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New Forms of Representation

Edited by Phanindra V. Wunnava

M.E. Sharpe
Armonk, New York
London, England
Union-Nonunion Wage Differentials and Macroeconomic Activity

Bradley T. Ewing and Phanindra V. Wunnava

This research identifies the differing responses of union and nonunion wages to shocks to real output growth, inflation, and the stance of monetary policy. The literature documents the existence of a union wage premium; however, previously the focus has primarily been at the microlevel and on whether or not a union worker receives greater compensation than an otherwise comparable nonunion worker (e.g., Freeman and Medoff 1984; Hirsch and Addison 1986; Lewis 1986; Wunnava and Ewing 1999, 2000). Research also links the wage differential to the stage of the business cycle (Wunnava and Hennesy 1991; Wunnava and Okunade 1996) and to industrial structure (Okunade, Wunnava, and Robinson 1992).

Theoretical macroeconomic models imply that the response of employment to changes in aggregate measures of economic activity depends on the degree to which wage and price rigidities exist (e.g., see Romer 1996). For example, in explaining the labor market dynamics of Keynesian-type models when wages are rigid relative to output prices, Sargent (1987) shows that employment rises with an increase in the price level. Sargent goes on to say that "sticky" money wages might occur in the long-term labor contracts, such as those often found in the union sector. Certainly, it is possible that the degree to which this stickiness exists differs by union and nonunion status, as well as by economic sector. Given the differences in compensation level of union and nonunion workers, and the link to the stage of the business cycle and industry, it is expected that the response of union-nonunion wage differentials to macroeconomic shocks may vary by industrial sector.

The relationship between the union-nonunion differential and macroeconomic factors is examined by computing generalized impulse response functions derived from the estimation of vector autoregression models. These response functions allow us to compare and contrast the effects of unanticipated changes in the macroeconomic factors on the wage differential within an industrial sector as well as between industrial sectors. An innovation to any of the variables may be interpreted as (unexpected) economic news. Clearly, firms and workers, and thus the wage gap, may be affected by movements in any of these macroeconomic variables. Knowledge of what leads to movements in the union-nonunion wage gap and how long shocks may last might be of concern to workers, firm owners and managers, as well as policy makers.

Macroeconomic Factors and the Union-Nonunion Wage Differential

An event (i.e., economic news) that affects either the union labor market or the nonunion labor market should influence the union-nonunion wage gap. Interestingly, Heywood and Deich (1987) studied the effect of unions on economic activity. According to their investigation, there is no evidence that unionism deters economic activity. At the aggregate level, the stage of the business cycle—whether the economy is in a growth period or recession—affects demand for labor. Therefore, news about downturns (downturns) in the economy should correspond to a general rise (fall) in labor demand. Haig (2001) and Neumark and Wachter (1995) discuss the behavior of union wages vis-à-vis nonunion wages during different stages of the business cycle. When the economy is near full employment, an increase in union wages may place upward pressure on nonunion wages due to the threat effect. During a recession, firms have an incentive to lay off high-paid union workers in an effort to lower costs. However, the accompanying increase in the pool of labor may lower nonunion wages. Consequently, changes in real output can be expected to affect the union-nonunion wage gap. Moreover, a number of papers have suggested that a relationship between aggregate economic activity and the union-nonunion wage gap. For example, Medoff (1979) and Wachter (1986) suggest a countercyclical wage gap based on theories of wage rigidities in long-term (multiyear) union contracts (Wunnava and Okunade 1996), as well as the prevalence of union seniority rules for assigning layoffs in recessions. Unless the recession threatens the very existence of the union itself, as per the "wage norm" hypothesis of Perry (1986) and Mitchell (1986), unions may
not concede to wage reductions. In contrast, Roess (1989) suggests that recessions induced by price shocks may widen the wage gap. Moreover, it is possible that the wage gap is unresponsive to the business cycle, either because firms are able to adjust their employment of labor and productive factors at relatively low cost or because nominal price and wage rigidities do not exist, as might be suggested by real business cycle models. Since union representation and strength varies by industry, the effects of change in real output on aggregate measures of the wage gap should be examined by the industrial sector (Okunade, Wannarai, and Robinson 1992).

