Trade Unions and the Rate of Change of Money
Wages in the United States

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The purpose of this paper is to examine a formal model of aggregate wage determination in the union and nonunion sectors of the U.S. economy in order to obtain testable implications concerning the effects of trade unions on the rate of change of money wages. These hypotheses are then tested with annual data spanning the period 1914-1959.

Since the early work of Phillips [30], a number of investigators have followed the lead of Hicks [16], Slichter [36] and others and interpreted their empirical work on wage inflation as reflecting, at least in part, the impact of trade unions on wage determination. Lipsey [24] argued that the significant effect of price changes on U.K. wages might reflect union forces, while Eckstein and Wilson [12], Eckstein [13], Perry [29], and Schultze and Tryon [35] interpreted the significant effect of profit rates on U.S. wage changes as evidence of union activity. More recently Pierson [31] has emphasized the empirical differences in the wage setting mechanism in the union and nonunion sectors of U.S. manufacturing.

1/ Hicks writes, for example, that: "It has never been the general rule that wage-rates have been determined simply and solely by supply and demand.... We get a better clue to actual behavior if we think of wages as being determined by an interplay between social and economic factors instead of being based on economic factors--and crude economic factors at that--alone." [16], pp. 389-404.
while Throop [40] has used cross-sectional data to support the inference of a trade-union induced upward shift of U.S. money wages during the 1950's. At the same time, however, a number of other investigators have questioned the interpretation offered by these authors. Kuh [21] has argued that the level of profits may be interpreted as a proxy for average labor productivity and that it is the latter variable which determines short-run money wage changes, while McGuire and Repping [25] have argued that much of the above work is based upon loosely specified models and is ambiguous of interpretation. This problem of interpretation clearly results because the variables used by the above authors are not measures unambiguously related to trade union activity. Hines [17], [18], [19], has recently attempted to overcome this objection by explicitly considering the effect of union-related variables on money wages in the U.K. Although somewhat dissimilar both in specification and conclusions, this paper follows Hines' lead in deriving and testing hypotheses explicitly relating variables measuring trade union activity to money wage changes in U.S. manufacturing.

Finally, a number of investigators, including Bowen and Berry [6], Perry [29], Eckstein [13], and Rees and Hamilton [32], have expressed their dissatisfaction with previous attempts to estimate any satisfactory stable equation relating money wage changes to market forces using annual U.S. data over an appreciable part of the post-1900 period.2/

2/For example, although Bodkin [5] stresses the role of the unemployment rate in determining money wage changes in the U.S., his consistently estimated wage-price system does not have a regression coefficient on this variable which is significant at conventional levels. See Rees and Hamilton [32].
This paper may also be interpreted therefore as an explicit attempt to incorporate and test a theory of institutional change as it affects the determination of money wages in the U.S. economy.\(^3\) In particular, if the theory of institutional change is correct, then the role of market forces in wage determination should be clarified.

The plan of the paper is as follows: In Section I a simplified formal model of wage behavior in the union and non-union sectors and the spill-overs between these sectors is proposed. Section II contains a discussion of the simultaneity problem and presents the basic estimated equations. Section III contains further evidence on the effects of trade unions on money wage changes, including a discussion of Hines' model [17], and an exploration of the historical reduced form trade-offs between wages, prices and the exogenous variables. Finally, Section IV contains a discussion of the implications and limitations of the results.

I. A Theoretical Framework

Since trade unions have organized only two-thirds of U.S. manufacturing industry to date, it has been widely recognized that if unions have had a general effect on the level of aggregate money wages then there must be some mechanism by which wages in the union and non-union sectors are related.\(^4\) First, certain variables which cause union wages to rise will

\(^3\) This method is consistent with Nerlove's [28] argument that "...the specter of 'structural change' which haunts econometricians is a red herring; that which is so called reflects our lack of ability to formulate theories of past development consistent with available data..."

