

# The union-nonunion wage differential over the business cycle

## Evidence from PSID \*

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There is some prior evidence [Lewis (1963), Mellow (1981), Freeman (1984)] to support the notion that union wage premium is counter-cyclical. This phenomenon may be a product of union wage rigidity (which results from: (a) long-term union contracts, and (b) union reliance on layoffs due to seniority dominance rather than work-share practices in response to worsening economic conditions) and efficiency wage considerations. This paper, within a pooled cross-section time-series framework, empirically tests for the presence of perceived counter-cyclical nature of the union–nonunion wage differential using the PSID data from 1979–1984. Results from this study do support the presence of counter-cyclical union wage premium during the time period studied.

### 1. Introduction

There has been limited research on the effects of the economy on the union–nonunion wage differential. Henry Gregg Lewis (1963) and Wesley Mellow (1981) have, however, investigated the cyclical trends of the union wage premium<sup>1</sup>. Lewis employed an aggregate model for the years 1920–1958 to compare the average wages of highly unionized and sparsely unionized sectors; and Mellow, using a First differencing technique based on CPS data, compared the change in the overall union–nonunion wage differential during a recessionary period (1974–1975) and during an economic expansion (1977–1978). Both found the union wage premium to be counter-cyclical. This paper

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<sup>1</sup> Other noteworthy studies include Richard Freeman's (1984) confirmation of Mellow's findings; and Freeman and Medoff's discussion, in chapter 7 of their book on unions (1984), of the effects of business cycles on hours of work and employment in both unionized and nonunionized industries. Interestingly enough, Michael D. Deich and John S. Heywood (1987) studied the effects of union on economic activity. According to their study, there is no evidence that unionism deters economic activity.

offers not only a more extensive investigation of fluctuation that Mellow detected in the wage differentials, but also a conversion of Lewis' aggregate time-series model to a micro pooled cross-section time-series format.

Of the theories which explain the fluctuations in the union-nonunion wage differential, the most prominent are union wage rigidity – as caused by long-term contracts and union reliance on lay-offs – and the presence of efficiency wages, of which nonunion shops pay less during periods of weak economic growth. This paper presents evidence from 1979 through 1984 using Panel Study of Income Dynamics (PSID) data that supports the notion of a counter-cyclical union wage premium.

## 2. Union wage rigidity and efficiency wages

Union wages are rigid for two reasons. First, the cost of negotiating forces employers to agree to long-term contracts<sup>2</sup>. These multi-year agreements, which cost of living agreements<sup>3</sup> (COLA's) often inflate, hold real wage growth steady<sup>4</sup>. Only if the unions agree to wage concessions, or if inflation outpaces the COLA constraints, will the real wage growth of union members slowdown. Second, union shops tend to rely more on lay-offs than on work-share practices<sup>5</sup>. Because senior workers possess the most powerful voice in union politics, it is customarily the union policy that junior workers, who are earning the lowest wages, suffer temporary job loss in periods of economic contraction<sup>6</sup>. Consequently, average real union wages may actually rise during a recession.

When there is a recession, there is a larger pool of unemployed workers seeking an ever shrinking number of jobs. Consequently, this competition forces growth in real wages to decline. On the other hand, unless the recession threatens the very existence of the union itself, unions will not concede to wage reductions<sup>7</sup>. Thus union wages hold steady and do not respond to short-term deviations in economic conditions. It is this rigidity in the face of a fluctuating general wage level which accounts for much of the variation in the union-nonunion wage differential. As the real nonunion wages vary in response to aggregate economic conditions and thus the nature of the labor market, so does (in the opposite direction) the union-nonunion wage differential.

It is important to note that the rigidity of the union contracts can work against the unions in a highly inflationary period. For instance, even the COLA's to which the unions consented in the late 1970's did not protect them from the 13.5% inflation during 1980. According to the U.S. Department of Labor real union wages and salaries declined by 2.7%, on average in 1980.

