Countercyclical Union Wage Premium? Evidence for the 1980s*

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Empirical results based on pooled male data from the Panel Survey of Income Dynamics indicate an overall union wage premium of about 11.92 percent for the 1980s. In response to fluctuations in local labor market conditions, proxied by the local unemployment rate, a much more flexible wage-setting process is found in the nonunion sector relative to the union sector. The long-term effect of unemployment on nonunion real wages suggests an approximate 0.6 percent decline for every one percentage point increase in unemployment, a statistically significant reduction, but the long-term effect of unemployment on real wages of union members is negligible. The union wage premium ranges between 11.6 to 12.3 percent for the sample years. Even though union wages are insensitive to short-run fluctuations in local labor market conditions, and are somewhat countercyclical in nature, widespread union wage concessions which occurred during the 1980s may now be exerting a downward pressure on union wages.

I. *Background*

Since Keynes's (1936) hypothesis on countercyclical wage movements, a number of researchers have investigated how wage rates vary with the business cycle. Empirical evidence on the direction and magnitude of the association is mixed. Neftci (1978), for instance, finds real wages to behave countercyclically for the period 1948–1971, while studies by Raisian (1983) and Keane et al. (1988) detect a procyclical pattern for the 1966–1981 period. The theoretical constructs of Barro and Grossman (1971), however, yield conditions under which changes in real wages could be procyclical, acyclical, or countercyclical.

The theories of wage rigidities in long-term union contracts and the dominance of union seniority (or "risk shifting") rules for assigning recessionary layoffs in unionized industries are often cited in support of countercyclical relative wage behaviors (Medoff, 1979; Wachter, 1986). But the union-nonunion wage gap is narrowed during prolonged recessions by employer resistance and in situations where a union wage increase would significantly reduce union employment. However, in an inflationary recession,[1] such as the 1974–1975 U.S. recession, the union-nonunion wage gap widens as other wages rise by a smaller amount or not at all (Rees, 1989, p. 73).

The true effect of the business cycle on the union-nonunion real wage gap may be biased in the presence of selectivity, particularly when the researcher uses cross-sectional data. Selectivity biases can be explained by theories of heterogeneity in skills or in other worker attributes and hedonic wages relating wage rates to the differential risks of layoffs in recessions (Keane et al., p. 1237). One implication is that high-wage workers stand greater risks of layoffs in nonunion, as opposed to union, settings. Thus, their wage gap should widen in downturns to generate a countercyclical differential wage behavior. Within the framework of the standard job-search model, in which high-wage workers (with higher stocks of wealth and reservation wages) can afford longer job searches, the likelihood increases that higher-wage nonunion workers would tend to remain unemployed longer (than low-wage earners) in a slump. Thus, this phenomenon also tends to generate a countercyclical overall union-nonunion wage differential. Currently, the research on the behavior of union-nonunion wage differential over the business cycle is somewhat dated,[2] and further research is important, especially in light of the well-publicized union wage concessions of the 1980s. Accordingly, we use the most recently available data series.

II. *Motivation, Methodology, and Data*

Panel or longitudinal data enables the researcher to avoid unmeasurable sample heterogeneity of earlier cross-sectional methods by concentrating on *changes* in measured variables for given individuals, under the assumption that unmeasured variables remain constant over time. Thus, the selectivity bias due to observed heterogeneity is minimized by fitting a (log of) real hourly wage (dependent variable: ln *W*) equation for a *pooled*[3] sample of male heads of households,[4] for the years 1981 through 1989 Panel Survey of Income Dynamics (PSID) microdata compiled by the University of Michigan. Control variables include human capital attributes [education dummies,[5] actual full-time labor market experience (experience square): *EXP* (*EXP2*), and tenure (tenure square): *TENURE* (*TEN2*)]; union status: *UNION* (=1 for union member, 0 otherwise); measure of business cycle[6] proxied by regional unemployment rate lagged (by two periods: *LUERATE2*, by one period: *LUERATE*) and current(*UERATE*), and union-unemployment rate interactions (lagged by two periods: *UNION**LUERATE2*, lagged by one period: *UNION**LUERATE* and current: *UNION**UERATE*). The model includes both lagged and current regional unemployment rates along with union status interactions specifically to test the hypothesis that, due to the multi-year contracts, unions resist wage cuts in response to worsening labor market conditions. Accordingly, it is logical to expect a *negative* sign for all the unemployment rate coefficients but a *positive* sign for all the union status-unemployment rate interactions. Other controls are for race (*WHITE*=1, 0 otherwise); marital status (*MARRIED*=1, 0 otherwise); regional (*SOUTH*=1, 0 otherwise); and occupational, and industry categories. The following is the empirical specification:

