] This forecasting method, which relates the change in price inflation to the level of the unemployment rate and other variables, can be derived from the kind of expected real wage growth relationship depicted in figures 1–3 along with an equation that relates price inflation to wage inflation and other variables.[15] Recently, there is evidence that typical short-run Phillips curve specifications have systematically overforecasted inflation.[16] Our results point toward the conclusion that this failure of the forecasts is most likely attributable to the part of the model linking price inflation to wage growth rather than to a change in the relationship between expected real wage growth and unemployment. This is consistent with the findings of Brayton et al. (1999), who show that including additional variables related to the markup of prices over wages helps to stabilize the Phillips curve.

Data

Our main results are based on two data sources. The first is the annual averages of the standard, monthly, state-level unemployment rates reported by the BLS. The second source is a measure of state-level, demographically adjusted wage growth that we construct from the micro data of the outgoing rotations of the *Current Population Survey* (CPS). The CPS, which is the source for such well-known statistics as the unemployment rate, is a monthly, nationally representative survey of approximately 50,000 households conducted by the Census Bureau.[17] Households in the CPS are in for four months, out for the following eight months, and then in again for four more months. Those in the fourth and eighth month of their participation are known as the outgoing rotation groups (ORG) and are asked some additional questions, including their earnings in the previous week. We compute an individual's hourly wage rate as the ratio of weekly earnings to weekly hours of work.[18] Pooled across the 12 months of the year, the ORGs yield an annual sample size of at least 150,000 households. They are available starting in 1979.

We summarize the individual-level wage data with an adjusted average wage for each state-year pair. These are obtained as state-year-specific intercepts in a regression of the natural logarithm of wages on demographic and educational characteristics:

$$1)\quad \omega_{ist} = w_{st} + x_{ist}\beta + \eta_{ist},$$

where ω_{ist} is the log of the wage for individual i in state s and year t. The vector, x_{ist}, of control characteristics is the same as that utilized by Blanchard and Katz (1997) and consists of a quartic in potential experience interacted with an indicator for sex, an indicator for marital status interacted with sex, a nonwhite indicator, a part-time indicator, and indicators for four educational attainment categories.[19] The estimated w_{st} coefficient is our measure of the adjusted log wage in state s and year t. Adjusted wage growth is $\Delta w_{st} = w_{st} - w_{st-1}$.

Figure 4 compares our ORG-based wage growth measure to four standard measures of annual wage growth. Three of the measures, AHE, Hourly Comp, and the ECI were discussed in the previous section. The fourth is a version of the ECI that is limited to the wage and salary components of employment cost. To facilitate comparison to the other measures, the ORG-based data in figure 4 are simple means, rather than the demographically adjusted figures discussed above. The correlation of our ORG-based measure is at least 0.72 with each of the other measures. This is about as high as the other measures are correlated with each other.

Close inspection of figure 4 suggests that our ORG-based measure is most similar to the ECI wages-only measure. This is true as well in figure 5, which plots the cumulative growth in the five measures since 1979.[20] The similarity of our ORG-based measure to the wages-only ECI likely reflects the fact that both measures capture only the value of wages and salaries. Neither reflects the value of benefits such as health insurance, whose relative growth rates have varied significantly over time. The AHE measure also excludes the value of benefits. Its divergence from the wages-only ECI and our ORG-based measure may be explained by its limitation to production and nonsupervisory workers.

The ORG data are our preferred source of state-level wage data. Their main attractions are large sample sizes and relatively rich associated demographic data. The lack of information on the value of benefits is a potential limitation. However, it seems plausible that the difference in growth rates

FIGURE 4

Nominal wage growth, 1980–99

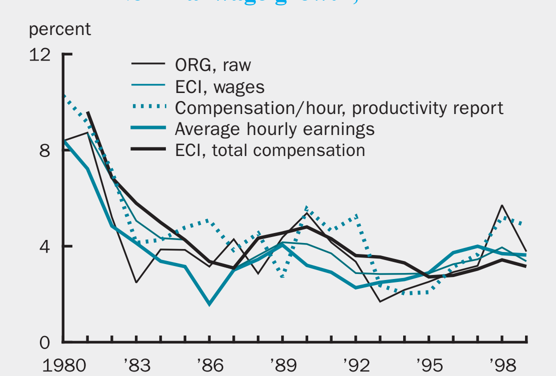
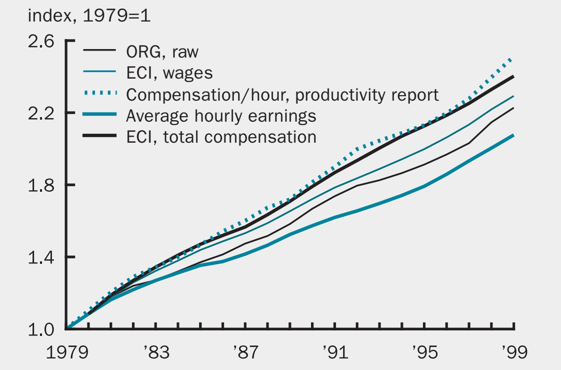


FIGURE 5

Cumulative nominal wage growth, 1980–99



between our measure and a more inclusive measure of total compensation is constant across states in a given year. If this is the case, as we explain further below, our estimates of the sensitivity of wage growth to unemployment will be unaffected. Nevertheless, to provide a check on the sensitivity of our results to the value of benefits, we also make use of the regional detail of the ECI. Unfortunately, the ECI is reported for only four regions, which severely limits the available degrees of freedom. Moreover, we did not have access to any micro data for the ECI, so we cannot demographically adjust the data.

Finally, another limitation of the ORG data is that they are not available prior to 1979, which might be considered a relatively short time series. Thus, in order to provide some evidence on the sensitivity of wage growth to unemployment in earlier years, we also use the annual demographic files from the March CPS. These contain responses to questions on earnings, weeks worked, and usual hours per week in the previous calendar year. Thus, a wage rate can be calculated as annual earnings divided by the product of weeks worked and usual hours per week.[21] These data are available in convenient electronic form starting in 1964, though prior to 1977, data from smaller states are not identified separately, reducing the number of degrees of freedom available.[22] Another drawback of the March data is the smaller sample size. Nationally, the sample is around 50,000 households, but for small states, samples can be as small as a few hundred households. This tends to make the associated wage measures quite volatile from year to year. In addition, we are forced to drop some of the early years of data because of unreasonably large changes in adjusted wages that we expect are the result of changes in sample design.