A second explanation for the counter-cyclical phenomenon is the use of efficiency wages in labor markets. According to Vroman, unions offer employers an inherent employee monitoring system for which they are compensated<sup>9</sup>. Conversely, nonunion shops pay an incentive or efficiency wage to encourage their employees to work harder and, in turn, avoid the high cost of implementing their own intricate monitoring system. During recessions, as the length of the average job search increases,

<sup>2</sup> According to Michael Wachter (1986), high transaction costs and labor legislation governing collective bargaining are the main causes for the existence of union wage rigidity.

<sup>3</sup> See Hendricks and Kahn (1983) about the impact of COLA clauses in the union contracts to combat the effects of inflation.

<sup>4</sup> Hendricks and Kahn in their seminal work on COLA's (1985) also predict that the union-nonunion wage differential in fact widens during the periods of unanticipated inflation.

<sup>5</sup> See James Medoff (1979).

<sup>6</sup> See Katherine Abraham and James Medoff (1983).

<sup>7</sup> This tendency is consistent with George Perry's (1980, 1986) concept of a wage norm, a generally accepted rate for wage increases. According to Perry, such a norm is insensitive to short-term cyclical fluctuations. See also Daniel Mitchell (1985, 1986) for the empirical evidence supporting the presence of wage norm relationship in mostly union sectors.

<sup>8</sup> See U.S. Department of Labor 'Employment Cost Index' and 'Consumer Price Index' releases.

<sup>9</sup> See S.B. Vroman (1989).

so does the cost of being fired. As a result, employees work harder to avoid losing their job and the need to pay efficiency wages declines. However, due to short-term wage rigidity, unions continue to collect their efficiency wages, which results in a higher union wage premium. In addition, Vroman argues that efficiency wages are more effective among the least compensated workers, who have yet to amass substantial savings. Accordingly, the disappearance of efficiency wages will cause relatively more shirking – and, in turn, firings – among the higher paid nonunion workers, driving downward the average nonunion wage and bolstering the union–nonunion wage differential.

### 3. Data and methodology

For estimation purposes, we used Panel Study of Income Dynamics (PSID) micro data for a period of six years (1979–1984), mainly to capture an entire business cycle. In order to avoid the heterogeneity biases discussed in Keane, Moffit, and Runkle<sup>10</sup>, the sample includes only white male participants who headed their households and were employed or on temporary lay-off throughout the 1979–1984 period. The data for these men is pooled<sup>11</sup> for the six years so that each of the 460 such individuals offers six observations (one for each year) to our total sample. To eliminate the effect of inflation on the level of overall wages, we deflate them with the consumer price index to put all hourly wages in 1982 dollars. Accordingly, our dependent variable is the log of real hourly wages ( $\ln RHW$ ). In addition to human capital variables (education, experience, experience squared; where  $\text{experience} = \text{age} - \text{education} - 6$ ), we employed industrial, occupational, regional, and union status ( $UNION = \text{unity for union members, and zero otherwise}$ ) dummies. Also, we included four other independent variables, which serve as barometers of economic activity.

The PSID survey reports the unemployment rate (percent) in each individual's county of residence. Although labor markets experience structural changes, over the 1979–1984 period the changes should be insignificant, allowing the unemployment rate to serve as a proxy for economic activity. Consequently, the first independent variable is the straight unemployment rate ( $UERATE$ ). It captures the immediate effects that the economy's vigor has on real wages. However, because even non-union wages are set at least six months – and probably one year – in advance, there is a delayed relationship between wages and economic growth. To capture this delayed effect, we use the lagged unemployment rate ( $LUERATE$ ) as another independent variable. Finally, in order to estimate the immediate and lagged effects of the unemployment rate on the union wage premium, we interact these independent variables with the union dummy variable, creating  $U*UERATE$  and  $U*LUERATE$ .

<sup>10</sup> See Michael Keane, Robert Moffit, and David Runkle (1988).

