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CHANGES IN THE UNION WAGE PREMIUM BY INDUSTRY

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A data appendix with additional results, and copies of the computer programs used to generate the results presented in the paper, are available from the authors at Department of Economics, Kansas State University, Waters Hall 327, Manhattan, KS 66506-4001.

ABSTRACT

Relying on CPS data, we compute estimates of the union wage premium, conditional on worker characteristics, for 32 private-sector industries over the period 1971-99. The dispersion of union premiums across industries has narrowed over time as high premiums have tended to fall and low premiums rise. At the aggregate level, the premium has drifted lower. When we model the union premium as a function of cyclical and structural variables and unmeasured industry characteristics, we find that COLA clauses reduce the cyclicality of the union premium and that increases in import penetration are strongly associated with rising union premiums. The effect of deregulation is mixed.

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Union workers receive higher pay than comparable nonunion workers, with the size of this premium varying over time. For example, Johnson (1984) estimates that the aggregate union premium ranged from 2 percent in 1945-49 to 38 percent in 1930-34. There is no reason, however, to anticipate equal movements of the premium in all industries. The effects of deregulation, heightened import competition, and technological change are felt unevenly across industries. At the aggregate level the union premium has become more stable,[1] but there is a question as to whether this recent stability is masking divergent trends at the industry level. For example, Linneman, Wachter, and Carter (1990) found evidence that the union premium was trending lower in some industries but higher in others, though they did not attempt to explain the reasons for different trends by industry.

The first purpose of this paper is to compute estimates of the union premium by industry and year that are more current and more detailed than those of Linneman, Wachter, and Carter. Their study provided estimates for ten industries over the period 1973-86.[2] Our estimates are for 32 industries and cover the years 1971-99. Second, we examine stability over time in the size of the union premium by industry and consider cyclical sensitivity and time trends. We also examine a previously untested proposition that higher COLA coverage reduces the cyclical responsiveness of the union premium. Finally, we explore the extent to which union wage premiums respond to such structural changes in the economy as economic deregulation and increased import competition. Although various theories suggest that increased import

[1]Filer, Hamermesh, and Rees (1996:501) conclude that "except for the period around 1970, the effect of unions on relative wages has remained essentially stable at between 15 and 20 percent since 1950 no matter what the macroeconomic conditions."

[2]More recent estimates for six private-sector industries are available in Hirsch and Macpherson (2000b).

penetration may cause the union premium to expand, prior research has failed to uncover a statistically significant linkage between imports and the union premium.

Estimating the Union Premium

In estimating the union wage premium, we adopt the methodology of Linneman, Wachter, and Carter (1990) and compute the union premium from Mincerian wage regressions that control for other characteristics of workers. Estimates are based on samples drawn from all Current Population Surveys (CPS) between 1971 and 1999 in which respondents were asked about union membership. As such, the data come from the March 1971, May 1973-81, and the Outgoing Rotation Groups of the 1983-99 surveys[3]—yielding estimates of the union wage premium for 27 years. We exclude from the regression samples agricultural workers and those who are self-employed, usually work one hour or less per week, or work without pay. We also drop observations with missing data on variables used in the analysis, leaving samples of 2,926,715 workers over the 27 years.

The wage equation underlying estimates of the union premium is:

$$(1) \qquad \ln w_{it} = \beta_t' X_{it} + \sum_j \alpha_{jt} I_{ijt} + \sum_j \pi_{jt} I_{ijt} U_{it} + \varepsilon_{it},$$

where w_{it} denotes the weekly wage of individual i in year t, X is a vector of control variables (specifically, education, experience and its square, and indicator variables for gender, black, other non-white, married, SMSA residence, part-time work, eight census divisions, and twelve occupations), I_j is an indicator variable for industry j, U is an indicator variable for union

[3]Information on union status was not collected in 1972 or 1982.

membership, and ε is the classical regression error. The coefficient π_{jt} captures the union wage premium in industry *j* in year *t*—the wage differential between union and nonunion workers conditional on measured characteristics of the worker. Estimates of the average union premium over the sample period are presented in the appendix for each of 32 two-digit industries, and estimates for individual years are available from the authors on request.[4] Figure 1 depicts the premium graphically for eight aggregate industries as well as for the private sector of the U.S. economy.

For the private sector, the average premium over our sample period is .155 log points, or 16.8 percent. Movements of this series through 1986 mirror the aggregate estimates of Linneman, Wachter, and Carter.[5] In both cases, the union premium reached its high in 1977, bottomed out in 1979, and then rebounded between 1979 and 1986. The one difference is that the aggregate estimates of Linneman, Wachter, and Carter are slightly lower because they

[4]For 1971, the premium could be estimated for only 30 industries because coding of the industry variable did not allow us to separate the "trucking," "air transportation," and "other transportation" industries.

[5]Our estimates are also comparable to those of Lewis (1986:179) based on his survey of 35 studies of the union wage effect. In particular, for the years when the studies overlap, the pattern of estimates is as follows:

Year	Present Study	Lewis
1971	.15	.15
1973	.15	.15
1974	.14	.14
1975	.18	.16
1976	.19	.18
1977	.20	.17
1978	.18	.17
1979	.13	.13

Likewise, our estimates parallel those of Curme and Macpherson (1991) over their sample period, 1979-88. They found that the premium increased between the subperiods 1979-81 and 1983-85 and edged lower over the final three years of the sample period.

included government workers, whose premium during their sample period averaged only 5.6 percent.

Our results confirm the relative stability of the aggregate premium observed by Linneman, Wachter, and Carter for the period 1973-86, although a slight downward drift in the premium is observed when the sample period is extended through 1999. (When we regress the aggregate premium, in log points, on a linear time trend, the coefficient of *TIME* is -.0010, significant at the 5 percent level.) Even so, over the years 1971-99, the aggregate premium fluctuated narrowly by historical standards (from 13 percent to 22 percent), and the value in 1999 was only 1.4 percentage point lower than twenty years earlier. Yet when the union premium is examined at a disaggregate level, diverging trends become evident, again consistent with the observations of Linneman, Wachter, and Carter.

As shown in Figure 1, there is a strong downward trend of the premium in construction and in wholesale and retail trade and a more modest and volatile decline in mining. In service and finance, the trend was lower initially and, after stalling in the early 1980s, a downward movement may have reemerged in the late 1990s. Upward trends are observed in durable-goods manufacturing, transportation, and communications and utilities. In nondurable-goods manufacturing, no trend is apparent. The trends depicted in Figure 1 for the sample period 1973-86 coincide with those reported by Linneman, Wachter, and Carter. In most cases the trends continued after 1986. The two exceptions are wholesale and retail trade, where the upward movement in the early period was reversed in the later period, and service and finance, where a downward trend was replaced by stability, at least through the mid-1990s.