Empirical results

Our analysis is based on a standard panel data statistical model for the response of wage growth to unemployment. That model can be written as

2) $$\Delta w^*_{st} = \alpha_s + \gamma_t + u_{st}\beta + \varepsilon_{st},$$

where Δw^*_{st} is the adjusted wage growth and u_{st} is the log of the average of the 12 monthly unemployment rates for state s in year t. The state-specific effects, α_s, control for additional characteristics that are constant across time within a given state. Such factors may include demographic and industrial mix variables, as well as differences across states in the generosity of social insurance and other factors that affect the natural rate of unemployment in a given state. The year-specific effects, γ_t, control for the level of expected inflation in year t, as well as for the effects of productivity and other variables that may affect wages to the extent that such variables are constant across states for a given year.

Year-specific effects may also control for the effects of the exclusion of the value of benefits from our ORG-based measure of wage growth. Specifically, suppose that equation 2 holds for a comprehensive measure of compensation growth that includes the value of benefits, and further that the difference between such a measure and our ORG-based measure of wage growth is constant across states for a given year. Then $\Delta w_{st} = \Delta w_{st}^* + g_t$ and equation 2 can be written as

3) $$\Delta w_{st} = \alpha_s + \gamma'_t + u_{st}\beta + \varepsilon_{st},$$

where $\gamma'_t = \gamma_t + g_t$. In this case, the lack of benefits information affects the estimates of the year effects, but not the estimate of β, the sensitivity of wage growth to unemployment.[23] Moreover, if we can identify the true wage growth averaged over all states for a year with a measure such as Hourly Comp, we can adjust the estimates of the year effects to be consistent with such data. That is, $g_t = \Delta w_t - \Delta w_t^*$ which is the difference between the ORG-based measure and hourly compensation for annual data.

Least-squares estimation of equation 3 is equivalent to least-squares estimation of

4) $$\Delta \tilde{w}_{st} = \tilde{u}_{st}\beta + \varepsilon_{st},$$

where $\Delta \tilde{w}_{st} = \Delta w_{st} - \overline{\Delta w_s} - \overline{\Delta w_t} + \overline{\Delta w}$ and $\tilde{u}_{st} = u_{st} - \bar{u}_s - \bar{u}_t + \bar{u}$ represent deviations from state-specific and year-specific means. That is,

$\overline{\Delta w_s}$ is the mean adjusted wage growth over all years in the sample for state s, $\overline{\Delta w_t}$ is the mean adjusted wage growth over all states for year t, and $\overline{\Delta w}$ is the overall mean of wage growth, and similarly for $\bar{u}_s$, $\bar{u}_t$, and $\bar{u}$.

Figure 6 is a scatter plot of $\Delta\tilde{w}_{st}$ versus $\tilde{u}_{st}$ and thus shows the nature of the evidence on which the cross-state approach draws. A loose, but clearly negative association is apparent in the data. As shown in the first column of the first row of table 1, the ordinary least squares estimate of the regression line in figure 6 has slope –0.042 with a standard error of 0.004. As in the previous scatter plots, the hyperbolic lines around the regression line represent confidence intervals for the mean wage growth associated with any level of the unemployment rate deviation. These are somewhat tighter than in the equivalent time-series scatter plots, reflecting the greatly increased degrees of freedom obtained by working with the state-level data.

Though the evidence of association seen in figure 6 is very strong, there is also a very wide scatter of points around the line. Clearly, a great many factors affect wages besides unemployment rates. Moreover, some of the very wild data points likely reflect substantial measurement error in the wage growth measure.

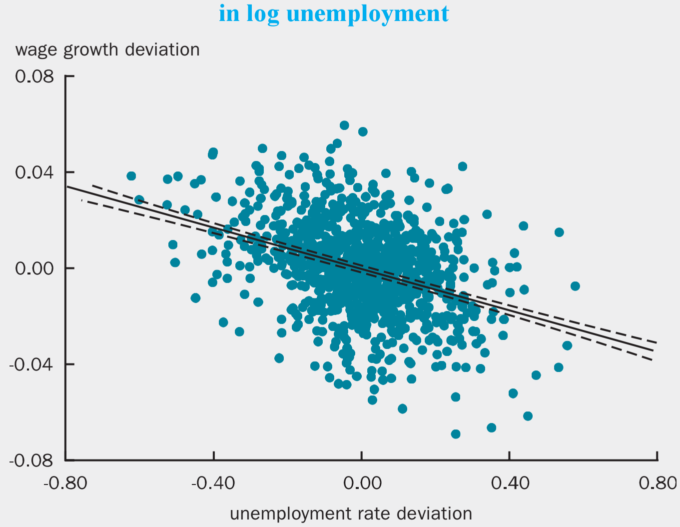
The second and third columns of table 1 present alternative estimation methods that reduce the influence of outliers. The second column simply weights the observations by state employment while the third column estimates the parameters using the biweight robust regression technique.[24] We prefer the latter method of estimation for its high degree of efficiency in the face of the kind of heavy-tailed data that we employ in this article. The first two digits of the estimates of the overall sensitivity of wage growth to unemployment are unaffected by choice of estimation method. However, consistent with its greater efficiency in the presence of outliers, the estimated standard errors from the robust regression technique are slightly smaller than those for ordinary or employment-weighted least squares.

Before examining how the estimates vary over time, it is informative to look more closely at the nature of the cross-state evidence. Figure 7 shows the 1999 level of unemployment in each of the 50 states and the District of Columbia. Rates varied from a low of 2.6 percent in New Hampshire to a high of 6.5 percent in the District of Columbia. But, as we have argued previously, the simple level of unemployment in the year may not be the best guide to the tightness of a state labor market. Average unemployment rates over the 1980–99 period varied from a low of 4.0 percent in South Dakota to a high of 10.2 percent in West Virginia. Much of this variation in states' average unemployment can be explained by slowly changing variables such as demographic composition, industry mix, and employment policies that do not necessarily affect optional wage growth.[25]

Figure 8 shows the deviations of 1999 state unemployment rates from their averages over the 1980–99 period. These relative unemployment indicators clearly differ a good deal from the standard measures shown in figure 7. For instance, the two extremes of 1999 unemployment, New Hampshire and the District of Columbia, are reasonably similar in terms of their deviations from their average rates, being 1.8 and 1.5 percentage points lower than their averages in 1999. In terms of unemployment deviations, the tightest labor market is Michigan's, where the 1999 unemployment rate of 3.7 percent is 4.7 points lower than its 1980–99 average of 8.4 percent. In contrast, the least tight labor market is in Hawaii where the current 5.5 percent unemployment rate is 0.4 points above its average over the last 20 years.[26] We find that such deviations from mean unemployment rates provide a superior guide to where labor markets are tight and, thus, that the raw unemployment rates seen in figure 7 can be somewhat misleading about where wage growth should be expected to be more rapid.

