Recent Evidence on the Relationship Between Unemployment and Wage Growth

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Abstract

The current expansion has delivered the lowest unemployment rates in decades, yet nominal wage growth has remained relatively contained. This suggests to some a shift in the historical relationship between unemployment and wage growth. We look across the states for more timely evidence of a change in this relationship. We find some evidence that the elasticity of real wage growth with respect to unemployment has fallen recently, a result that is not due to a compositional shift toward college-educated workers. However, evidence of a weakened relationship is itself weak, depending on inherently arbitrary decisions about when a shift may have occurred. In addition, we find that levels of real wage growth associated with high, medium, and low unemployment have remained relatively constant.

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I. Introduction

The long economic expansion of the 1990s 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.[1] 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 paper, we look across the states for more timely evidence of a change in the relationship between unemployment and real wage growth. 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. Thus, it may be possible to identify changes in that response that would take many years of time-series data to uncover. Previous work has demonstrated a relationship between unemployment and real wage growth across states that is analogous to that in time-series data.[2] 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

[1] Speculation about a change in the relationship between unemployment and wage growth has taken a number of forms, not all of which have been well reasoned. 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 for real, not nominal, wage growth (e.g., Friedman 1968). Among the few papers offering evidence of a break down in the unemployment – wage growth relationship is a study by Lehrman and Schmidt (1999) suggesting that the level of unemployment across states is not now related to wage growth. Footnote 14 below documents why we believe those authors' results differ from ours.

[2] 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

expectations are constant across states, differences in wage growth across states in a given year are unrelated to inflation expectations. Similarly, to the extent that other variables, such as productivity growth, that affect wage growth are constant across states in a given year, comparisons of states' wage growth rates are also unaffected by these variables.

Our empirical work confirms the negative cross-state correlation between unemployment and wage growth. We also find that the elasticity of wages with respect to unemployment has fallen recently, 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 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.

We also briefly examine how the level of wage growth associated with 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.

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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.

Finally, because 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 also investigate a number of alternative measures of labor market tightness. 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.

Recently, there is evidence that typical short-run Phillips curve specifications have systematically overforecasted inflation. 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 variables related to the markup of prices over wages helps stabilize the Phillips curve.

II. Background

Some shift in the relationship between unemployment and real wage growth in the 1990s 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.[3]

A rough indication of the time-series evidence on this question can be gleaned from figure 1, a scatter plot of annual data on the excess of hourly compensation growth over the previous year's

[3] Aaronson and Sullivan (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.

CPI inflation versus the natural logarithm of the annual unemployment rate.[4] The relationship depicted in figure 1 is analogous to the wage equations in some macroeconometric models. It can be motivated by assumptions that wages are set to exceed expected inflation by an amount that depends on the unemployment rate and expected inflation is equal to the level of inflation in the previous year.[5]

The figure shows that there is a loose, but reasonably clear, negative correlation between unemployment and wage growth in excess of lagged inflation. The least squares regression line shown in the figure slopes downward with an elasticity of -0.055 and an estimated standard error of 0.009.[6] A line connecting the values from 1992 to 1999 highlights the data for the current expansion, when the unemployment rate was falling from 7.5 percent to 4.2 percent. As can be seen, the growth of hourly compensation 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.

However, such departures of wage growth from expectations 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 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.

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[4] The compensation data is from the BLS's productivity and cost release. Results using the BLS's average hourly earnings and employment cost index (ECI) series are somewhat similar, although the latter series is not as tightly associated with unemployment. We use the productivity report's measure because it covers wage and nonwage forms of compensation and goes back farther than the ECI. Abraham et al (1999) discusses the differences in these wage series.

[5] Blanchard and Katz (1997) discuss the relationship between the kind of time-series evidence depicted in figure 1 and the cross-state evidence that is the main focus of this article. Of course, wage equations in actual macroeconometric models are considerably more elaborate than what is represented in the figure. In particular, they use higher frequency data, allow for more complicated dynamics, and include other variables, such as the level of productivity. Blanchard and Katz note that these other variables are often found to have little impact on wage growth forecasts.