In addition to presenting estimates of the union premium for a more finely disaggregated set of industries, the appendix also provides the estimated coefficients of a linear time trend when the premium was estimated as a function of this variable. The table reveals that upward and downward trends have been equally common. In sixteen of the industries, the trend is statistically significant at the 5 percent level—rising in half of these industries and falling in the other half. The trends are not completely offsetting, however, because of larger employment shares in industries experiencing significant downward trends in the premium.

There is also evidence of a convergence of the union premium across industries. The premium has generally declined in industries for which the premium was larger than average in the 1970s (e.g., construction and wholesale and retail trade) while rising in industries that started with a low premium (durable-goods manufacturing, communications and utilities).[6,7] To test the degree and significance of this convergence, for each year we computed the standard deviation of the union premium across the 32 industries and then fitted the following regression of this standard deviation on a time trend:[8]

$$(2) \qquad \underset{}{STD.DEV_t} = \underset{(.0032)}{.1191} - \underset{(.0002)}{.0027}\, TIME_t + e_t, \qquad R^2 = .860, \qquad N = 26,$$

[6]Linneman, Wachter, and Carter (p. 51) appear to reach a different conclusion: "The already high-premium industries have generally been increasing their wage premiums." Although the authors do not provide a basis for this claim, their conjecture turns out to be unimportant to their central finding—that unions have been losing employment share in industries with high union premiums.

[7]As a referee observed, one factor that may be contributing to a convergence of the union premium is the increased concentration of unions over time. As union membership has become more concentrated among a relatively small number of union organizations, the number of bargaining styles and targets has apparently diminished, resulting in less variability of the union premium across industries.

[8]This experiment excludes 1971 because of the different industry coding that year.

where standard errors are listed in parentheses and e_t denotes the regression residual. The regression indicates a significant narrowing of differences in the union premium across industries. Figure 2, which plots the time series of the standard deviation of the union premium along with the lowess smoother of the scatter points, further illustrates the compression of the union premium over time. Over the sample period, the standard deviation of the union premium across industries declined by 48 percent from 1973-77 (five-year average standard deviation = .114) to 1995-99 (average = .059). As indicated by the figure, the reduction in dispersion is less obvious during the first half of the sample period (which coincides with the period examined by Linneman, Wachter, and Carter) than in the latter half.

Modeling Changes in the Union Premium

Having documented substantial and significant changes in the union premium by industry, we next explore reasons for these movements, starting with cyclical considerations.

Cyclicality, Inflation, and COLA Coverage

Because of greater rigidity of union wages, the union wage premium varies over the business cycle.[9] Long-term contracts limit the influence of cyclical fluctuations on union wages. In addition, compared to the nonunion sector, when market demand deteriorates unions are more likely to accept reduced employment, especially of junior workers, in order to preserve wages. Conversely, when markets strengthen, union wages rise more modestly than nonunion wages. Thus, union wages are more immune to the ups and downs of the business cycle than are nonunion wages.

The degree of cyclicality of the union premium depends on inflationary considerations and the nature of union contracts. When inflation is greater than anticipated, cyclicality is amplified. Whereas the union premium contracts as labor markets tighten, that contraction is even greater when inflation exceeds expectations (because nonunion wages are more responsive to the higher inflation). Of course, the effects of strengthening markets and increased inflation can be offset partially through cost-of-living adjustments (COLAs), which make union wages more responsive to fluctuations in aggregate demand. Therefore, while the union premium shrinks during economic upturns, the magnitude of the compression is likely to be inversely related to the degree of wage indexation in the union sector. In fact, various authors have not only argued that cyclical sensitivity of the union premium declined in the 1970s and 1980s, but they have also suggested that a likely reason for this diminished sensitivity is the increased percentage of the unionized work force covered by COLA clauses (Moore and Raisian 1980, Hendricks 1981, Ashraf 1990). Although such conjectures appear plausible, and warrant investigation, the effect of COLA coverage on the cyclicality of the union premium remains untested.

In addition to dampening the cyclicality of the union premium, COLA coverage may also be linked directly to the premium. Indeed, increases in the premium in the 1970s and the early 1980s are often attributed to higher COLA coverage.[10] One reason, as Mitchell (1980:149) observed, is that "the ability of a union to obtain an escalator clause may be correlated with its

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[9]Among empirical studies to document the influence of the business cycle on union-nonunion wage differentials are Lewis (1963, 1986), Moore and Raisian (1980), Hendricks (1981), Pencavel and Hartsog (1984), and Wunnava and Honney (1991). See McDonald and Solow (1985) for a classic theoretical treatment.

[10]For example, see Mitchell (1989); Wachter and Carter (1989); Linneman, Wachter, and Carter (1990); and Ashraf (1990).

relative bargaining power," a proposition supported by empirical analysis.[11] To the extent COLA coverage proxies for union bargaining power, increased COLA coverage is expected to result in a widening union premium. Finally, apart from influencing cyclicality of the union premium, deviations of actual and expected inflation may have a direct effect on the union premium. Other things equal, the premium should be lower in periods in which unions underestimate the extent of inflation.

Structural Shifts

The union premium is also likely to be influenced by two major structural changes: economic deregulation and increased international trade. Both changes affect competition in product markets and, in turn, in labor markets. Whether the union wage premium rises or falls depends on the relative response of union and nonunion wages.

Consider first the effects of deregulation. Where deregulation removes entry barriers and introduces price competition, the result is likely to be lower union rents and a compression of the union premium. This appears to be what occurred in the trucking industry. Following deregulation, low-cost firms entered the industry, trucking rates fell, and unions lost market share. Union wages declined, but nonunion wages were little affected. The union premium, which had been unusually large, apparently because of regulatory rent creation (Moore 1978), fell to a level more typical of the overall economy (Rose 1987, Hirsch 1988, 1993).

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[11]COLA coverage has been found to be positively and significantly related to such proxies of union power as size of the bargaining unit, industry unionization rate, and the inverse of the number of unions in the industry. See Ehrenberg, Danziger, and San (1983); Hendricks and Kahn (1983); Prescott and Wilton (1992); and Ragan and Bratsberg (2000).

In other industries, however, regulation apparently created no rents for unions to capture or actually dissipated rents.[12] Accordingly, Hendricks (1994) argued that the effect of deregulation varies from industry to industry depending on such factors as whether or not regulation created rents for unions, the extent to which such rents were shared with nonunion workers (by nonunion employers who felt threatened by potential unionization), and the effect of deregulation on union power (resulting from expansion of the nonunion sector, elimination of mutual aid pacts, switching from national to regional bargaining, etc.).

Hirsch and Macpherson (1998, 2000a) reached a similar conclusion and further found evidence that the adjustment to deregulation takes different paths in different industries. For example, in trucking, union wages fell immediately and appreciably following deregulation whereas nonunion wages, already at or near competitive levels, showed little change. As a result, the union premium shrank quickly following deregulation. In contrast, in the airline industry wages of nonunion workers eroded following deregulation, as the substantial rents they earned under regulation dissipated. Unions, however, retained considerable bargaining power because of high industry unionization and the ease with which they could pressure airlines that operate out of hubs. Consequently, airline unions were more effective in preserving high wages, and the union wage advantage expanded in the years following deregulation. Thus, deregulation affected the union premium differently in the two industries, and the pace of adjustment also differed.