FIGURE 6

Deviation in wage growth versus deviation in log unemployment



Note: Dashed lines indicate 90 percent confidence bands.

TABLE 1

State wage curve elasticities

	OLS	WLS	Robust
Log unemployment rate	–0.042* (0.004)	–0.042* (0.004)	–0.042* (0.003)
Adjusted R-squared	0.467	0.550	0.463
Unemployment rate, 1980–84	–0.047* (0.005)	–0.049* (0.005)	–0.045* (0.005)
Unemployment rate, 1985–89	–0.046* (0.005)	–0.046* (0.005)	–0.044* (0.005)
Unemployment rate, 1990–94	–0.038* (0.006)	–0.040* (0.007)	–0.039* (0.006)
Unemployment rate, 1995–99	–0.032* (0.007)	–0.030* (0.006)	–0.033* (0.006)
F test p-statistic:			
UR, 1980–94=UR, 1995–99	0.082	0.027	0.086
UR, 1980–84=UR, 1995–99	0.059	0.011	0.074
UR, 1985–89=UR, 1995–99	0.058	0.037	0.092
UR, 1990–94=UR, 1995–99	0.435	0.218	0.395
Adjusted R-squared	0.469	0.552	0.461

*significant at the 5 percent level.

Notes: OLS is the ordinary least squares estimate. WLS is the observation weighted by state employment. UR is the unemployment rate. All regressions include state and year fixed effects. The last column includes industry and occupational composition controls. Robust standard errors are in parentheses. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period. The F test measures are calculated by the log of the unemployment rates times the dummy variable for one period being held equal to the log of the unemployment rate times the dummy variable for another period. ORG wage data are adjusted for education, experience, gender, race, and full time status.

Sources: Authors' calculations using data from the U.S. Department of Labor, Bureau of Labor Statistics for the unemployment rate and from the U.S. Department of the Census, *Current Population Survey*, for the weighted averages from the ORG for industry, occupation, and union composition.

Table 1 also shows estimates of the response of wage growth to unemployment for four five-year periods. The results suggest that wage growth has become somewhat less sensitive to unemployment in the 1990s. The robust regression methodology yields estimates of –0.045 and –0.044 for the early and late 1980s. The coefficient estimate for the early 1990s fell to –0.039, and that for the late 1990s was –0.033. Of course, even in the late 1990s, the estimates in table 1 are highly statistically significant, with t-statistics of around five. There is modestly strong evidence that the coefficient has changed over time. The F statistics shown in the table imply that the hypotheses that the 1995–99 coefficient is the same as the 1980–84, 1985–89, and the 1980–94 averages can be rejected at the 10 percent level, but not at the 5 percent level. The hypothesis that the 1995–99 coefficient is the same as the 1990–94 coefficient cannot be rejected at any standard confidence level.

FIGURE 7

1999 average unemployment rate

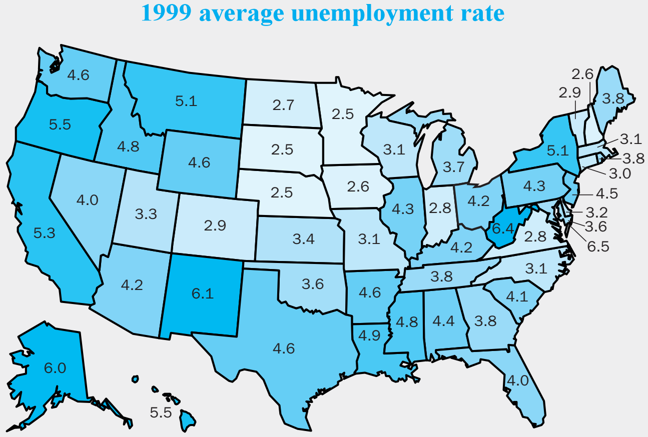


Figure 9 shows the result of estimating a separate slope for each year of the sample. Such estimates are based on the model

$$5)\quad \Delta w_{st} = \alpha_s + \gamma_t + u_{st}\beta_t + \varepsilon_{st},$$

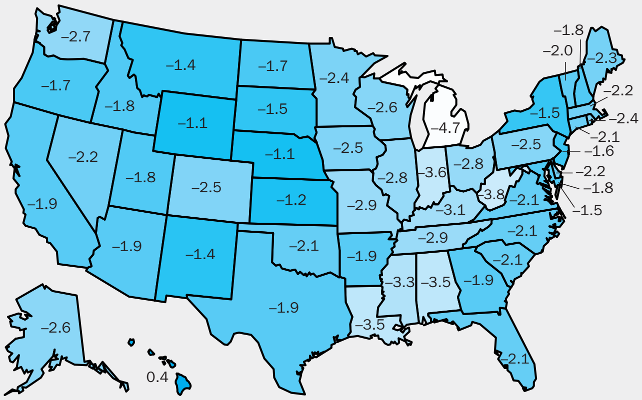
which continues to impose a common state effect, but allows the intercept and slope to vary freely over the sample period. Robust estimates of the slopes by year are plotted in figure 9 along with 90 percent confidence intervals. Since each data point is essentially estimated from 51 rather noisy observations, the confidence intervals tend to be somewhat wide. Still, all 20 coefficients are statistically significant at the 5 percent level.

The pattern of estimates shown in figure 9 leads us to view the evidence of a systematic drop in the magnitude of the coefficient as somewhat weak. The magnitude of the elasticity has decreased in recent years, with 1998 having the single smallest coefficient. But as recently as 1994 and 1995 the coefficient was about as large as it ever has been. And there have been previous years—1985 and 1993—in which the coefficient has declined, only to increase again subsequently.