<sup>11</sup> Since the cross-section OLS estimates tend to overstate the union wage effects due to possible correlation between unmeasured productivity and union status [see Lewis (1986)], this paper employs pooled cross-section time-series framework to avoid unmeasurable sample heterogeneity by concentrating on changes in measured variables for given individuals, under the assumption that unmeasured variables remain constant over time. Pooling the data over both cross-section and time-series requires non-OLS estimation procedures because of potential correlation among the disturbances. The standard error components model attempts to account for these correlations, raising the asymptotic efficiency of the estimates. This is done by breaking the error term ( $\epsilon_{it}$ ) in eq. (1) into three separate components, an individual component ( $v_i$ ), a time component ( $e_t$ ), and a component accounting for the possibility that disturbances may be peculiar to an individual at a specific point of time ( $z_{it}$ ). The variants –  $v_i$ ,  $e_t$ ,  $z_{it}$  – are assumed to be independent of each other as well as independent of the independent variables included in the model. The best linear unbiased estimate for the coefficient vector  $\beta$  when the variance components are known is the generalized least squares estimates. However, when the variance components are unknown (which is the case in the present context), the GLS estimates cannot be computed. Instead, the modified (feasible) GLS estimates are computed based on the 'fitting-of-constant' method of Searle (1971).

As the unemployment rate increases, real wages in general fall as competition for jobs becomes more intense. At the same time, however, the union wage premium should increase, because, while union wages are rigid, real nonunion wages decline in response to this increased competition and to a waning necessity for employers to pay efficiency wages. Therefore, we anticipate that *UERATE* and *LUERATE* will assume negative signs, and that the coefficients for  $U * UERATE$  and  $U * LUERATE$  will be positive. Accordingly, the following is our empirical model<sup>12</sup>:

$$\begin{aligned} \ln RHW_{it} = & \beta_0 + \beta_1(\text{Education})_{it} + \beta_2(\text{Experience})_{it} + \beta_3(\text{Experience Squared})_{it} \\ & + \beta_4(\text{UNION})_{it} + \beta_5(\text{LUERATE})_{it} + \beta_6(\text{UERATE})_{it} \\ & + \beta_7(U * \text{LUERATE})_{it} + \beta_8(U * \text{UERATE})_{it} \\ & + (\text{Vector of Industrial Dummies}) + (\text{Vector of Occupational Dummies}) \\ & + (\text{Vector of Regional Dummies}) + \epsilon_{it}, \end{aligned} \quad (1)$$

where

$i = 1, 2, \dots, 460$  (cross-section), and  $t = 1979, 1980, \dots, 1984$  (time-series).

#### 4. Results and conclusions

The MGLS results of the micro, pooled, cross-section, time-series regression appear in table 1. Each of the independent variables adhere to a priori expectations. In addition, the standard human capital variables and the lagged unemployment rate are all significant at 1%. The union interaction with the unemployment rate ( $U * UERATE$ ) and the unemployment rate itself are significant at 10%. The only independent variable which appears insignificant is the union interaction with the lagged unemployment rate.

These results support the notion of a counter-cyclical wage differential, and highlight its immediate response to changes in the pace of economic growth. The results also indicate that the previous year's unemployment rate is more important than that of the present situation in determining real wage levels. This indication is understandable, given that most wages – and in particular union wages – are set many months or years in advance. On the other hand, the insignificance of the lagged unemployment rate's interaction with the union dummy variable demonstrates that the spread between union and nonunion wages is responsive primarily to present economic activity. Evidently, union wages rise relative to nonunion wages only during periods of weakened economic activity, even though wages in general react to labor market slack in both a timely and a lagged sense.

These results show that the relatively more perpetual contract negotiation within the nonunion sector and the disappearance of its efficiency wages – which cause a disproportionate number of job eliminations among the senior nonunion wage earners – drive the real wages of nonunion workers relatively lower. Because there appears to be no lagged effect of the economy on the real wages of union workers relative to those of their nonunion counterparts, as expected it indicates that nonunion wages are more responsive to economic conditions<sup>13</sup>.

<sup>12</sup> Wunnava and Okunade (1991) employed a similar specification in comparing cross-sectional and panel (pooled) union wage effects. Specifically their model controls for neither the lagged unemployment rate nor its interaction with union status.

Table 1  
MGLS regression results <sup>a</sup> (t-values in parentheses). Dependent variable: Ln real hourly wages (ln *RHW*).