Moreover, the theoretical basis for the relationship depicted in figure 1 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 does not 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 figure 1 with a standard, aggregate labor supply curve. This is analogous to the relationship in figure 1, 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.

Other models go beyond the simple competitive framework to allow involuntary unemployment and the unemployment rate to be related to wages. For example, 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 (Mortensen and Pissarides 1994). Alternatively, efficiency wage arguments like Shapiro and Stiglitz (1984) and Salop (1979) generate a link between unemployment and wages 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.

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[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.

Even in search and efficiency wage models, the standard unemployment rate may not be the variable most directly related to wages (Blanchard and Katz 1997). Rather, in both classes of models, the exit rate from 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 relatively 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.[7] Ultimately, which measure best captures the labor market forces influencing wages is an empirical question, the answer to which could change over time.

III. Data

Our main results are based on two data sources. The first is the annual averages of the BLS's monthly, state-level unemployment rates. The second source is a measure of state-level, demographically adjusted wage growth that we construct from the outgoing rotations (ORG) of the Current Population Survey (CPS). We compute an individual's hourly wage as the ratio of weekly

[7] 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, in Canadian data, Jones and Riddel (1999) show that 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.

earnings to weekly hours of work.[8] Pooled across the 12 months of the year, the ORGs yield an annual sample size of at least 150,000 households. They are available from 1979 to 1999.

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:

$$(1) \qquad w_{\mathbf{ist}} = \mathbf{w}_{\mathbf{st}} + \mathbf{x}_{\mathbf{ist}} b + h_{\mathbf{ist}} ,$$

where $w_{\mathbf{ist}}$ is the log of the wage for individual i in state s and year t. The vector, $\mathbf{x}_{\mathbf{ist}}$, 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.[9] The estimated $\mathbf{w}_{\mathbf{st}}$ coefficient is our measure of the adjusted log wage in state s and year t.[10]

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 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.

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

[8] We drop observations on workers whose computed wage is less than 50 cents per hour or more than \$100 per hour.
[9] 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 year effects are included in the estimation.
[10] The correlation of our ORG measure is at least 0.72 with other aggregate wage measures, including average hourly earnings, ECI, and compensation per hour. This is about as high as the other measures are correlated with each other.

on the sensitivity of wage growth to unemployment in earlier years, we also use the annual demographic files from the March CPS.[11] 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.[12] These data are available starting in 1964, though prior to 1977, data from smaller states are not identified separately. A 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. 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 because of changes in sample design.

IV. Empirical results

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

(2) $$\Delta \mathbf{w}^*_{\mathbf{st}} = \alpha_{\mathbf{s}} + \gamma_{\mathbf{t}} + \mathbf{u}_{\mathbf{st}}\beta + \epsilon_{\mathbf{st}},$$

where $\Delta \mathbf{w}^*_{\mathbf{st}}$ is adjusted wage growth and $\mathbf{u}_{\mathbf{st}}$ is the log of the unemployment rate for state s in year t. The state-specific effects, $\alpha_{\mathbf{s}}$, control for characteristics that are constant across time within a state. The year-specific effects, $\gamma_{\mathbf{t}}$, 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 in 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 if the difference is constant across states for a given year. Then $\Delta \mathbf{w}_{\mathbf{st}} = \Delta \mathbf{w}^*_{\mathbf{st}} + \mathbf{g}_{\mathbf{t}}$ and equation 2 can be written as

(3) $$\Delta \mathbf{w}_{\mathbf{st}} = \alpha_{\mathbf{s}} + \gamma_{\mathbf{t}} + \mathbf{u}_{\mathbf{st}}\beta + \epsilon_{\mathbf{st}},$$

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

where $g_t = g_t + \mathbf{g_t}$. In this case, the lack of benefits information affects the estimates of the year effects, but not the estimate of b.[13] Moreover, if we can identify the true wage growth averaged over all states for a year with a measure such as hourly compensation from the productivity report, we can adjust the estimates of the year effects to be consistent with such data (i.e. $\mathbf{g_t} = \overline{D\mathbf{w_t}} - \overline{D\mathbf{w_t^*}}$).