The drop in coefficients in table 1 is also dependent on the imposition of a constant elasticity functional form. Such

FIGURE 8

Unemployment deviation, 1999 average minus 1980–99 average



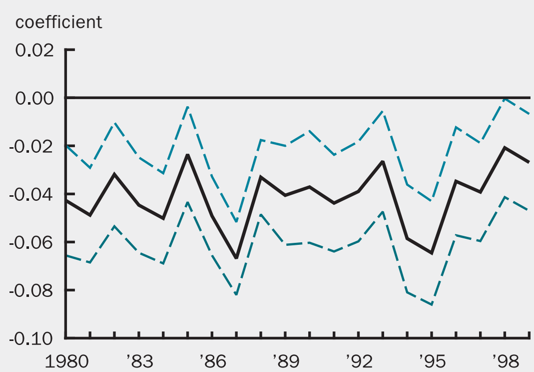
Rather, the late 1980s appears to be the period that was different, having a higher estimated coefficient than the other three periods. We prefer the constant elasticity specification of table 1 because of the better fit to the data, but the results of table 2 reinforce our view that the evidence of a decline in the sensitivity of wage growth to unemployment is rather weak.

Table 3 explores the sensitivity of the results in table 1 to alternative specifications. These all employ the robust regression methodology, but change other aspects of the specification. The first column shows the slope coefficients when we include additional variables measuring the fraction of workers in the various one-digit industries and occupations. Such variables may control for variation across states in productivity growth and other factors that determine wage growth. The coefficients tend to be smaller in magnitude than those in table 1, but the conclusions one would draw are similar; while the coefficient for the late 1990s is somewhat smaller, it is still highly statistically significant.

a form implies that the difference between unemployment rates of 3 percent and 4 percent is equivalent to the difference between rates of 6 percent and 8 percent. If instead, absolute differences in unemployment rates have the same effect on wage growth no matter how high or low they are, then the specification estimated in table 1 will force the coefficient for recent years, when unemployment has been relatively low, to fall, even if there has been no change in the relationship between wage growth and the level of unemployment. Table 2, which contains estimates based on a common slopes, rather than common elasticities, specification, contains some evidence in support of this hypothesis. Specifically, with a common slopes specification, there is no evidence of a decline in the sensitivity of wage growth to unemployment.

FIGURE 9

Annual unemployment rate coefficients full sample, robust regression



Note: Dashed lines indicate 90 percent confidence bands.

TABLE 2

State wage curve elasticities, alternative labor market indicators

	Log wage on level of labor market condition
Unemployment rate, 1980–84	–0.0053* (0.0007)
Unemployment rate, 1985–89	–0.0068* (0.0007)
Unemployment rate, 1990–94	–0.0063* (0.0010)
Unemployment rate, 1995–99	–0.0064* (0.0012)
F test p-statistic:	
UR, 1980–94=UR, 1995–99	0.751
UR, 1980–84=UR, 1995–99	0.358
UR, 1985–89=UR, 1995–99	0.760
UR, 1990–94=UR, 1995–99	0.968
Adjusted R-squared	0.450

*Significant at the 5 percent level.

Notes: UR is the unemployment rate. Regression includes state and year fixed effects and is estimated using robust regression. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period. The F test measures are calculated by the log of the unemployment rates times the dummy variable for one period being held equal to the log of the unemployment rate times the dummy variable for another period.

TABLE 3

State wage curve elasticities, alternative estimates

	Industry and occupation controls	Lag unemployment rate	No fixed effects	No year fixed effect	No state fixed effect	Raw wage change data
Unemployment rate, 1980–84	–0.042*	–0.036*	–0.016*	–0.034*	–0.024*	–0.044*
	(0.006)	(0.006)	(0.002)	(0.003)	(0.005)	(0.006)
Unemployment rate, 1985–89	–0.038*	–0.038*	–0.028*	–0.048*	–0.030*	–0.042*
	(0.005)	(0.005)	(0.003)	(0.004)	(0.004)	(0.005)
Unemployment rate, 1990–94	–0.034*	–0.027*	–0.028*	–0.048*	–0.014*	–0.034*
	(0.006)	(0.006)	(0.003)	(0.004)	(0.005)	(0.006)
Unemployment rate, 1995–99	–0.028*	–0.030*	–0.029*	–0.052*	–0.012*	–0.038*
	(0.006)	(0.006)	(0.003)	(0.004)	(0.005)	(0.006)
F test p-statistic:						
UR, 1980–94=UR, 1995–99	0.093	0.497	0.026	0.596	0.025	0.673
UR, 1980–84=UR, 1995–99	0.050	0.377	0.000	0.014	0.052	0.392
UR, 1985–89=UR, 1995–99	0.138	0.212	0.701	0.609	0.002	0.521
UR, 1990–94=UR, 1995–99	0.380	0.682	0.366	0.609	0.752	0.639
Adjusted R-squared	0.469	0.434	0.155	0.178	0.449	0.460

*Significant at the 5 percent level.

Notes: UR is the unemployment rate. All regressions include state and year fixed effects, unless noted, and are estimated using robust regression. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period. The F test measures are calculated by the log of the unemployment rates times the dummy variable for one period being held equal to the log of the unemployment rate times the dummy variable for another period. See text for more details.

The next column in table 3 uses the unemployment rate from the year before rather than the current year. This lowers the coefficients. The decline in the recent period is smaller, however. The next three columns explore the sensitivity of the results to the inclusion of fixed effects. Leaving out year effects makes the coefficients larger in magnitude, reflecting the fact that years with lower unemployment have had higher than average wage growth. Leaving out state effects significantly weakens the results, which reflects the fact that states with higher than average mean unemployment rates tend to have higher mean wage growth. Leaving out both kinds of fixed effects produces weak results as well. Both kinds of fixed effects are statistically significant according to the usual F statistic. Thus we prefer the specification estimated in table 1, and view the other results as indicating the effects of various forms of specification errors. Finally, using the raw wage growth data instead of the demographically adjusted wage growth figures has a relatively small effect on the results.

As we have noted, Lehrman and Schmidt (1999) report no evidence of a cross-state association between unemployment and wage growth. Lehrman and Schmidt use the ORG files to estimate state-specific wage growth between the first quarters of 1995 and 1998, computing mean wage growth for four "quartiles" of the unemployment distribution in the first quarter of 1998. They find little or no association between unemployment quartile and wage growth.