<i>Intercept</i>	<i>Education</i>	<i>Experience</i>	<i>Experience Squared</i>	<i>UNION</i>	<i>LUERATE</i>	<i>UERATE</i>
0.70367 (4.56)	0.03237 (9.58)	0.01216 (4.60)	-0.00022 (4.33)	0.07772 (3.00)	-0.00521 (2.75)	-0.00296 (1.60)
[UNION interactions]						
<i>U*LUERATE</i>	<i>U*UERATE</i>					
0.00089 (0.36)	0.00415 (1.55)					
<i>Variance component estimates</i>						
$\hat{\sigma}_0^2$ (estimated cross-section error variance) = 0.03167						
$\hat{\sigma}_e^2$ (estimated time-series error variance) = 0.00315						
$\hat{\sigma}_z^2$ (estimated purely random error variance) = 0.01221						
Adj. $r^2$ = 0.4403 <i>F</i> -ratio = 51.25 Effective sample <sup>b</sup> : 2300						
<i>Partial derivative of lnRHW with respect to UNION</i> <sup>c</sup>						
1980	1981	1982	1983	1984	1980-1984	
0.1115 (8.23)	0.1140 (8.96)	0.1244 (9.27)	0.1237 (9.33)	0.1159 (8.75)	0.1172 (9.43)	
<i>Effect of unemployment on wages</i>						
Non-union workers <sup>d</sup>		-0.00817 (3.64)				
Union workers <sup>e</sup>		-0.00313 (0.71)				

<sup>a</sup> Other controls include dummies for industries, occupations, and regions. Full regression results are available upon request.

<sup>b</sup> Due to the inclusion of lagged unemployment rate (*LUERATE*) we lost one observation per individual in our sample.

<sup>c</sup> Evaluated at the variable means.

<sup>d</sup> Sum of *LUERATE* and *UERATE* coefficients.

<sup>e</sup> <sup>d</sup> plus *UNION* interaction coefficients.

In order to determine the overall union effect on wages during the six years of our sample, we evaluated the partial derivative of ln*RHW* with respect to *UNION*. This comprehensive effect of unions on wages is in accordance with a priori expectations (see the bottom half of table 1). That is, it is greatest during 1982, a year in which real GNP declined by 2.5%, and generally adheres to the expectation that the greater the percent change in real GNP, the smaller the union wage premium. At first blush, however, this relationship appears not to hold during 1980 and 1981. It may be a mistake, however, to focus on the business cycle, instead of the cycles in employment itself; due to changes in productivity, the employment peaks and troughs are infrequently coincidental with those of real GNP. Table 2 shows how these measures of the economy deviated during our sample period. This table shows that, if we were to use employment as a gauge of economic activity, we would expect the union wage premium to grow from 1980 to 1982, and then to deteriorate through 1984, which is exactly what our partial derivative evaluations show.

<sup>13</sup> In fact the total effect of lagged and current unemployment rate on wages computed separately for nonunion and union workers indicate that the nonunion wages are relatively much more responsive than union wages. See the bottom of table 1, especially the endnotes d and e.

Table 2  
Major economic indicators.

Year	Real GNP (percent change)	Nonfarm employment (Change in '000)	Civilian unemployment rate (Percent)
1979	2.48%	2061	5.85
1980	-0.16	233	7.17
1981	1.93	-68	7.62
1982	-2.55	-2106	9.71
1983	3.57	3487	9.60
1984	6.78	3947	7.51

Source: U.S. Department of Commerce, U.S. Department of Labor.

However, comparing two non-successive years tends not to support our hypothesis. That is, even though the economy added more jobs (or grew more quickly) in 1984 than it did in 1980, the wage premium was smaller in 1980. We believe that this apparent contradiction results because the labor market was not as structurally stable during the 1979-1984 period as we had anticipated<sup>14</sup>. For instance, the participation rate among the female work force climbed by 2.7 percentage points, while that of the male labor force fell by 1.5 percentage points<sup>15</sup>. These changes cause slack in the labor markets that neither percent change in real GNP nor change in payroll employment capture. For this reason, it appears that the union wage premium is more responsive to the economy's deviation from its potential growth than it is to absolute fluctuations in its growth rate.

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<sup>14</sup> However, this phenomenon is somewhat consistent with the predictions made by Hendricks and Kahn (1985) concerning the effects of unanticipated inflation on the union-nonunion wage differential due to COLA's.

<sup>15</sup> See U.S. Department of Labor "Employment Situation" release.

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