Analogous to the situation for regulation, the effect on the union premium of increased international competition is theoretically ambiguous, may take years to complete, and may differ

[12] For example, Hendricks (1994) claimed that regulation eroded rents in the railroad industry by forcing railroads to maintain unprofitable lines and preventing mergers. Once the industry was deregulated, railroads

by industry. To the extent imports reduce demand for domestic output in an industry and, in turn, demand for domestic labor, wages might be expected to fall for both union and nonunion workers, assuming that the supply of nonunion labor to the industry is not perfectly elastic. Whether the union premium grows or contracts depends on the relative changes in union and nonunion wages. If increased import competition eats into union rents, the union premium may erode. On the other hand, prior analysis of cyclical changes in the union premium indicates that union wages are more rigid than nonunion wages. If union wages are less responsive than nonunion wages to demand shifts resulting from foreign competition, the union premium might be expected to widen as imports rise.

A widening wage premium can also be explained by the model of Lawrence and Lawrence (1985). In this model, imports permanently decrease the demand for domestic output (the end-game case) or at least slow output growth (the slow game). Assuming long life of capital, the net effect is to reduce the elasticity of labor demand by limiting the ability of firms to substitute capital for labor. When the elasticity of substitution between capital and labor is one, or when the effect of the lower elasticity of labor demand dominates the effect of reduced output, unions raise wages and expropriate rent from capital. Lawrence and Lawrence argue that their model explains why unions in such industries as steel and automobiles have responded to increased foreign competition by raising wages and inflating the union premium.

In the model of Staiger (1988), increased import competition reduces the share of union employment in labor-intensive firms and increases the share in capital-intensive firms. The greater capital intensity reduces the elasticity of derived demand for union labor, allowing the rent-maximizing union to raise the wage premium. Again, the result is a positive relationship

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abandoned unprofitable lines; labor productivity soared; and, unlike trucking, rates generally increased.

between import competition and the union premium.[13] In contrast, international competition has an ambiguous effect in the model of Grossman (1984), depending on such factors as the elasticity of substitution between capital and labor.

Empirical studies to date have not attempted to directly estimate the effect of expanded trade on the union premium, so the effect must be inferred by comparing the estimated responsiveness of union and nonunion wages. Belman (1988), Freeman and Katz (1991), and Bertrand (1998) found a negative correlation between import penetration and wages of both union and nonunion workers. In contrast, Macpherson and Stewart (1990) concluded that imports reduce wages only for union workers,[14] although they acknowledged that their results might not be valid beyond their sample period, 1975-81.[15] In none of these studies is there evidence of a statistically significant difference in the responsiveness of union and nonunion wages, nor is there a consistent pattern of parameter estimates across studies. Therefore, the effect of increased imports on the union premium remains uncertain, both theoretically and empirically.

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[13] Naylor (1998) reaches a similar conclusion based on a framework that incorporates union-firm bargaining into a trade model with imperfectly competitive markets. In this setting, the labor union responds to increased economic integration by setting higher Nash equilibrium wages.

[14] Actually, when they expanded the empirical model to allow for differential effects by level of unionization, they found that imports reduce union wages only in industries with low unionization. For industries with average or above-average unionization, their results indicate a positive effect of imports on union wages.

[15] The 1975-81 sample period predates the surge in U.S. imports and therefore cannot be expected to capture the strong international pressures that have developed since that time. Whereas the import penetration ratio for manufacturing increased from 7 percent in 1975 to 21 percent in 1999, it rose by only 2 percentage points between 1975 and 1981.

It is also important to recognize that the pattern of import penetration and union premiums across industries may differ from the longitudinal relationship between imports and the union premium within an industry. Industries with high imports are likely to differ from other industries in dimensions other than imports (union strength, managerial opposition to unions, rent sharing, profitability, etc.). For example, industries with strong competition from foreign producers may generally be less profitable and have weaker unions. In that case, imports and the union premium may be inversely related across industries at a given point in time. Even so, *increases* in imports may be associated with *increases* in the union premium—either because of greater rigidity of union wages or for reasons spelled out in the models of Lawrence and Lawrence, Staiger, and others. In that event, the relationship between imports and the union premium will be positive in time-series analysis regardless of any cross-sectional correlation. For this reason, the effect of rising imports can be estimated reliably only by tracking changes in imports over time.

Endogeneity Issues

Before estimating the relationship between imports and the union premium, we address the issue of causality. Although there are theoretical reasons to anticipate that changes in imports lead to changes in the union premium, it is also possible that imports respond to the union premium. Endogeneity issues also arise in studies (such as Linneman, Wachter, and Carter) that examine whether increases in the union premium reduce the unionization rate.[16] A related issue is whether intertemporal changes in union density lead to changes in the union premium in the industry. Although cross-sectional analysis often shows a positive correlation

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[16]See Hirsch and Schumacher (2001) for a discussion of factors contributing to the decline in private-sector unionization.

between union density and union premium, it is possible that some other factor leads to both high unionization in the industry and a high unionization premium and that changes over time in the two variables are not related. Indeed, if falling unionization results in a lower union premium, it is puzzling that the union premium has held up so well despite the dramatic erosion of union density.

To test for causality among the three variables—union premium, density, and the import penetration ratio—we alternatively regressed each variable on lagged values of itself and lagged values of the two other variables, controlling for a quadratic time-trend and industry fixed effects (see Table 1). Results indicate that imports Granger-cause both the union premium and density (the import penetration ratio is statistically significant in the premium and density regressions) and that the union premium Granger-causes density (the premium is significant in the density regression). There is no indication, however, that the premium or density causes imports or that density causes the premium. Accordingly, the union premium is estimated as a function of an exogenous import ratio. Causality tests also justify treating the union premium and imports as exogenous in analyses of union density.[17]

[17]To allow for differential effects across industries, we also performed the Granger causality analysis at a more disaggregated level (for durable-goods manufacturing, nondurable-goods manufacturing, and nonmanufacturing, as well as for each of the 32 individual industries). Not surprisingly, the statistical evidence weakens in such an experiment, and causal relationships significant at the aggregate level are not always significant at a disaggregated level. In the union premium equation the coefficient of imports is statistically significant in both durable and nondurable manufacturing industries (as in Table 1), but in the union density equation the coefficient of imports is significant only for durable manufacturing industries. Similarly, in the density equation the coefficient of union premium, though still significant in nonmanufacturing industries, loses significance in manufacturing. Importantly, however, results at the disaggregated level do not indicate that the premium or unionization Granger-causes imports, which again suggests that imports are exogenous in premium and union density equations.