$$\begin{aligned}\ln W_{it} = {} & a_0 + [\text{vector of } EDUCATION \text{ dummies}] + b_1(EXP)_{it} \\ & + b_2(EXP2)_{it} + b_3(TENURE)_{it} + b_4(TEN2)_{it} + b_5(UNION)_{it} \\ & + b_6(LUERATE2)_{it} + b_7(LUERATE)_{it} + b_8(UERATE)_{it} \\ & + b_9(UNION{*}LUERATE2)_{it} + b_{10}(UNION{*}LUERATE)_{it} \\ & + b_{11}(UNION{*}UERATE)_{it} + b_{12}(WHITE)_{it} + b_{13}(MARRIED)_{it} \\ & + b_{14}(SOUTH)_{it} + [\text{vector of } OCCUPATION \text{ and } INDUSTRY \text{ dummies}] \\ & + e_{it},\end{aligned}$$

where $i = 1,2, \ldots, 889$ (cross-sectional units), and $t = 1983, 1984, \ldots, 1989$ (time-series units).

III. *Empirical Results and Conclusions*

Table 1 presents the main regression results. Human capital[7] and other demographic[8] variables behave as expected. As predicted, both lagged and current unemployment rate coefficients are negative (because wages in general are depressed due to worsening labor market conditions) and their interactions with union status are indeed positive (i.e., unions resist wage cuts[9] in such circumstances and also union contracts are multi-year and, for the most part, are insulated from any short-run fluctuations). From equation (1) the union effect can be obtained by partially differentiating $\ln W$

Table 1

The University of Michigan's Panel Survey of Income Dynamics (PSID) Pooled (1981–1989) Regression Estimates of Cross-sectionally Correlated and Time-wise Autoregressive Model

(Dependent Variable: natural log of real wages)

Variable (Description)	Estimate	t-ratio
Constant	1.6896	23.536
Schooling Dummies (omitted HS dropouts)		
EDUC1 (High School)	.179	17.687
EDUC2 (Some College)	.273	28.325
EDUC3 (College)	.522	44.235
EDUC4 (College+)	.624	43.273
EXP (Yrs of full-time actual experience)	0.5640E-02	10.177
EXP2 (EXP^2)	-0.423E-04	-4.067
TENURE (# of months at present employer)	0.987E-03	20.206
TEN2 ($TENURE^2$)	-0.903E-06	-16.877
UNION (=1 if union member, 0 otherwise)	0.542E-01	3.687

Table 1 — *Continued*

Variable (Description)	Estimate	*t*-ratio
Regional Unemployment Rate		
LUERATE2 (Lagged by two periods)	-0.174E-02	-3.186
LUERATE (Lagged by one period)	-0.309E-02	-5.957
UERATE (Current)	-0.195E-02	-4.258
Union-unemployment interactions		
(*UNION**LUERATE2*)	0.326E-02	3.006
(*UNION**LUERATE*)	0.292E-02	2.732
(*UNION**UERATE*)	0.116E-02	1.199
WHITE (=1 if white, 0 otherwise)	0.134	17.778
MARRIED (=1 if married, 0 otherwise)	0.325E-01	4.352
SOUTH (=1 if from South, 0 otherwise)	-0.773E-01	-12.107
Occupation dummies (omitted private household workers)		
OCC1 (Professional)	0.285	4.056
OCC2 (Technical)	0.276	3.922
OCC3 (Managers)	0.227	3.194
OCC4 (Administrators)	0.228	3.261
OCC5 (Sales)	0.178	2.537
OCC6 (Clerical)	0.152	2.160
OCC7 (Craftsmen)	0.6941E-01	0.983
OCC8 (Non-transport operators)	0.107	1.509
OCC9 (Transport operators)	-0.901	-3.537
OCC10 (Non-farm laborers)	-0.259	-2.530
OCC11 (Service)	0.776E-01	1.099
Industry dummies (omitted public administration workers)		
IND1 (Forestry/Fisheries)	-0.150	-2.928
IND2 (Mining)	0.831E-01	2.717
IND3 (Construction)	-0.196E-01	-1.912
IND4 (Manufacturing)	0.334E-01	4.302
IND5 (Transportation)	0.742E-01	8.232
IND6 (Wholesale)	-0.357E-01	-3.606
IND7 (Retail)	0.569E-01	4.157
IND8 (Finance/Real estate)	-0.280E-01	-2.280
IND9 (Repair services)	-0.268	-6.956
IND10 (Personal services)	-0.432E-01	-1.558
IND11 (Entertainment)	-0.150	-15.155