Table 1 reports estimates of b from equation (3). As shown in the first column of the first row of table 1, the ordinary least squares estimate is -0.042 with a standard error of 0.004. 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. 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. 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.

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 to -0.033 in the late 1990s. Of course, even in the late 1990s, the estimates in table 1 are statistically

[12] Usual weekly hours is not available prior to 1976. In its place, we use hours worked in the week prior to the survey.
[13] 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.

significant, with t-statistics around five. But there is 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, although 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 2 shows the result of estimating a separate slope for each year of the sample. Such estimates are based on the model

$$(4) \qquad \Delta \mathbf{w}_{\mathbf{st}} = \alpha_{\mathbf{s}} + \gamma_{\mathbf{t}} + \mathbf{u}_{\mathbf{st}}\beta_{\mathbf{t}} + \varepsilon_{\mathbf{st}} ,$$

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 2 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 2 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. 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. The fourth column of table 1, 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. Rather, the late 1980s appear to be the period that was different, having a higher estimated coefficient than the other three periods. We prefer the constant elasticity specification because of the better fit to the data, but the results in column 4 reinforce our view that the evidence of a decline in the sensitivity of wage growth to unemployment is weak.

Table 2 explores the sensitivity of the results to alternative specifications. The first column shows the slope coefficients when we include additional variables measuring the fraction of workers in one-digit industries and occupations. Such variables may control for state variation 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.

The next column in table 2 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.[14]

[14] Lerman and Schmidt (1999) report no evidence of a cross-state association between unemployment and wage growth. They 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

One possible explanation for the falling coefficient on unemployment 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.[15] The results in the last two columns of table 2, however, show that this is not the case. The decline in coefficients is seen both for noncollege and college workers.[16] Something other than a compositional shift towards college workers explains the lower late-1990s coefficients on unemployment.

Table 3 shows estimates of our basic specification using the March CPS data. The results shown for five year intervals between 1964 and 1998 suggest a quite stable relationship between unemployment and wage growth, with elasticity estimates generally near -0.035 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.

Finally, as a robustness check of the importance of benefits on the real wage-unemployment relationship, we also looked at the regional ECI data. Because these data are available for only four regions and date back to only 1983, there are many fewer degrees of freedom. However, the ECI does allow us to look directly at the association between total compensation and unemployment. While the results for ECI wages and salaries are relatively similar to those in table 1, for total compensation, the coefficient for the most recent five-year period is small and not statistically

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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 2 using lagged unemployment rates suggest that the match of the time periods of unemployment and wage growth matters to the estimates. Lehrman and Schmidt also use data on unemployment in 1998, which figure 2 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, 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.

[15] Furthermore, Solon et al (1994) argue that, in aggregate time-series data, compositional bias arises because low-skill workers are overweighted during expansions, leading to a downward bias in the procyclicality of real wages.

significant. However, looking closely at the individual observations suggests 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. Therefore, given how little regional variation underlies the data, 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 4 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 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 compensation 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 $\sum \alpha_s = \mathbf{0}$, the predicted mean ORG-based adjusted wage growth associated with log unemployment rate $\overline{\mathbf{u}_t}$ for year t is $\overline{\Delta \mathbf{w}_t} = \gamma_t + \overline{\mathbf{u}_t}\beta_t$, and the predicted mean hourly compensation growth is $\overline{\Delta \mathbf{w}_t^*} = \gamma_t' + \overline{\mathbf{u}_t}\beta_t$. The predicted amount by which the growth of hourly compensation exceeds the growth in business sector prices, which is a reasonable measure of real wage growth, is $\overline{\Delta \mathbf{w}_t^*} - \Delta \mathbf{p}_t = \gamma_t' - \Delta \mathbf{p}_t + \overline{\mathbf{u}_t}\beta_t$, where $\Delta \mathbf{p}_t$ is the change in the log average price deflator for the business sector. Table 4 shows the predicted average real wage growth

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[16] This pattern emerges when we use education-specific unemployment rates as well.