Empirical Evidence

In this section, we exploit the panel of union-premium estimates for each industry and examine the empirical linkages between the estimated premium and cyclical and structural factors. Although other studies have also attempted to explain changes in the union premium by industry (Linneman, Wachter, and Carter 1990:50) and by race and occupation (Moore and Raisian 1980:126) based on a similar two-step methodology, such an approach has been less common than directly estimating wage equations for union and nonunion workers. The two-step approach has also been less successful, perhaps because of the short panels used in past studies. Conceptually, however, the two-step approach offers the important advantage that it avoids aggregation bias in standard errors arising from inclusion of industry-level covariates in micro-level regressions (Moulton 1990).[18]

To capture the factors influencing the union premium described in the previous section, we specify the regression model underlying the empirical analysis as follows:

$$\text{(3)} \qquad \hat{\pi}_{jt} = \gamma_0 + \gamma_1 UR_t + \gamma_2 UR_t * (i - i^e)_t + \gamma_3 UR_t * COLA_{jt}$$
$$+\gamma_4 COLA_{jt} + \gamma_5 (i - i^e)_t + \gamma_6 DEREG_{jt} + \gamma_7 IMPORT_{jt} + u_{jt},$$

where $\hat{\pi}_{jt}$ denotes the estimated union premium for industry j in year t (listed in the appendix), UR is the aggregate unemployment rate, $(i\text{-}i^e)$ the difference between actual and expected inflation, $COLA$ the fraction of union workers in the industry covered by COLA clauses at the

[18]Neumark and Wachter (1995) use a similar two-step methodology in their study of union effects on nonunion wages.

end of the year, *DEREG* an indicator variable for whether or not the industry has been deregulated, *IMPORT* the import penetration ratio of the industry (calculated as the ratio of imports to industry shipments), and *u* the regression error.[19]

To account for the multi-year duration of most union contracts, values of the independent variables except for *DEREG* are averaged over the most recent three years.[20] For deregulation we estimate the long-run effect with a dichotomous variable (signifying the post-regulation period); and, to allow for a differential short-run effect, we also add a second term. For each industry we alternatively include a variable that indicates that deregulation occurred in

[19]We collected unemployment rates from the U.S. Bureau of Labor Statistics (http://www.bls.gov), data on COLA coverage by industry from the *Monthly Labor Review* (various issues), and the inflation data from the *Livingston Survey* (http://www.phil.frb.org/econ/liv/). The dummy variable for deregulation was set to unity starting with 1978 for the airline industry, 1979 for trucking (Hirsch 1988, 1993), 1980 for railroads and for finance, insurance, and real estate, and 1984 for communications. To construct the import penetration variable, we tabulated imports and industry shipments through 1994 from Feenstra (1996) and thereafter from the U.S. Bureau of the Census, *U.S. Merchandise Trade, series FT900 (December)* and *Manufactures' Shipments, Inventories, and Orders* (http://www.census.gov). Industry COLA rates are not available until 1973, but aggregate COLA rates, which are available, are similar in value for the years 1971-73, so we used the 1973 industry value as a proxy for the 1971-73 average COLA rate in the industry. COLA rates were last published for the 1995 year, at both the industry and aggregate level. For 1996-99, we imputed values of COLA based on a regression using 1973-95 data to estimate COLA coverage as a function of inflation uncertainty, union density, a dummy variable for deregulation, and an industry fixed effect. Inflation uncertainty is proxied by the standard deviation of inflation forecasts contained in the *Livingston Survey*. See Ragan and Bratsberg (2000) for an analysis of determinants of COLA coverage within an industry. Dropping the years with imputed data from the sample does not affect any of the conclusions drawn in this section.

[20]Over the sample period the mean duration of union contracts was approximately 33.5 months (based on data from the U.S. Bureau of Labor Statistics, *Current Wage Developments*, various issues).

the current year or prior year (*DEREG2*), occurred in the current year or prior three years (*DEREG4*), occurred within the past five years (*DEREG6*), or occurred within the past seven years (*DEREG8*). For imports, differential effects and differential lags are allowed for durable-goods manufacturing and nondurable-goods manufacturing. In the baseline regression, we impose a two-year lag in both sectors, so that the three-year average value is calculated from imports two, three, and four years ago. Finally, because the dependent variable of equation 3 is itself estimated, we apply weighted least squares to account for heteroscedasticity of the regression model.[21]

Diagnostic checks of treatment of the regression error reject the OLS specification in favor of a model that includes industry fixed effects.[22] Furthermore, tests of the time-series properties of the error term suggest strong industry-specific autoregressive processes in the

[21]We use the industry observation counts from the first-step regressions as weights, but obtain similar results when we weigh the regression by the inverse of the variance of the premium estimate from the first-step regression. We present results based on observation count-weighted regressions because parameter estimates correspond more closely to population averages.

[22]A Hausman specification test yields a chi-squared (7 degrees of freedom) test statistic of 171.12 and a Wald test of joint significance of industry effects, an F (31 and 793 degrees of freedom) test statistic of 99.09. Critical values at the one-percent level are 18.47 and 1.71, respectively. Consistent with the recommendation of Lewis (1986) and the approach of Linneman, Wachter, and Carter (1990), we do not formally account for selectivity of workers into the union sector when estimating the union premium (see Lee (1978) and Heckman (1990) for discussion). But to the extent that industry-level biases are constant over time, they are captured by the industry fixed effects of the second-step regression.

regression error.[23] Our empirical specification therefore allows for the following structure of the error term of equation 3:

(4) $u_{jt} = v_j + \theta_{jt}$, and

(5) $\theta_{jt} = \rho_j \theta_{jt-1} + \omega_{jt}$,

where *v* denotes the industry fixed effect, ρ_j is the autocorrelation coefficient of regression errors of industry *j* (after netting out the fixed effect), and ω is white noise.

Cyclical Factors

Table 2 presents feasible least squares estimates of the parameters of equation 3. As robustness checks, the table lists results from five versions of the empirical model—four that are estimated controlling for industry fixed effects and, for comparison, one estimated without industry fixed effects. Regardless of specification, the table shows a positive and statistically significant coefficient of the lagged unemployment rate—indicating strong counter-cyclical movements in the union wage premium. For example, according to parameter estimates in column 2, a one percentage point increase in the unemployment rate raises the union premium by 1.9 percentage points when actual inflation meets expectations and union workers are not covered by COLA clauses. This result is consistent with the notion that union wages are more rigid over the business cycle than nonunion wages.

[23]A likelihood ratio test of no autocorrelation versus industry-specific AR(1) processes yields a chi-squared (32 degrees of freedom) test statistic of 499.89. The critical value at the one-percent level is 53.49. Similarly, based on a likelihood ratio test we reject the intermediate specification of a common AR(1) process for all 32 industries.