The results above may explain some of the difference between their results and ours. Lehrman and Schmidt use the unemployment rate for only the last quarter of the period, rather than the average over the whole period. The results in table 3 using lagged unemployment rates suggest that the match of the time periods of unemployment and wage growth may matter. Lehrman and Schmidt also use data on unemployment in 1998, which figure 9 says provides the weakest results of any year. Moreover, they only look at a single cross-section of data and so cannot control for state-specific fixed effects which table 3 shows is important. Finally, fitting a nonlinear specification seems to us to be asking a lot of 51 noisy observations. Clearly, figure 6 shows that there is a wide scatter around what is still a highly significant negative relationship. Thus, it would be quite surprising to see a clean pattern of means across quartiles when each of those means was estimated with only 12 or 13 observations.

One possible explanation for the falling coefficient on unemployment in table 1 is the changing nature of the work force. For instance, it is has been previously shown that wage growth among college-educated workers is less sensitive to unemployment than that among other workers. Thus the increasing share of college-educated workers could cause a decline in the unemployment coefficient of the kind seen in table 1. The results in table 4, however, show that this is not the case. The decline in coefficients is seen both for noncollege and college workers. Something other than a compositional shift towards college workers explains the lower late-1990s coefficients on unemployment.

Table 5 shows estimates of our basic specification using the March CPS data. As we noted, the advantage of this dataset is that it is available for earlier periods. Its disadvantage is that its wage measures are noisier, being based on a sample one-third as large as the ORG data. The results shown for five year intervals between 1964 and 1998, the last available data, suggest a quite stable relationship between unemployment and wage growth, with elasticity estimates generally near –0.03 except for the 1984 to 1988 period when the elasticity was estimated to be –0.045. Moreover, the F-statistics indicate that even the latter estimate is not statistically different from the estimate for the most recent period. The coefficients in table 5 are, however, somewhat lower than those in table 1. This must reflect differences in the nature of the March CPS wage measure, which is based on the previous calendar year, rather than the previous week.

Table 6 reports results obtained from the regional ECI data both for wages and salaries only and for total compensation. Because these data are available for only four regions, there are many fewer degrees of freedom. The first and third columns show results for periods similar to those shown in table 1.[27] These results for wages and salaries are relatively similar to those in table 1, except in the first period, when the data may have been somewhat suspect due to the newness of the series. However, for total compensation, the coefficient for the most recent five-year period is small and not statistically significant. Looking closely

TABLE 4

State wage curve elasticities, by education

	Noncollege sample	College sample
Unemployment rate, 1980–84	–0.047* (0.006)	–0.038* (0.011)
Unemployment rate, 1985–89	–0.046* (0.005)	–0.037* (0.009)
Unemployment rate, 1990–94	–0.039* (0.006)	–0.034* (0.011)
Unemployment rate, 1995–99	–0.035* (0.006)	–0.027* (0.011)
F test p-statistic:		
UR, 1980–94=UR, 1995–99	0.111	0.371
UR, 1980–84=UR, 1995–99	0.083	0.395
UR, 1985–89=UR, 1995–99	0.087	0.397
UR, 1990–94=UR, 1995–99	0.531	0.594
Adjusted R-squared	0.424	0.202

*Significant at 5 percent level.
Notes: UR is the unemployment rate. All regressions include state and year fixed effects and, unless noted, are estimated using robust regression. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period. The F test measures are calculated by the log of the unemployment rates times the dummy variable for one period being held equal to the log of the unemployment rate times the dummy variable for another period.

TABLE 5

State wage curve elasticities Wage growth: March CPS, 1964–98

	March CPS
Unemployment rate, 1964–68	–0.028* (0.013)
Unemployment rate, 1969–73	–0.026 (0.014)
Unemployment rate, 1974–78	–0.033* (0.012)
Unemployment rate, 1979–83	–0.030* (0.009)
Unemployment rate, 1984–88	–0.045* (0.007)
Unemployment rate, 1989–93	–0.028* (0.009)
Unemployment rate, 1994–98	–0.030* (0.009)
F test p-statistic:	
UR, 1964–93=UR, 1994–98	0.656
UR, 1964–68=UR, 1994–98	0.906
UR, 1969–73=UR, 1994–98	0.813
UR, 1974–78=UR, 1994–98	0.845
UR, 1979–83=UR, 1994–98	0.977
UR, 1984–88=UR, 1994–98	0.164
UR, 1989–93=UR, 1994–98	0.891
Time period	1964–98
Adjusted R-squared	0.614

*Significant at the 5 percent level.
Notes: UR is the unemployment rate. All regressions include state and year fixed effects and are estimated using robust regression. The unemployment rate is from the BLS for 1978–98 and state UI records for 1964–77. Some states are not uniquely identified in the March CPS prior to 1977. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period. The F test measures are calculated by the log of the unemployment rates times the dummy variable for one period being held equal to the log of the unemployment rate times the dummy variable for another period.

TABLE 6

State wage curve elasticity
Wage growth: Employment Cost Index, 1983–99

	Wages and salaries		Total compensation	
	5-year intervals	3-year intervals	5-year intervals	3-year intervals
Unemployment rate, 1983–84	–0.007 (0.005)		0.006 (0.005)	
Unemployment rate, 1985–89	–0.030* (0.005)		–0.030* (0.006)	
Unemployment rate, 1990–94	–0.039* (0.010)		–0.019** (0.011)	
Unemployment rate, 1995–99	–0.025* (0.007)		–0.008 (0.009)	
Unemployment rate, 1983–85		–0.019* (0.005)		–0.006 (0.007)
Unemployment rate, 1986–88		–0.031* (0.005)		–0.028* (0.007)
Unemployment rate, 1989–91		–0.035* (0.011)		–0.050* (0.013)
Unemployment rate, 1992–94		–0.031* (0.010)		–0.027* (0.013)
Unemployment rate, 1995–97		–0.015** (0.009)		–0.010 (0.011)
Unemployment rate, 1998–99		–0.044* (0.009)		–0.028* (0.011)
Adjusted R-squared	0.757	0.801	0.768	0.784

*Significant at the 5 percent level.
**Significant at the 10 percent level.
Notes: UR is the unemployment rate. All regressions include state and year fixed effects and are estimated using robust regression. The Employment Cost Index (ECI) is aggregated to four regions—East, South, Midwest, and West. Therefore, the sample includes four regions over 17 years, or 68 observations. ECI data are not demographically adjusted. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period.

at the individual observations suggests, however, that a very small number of data points are driving this result. Moreover, when we break the data into three-year intervals, the results suggest less evidence of a drop in the sensitivity of total compensation growth to unemployment. Given how little regional variation underlies the data in table 6, we consider the consistency of the results with those in table 1 to be reasonably good.