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 $D(\mathbf{w^*/p})$. According to the above, that unemployment rate is $\mathbf{u^*} = [D(\mathbf{w^*/p}) - (g_t - D\mathbf{p_t})]/b_t$. The last column of table 4 shows the values of this quantity corresponding to the mean real wage growth rate over 1980-99, which was about 1.5 percent per year. That unemployment rate was nearly 7 percent in the early 1980s, but since then has been relatively constant at about the 6 percent level that we estimate for the late 1990s. We view the results in table 4 as confirming the relatively stable relationship between real wage growth and unemployment.

Finally, we argued previously that other labor market variables might predict wage growth better than the 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 5 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, including an ORG-based unemployment rate, a measure of unemployment that includes those who say they want a job regardless of whether they have recently searched, a broader unemployment rate that 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 recent decline in coefficient magnitude that is seen in table 1. 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. However, this may reflect the introduction of computer-aided interviewing technology with the 1994 CPS redesign, which had the effect of introducing a series break in short-term unemployment measures.

The results in table 5 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 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. 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.

V. Conclusion

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 and other labor market measures 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.

Our results have implications for work on inflation forecasting. Traditional short-run, or expectations-augmented, Phillips curve methodologies have tended to overpredict the change in inflation in recent years (Brayton et al. 1999). 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.

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Table 1
State Wage-Unemployment Elasticities [1]

	OLS	WLS	Robust	Robust
Unemployment rate	-0.042 *	-0.042 *	-0.042 *	-0.0059 *
	(0.004)	(0.004)	(0.003)	(0.0005)
Adjusted R-squared	0.467	0.550	0.463	0.450
Unemployment rate, 1980-84	-0.047 *	-0.049 *	-0.045 *	-0.0053 *
	(0.005)	(0.005)	(0.005)	(0.0007)
Unemployment rate, 1985-89	-0.046 *	-0.046 *	-0.044 *	-0.0068 *
	(0.005)	(0.005)	(0.005)	(0.0007)
Unemployment rate, 1990-94	-0.038 *	-0.040 *	-0.039 *	-0.0063 *
	(0.006)	(0.007)	(0.006)	(0.0010)
Unemployment rate, 1995-99	-0.032 *	-0.030 *	-0.033 *	-0.0064 *
	(0.007)	(0.006)	(0.006)	(0.0012)
F test p-statistic:				
UR, 1980-94=UR, 1995-99	0.082	0.027	0.086	0.751
UR, 1980-84=UR, 1995-99	0.059	0.011	0.074	0.358
UR, 1985-89=UR, 1995-99	0.058	0.037	0.092	0.760
UR, 1990-94=UR, 1995-99	0.435	0.218	0.395	0.968
Adjusted R-squared	0.469	0.552	0.461	0.450
Log of unemployment rate	yes	yes	yes	no

Notes:

Table 2
State Wage-Unemployment Elasticities [1]
Alternative Estimates

	Industry and Occupation Controls	Lag Unemployment rate	No Fixed Effects	No Year Fixed Effect	No State Fixed Effect	Noncollege Sample	College Sample
Unemployment rate, 1980-84	-0.042 *	-0.036 *	-0.016 *	-0.034 *	-0.024 *	-0.047 *	-0.038 *
	(0.006)	(0.006)	(0.002)	(0.003)	(0.005)	(0.006)	(0.011)
Unemployment rate, 1985-89	-0.038 *	-0.038 *	-0.028 *	-0.048 *	-0.030 *	-0.046 *	-0.037 *
	(0.005)	(0.005)	(0.003)	(0.004)	(0.004)	(0.005)	(0.009)
Unemployment rate, 1990-94	-0.034 *	-0.027 *	-0.028 *	-0.048 *	-0.014 *	-0.039 *	-0.034 *
	(0.006)	(0.006)	(0.003)	(0.004)	(0.005)	(0.006)	(0.011)
Unemployment rate, 1995-99	-0.028 *	-0.030 *	-0.029 *	-0.052 *	-0.012 *	-0.035 *	-0.027 *
	(0.006)	(0.006)	(0.003)	(0.004)	(0.005)	(0.006)	(0.011)
F test p-statistic:							
UR, 1980-94=UR, 1995-99	0.093	0.497	0.026	0.596	0.025	0.111	0.371
UR, 1980-84=UR, 1995-99	0.050	0.377	0.000	0.014	0.052	0.083	0.395
UR, 1985-89=UR, 1995-99	0.138	0.212	0.701	0.609	0.002	0.087	0.397
UR, 1990-94=UR, 1995-99	0.380	0.682	0.366	0.609	0.752	0.531	0.594
Adjusted R-squared	0.469	0.434	0.155	0.178	0.449	0.424	0.202

Notes:
*=significant at 5 percent level
[1] UR= unemployment rate. All regressions include state and year fixed effects, unless noted, and are estimated using robust regression.

Table 3
State Wage-Unemployment Elasticities [1]
Using the March CPS, 1964-98

Unemployment rate, 1964-68	-0.034	*
	(0.013)	
Unemployment rate, 1969-73	-0.029	*
	(0.014)	
Unemployment rate, 1974-78	-0.039	*
	(0.012)	
Unemployment rate, 1979-83	-0.036	*
	(0.009)	
Unemployment rate, 1984-88	-0.045	*
	(0.007)	
Unemployment rate, 1989-93	-0.036	*
	(0.009)	
Unemployment rate, 1994-98	-0.036	*
	(0.010)	
F test p-statistic:		
UR, 1964-93=UR, 1994-98	0.834	
UR, 1964-68=UR, 1994-98	0.900	
UR, 1969-73=UR, 1994-98	0.644	
UR, 1974-78=UR, 1994-98	0.879	
UR, 1979-83=UR, 1994-98	0.990	
UR, 1984-88=UR, 1994-98	0.482	
UR, 1989-93=UR, 1994-98	0.971	
Time period	1964-1998	
Adjusted R-squared	0.414	

Notes:
*= significant at the 5 percnet level.
[1] UR=unemployment rate. Regression includes state and year fixed effects and is estimated using robust regression. The unemployment rate is from the BLS for 1978-99 and state UI records for 1964-77. Some states are not uniquely identified in the March CPS prior to 1977. Dates reference the previous year's CPS earnings question. For example, wage data for 1998 is from the March 1999 CPS.

Table 4
Wage Growth Function

	Slope	Average Intercept: Raw	Average Intercept: Adjusted	Real Wage growth Associated With Unemployment Rate of: 4%	6%	8%	Unemployment Rate Consistent with 1980-99 Average Real Wage Growth
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	-0.038	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)

Table 5
State Wage-Unemployment Elasticities [1]
Alternative Labor Market Indicators

	BLS Unemp. rate	ORG Unemp. rate	Unemp. plus NILF who want job	Unemp. plus NILF who want job plus PT for econ reasons	White, male, age 25-54 unemp. rate	Emp-pop ratio[2]	Exit rate out of unemp.[3]	Unemp. 0-5 weeks[3]	Unemp. 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=1995-99	0.086	0.023	0.024	0.002	0.055	0.936	0.000	0.001	0.020
UR, 1980-84=1995-99	0.074	0.022	0.015	0.001	0.058	0.436	0.000	0.002	0.043
UR, 1985-89=1995-99	0.092	0.024	0.026	0.003	0.094	0.897	0.002	0.000	0.034
UR, 1990-94=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

Notes:
*=significant at 5 percent level.
[1] UR=unemployment rate. All regressions includes state and year fixed effects and are estimated using robust regression.
[2] Detrended.
[3] 1994 is excluded.

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Figure 2

Annual Log Unemployment Rate Coefficients