Our parameter estimate of the sensitivity of the union wage premium to the unemployment rate is strikingly similar to estimates from prior studies. Indeed, evaluated at the sample averages of the interaction terms,[24] the estimated impact of unemployment on the union premium from the longitudinal specifications in columns 1-4 falls in the middle of the .010 to .020 range reported by Lewis (1986:154). Moreover, the estimates are very similar to the negative of the partial correlation coefficient between the unemployment rate and wages reported in recent panel-data based studies of movements of real wages over the business cycle (Bils 1985; Solon, Barsky, and Parker 1994).[25] The implication is that the pro-cyclical behavior of real wages uncovered in the literature derives largely from the nonunion sector and not from the union sector.

As predicted, cyclicality of the union premium increases when inflation is greater than anticipated (the coefficient of *Unempl Rate*$*(i-i^e)$ is positive). Also as expected, unions reduce the cyclical sensitivity of the wage premium through COLA clauses. When the regression specification includes industry fixed effects, empirical results show that COLA coverage for all union workers in the industry would eliminate much of the cyclical sensitivity of the union premium (reducing the effect of unemployment swings by between 44 and 60 percent depending on specification). Our longitudinal evidence therefore supports the conjectures of Moore and

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[24]Over the sample period, actual inflation exceeded expectations by .40 percentage point, and the fraction of union workers with COLA clauses averaged .38.

[25]See Abraham and Haltiwanger (1995) for a review of this literature. Abraham and Haltiwanger conclude that estimates of business cycle effects on wages are sensitive to composition bias because "less-skilled workers account for a smaller share of employment at business cycle troughs than at business cycle peaks" (p. 1243). If composition bias from changes in unobservable characteristics of workers is more relevant for the nonunion sector, the findings uncovered in Table 2 may be understated.

Raisian (1980) and Hendricks (1981) that reduced cyclical sensitivity of the union premium observed in the 1970s was related to the growth of COLA coverage among union workers during that period.

Another implication of these results is that cyclical sensitivity increased in the late 1980s and the 1990s when COLA coverage declined. This is exactly what Grant (2001) found when he compared the cyclical sensitivity of the union premium in the periods 1975-81 (when 59.0 percent of union workers were covered by COLA clauses) and 1983-93 (when the figure was 41.8 percent), though he did not address the role played by COLA clauses.[26]

The union premium in an industry is also positively related to the level of COLA coverage in the industry. Our interpretation of this relationship is that the independent variable proxies for union bargaining strength, and that the positive coefficient of *COLA* reflects the higher wage premium of trade unions in periods with favorable bargaining conditions. Finally, the premium is smaller when inflation is greater than expected.

Structural Factors

The effect of deregulation is initially estimated on the assumption that long-run and short-run effects are the same (the full effect of deregulation is felt immediately). When the effect of deregulation is constrained to be the same in all industries (column 1), no effect is uncovered—the estimated coefficient is insignificant and nearly zero in value. But when we allow for different effects in each of the five industries that were deregulated, deregulation is estimated to significantly raise the union premium in telecommunications, consistent with

[26]Grant found that union wages were less responsive to the current unemployment rate in the latter time period—which we attribute to reduced COLA coverage. An alternative explanation is that there was a structural shift in cyclicality in 1983 or a trend component to cyclicality. We tested for these possibilities but found no evidence of either.

Hendricks (1998), and reduce the premium in the financial sector. This uneven pattern of results supports the insights of Hendricks (1994) and Hirsch and Macpherson (1998, 2000a) that the effect of deregulation differs by industry. The coefficients of the remaining deregulation variables are insignificant, but the pattern is consistent with the estimates of others—positive for railroads (Hendricks 1994) and airlines (Hirsch and Macpherson 2000a) and negative for trucking (Rose 1987; Hirsch 1988, 1993; Hirsch and Macpherson 1998).[27]

Next, a variable is added to the regression in an attempt to measure separately the short-run effect of deregulation. When the short-run effect is constrained to be of the same duration in each industry, the best fit, as judged by the log-likelihood value, is for the variable *DEREG4*, which assumes that the long-run effect begins in the fourth year following deregulation (see column 3). But with the exception of the finance industry, standard errors of the estimated coefficients are high, leaving us uncomfortable saying anything about differential lags by industry. In the remainder of the analysis, we present both long-run and short-run estimates.

The evidence in Table 2 also shows that the union wage premium relates importantly to the level of import penetration of an industry. However, the sign of the relationship depends crucially on whether or not the regression specification includes industry fixed effects.

[27]Because we include in the trucking industry both the for-hire sector, which was regulated by the ICC, and the private-carrier sector, which was largely exempt from ICC regulation and which has been found to be unaffected by deregulation (Hirsch and Macpherson 1998), our estimate understates the effect of deregulation in the for-hire sector. This fact may also account for the insignificance of our coefficient.

According to the longitudinal evidence in columns 1-4, imports *raise* the union premium.[28] For example, results in column 1 suggest that a 10 percentage point increase in the import penetration ratio leads to a highly significant 2.0 percentage point increase in the union premium. Over the sample period, higher imports are estimated to have increased the union premium in manufacturing by 2.7 percentage points.

In contrast, results in column 5 (which are based on a specification that omits industry fixed effects) suggest that in a cross-section of industries there is a significant *negative* relationship between the union premium and the level of import penetration. The implication is that, across industries, import penetration is correlated with unmeasured characteristics that depress the union premium and that failure to include industry effects results in a severe negative bias in the coefficient of the import penetration ratio.[29] To obtain unbiased estimates, it is necessary to control for industry characteristics and to examine movements in import penetration over time. When this is done, imports are seen to increase the union premium, consistent with the models of Lawrence and Lawrence (1985), Staiger (1988), and others. The results of column 4 further indicate that the effect of imports is greater in durable-goods manufacturing than in nondurable-goods manufacturing.

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[28]The import penetration variable is lagged two years, which produces a better fit than current, one-year lagged, or three-year lagged values, although coefficient estimates are comparable. We find no indication that the lag structure of import penetration differs by industry.

[29]The only other substantive difference between cross-sectional and panel estimates is for deregulation. Cross-sectional results point to an *increase* in the union premium in trucking following deregulation. This result, at odds with prior studies, can be explained with two observations: (1) When industry effects are dropped, "trucking deregulation" largely captures the industry effect and (2) the union premium in trucking is high. This anomalous

When we add a quadratic time trend to the regression specification, the coefficient of $TIME^2$ is negative and statistically significant (column 4). Based on parameter estimates of the two time variables, the union premium has trended lower since the late 1970s. Allowing for a time trend magnifies the estimated impact of imports in durable-goods manufacturing but has little effect on other coefficient estimates.