Thus far, our results have been limited to showing how the sensitivity of wage growth to unemployment has varied over time. Table 7 shows, in addition, how the level of wage growth associated with any level of unemployment has varied over time. Such quantities depend on both the estimated slope coefficients, β_t, and the year effects, γ_t. The values shown in table 7 are based on the specification of table 1 in which slopes are constant for each five-year period. The values in the column labeled Average Intercept–Raw are the average of the five-year effects (γ_ts) estimated for the period. The adjusted values in the next column are our estimates of the γ_t', the values that would correspond to the more comprehensive Hourly Comp wage growth measure. The intercept values are somewhat difficult to interpret because they potentially capture the effects of a number of variables. However, the fact they have fallen over time is consistent with the notion that they capture changes in expected inflation.

Given the normalization that $\sum \alpha_s = 0$, the predicted mean ORG-based adjusted wage growth associated with log unemployment rate $\bar{u}_t$ for year t is $\overline{\Delta w_t} = \gamma_t + \bar{u}_t\beta_t$, and the predicted mean Hourly Comp growth is $\Delta w_t^* = \gamma_t' + \bar{u}_t\beta_t$. To obtain estimates of predicted real wage growth, we subtract the rate of price inflation. In particular, the predicted amount by which the growth of Hourly Comp exceeds the growth in business sector prices, which is a reasonable measure of real wage growth, is $\overline{\Delta w_t - \Delta p_t} = \gamma_t' + \Delta p_t + \bar{u}_t\beta_t$, where Δp_t is the change in the log average price deflator for the business sector. Table 7

TABLE 7

Wage growth function

Period	Slope	Average Intercept		Real wage growth associated with unemployment rate of			Unemployment rate consistent with 1980–99 average real wage growth
		Raw	*Adjusted*	*4%*	*6%*	*8%*	
1980–84	–0.047	0.143	0.165	4.1	2.2	0.8	6.9
	(0.005)	(0.011)	(0.011)	(0.4)	(0.2)	(0.1)	(1.7)
1985–89	–0.046	0.111	0.122	3.1	1.2	–0.1	5.6
	(0.005)	(0.008)	(0.008)	(0.2)	(0.1)	(0.2)	(1.0)
1990–94	–0.0381	0.097	0.107	2.8	1.3	0.2	5.7
	(0.006)	(0.010)	(0.010)	(0.3)	(0.1)	(0.1)	(1.6)
1995–99	–0.032	0.084	0.086	2.8	1.5	0.6	6.0
	(0.006)	(0.009)	(0.009)	(0.5)	(0.2)	(0.2)	(1.6)

shows the predicted average real wage growth calculated in this manner for unemployment rates of 4 percent, 6 percent, and 8 percent. For an unemployment rate of 4 percent, predicted real wage growth dropped between the early and late 1980s, but has been reasonably constant since then. Our estimates currently predict real wage growth of 2.8 percent when the unemployment rate is 4 percent, about its current value. The predicted real wage growth rates associated with 6 percent and 8 percent unemployment also fell between the early and late 1980s, and since then have been fairly constant. The 0.6 percent level of wage growth predicted for 8 percent unemployment in the last period has, however, returned to about its level for the early 1980s.

One can also ask what level of unemployment is predicted to deliver a particular rate of real wage growth, say $\Delta(w^*/p)$. According to the above, that unemployment rate is $u^* = [\Delta(w^*/p) - (\gamma_t - \Delta p_t)]/\beta_t$. The last column of table 7 shows the values of this quantity corresponding to the mean real wage growth rate over the 1980–99 period, which was about 1.5 percent per year. That unemployment rate was nearly 7 percent in the early 1980s, but has been relatively constant since then at about the 6 percent level that we estimate for the late 1990s. We view the results in table 7 as confirming the relatively stable relationship between wage growth in excess of inflation and unemployment.

We argued previously that there might be labor market variables that predict wage growth better than the standard civilian unemployment rate. The recent drop in the coefficient on unemployment seen in table 1 might even reflect a misspecification in which unemployment is proxying for a more appropriate measure of labor market conditions. The drop in the unemployment coefficient might then be due to a lower correlation of unemployment with the preferred variable, which could have a stable relationship to wage growth. The results in table 8 suggest, however, that the decline in the coefficients in table 1 are not due to the unemployment rate becoming a poorer proxy for a superior measure of labor market tightness. The table shows the results of replacing the unemployment rate with several other measures of labor market conditions. These include an unemployment rate calculated from the ORG data, a measure of unemployment that includes all nonemployed workers who say they want a job regardless of whether they have recently searched, an even broader unemployment rate that also includes those who work part-time for economic reasons, a narrower measure that includes only white males between the ages of 25 and 54, the employment-to-population ratio, a measure of the exit rate out of unemployment, the fraction of the labor force unemployed five or fewer weeks, and the portion of the labor force unemployed 15 or more weeks. Virtually all the measures show the decline in coefficient magnitude in the most recent period that we see in table 1 for the unemployment rate. The drop off in the sensitivity of wage growth is especially significant for the exit rate out of unemployment and the rate of short-term unemployment. This may reflect the introduction of computer-aided interviewing technology with the 1994 CPS redesign, which had the effect of introducing a break in the series on short-term unemployment.