The expected rate of inflation affects the real wage and would, therefore, affect employment decisions. In the presence of nominal wage rigidity, an inflation shock lowers the real wage. If constraints such as contracts and so on make union compensation less flexible than wages, the fall in the real wage of nonunion workers will exceed that of the union workers and the wage gap will widen. Note that the widening may occur with a lag in the presence of contracts and employment wage agreements or when wages are set at the beginning of the period, as in Sargent's (1987) depiction of the Keynesian model. Over time, as contracts are renegotiated and new wage agreements are made, the money wage is expected to adjust upward and the equilibrium real wage is restored. Furthermore, there is another avenue in which inflation shocks may affect the real wage. Unanticipated inflation, by creating volatility and uncertainty in price changes, may restrict production activity and, thus, firm hiring. Union firms may have less ability to optimally adjust employment levels due to seniority or layoff rules and contract provisions. Consequently, if low wage (short tenure) union workers and nonunion workers are the first to be let go, then the wage gap should widen. Moreover, if it takes some time for the price uncertainty to be resolved, perhaps as economic information is revealed and processed by agents, then the response of the wage gap to the inflation shock may persist for a number of periods. Hendricks and Kahn (1983), in their seminal work on cost of living agreements, also predict that the union-nonunion wage differential widens during the periods of unanticipated inflation.

Thoebbecke (1997) and Ewing (2001) argue that money may have real effects and that monetary policy may represent a significant source of business cycles. Tighter monetary policy tends to reduce aggregate demand through an interest rate effect and, in the presence of rigidities, output falls and employment is affected. In general, a rise in the federal funds rate places upward pressure on rates. In the short run, the Federal Reserve's actions may have a more pronounced effect on nonunion wages than on union wages. If nonunion wages fall relatively more than union wages, the differential becomes wider. This might be the case if union firms face restrictions on their ability to optimally adjust employment levels. Thus, it is expected that the wage gap will rise with a sudden monetary tightening, and the response will be more pronounced in those sectors that are sensitive to interest rate movements.

As the previous discussion attests, the union-nonunion wage gap may be linked to macroeconomic factors. This chapter adds to the literature on unions and unemployment and macroeconomic activity by providing insight into the response of the wage gap to innovations in real output growth, monetary policy, and inflation.

A Simple Reduced-form Model of the Union-Nonunion Wage Differential

In this section, we briefly outline a reduced-form model of the union-nonunion wage differential derived from general specifications of supply and demand in the market for union and nonunion workers. We specify supply and demand in the nonunion worker market as follows:

\[ N^d = N^d(X^S, w^j) + \phi^d \]
\[ N^d = N^d(X^{ID}, w^j, w_j - w_p) + \phi^{MD} \]

where \( X^S \) and \( X^{ID} \) are vectors of exogenous variables that affect the supply and demand for nonunion workers, respectively, \( w^j \) is the union (log) real wage and \( w_j \) is the union (log) real wage. \( \phi^d \) and \( \phi^{MD} \) are shocks to supply and demand for nonunion workers that are assumed to have zero mean and are uncorrelated.

Similarly, the supply and demand for union workers may be represented as:

\[ L^s = L^s(X^S, w^j) + \phi^s \]
\[ L^s = L^s(X^{ID}, w^j, w_j - w_p) + \phi^{MD} \]

where \( X^S \) and \( X^{ID} \) are vectors of exogenous variables that affect the supply and demand for union workers. \( \phi^s \) and \( \phi^{MD} \) are shocks to supply and demand that are assumed to have zero mean and are uncorrelated.

The underlying structural equations can be solved to obtain the reduced-form equation for the union-nonunion wage differential:

\[ GAP = (w_j - w_p) = \beta(X^S, X^{ID}, X^{SS}, X^{ID}) + E \]

The size of the wage gap will respond to changes in the exogenous variables that affect supply and demand in the markets for union and nonunion workers. Based on the reasons given earlier, we treat real output growth, inflation, and the stance of monetary policy as these variables.
The Data

Shocks to real output, monetary policy, and inflation are examined over the period 1976Q3 through 2001Q1 to see how union-nonunion wage differentials respond to innovations in these macroeconomic variables. Following the work of Berman and Blinder (1992), Thorbecke (1997), and Ewing (2001), we use changes in the federal funds rate as a proxy for the stance of monetary policy. The consumer price index for all urban consumers is used to compute the inflation rate (Park and Ratti 2000). Real economic activity is gauged by the growth rate in real gross domestic product. We use the Employment Cost Index (ECI) series for wages and salaries of (private industry) union workers and nonunion workers to construct the union-nonunion wage gaps. Wage gaps are computed for total private industry, goods-producing industries, manufacturing industries, nonmanufacturing industries, and service-producing industries. Each ECI index is seasonally adjusted. The five wage gaps are defined as the log difference between wages and salaries of union workers and nonunion workers. Thus, the quarterly data consist of changes in the federal funds rate, growth in real gross domestic product, consumer price inflation, and five union-nonunion wage differentials. All data were extracted from the Econometric database. Table 6.1 provides more detailed information on data and variable definitions.