Recently, a growing literature has examined the impact of international competition on the U.S. wage structure (see Burtless (1995) and Freeman (1995) for summaries of arguments and findings). Of particular relevance, Borjas and Ramey (1994, 1995) concluded that the surge in net imports in durable-goods manufacturing industries contributed importantly to the rising U.S. wage inequality during the 1980s. In Figure 3 we plot the relationship between the import penetration ratio and the union premium in the critical durable-goods sector. With the exception of the expansionary periods of the late 1970s and the late 1980s (for which the preceding results predict a decline of the union premium), the plot shows a marked positive relationship between the two measures. In fact, from the early period of our sample (1971-75) until the latest five-year period (1995-99), imports of durable goods grew from 8.0 to 25.9 percent of domestic production and the union premium in the durable-goods sector increased from 8.5 to 14.8 percent. Our results therefore indicate that, in addition to raising skill differentials in durable-goods manufacturing as documented by Borjas and Ramey, imports have also promoted wage inequality by significantly increasing the returns to union membership.[30]

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finding underscores the need to control for industry fixed effects when studying intertemporal changes in the union premium.

[30]In supplementary regressions we examined a companion issue: Do unions that raise the union premium pay a price in terms of a greater loss in union membership, as Linneman, Wachter, and Carter (1990) find, and does increased foreign competition accelerate declines in density (apart from the indirect effect on wages)? Union

Summary and Conclusions

The present study estimates and analyzes the union wage premium for 32 private-sector industries over the period 1971-99. In the first stage of the study, CPS data are used to estimate the wage advantage of union workers relative to nonunion workers conditional on measured characteristics. The second stage consists of regression analysis that accounts for industry fixed effects and autoregressive processes and explains movements in the union premium over time as a function of cyclical and structural variables.

At the aggregate level the estimated premium ranged from 13 to 22 percent, trending modestly lower during the sample period. At the industry level, union premiums have often moved in opposite directions. In such industries as durable-goods manufacturing and

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density in industry *j* in year *t* was estimated, using a grouped probit specification, as a function of the union premium and import penetration in the industry the prior year, a time trend, and industry fixed effects. In results available on request, we find evidence that a higher premium reduces density only in the nonmanufacturing sector. Unionization is also negatively and significantly correlated with import penetration, especially in durable-goods manufacturing.

Finally, we approximate the contributions of the union premium and foreign competition to the decline in union density, so that we can compare our results with those of Linneman, Wachter, and Carter, who simulate the change in density over their sample period (1973-86) for each of their industries and then use 1986 industry shares to project the change in density over their sample. Based on premium and density variables comparable to ours, they estimate that changes in union premiums at the industry level account for 13.9 percent of the decline in density over their sample period. When we replicate their simulation based on our sample period, using our more narrow industrial classifications and 1999 industry shares, we find no evidence that changes in union premiums have contributed to the overall decline in unionization. To the contrary, given the net reduction in the union premium over the sample period, changes in union premiums have had a small positive effect on union density. On the other hand, increased import penetration is estimated to account for 6-7 percent of the decline in union density over the sample period.

communications and utilities, the union premium has trended upward, while unions in other industries (e.g., construction and wholesale and retail trade) have seen their premiums erode. The general pattern has been for union premiums that were initially low to rise and for high premiums to fall. As a consequence, differences in union premiums across industries have diminished significantly over the sample period.

The union wage premium narrows as labor markets strengthen, but the extent of this narrowing depends on deviations of actual and expected inflation and on the fraction of union workers who are covered by cost-of-living adjustments. Cyclicality of the union premium increases when inflation is greater than anticipated. On the other hand, COLA clauses, by linking union wages to market conditions, significantly lessen the cyclical sensitivity of the union premium. In the absence of COLA provisions, and assuming actual inflation equals expected inflation, each one percentage point reduction in the unemployment rate is associated with a 1.9 percentage point decline in the union premium. COLA clauses eliminate 44 to 60 percent of these cyclical movements of the premium.

Apart from their influence on cyclicality, COLA clauses are also directly related to the union premium. We attribute this relationship to the correlation, previously uncovered, between union bargaining power and the extent of COLA coverage. The union premium is also smaller when inflation is greater than anticipated.

The effect of deregulation varies by industry, consistent with the predictions of other authors. For the telecommunications industry, deregulation is associated with a significant increase in the union premium. For other industries, the effect is negative or insignificant.

Increased import penetration has been strongly associated with a rising premium, a finding consistent with various theories but not previously documented in empirical analyses.

Even though the cross-sectional correlation between import penetration and the union premium is negative (as one would expect if industries with high imports have weaker unions or lower profits), changes in import penetration over time are strongly and positively linked to changes in the union premium. Over the sample period, increased foreign competition is estimated to have lifted the union premium in manufacturing by nearly 3 percentage points.

In summary, the present study provides estimates of the union premium by detailed industry, compares aggregate estimates with those of other studies, and uncovers significant determinants of changes in union premiums over time. Perhaps the most important finding of the study is that increases in import penetration have been associated with higher union premiums.

Appendix

Average Union Wage Premium and Trend in Union Premium by Industry

Industry	Premium (Std.Dev.)	Trend (Std.Err.)	Industry	Premium (Std.Dev.)	Trend (Std.Err.)
Mining	.069	-.0020*	Fabricated Metal	.121	-.0004
	(.050)	(.0011)		(.032)	(.0008)
Construction	.316	-.0053***	Industrial Metal	.090	-.0004
	(.053)	(.0007)		(.030)	(.0007)
Food Products	.156	-.0028***	Electronic Equipment	.070	.0023***
	(.035)	(.0006)		(.038)	(.0008)
Tobacco	.168	.0048	Transportation Equipment	.150	.0048***
	(.151)	(.0035)		(.053)	(.0008)
Textiles	.021	.0006	Instruments	.019	.0011
	(.068)	(.0016)		(.080)	(.0019)
Apparel	.064	-.0022**	Misc. Manufacturing	.117	.0041***
	(.038)	(.0008)		(.065)	(.0013)
Lumber	.148	-.0008	Railroads	.130	.0037**
	(.061)	(.0014)		(.079)	(.0017)
Furniture	.076	.0016	Trucking	.275	-.0028**
	(.058)	(.0013)		(.057)	(.0012)
Paper	.121	.0033*	Air Transportation	.257	.0010
	(.078)	(.0017)		(.087)	(.0021)
Printing	.294	-.0064***	Other Transport	.274	-.0013
	(.079)	(.0014)		(.076)	(.0018)
Chemicals	.027	.0023***	Communications	.092	.0048***
	(.037)	(.0007)		(.062)	(.0011)
Petroleum	.040	.0042**	Utilities	.118	.0001
	(.087)	(.0019)		(.037)	(.0009)
Rubber	.139	.0010	Wholesale Trade	.134	-.0013
	(.044)	(.0010)		(.044)	(.0010)
Leather	.102	.0055**	Retail Trade	.236	-.0065***
	(.101)	(.0021)		(.072)	(.0011)
Stone, Clay, and Glass	.138	-.0017	Finance	.059	-.0046***
	(.058)	(.0013)		(.066)	(.0013)
Primary Metal	.091	-.0001	Services	.133	-.0024***
	(.044)	(.0010)		(.033)	(.0006)