\(^4\) See, for example, Rees [33].
not typically have a direct influence on wage changes in the non-union sector; but the mere fact that union wage rates have risen may, for a variety of reasons, cause non-union wages to rise. Second, the desire and ability of unions to secure large wage increases for their members will depend in part on the magnitude of the union wage relative to the non-union wage. Combining these assumptions and some further hypotheses yields a number of testable implications concerning aggregate wage behavior.

A. Behavioral Assumptions

Conventional static theory suggests that the "typical" trade union will try to obtain a wage rate which equates at the margin an internally and politically determined trade-off between employment and wages with the net derived demand for unionized labor. The latter function will in turn depend upon the extent of substitution between the output produced by union and nonunion firms and hence upon the nonunion wage. A result of the conventional theory is thus that the desired union wage ($\bar{W}_u^*$) from the point of view of the union will be determined as proportionately larger than the

5/ These are essentially the views of Slichter [36].

6/ Most of the studies of wage inflation referred to above assume, explicitly or implicitly, that there are no monetary constraints on the inflationary process, a caveat that also applies to the investigation in this paper.

7/ See, for example, the formulations by Cartter [7, Chapt. 8] or Throop [40, pp. 60-81].
current nonunion wage ($\bar{w}_n$), i.e.

$$\bar{w}_u^* = a \bar{w}_n,$$

with $a > 1$. In logarithms (with all variables expressed this way henceforth) this is:

(1) $\ln \bar{w}_u^* = \ln \bar{w}_n + a, \; a > 0$. 8/

At the same time, however, there will be a tendency for the desired nonunion wage ($\bar{w}_n^*$), from the point of view of the nonunion employer, to be determined as some proportion (most likely less than unity) of the current union wage ($\bar{w}_u$), i.e.

(2) $\bar{w}_n^* = \bar{w}_u + b, \; b < 0$

The justification for this relationship requires further comment:

1. The existence of unionism in one sector of the economy will have an effect on the desired wage in the nonunion sector because of the threat which a large union/nonunion wage differential poses to the future organization (and subsequent rise in wages) of nonunion employers. It is easy to show, under a plausible set of assumptions, that the profit maximizing nonunion employer will always consider the level of the union wage

---

8/ Empirical evidence that a relationship such as (1) has held in the U.S. over a long period of time may be deduced from the careful summary and analysis by Lewis [23].
rate in determining his own optimal wage offer. Although it does not appear to have been systematically tested, there is substantial casual evidence that the threat of unionism has historically been an important factor in the determination of nonunion wages.

9/ The wage which a nonunion employer expects to pay is

\[ E(W) = pW_u + (1-p)W_n, \]

where \( p \) is the probability of organization and must surely depend positively on \( W_u / W_n \), an index of the success of trade unions in the industry. Thus \( p = p(W_u / W_n) \), \( p' > 0 \), and hence \( dp/dW_n < 0 \). Since there is a one-to-one correspondence between profits and the wage rate, the optimum nonunion wage may be determined by differentiating (a) and solving to obtain

\[ W_n^* = W_u + (1-p)/dp/dW_n, \]

assuming the second order condition is satisfied. See Lewis [23] for a similar argument and Rosen [34] for an interesting development and discussion of the implications of this result as well as some suggestive cross-sectional empirical work. As Rosen points out, \( p \) must also depend on the extent of union organization so that the impact of the threat effect is likely to vary over time and to be highly correlated with the growth in trade union membership.