Table 6.2 presents descriptive statistics for the variables. Somewhat surprisingly, the largest mean wage gap is found in service-producing industries, while the smallest is in manufacturing. Table 6.3 (see p. 154) shows the associated estimated (contemporaneous) correlation matrices. Generally speaking, union-nonunion wage gaps are negatively correlated with changes in real output and positively correlated with inflation. Monetary policy changes are negatively correlated with the wage gaps in nonmanufacturing and service-producing industries as well as with (total) private industry. In contrast, monetary policy is positively correlated with the wage gaps in manufacturing and goods-producing industries.

Tests of Stationarity

The proper specification of a vector auto-regression (VAR) model depends on the univariate properties of the variables under investigation. In particular, it is important to ascertain the data-generating process of each series. The purpose of this section is to make a distinction between a trend-stationary process and a unit-root process. In the former case, the (perhaps detrended) level of a series would be appropriate to use in the VAR, while if the series has a unit root, it is necessary to first-difference the series to render a station-
Table 6.3
Estimated Correlation Matrices

Panel A: Prusa industry

<table>
<thead>
<tr>
<th></th>
<th>GAPPI</th>
<th>GROWTH</th>
<th>MPOLICY</th>
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Panel B: Goods-producing industry

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<tr>
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Panel C: Manufacturing industry

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<th>MPOLICY</th>
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Panel D: Nonmanufacturing industry

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Panel E: Service-producing industry

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<td>0.2565</td>
<td>1.000</td>
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Table 6.4
Tests of Stationarity

<table>
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<tr>
<th>Variable</th>
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<th>Phillips-Perron</th>
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<tr>
<td>GAPMIPF</td>
<td>-6.27a</td>
<td>-0.31p</td>
</tr>
<tr>
<td>GAPMIPMF</td>
<td>-5.32a</td>
<td>-0.62p</td>
</tr>
<tr>
<td>GAPPI</td>
<td>-5.89a</td>
<td>-0.09p</td>
</tr>
<tr>
<td>GAPPSPF</td>
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<td>-0.79p</td>
</tr>
<tr>
<td>GROWTH</td>
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<td>-7.41p</td>
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<tr>
<td>MPOLICY</td>
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<tr>
<td>INF</td>
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<td>-8.56p</td>
</tr>
</tbody>
</table>

Notes: Superscripts a, b denote significance at the 1 percent and 10 percent levels based on critical values in MacKinnon (1991). Δ denotes the first-difference operator. One lag was used in the augmenting term, as suggested by Akaike’s information criterion, and was sufficient to ensure the absence of autocorrelation in the residuals.

Dickey-Fuller (ADF) test is used to check for the presence of unit roots and is based on the ordinary least squares regression of equation 1.

\[
\Delta X_t^\alpha = \rho_1 + (\rho_2 - 1) X_{t-\alpha}^\alpha + \rho_f + \sum \alpha_t \Delta X_t^\alpha + \epsilon_t
\]

where \( X^\alpha \) is the individual series under investigation, \( \Delta \) is the first-difference operator, \( t \) is a linear time trend, \( \epsilon_t \) is a covariance stationary random error, and \( \alpha \) is determined by Akaike’s information criterion to ensure serially uncorrelated residuals. The null hypothesis is that the variable is non-stationary at the same order as the differencing operator and is statistically significant. The finite sample critical values for the ADF test developed by MacKinnon (1991) are used to determine statistical significance.