Note: The column labeled "premium" lists the average union premium estimate over the sample period (1971-99), and the column labeled "trend" presents the coefficient from a regression of the union premium on a trend variable.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

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Table 1. Causality Tests

	Dependent Variable:					
	Union Premium		Union Density		Import Penetration	
	(1)	(2)	(3)	(4)	(5)	(6)
Premium(t-1)	.6016***	.5655***	-.0230**	-.0410***	.0311	.0349
	(.0282)	(.0364)	(.0116)	(.0150)	(.0220)	(.0226)
Premium(t-2)		.0524		.0269*		.0019
		(.0348)		(.0143)		(.0223)
Density(t-1)	-.0670	-.0019	.7883***	.7450***	.0092	.0035
	(.0446)	(.0858)	(.0184)	(.0353)	(.0313)	(.0339)
Density(t-2)		-.0644		.0492		.0221
		(.0794)		(.0327)		(.0325)
Import(t-1)	.0864***	-.1436	-.0307**	-.0018	1.0434***	1.2616***
	(.0331)	(.1205)	(.0136)	(.0496)	(.0129)	(.0555)
Import(t-2)		.2378*		-.0297		-.2305***
		(.1212)		(.0499)		(.0570)
R^2	.8931	.8941	.9892	.9892	.9837	.9842
P-values from F-tests:						
$\beta_{\text{Prem(t-1)}} = \beta_{\text{Prem(t-2)}} = 0$				.0236		.2703
$\beta_{\text{Density(t-1)}} = \beta_{\text{Density(t-2)}} = 0$		.2811				.7317
$\beta_{\text{Import(t-1)}} = \beta_{\text{Import(t-2)}} = 0$		.0062		.0732		

Notes: Sample size is 832 in columns 1-4 and 520 in columns 5 and 6, which are restricted to manufacturing industries. Standard errors are reported in parentheses. Regressions also include a quadratic time trend and industry fixed effects. In columns 1-4, observations are weighted by the industry observation count of the first-step regression.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

Table 2. Determinants of Union Wage Premium

	(1)	(2)	(3)	(4)	(5)
Unemployment Rate	.0187***	.0187***	.0192***	.0166***	.0117***
	(.0017)	(.0017)	(.0017)	(.0017)	(.0041)
Unempl Rate*$(i-i^e)$	.0024***	.0026***	.0023***	.0033***	.0025
	(.0009)	(.0009)	(.0008)	(.0008)	(.0025)
Unempl Rate*COLA	-.0094**	-.0092**	-.0085**	-.0099***	-.0128*
	(.0037)	(.0038)	(.0038)	(.0036)	(.0074)
COLA	.0749**	.0763**	.0730**	.0526*	.0627
	(.0312)	(.0313)	(.0313)	(.0304)	(.0548)
$(i-i^e)$	-.0165**	-.0182***	-.0154**	-.0228***	-.0188
	(.0066)	(.0065)	(.0064)	(.0064)	(.0185)
Deregulation	-.0214				
	(.0165)				
Dereg Rail		.0329	.0416	.0636	-.0137
		(.0905)	(.0917)	(.0862)	(.0406)
Dereg4 Rail			-.0247	-.0303	.0285
			(.0563)	(.0552)	(.1186)
Dereg Trucking		-.0716	-.0652	-.0499	.1158***
		(.0560)	(.0599)	(.0541)	(.0190)
Dereg4 Trucking			-.0455	-.0443	-.0251
			(.0506)	(.0508)	(.1357)
Dereg Air		.0554	.0643	.0645	.1100***
		(.1262)	(.1205)	(.1117)	(.0286)
Dereg4 Air			-.1007	-.0949	-.0735
			(.0804)	(.0807)	(.1997)
Dereg Communications		.0752**	.0798***	.0880***	-.0225
		(.0316)	(.0305)	(.0341)	(.0232)
Dereg4 Communications			-.0086	-.0083	.0029
			(.0201)	(.0218)	(.0433)
Dereg Finance		-.0614***	-.0573***	-.0334*	-.1115***
		(.0191)	(.0175)	(.0184)	(.0104)
Dereg4 Finance			-.0439***	-.0587***	-.0474
			(.0113)	(.0111)	(.0359)
Import Penetration	.1955***				
	(.0337)				
Import Penetration		.2048***	.2183***	.3097***	-.1231***
Durables		(.0427)	(.0433)	(.0459)	(.0432)
Import Penetration		.1655***	.1701***	.1839***	-.1931***
Nondurables		(.0513)	(.0504)	(.0477)	(.0439)
Time				.0007	-.0001
				(.0015)	(.0024)
$Time^2/100$				-.0100**	-.0078
				(.0041)	(.0069)
Industry Fixed Effects	Yes	Yes	Yes	Yes	No

Notes: Sample size is 832. Standard errors are reported in parentheses. Dereg4 equals unity if deregulation of industry took place within last four years; $(i-i^e)$ denotes the difference between actual and expected inflation. Regressions are estimated with industry-specific AR(1) process in error term (except column 5). Each observation is weighted by the industry observation count of the first-step regression.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