BUSE–R^2	.621
F-Value	252.713
Sample Size	6,223

From equation (1) the union effect can be obtained by partially differentiating ln$W$ with respect to *UNION*:

$$(\partial \ln W / \partial UNION) = b_5 + b_9(LUERATE2)_{it} + b_{10}(LUERATE)_{it} + b_{11}(UERATE)_{it}. \quad (2)$$

Evaluating (2) at the sample means, the overall union wage premium is about 11.92 percent, and the range for the sample years is between 11.6 to 12.3 percent.[10] Negative and statistically significant unemployment rate variables (i.e., *LUERATE2*, *LUERATE*, and *UERATE*) may indicate a much more flexible wage-setting process in the nonunion sector (relative to union sector) in response to local labor market conditions. As expected, all the union-unemployment-rate interaction variables are positive which is consistent with the argument that union wages are insensitive to short-run fluctuations of local labor-market conditions. Thus, it may be the general economy-wide union wage concessions[11] that lead to only a 12 percent union wage premium for the 1980s, in spite of union efforts to oppose wage cuts due to worsening local labor market conditions. The long-term effect of unemployment on nonunion real wages yields a 0.6 percent decline[12] for every one percentage point increase in unemployment and is statistically significant. On the other hand, the long-term effect of unemployment on real wages of union members[13] is negligible. These findings are very similar to those of Wunnava and Honney (1991).

NOTES

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[1]See Hendricks and Wallace (1983) about COLA clauses in union contracts in combating the effects of inflation.

[2]Some of the previous studies are by Lewis (1963), utilizing an aggregate model based on 1920-1958 time series data; Pencavel and Hartsog (1984), using annual Current Population Survey (CPS) data for years 1920 through 1980 to extend Lewis's work; Moore and Raisian (1980), employing Panel Study of Income Dynamics (PSID) stacked data for years 1967 through 1974; Moore and Raisian (1987), using a similar methodology as their earlier study estimated union-nonunion differentials exclusively for public administration, educational, and private sectors based on PSID 1970 through 1979, and CPS sample for years 1979 and 1983; Mellow (1981), fitting a differenced model to 1974-1975 and 1977-1978 micro CPS data; Mincer (1983), using almost similar methodology as Mellow to National Longitudinal Survey (NLS) panels for Young Men (YM) and Older Men for 1969 and 1971, and PSID 1968 and 1978 (he treated the PSID for 1968-1978 as a time series of adjacent year pairs) data. Using the same framework as Mellow and Mincer, Freeman (1984) utilized CPS for the year pair 1974 and 1975, PSID for the year pair 1970 and 1979, NLS, YM, for the year pair 1970 and 1978, and Quality of Employment Survey (QES) for the year pair 1973 and 1977; Lewis (1986), surveying about 200 post-1963 studies which employed mostly the above data sets; and Wunnava and Honney (1991), based on 1979-84 PSID pooled data. Even though the above studies differ in specification, sample period, data, and econometric treatment — the common theme is that union wage premium is positive, ranges between 10 to 25+ percent across different demographic groups, and is somewhat countercyclical in nature.

[3]By assuming a cross-sectionally heteroscedastic (also mutually correlated) and time-wise autoregressive model, a GLS-based estimation method can be used to obtain efficient estimates. Subjecting the observations

to two transformations, one designed to remove autocorrelation and the other to remove heteroskedasticity, yields a disturbance term that is asymptotically nonautoregressive and homoscedastic. Briefly, OLS is applied to obtain the original regression residuals; then these are used to perform transformations so that the errors are asymptotically nonautoregressive and homoscedastic. For details see Kmenta (1986, pp. 618-22). The *CORCOEF* option available under the *POOL* command of SHAZAM (White, 1993) statistical software is used for estimation purposes.