An alternative unit root test developed by Phillips and Perron (1988) allows for weak dependence and heterogeneity in the error term and is robust to a wide range of serial correlation and time-dependent heteroskedasticity. The test is based on the following regression:

\[
X_t^\alpha = \eta_0 + \eta_1 (t - T/2) + \lambda X_{t-\alpha}^\alpha + \nu_t
\]

where \( (t - T/2) \) is the time trend with \( T \) representing the sample size and \( \nu_t \) is the error term. The null hypothesis of a unit root, \( H_0: \lambda = 1 \), is tested against the alternative hypothesis that \( X_t^\alpha \) is stationary around a deterministic trend \( (H_1: \lambda < 1) \). As in the ADF test, MacKinnon critical values may be used to determine statistical significance for the Phillips-Perron test.
The results of the unit root tests are presented in Table 6.4. The first-difference of each wage gap measure is stationary and, thus, a shock to a change in the union-nonunion wage gap will revert to the mean. Consistent with previous research, such as Ewing (2001) and Park and Ratti (2000), GROWTH, MPOLICY, and INF are all found to be stationary series.\textsuperscript{3}

vector Autoregression and Generalized Impulse Response Analysis

Dynamic analysis of VAR models can be conducted using innovation accounting methods, such as impulse response functions. However, this method has been criticized because results from impulse response analysis are subject to the "orthogonality assumption" and may differ markedly depending on the ordering of the variables in the VAR (Lütkepohl 1991). To overcome this problem, we employ the "generalized" impulse response function developed by Pesaran and Shin (1998) and Koop, Pesaran, and Potter (1996). This method is not sensitive to the ordering of the variables in the VAR. Ewing, Lemieux, and Malik (2002) provide additional explanations on the use of this method. Pesaran and Shin (1998) describe the generalized impulse response analysis in the following way.\textsuperscript{6} Consider the infinite moving average representation of the VAR:

\[
x_t = \sum_{j=1}^{\infty} \Phi_j \mu_{t-j}
\]

where \(x_t\) is an \(m\times 1\) vector of the variables under investigation, \(A_j = \Phi_j \Phi_{j-1} \cdots \Phi_2 \Phi_1\), and \(A_0 = \Phi_0 = 1\), and \(A_j = \Phi_j \Phi_{j-1} \cdots \Phi_2 \Phi_1\) is for \(j < 0\).

Let us denote the generalized impulse response function (G) for a shock to the entire system, \(u_t\), as:

\[
G_j = E(x_{t+|j|} | u_t = u, \text{C}_t^p, \text{C}_{t-1}^p) - E(x_{t+|j|} | \text{C}_t^p, \text{C}_{t-1}^p)
\]

where the history of the process up to period \(t+1\) is known and denoted by the information set \(\text{C}_t^p\). Assume \(u_t = N(0, \Sigma)\), and \(E(u_t = u) = \delta_t = \delta_t(\sigma_x^{-1})\delta_t\), where \(\delta_t = \delta_t(\sigma_x^{-1})\delta_t\) denotes a standard error shock. Furthermore, \(\delta_t\) is \(m\times 1\) with the 0th element equal to one and all other elements equal to zero. The generalized impulse response function for a one-standard deviation shock to the \(i\)th equation in the VAR model on the \(j\)th variable at horizon \(h\) is:

\[
G_{i,j,h} = \{\delta_t(\sigma_x^{-1})\delta_t\} / (\sigma_x^{-1})_{i,j}, i,j = 1,2, \ldots, m
\]
The responses of ΔGAPGP to macroeconomic shocks are shown in Figure 6.2. Similar to the case of private industry, a real output growth shock does not significantly affect the change in the goods-producing wage gap. However, the response of ΔGAPGP to a monetary policy shock is actually negative and significant one quarter after the shock then, as expected, becomes positive and significant for two quarters. This suggests that the Federal Reserve’s actions can affect the union-nonunion wage gap in the goods-producing sector. Moreover, the unexpected monetary tightening leads to an observed “cycling” of the wage gap. This type of response to monetary shocks of economic aggregates is found in many macroeconomic models that incorporate expectations that rely on a standard IS-LM framework with predetermined prices. A significant impulse response to MPOLICY sug-
Figure 6.3 Manufacturing Industries

Notes: Horizon is measured in quarters. Δ denotes first-difference operator.