10/ First, there are a number of well-known examples of this phenomenon available. Slichter [36] reports that in an effort to halt the spread of unionism the U.S. Steel Corporation granted its "company unions" a 10% wage increase in late 1936 even though about one-sixth of the U.S. labor force was unemployed at the time and the consumers' price index had been stable for a year. Levinson, in his carefully documented and recent case study of the U.S. lumber and paper industries reports a different type example: "The close geographic, technological, and corporate relationships between the lumber and paper industries led to a situation, however, in which the political-power pressures in the former led to important 'carryover' effects into the latter because of the fear of both the conservative paper union leaders and the paper companies that their traditionally peaceful bargaining relationship might be seriously threatened if the large improvements in lumber benefits were not matched or exceeded" [22, pp. 261-264]. Second, it should be recognized that the threat-effect hypothesis has implications for a number of observable variables other than wages. For example, suppose that the strength of the threat effect is correlated with the growth of trade union membership and that unorganized employers have other legal weapons in their "portfolio" for combatting the spread of unionism aside from wage increases. Then the (Continued on following page.)
2. To the extent that the quality of the work force in an industry is positively related to the wage rate in that industry, nonunion firms will be forced to raise \( \hat{w}_n \) when \( w_u \) rises or suffer a decline in average labor efficiency.\(^{11}\)

3. Finally, even if workers in the nonunion sector are excluded from potential employment in the union sector, the morale and personal productivity of nonunion workers may fall if \( (w_u - w_n) \) becomes very large.\(^{12}\)

Now the proportionate wage change in the union sector at a point in time will consist of two components: (1) that which is due to the operation of market forces and independent union pushfulness, and (2) that which is due to \( w_u \) "catching up" with \( w^*_u \). Symbolically, it is hypothesized that

\[
(3) \quad \dot{w}_u = \alpha(w^*_u - w_u) + \epsilon(z_u), \quad \alpha > 0,
\]

where \( \alpha \) is an adjustment coefficient representing the rate at which \( w_u \) catches

(Continued from previous page.)

extent of use of these costly alternative weapons should be correlated with trade union membership growth if the threat effect is truly important. One weapon historically available in the U.S. was the employee representation plan (so-called by employees) or company union (so-called by trade unions). It is well known that these plans first appeared in the U.S. in the period 1914-1920, a period in which union membership nearly doubled. See Paul H. Douglas \[11\]. Further, employees covered by such plans declined 18% in the period 1928-32 and increased over 100% between 1932-35 while union membership moved similarly. See \[26\] and \[27\]. Employee representation plans were made illegal by the passage of the Wagner Act in 1935.

\(^{11}\) See Ashenfelter and Johnson \[2\] for evidence that average labor quality is related to the wage rate.

\(^{12}\) This argument is due to Slichter \[36\]. Also see Behman \[4\] for a similar position. It is, of course, very difficult to obtain evidence on this issue.
up with $W_u^*$ and $Z_n$ is a vector of variables influencing the size of independent wage changes in the unionized sector. The proportionate wage change in the nonunion sector may similarly be viewed as the sum of independent and "spillover" components, so that

$$
\dot{W}_n = \beta(W_n^* - W_n) + \psi(Z_n), \quad \beta > 0,
$$

where $Z_n$ is a vector of variables which influences $\dot{W}_n$ independently of the size of $W_n$ relative to $W_u$ and presumably reflects primarily market forces.

Substitution of (1) and (2) into (3) and (4) gives the following system of two linear differential equations:

$$
\begin{align*}
\dot{W}_u &= -\alpha W_u + \alpha W_n + \alpha a + \psi(Z_u) \\
\dot{W}_n &= \beta W_u - \beta W_n + \beta b + \psi(Z_n).
\end{align*}
$$

Since it is not possible to obtain solutions for $W_u$ and $W_n$ for this system in the usual way, it is necessary to convert the equations into a system in terms of $\dot{W}_u$ and $\dot{W}_n$. Differentiating (on the assumptions that $Z_n$ and $b$ remain constant over the interval under consideration) (4) with respect to time to get

$$
\ddot{W}_n = \beta \dot{W}_u - \beta \dot{W}_n,
$$

and substituting (5) into (7) and the expression for $\dot{W}_u$ from (6) into the result gives

$$
\ddot{W}_n = -(\alpha + \beta)\dot{W}_n + \alpha \beta (a + b) + [\alpha \psi(Z_n) + \beta \psi(Z_u)].
$$