[4]To avoid possible biases due to gender discrimination only male heads of household have been selected. Our main objective is to investigate the range of union effects on wages for the 1980s circumventing the heterogeneity bias of cross-sectional estimates (for example, estimates based on CPS data). This is possible if, and only if, enough *consistent* time-series data are available on given individuals who can be followed over time. Currently, the only available longitudinal data set containing information on wages, union status, standard human capital, and demographic variables *for each and every year* during the 1980s is the Panel Survey of Income Dynamics micro data. The other available longitudinal data — NLS (National Longitudinal Survey) — are not surveyed every year, and also do not contain accurate information regarding union status and unemployment rates (crucial variables to implement the proposed methodology). The same cross-section of 889 individuals with consistent data were followed throughout sample years. Out of the 889 set of same individuals, the number of union members ranged from 229 to 262 through the sample years. Further examination of the data indicated that "always union members" were 123, and "always nonunion members" were 546. Hence, in the sample there were 220 (i.e., 889 - [123 + 546]) individuals who made at least one switch in their union status. Due to the inclusion of the unemployment rate lagged by two time periods in the model, two observations per individual in the sample have been lost. The time-series component therefore, is 1983 through 1989 (i.e., 7 years) — resulting in a pooled sample of 6,223 ($=889 \times 7$). Wages are expressed in 1981 dollars.

[5]A vector of four education dummies consisting of highschool graduates, some college, college, and advanced degrees are included in the model; the omitted category is high-school dropouts.

[6]Cain (1979) makes a very strong case for using unemployment statistics as an indicator of cyclical performance of the economy. Lilien (1982) shows that much of the overall cyclical unemployment represents sectoral shifts in demand during the 1970s. Raisian (1983) uses the unemployment rate in the industry in which the individual worked minus the average unemployment rate for that industry as a cyclical proxy.

[7]The returns to education seem to increase with added years of schooling. The experience-earnings and tenure-earnings profiles are nonlinear in nature with correct signs. All human capital variables are statistically significant.

[8]White males enjoy a 14.33 percent wage advantage over their nonwhite counterparts; married men have a 3.3 percent wage premium relative to single men; and southern males, on average, are paid about 8 percent less than males from other regions of the country.

[9]Due to the dominance of union seniority rules, unions may rely mostly on recessionary layoffs rather than wage cuts. See Medoff (1979) and Wachter (1986).

[10]$[.0541 + .00325(\overline{LUERATE2} = 7.81) + .00292(\overline{LUERATE} = 7.618) + .0016(\overline{UERATE} = 7.31)] = .1127$ (standard error for this expression is .006106, resulting in a t-statistic of 18.457 — the actual numerical bootstrapping of the standard error can be obtained on request). Given the log-linear model, the union wage premium $(e^{0.1127} - 1) = .1192$ or 11.92 percent. To derive the annual union wage effects, equation (2) can also be evaluated at the year specific sample means. The percent of union wage premium based on year specific sample means are: 1983, 11.8; 1984, 11.7; 1985, 11.6; 1986, 12.3; 1987, 12.3; 1988, 12.1; and 1989, 12.2. This trend is consistent with the argument that unions are unable to hold on to their wage gains of 20+ percent from the 1970s and are forced to make wage concessions (see Bell, 1989). Also fading coverage of COLA provisions in the majority of union contracts due to easing inflationary pressure on the economy may be a contributing factor (see *Monthly Labor Review*, January issues for 1982 through 1989). Interestingly, in a longitudinal study using only one-period lagged unemployment rate, current unemployment rate, and union interactions (based on PSID covering 1980-1984 period) along with other standard controls, Wunnava and Honney (1991) showed the overall union wage premium to be about

12.43 percent, and the range in their sample years to be between 11.8 and 13.24 percent. The specification proposed in equation (1) seems to be far superior to one proposed by Wunnava and Honney (1991). The F-test to test the composite null hypothesis: $b_6 = b_9 = 0$ is statistically significant — F statistic = 6.51 (with 2, 6182 degrees of freedom — yielding a significance level of .00151).

[11]Bell (1995), using firm-specific data for years 1980 through 1987, demonstrated that union wage concessions were most likely in small firms, in firms paying high wages, and in firms with relatively low union coverage.