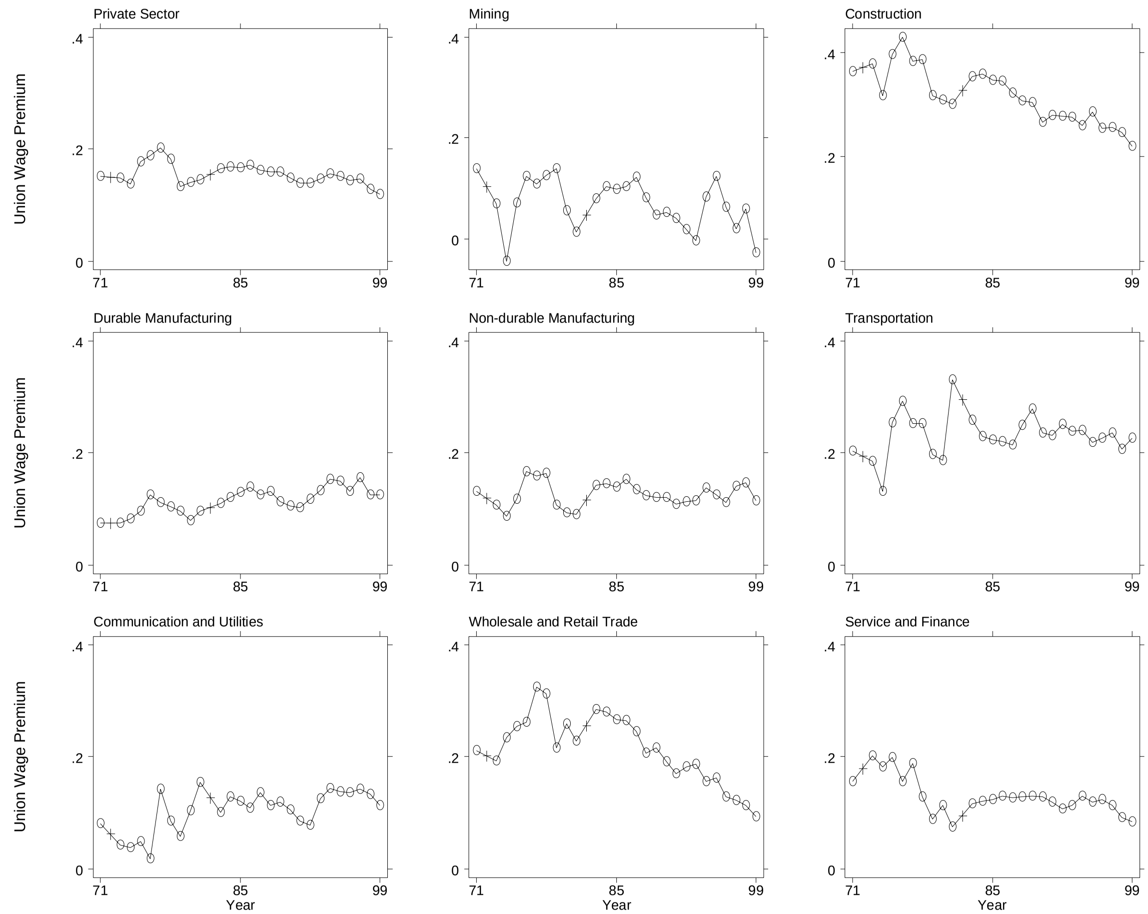

Figure 1. Trends in the Union Wage Premium

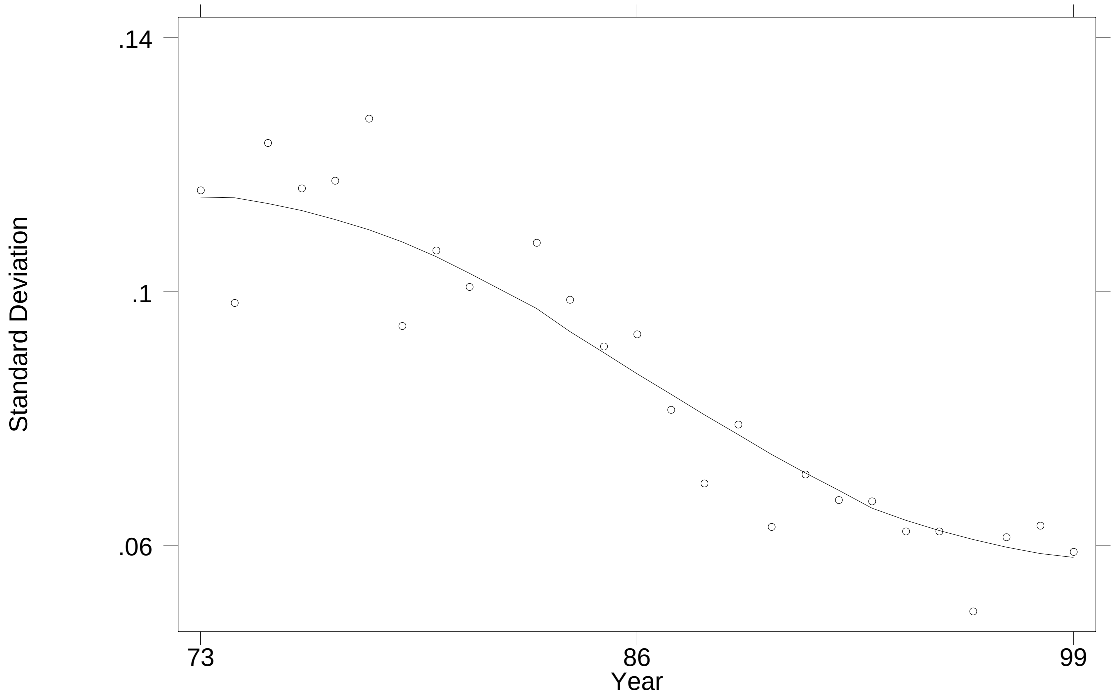

Figure 2. Trend in the Standard Deviation of the Union Premium Across Industries

Notes: The standard deviation is calculated from union premiums of 32 two-digit industries. The solid curve depicts the lowess smoother of the scatter points using a bandwidth of 1.0.

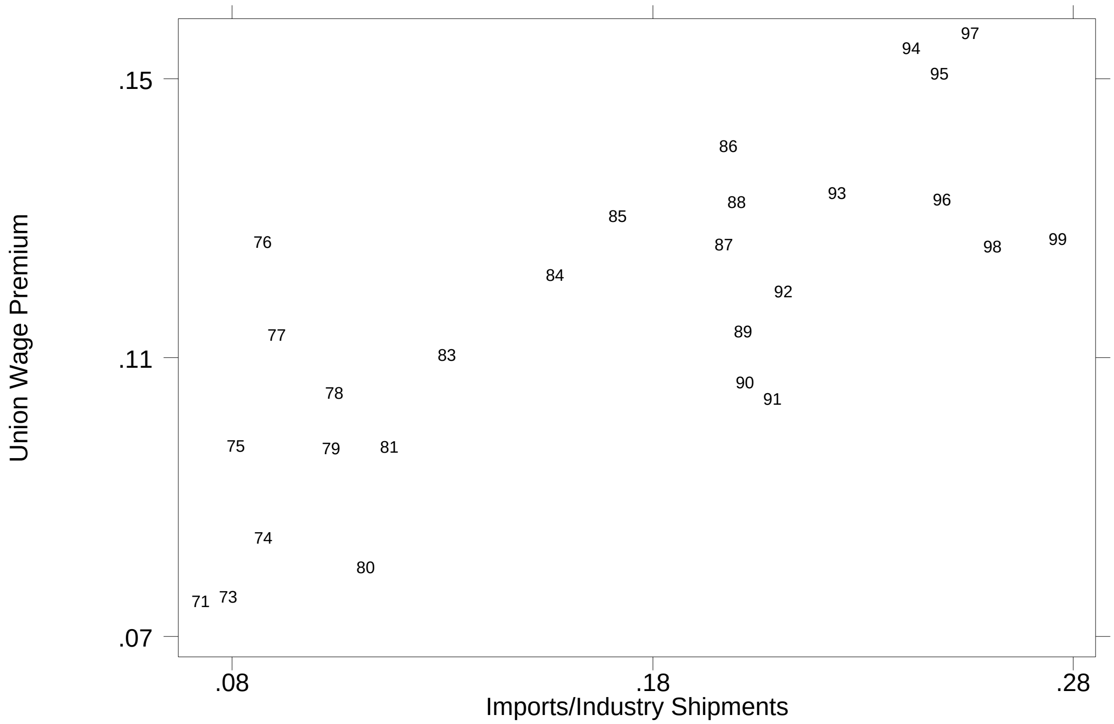

Figure 3. Union Wage Premium and Import Penetration, Durable-Goods Manufacturing