Figure 6.3 presents the responses of change in the manufacturing sector wage gap to the macroeconomic shocks. The responses are quite similar to those found in the goods-producing sector with few exceptions. A shock to GROWTH has no effect, while a shock to MPOLICY has a positive effect following a two-quarter lag. The MPOLICY effect then lasts for about four to five quarters (i.e., up to about seven quarters following initial impact). As in the case of the goods-producing wage gap, we attribute the response of the manufacturing wage gap to unanticipated monetary tightening to these firms' interest rate sensitivity. Like the goods-producing sector, the response of ΔGAPMF to an inflation shock is positive and significant following a one-quarter lag. The response is a bit stronger than that found in the goods-producing sector but persists for about the same length of time. Similar to the other wage gaps, the own impulse response lasts for one quarter.

Figure 6.4 Nonmanufacturing Industries

Notes: Horizon is measured in quarters. Δ denotes first-difference operator.

Figure 6.4 shows how ΔGAPMF responds to macroeconomic shocks. In contrast to the other sectors, the nonmanufacturing wage gap falls with a real output shock. In particular, the response is negative and significant one quarter after the shock and remains significant for about one quarter. This finding is consistent with the countercyclical wage gap theories. No significant response is found for MPOLICY, suggesting that the labor market actions of firms in this sector are relatively insensitive to interest rate changes. The wage gap responds positively and significantly to an inflation shock, after a two-quarter lag. The inflation effect, while smallest in magnitude compared to the other sectors, persists for around four quarters.
quartets. As with the other wage gaps, the own impulse response lasts for one quarter.

The response to macroeconomic shock of changes in the service-producing union-nonunion wage gap is presented in Figure 6.5. Shocks to GROWTH and to MPOLICY are insignificant, while an inflation shock is significant and positive following a two-quarter lag. The inflation effect persists for only about three quarters and is the shortest in duration of all the sectors. The relatively faster dissipation of inflation shocks suggests that the firms in the service-producing sector exist in a competitive market environment, in which price changes are absorbed into wages more quickly. Consistent with the other wage gaps, the own impulse response for this sector lasts for just one quarter.

Conclusions

This chapter examined and documented the response of union-nonunion wage differentials to shocks in three key macroeconomic variables using the newly developed technique of generalized impulse response analysis. The technique is robust in terms of the choice of ordering variables in the VAR, thus one can accurately examine and compare both the severity and extent of shocks to these variables on the wage gap. The results add to the literature on the relationship between the macroeconomy and the union-nonunion wage gaps.

The results can be summarized as follows. For each sector as well as the total private industry, an inflation shock leads to a widening of the wage gap and occurs after a short lag and may last for several quarters. Generally speaking, a monetary policy shock is associated with a wider gap, which appears only after a couple of quarters, in each sector (and overall) except in nonmanufacturing and service-producing industries. Finally, growth shocks are found to be significant and negative only in the nonmanufacturing sector.

Notes

1. This simple analysis assumes diminishing marginal product of labor and no adjustment costs.
2. This chapter employs the recently developed econometric technique of generalized impulse response analysis (Koop, Pesaran, and Potter 1996; Pesaran and Shin 1998).
3. The raw data set begins before this date, but due to data transformations (e.g., growth rates), the usable or adjusted sample period begins in 1976Q3.
4. The results for INS are not as clear as those for the other variables. However, given the findings of Engle (1982), the Phillips-Perron test is probably more appropriate than the ADF for the case of inflation.
5. It was determined that the variables were not cointegrated. The results of Johansen-Sælund cointegration tests, which allowed for deterministic trends in the (levels of the variables), are available on request.
6. For a more detailed discussion, including proofs, see Pesaran and Shin (1998). Additional background material on the development of generalized impulse response analysis can be found in Koop, Pesaran, and Potter (1996).
7. The traditional orthogonalized impulse response employs a Cholesky decomposition of the positive definite matrix covariance matrix, $E$ of the shocks ($e_t$).
8. Pesaran and Shin state that "generalized impulse responses are unique and fully take account of the historical pattern of correlations observed amongst the different shocks" (1998, 20). Thus, they caution against using orthogonalized responses, since there is generally no clear guidance as to which of many possible parameterizations to employ. Note that generalized and orthogonalized impulse responses coincide only when the covariance matrix is diagonal.
9. LL is the system log-likelihood from the VAR and LL is computed as the sum of the log-likelihood values from the individual equations in the VAR.

10. Significance is determined by the use of confidence intervals representing plus-minus two standard deviations. See Runkle (1987) for a discussion on confidence intervals.

References