The results in table 8 suggest that the standard unemployment rate is not the only measure that might be used to judge the tightness of labor market conditions. Judging by the standard R-squared measure, several variables predict wage growth about as well as the unemployment rate. Indeed, the rate of long-term unemployment actually does very slightly

TABLE 8

State wage curve elasticities, alternative labor market indicators

	BLS unempl Rate	ORG unempl rate	Unempl plus NILF who want job	Unempl plus NILF who want job plus PT for econ reasons	White male age 25–54 unempl rate	Empl-pop ratio[a]	Exit rate out of unempl[b]	Unempl 0–5 weeks[b]	Unempl 15+ weeks
Unemployment rate, 1980–84	–0.045*	–0.043*	–0.050*	–0.058*	–0.024*	0.194*	0.036*	0.025*	–0.022*
	(0.005)	(0.005)	(0.006)	(0.007)	(0.004)	(0.029)	(0.007)	(0.008)	(0.003)
Unemployment rate, 1985–89	–0.044*	–0.042*	–0.047*	–0.051*	–0.023*	0.173*	0.025*	–0.036*	–0.022*
	(0.005)	(0.005)	(0.005)	(0.005)	(0.003)	(0.028)	(0.007)	(0.007)	(0.003)
Unemployment rate, 1990–94	–0.039*	–0.035*	–0.039*	–0.038*	–0.021*	0.164*	0.022*	–0.001	–0.020*
	(0.006)	(0.006)	(0.007)	(0.007)	(0.004)	(0.030)	(0.006)	(0.008)	(0.003)
Unemployment rate, 1995–99	0.033*	–0.027*	–0.031*	–0.029*	–0.016*	0.176*	0.001	0.003	–0.014*
	(0.006)	(0.006)	(0.006)	(0.007)	(0.004)	(0.030)	(0.002)	(0.006)	(0.003)
F test p-statistic:									
UR, 1980–94=UR, 1995–99	0.086	0.023	0.024	0.002	0.055	0.936	0.000	0.001	0.020
UR, 1980–84=UR, 1995–99	0.074	0.022	0.015	0.001	0.058	0.436	0.000	0.002	0.043
UR, 1985–89=UR, 1995–99	0.092	0.024	0.026	0.003	0.094	0.897	0.002	0.000	0.034
UR, 1990–94=UR, 1995–99	0.395	0.252	0.337	0.274	0.246	0.607	0.001	0.637	0.138
Adjusted R-squared	0.461	0.453	0.448	0.457	0.438	0.409	0.413	0.412	0.466

*Significant at the 5 percent level.

[a]Detrended.

[b]1994 is excluded.

Notes: UR is the unemployment rate. ORG is the outgoing rotation groups. BLS indicates U.S. Bureau of Labor Statistics. NILF is not in labor force. PT indicates part time. All regressions include state and year fixed effects and are estimated using robust regression. The unemployment rate for each period is measured by the log of the unemployment rate times a dummy variable for the time period. The F test measures are calculated by the log of the unemployment rates times the dummy variable for one period being held equal to the log of the unemployment rate times the dummy variable for another period.

better. The two broader measures of unemployment, which include all of those who say they want a job and those workers plus those who are involuntarily part-time, come reasonably close to matching the predictive power of the standard unemployment rate, while the narrower measure that is limited to prime-age white males does less well. Perhaps somewhat surprisingly, the measures that may be more closely connected to theory, the employment-to-population ratio and the exit rate from unemployment, are among the least well performing measures, though in the latter case this may be due to breaks in the data series that may, with some work, be repairable. A fully satisfactory comparison of the forecasting abilities of the various labor market variables would require the use of higher frequency data, more elaborate dynamics, and some attention to the out-of-sample properties of the forecasts. We regard the results in table 8 as suggesting that such work may be quite fruitful.

Conclusion

In this article, we have shown that the negative cross-state correlation between unemployment and wage growth persists even in recent data. We find some evidence of a decline in the sensitivity of wage growth to unemployment in the late 1990s. But, we regard that evidence as being somewhat weak because it is dependent on exactly when the line between periods is drawn and whether the relationship is modeled as one in which percentage or absolute differences in unemployment rates have constant effects on wage growth.

Of course, the relationship between unemployment and wage growth is a loose one. Unemployment is only one of many factors that affect wage growth, so that looking at a small number of states or years, differences in unemployment rates may not always provide a good prediction of differences in wage growth. But with enough data, the relationship between unemployment and wage growth emerges fairly clearly and does not appear to be dependent on any arbitrary details of our analysis.

We also find that several other labor market indicators predict wage growth about as well as the standard civilian unemployment rate. Refining such measures and studying their forecasting abilities more systematically may be a fruitful area for further research.

Finally, our results may have implications for work on inflation forecasting, an important component in the monetary policy process. Traditional short-run, or expectations-augmented, Phillips curve methodologies have tended to overpredict the change in inflation in recent years.[28] That methodology depends upon both the relationship between unemployment and expected wage growth and the relationship between wage growth and price inflation. Given the many fundamental changes that may be affecting the labor market, it is natural to look for a change in the relationship between unemployment and wage growth. But, our finding that the cross-state relationship between unemployment and wage growth has been relatively stable suggests that more attention be given to the link between wage growth and price inflation as the source of instability in the short-run Phillips curve. This seems consistent with findings such as those in Brayton et al. (1999) that adding variables to account for variation in the markup of prices over wages may be the most attractive way to stabilize the relationship between unemployment and changes in price inflation.

NOTES

[1]Friedman (1968) and Phelps (1973) are classic statements of this point.

[2]In the years since Phillips' (1958) paper, the correlation between nominal wage growth and unemployment has been close to zero in U.S. data.

[3]Abraham et al. (1999) discuss the differences in these wage measures.

[4]Blanchard and Katz (1997) discuss the relationship between the kind of time-series evidence depicted in figures 1–3 and the cross-state evidence that is the main focus of this article.

[5]Blanchard and Katz (1997) note that, empirically, these other variables are often found to have little impact on wage growth forecasts.

[6]These were computed under the usual ideal assumptions that error terms are uncorrelated and of constant variance, and thus may be somewhat optimistic. The hyperbolic lines around the regression line represent 90 percent confidence intervals for the expected level of wage growth in excess of inflation at a given level of log unemployment.

[7]Aaronson and Sullivan (1998, 1999) discuss the implications for wages of a drop in job security. Katz and Krueger (1999) discuss reasons for a drop in the natural rate of unemployment.

[8]See, for example, Mortensen and Pissarides (1994).

[9]See, for example, Shapiro and Stiglitz (1984) and Salop (1979).

[10]Blanchard and Katz (1997) provide a cogent discussion of these issues.

[11]Castillo (1998) shows that in U.S. data, those outside the labor force who want a job are less attached to the labor market than unemployed workers. However, Jones and Riddel (1999) show that in Canadian data, those out of the labor force who report wanting a job are closer to the unemployed than to others who are out of the labor force, in terms of their subsequent probabilities of employment.