[12]The long-term effect of unemployment is derived by the sum of the estimated coefficients of *LUERATE2, LUERATE,* and *UERATE* (i.e., $b_6 + b_7 + b_8 = -.006$). The estimated standard error for the sum of these coefficients is .000928 (actual numerical derivation can be obtained on request), resulting in a t-value of -7.32.

[13]The sum of the unemployment rate and the union unemployment rate interaction coefficients (i.e., $b_6 + b_7 + b_8 + b_9 + b_{10} + b_{11}) = .00056$ (with a t-value of .187). The actual numerical derivation of the estimated standard error can be obtained on request.

REFERENCES

Barro, Robert J., and Herschel I. Grossman. "A General Disequilibrium Model of Income and Employment." *American Economic Review* 61 (March 1971): 82-93.

Bell, Linda. "Union Concessions in the 1980s." *Federal Reserve Bank of New York Quarterly Review* 14 (Summer 1989): 44-58.

__________. "Union Concessions in the 1980s: The Importance of Firm-Specific Factors." *Industrial and Labor Relations Review* 48 (January 1995): 258-75.

Cain, Glen. "The Unemployment Rate as an Economic Indicator." *Monthly Labor Review* 102 (March 1979): 24- 35.

Freeman, Richard. "Longitudinal Analysis of the Effects of Trade Unions." *Journal of Labor Economics* 2 (January 1984): 1-26.

Hendricks, Wallace, and Lawrence Kahn. "Cost of Living Clauses in Union Contracts: Determinants and Effects." *Industrial and Labor Relations Review* 36 (April 1983): 447-60.

Keane, Michael, Robert Moffitt, and David Runkle. "Real Wages over the Business Cycle: Estimating the Impact of Heterogeneity with Micro Data." *Journal of Political Economy* 96 (December 1988): 1232-66.

Keynes, John. *The General Theory of Employment, Interest, and Money.* London: Macmillan, 1936.

Lewis, H. Gregg. *Unionism and Relative Wages in the United States.* Chicago: University of Chicago Press, 1963.

__________. *Union Relative Wage Effects: A Survey.* Chicago: University of Chicago Press, 1986.

Lilien, David. "The Cyclical Pattern of Temporary Layoffs in United States Manufacturing." *Review of Economics & Statistics* 62 (February 1980): 24-31.

Medoff, James. "Layoffs and Alternatives Under Trade Unions in United States Manufacturing." *American Economic Review* 69 (June 1979): 380-95.

Mellow, Wesley. "Unionism and Wages: A Longitudinal Analysis." *Review of Economics & Statistics* 63 (February 1981): 43-52.

Mincer, Jacob. "Union Effects: Wages, Turnover, and Job Training." *Research in Labor Economics* (Supplement 2, 1983): 217-52.

Moore, William, and John Raisian. "Cyclical Sensitivity of Union/Nonunion Relative Wage Effects." *Journal of Labor Research* 1 (Spring 1980): 115-32.

__________. "Union-Nonunion Wage Differentials in the Public Administration, Educational, and Private Sectors: 1970- 1983." *Review of Economics & Statistics* 69 (November 1987): 608-16.

Neftci, Salih N. "A Time-Series Analysis of the Real Wages-Employment Relationship." *Journal of Political Economy* 86, Pt. 1 (April 1978): 281-91.

Pencavel, John, and Catherine Hartsog. "A Reconsideration of the Effects of Unionism on Relative Wages and Employment in the U.S., 1920-80." *Journal of Labor Economics* 2 (April 1984): 193-232.

Raisian, John. "Contracts, Job Experience, and Cyclical Labor Adjustments." *Journal of Labor Economics* 1 (April 1983): 152-70.

Rees, Albert. *The Economics of Trade Unions.* 3rd ed. Chicago: University of Chicago Press, 1989.

Wachter, Michael. "Union Wage Rigidity: The Default Settings of Labor Law." *American Economic Review* 76 (May 1986): 240-44.

White, Kenneth J. *Shazam User's Reference Manual* (Version 7.0) New York: McGraw-Hill, 1993.

Wunnava, Phanindra V., and John K. Honney. "The Union-Nonunion Wage Differential over the Business Cycle." *Economics Letters* 37 (1991): 97-103.