[12]An important reference is Blanchflower and Oswald (1994), who document a cross-sectional relationship between unemployment and wages in a number of countries over a number of periods. Blanchflower and Oswald interpret their results as a relationship between unemployment and the level of wages because in their statistical models for the wage level, lagged wages are estimated to have small coefficients. We agree, however, with Blanchard and Katz (1997) and Card and Hyslop (1996) that these low estimates are the result of substantial measurement error in Blanchflower and Oswald's wage measures as well as their inappropriate use of annual, rather than hourly earnings. We find that in models employing hourly wage measures obtained from samples large enough to minimize measurement error, the coefficient on lagged wages is quite close to unity. Thus, the relationship is best thought of in terms of wage growth rather than wage levels. Roberts (1999) and Whelan (1999) show that the form of the micro-data relationship may not matter for the form of aggregate inflation dynamics.

[13]Results on wage growth across states are a small part of Lehrman and Schmidt's (1999) lengthy study. The description of the empirical analysis in Zandi (2000) is not particularly detailed, but his results appear to be consistent with our findings. Zandi concludes that the Phillips curve is "alive and kicking." Whether this follows from his or our evidence depends, however, on what one means by the "Phillips curve." If one means that expected wage growth is related to unemployment, we agree with his conclusion. However, as we discuss below, if the Phillips curve is taken to be the short-run, or expectations-augmented, relationship between unemployment and changes in price inflation, his conclusion doesn't necessarily follow from his results.

[14]See, for example, Gordon (1997).

[15]See, for example, Blanchard and Katz (1997).

[16]See Brayton et al. (1999).

[17]Until 1996, there were approximately 60,000 households in the survey.

[18]We drop observations on workers whose computed wage is less than 50 cents per hour or more than $100 per hour.

[19]Blanchard and Katz (1997) estimate separate regression models for each year of data while we estimate a single, pooled regression. This makes no appreciable difference to the results when, as in the models we estimate, year effects are included in the estimation.

[20]The ECI compensation series is scaled to equal the ORG measure in 1982, the first year it is available.

[21]Prior to 1976, data on usual weekly hours is not available; in its place we use data on hours worked in the week prior to the survey.

[22]In our analysis of the March data, unemployment rates before 1978 are obtained from state unemployment insurance claims data.

[23]This argument goes through more generally if the difference between the ORG wage growth measure and an ideal wage growth measure has an error components structure that is limited to a year effect, a state effect, and an error term that is uncorrelated with unemployment.

[24]We use the default tuning parameters in the Stata statistical procedure. (These control the rate at which outliers are downweighted.) See Stata Corporation (1999) for a description of the technique.

[25]In a cross-sectional regression, such variables explain about 60 percent to 75 percent of the variation in state average unemployment rates.

[26]Hawaii is the only state for which 1999 was an above-average year for unemployment.

[27]The regional ECI data are not available before 1983.

[28]See, for example, Brayton et al. (1999).

REFERENCES

Aaronson, Daniel, and Daniel G. Sullivan, 1999, “Worker insecurity and aggregate wage growth,” Federal Reserve Bank of Chicago, working paper, No. 99-30.

__________, 1998, “The decline of job security in the 1990s: Displacement, anxiety, and their effect on wage growth,” *Economic Perspectives*, Federal Reserve Bank of Chicago, Vol. 22, No. 1, First Quarter, pp. 17–43.

Abraham, Katherine, James Spletzer, and Jay Stewart, 1999, “Why do different wage series tell different stories?,” *American Economic Review Papers and Proceedings*, Vol. 89, No. 2, pp. 34–39.

Blanchard, Olivier, and Lawrence F. Katz, 1997, “What we know and do not know about the natural rate of unemployment,” *Journal of Economic Perspectives*, Vol. 11, No. 1, Winter, pp. 51–72.

Blanchflower, David G., and Andrew J. Oswald, 1994, *The Wage Curve*, Cambridge, MA: MIT Press.

Brayton, Flint, John M. Roberts, and John C. Williams, 1999, “What’s happened to the Phillips curve?,” Board of Governors of the Federal Reserve System, working paper, September.

Card, David, and Dean Hyslop, 1996, “Does inflation grease the wheels of the labor market,” National Bureau of Economic Research, working paper, No. 5538.

Castillo, Monica D., 1998, “Persons outside the labor force who want a job,” *Monthly Labor Review*, Vol. 121, No. 7, July, pp. 34–42.

Friedman, Milton, 1968, “The role of monetary policy,” *American Economic Review*, Vol. 58, No. 1, pp. 1–17.

Gordon, Robert J., 1997, “The time-varying NAIRU and its implications for economic policy,” *Journal of Economic Perspectives*, Vol. 11, No. 1, Winter, pp. 11–32.

Jones, Stephen, and Craig Riddell, 1999, “The measurement of unemployment: An empirical approach,” *Econometrica*, Vol. 67, No. 1, pp. 147–161.

Katz, Lawrence, and Alan Krueger, 1999, “The high pressure U.S. labor market of the 1990s,” *Brookings Papers on Economic Activity*, Vol. 1, chapter 1.

Lehrman, Robert, and Stefanie Schmidt, 1999, “An overview of economic, social, and demographic trends affecting the U.S. labor market,” Urban Institute, Washington, working paper.

Mortensen, Dale, and Christopher Pissarides, 1994, “Job creation and job destruction in the theory of unemployment,” *Review of Economic Studies*, Vol. 61, No. 3, pp. 397–415.

Phelps, Edmund, 1973, *Inflation Policy and Unemployment Theory*, New York: W. W. Norton.

Phillips, A. William, 1958, "The relationship between unemployment and the rate of change of money wage rates in the United Kingdom, 1861–1957," *Economica*, Vol. 25, November, pp. 283–299.

Roberts, John M., 1997, "The wage curve and the Phillips curve," Board of Governors of the Federal Reserve System, staff working paper, No. 1997-57.

Salop, Steven, 1979, "A model of the natural rate of unemployment," *American Economic Review*, Vol. 69, No. 1, pp. 117–125.

Shapiro, Carl, and Joseph Stiglitz, 1984, "Equilibrium unemployment as a worker discipline device," *American Economic Review*, Vol. 74, No. 3, pp. 433–444.

Stata Corporation, 1999, *Stata Statistical Software: Release 6.0*, College Station, TX: Stata Corporation.

Whelan, Karl, 1997, "Wage curve vs. Phillips curve: Are there macroeconomic implications?" Board of Governors of the Federal Reserve System, staff working paper, No. 1997-51.

Zandi, Mark, 2000, "The Phillips curve: Alive and kicking," *The Dismal Scientist*, available on the Internet at www.dismal.com, January 21.