---

13/ This is because $\begin{bmatrix} -\alpha & \alpha \\
\beta & -\beta \end{bmatrix}$ is singular.
An expression similar to (8) may be obtained for \( W_u \), so that the steady state rates of growth of the union and nonunion wage rates are

\[
\dot{w}_n^u = \dot{w}_n^o = \frac{1}{(\alpha + \beta)} \left[ \alpha \psi(Z_u) + \alpha \psi(Z_n) + \beta \psi(Z_u) \right].
\]

On this view of aggregate wage determination the steady-state rates of growth of union and nonunion wage rates are equal. Further, there are two components to the long-run wage inflation process: (1) that which is due to the independent forces \( \psi(Z_u) \) and \( \psi(Z_n) \), and (2) that which is due to \( \alpha + \beta \), the forces affecting desired union/nonunion wage differentials. The sign of \( \alpha + \beta \) depends upon whether nonunion employers think the union/nonunion wage differential should be larger or smaller than unions think it should be. If the latter, the spillover mechanism adds an inflationary component to wage changes and if the former it adds a deflationary component. Finally, if the two parties agree, \( \alpha = -\beta \), and the steady-state wage change depends only upon \( \psi(Z_u) \) and \( \psi(Z_n) \).

From (8) the solution for \( W_n \) is

\[
(9) \quad \dot{W}_n = \dot{w}_n^o (1 - e^{-(\alpha + \beta) t}) + \dot{w}_n^o e^{-(\alpha + \beta) t},
\]

where \( \dot{w}_n^o \) is the rate of growth of the nonunion wage rate in some initial time period. The time rate of change of the logarithm of the union/nonunion relative wage, \( D = W_u - W_n \), is obtained from (5) and (6) as

\[
(10) \quad D = -((\alpha + \beta) D + (\alpha \psi - \beta \psi) + [\psi(Z_u) - \psi(Z_n)]
\]

\[\text{Note that Rosen's theoretical argument [34] may be interpreted as implying that there are forces tending to make } \alpha = -\beta \text{ in full equilibrium.}\]
so that

\[ D = D^e \left(1 - e^{-(\alpha + \beta) t}\right) + D^o e^{-(\alpha + \beta) t}, \]

where

\[ D^e = \frac{1}{(\alpha + \beta)} \left[ \psi(Z_u) - \psi(Z_n) + (\alpha a - \beta b) \right] \]

is the steady state union/nonunion relative wage, and \( D^o \) is the size of the differential in some initial time period. From (11) it can be seen that even though the time paths of \( \psi(u) \) and \( \psi(n) \) are not determined by this model the path of the wage differential is. Further, from (12) the equilibrium wage differential may be interpreted as a weighted average of the parameters \( \alpha \) and \( \beta \). For present purposes, however, empirical interest is centered on the proportionate change of the aggregate money wage rate over some finite interval. For convenience of exposition define the aggregate wage as the weighted geometric mean of the union and the nonunion wage rates so that its logarithm is

\[ \ln W = \ln W_u + \ln W_n (1-U), \]

where \( U \) is the ratio of employed union members (T) to total employment (E). Then

\[ \dot{\ln W} = \dot{\ln W}_u + \dot{\ln W}_n (1-U) - \ln W_u \]

\[ = \dot{\ln W}_n + \dot{\ln D} U + \dot{\ln D}, \]

which, on the convenient assumption that the rates of growth of \( T \) and \( E \) are constant over the short interval under consideration at \( g_T \) and \( g_E \) respectively becomes

\[ \dot{\ln W} = \dot{\ln W}_n + \dot{\ln D} \gamma e^{\lambda T} + \gamma \dot{\ln D} \gamma e^{\lambda T}, \]
where \( \lambda = g_t - g_e \) and \( U^O \) is the proportion of employment unionized in the initial period. Substituting (9), (10), and (11) into (13) and then integrating over a finite interval of length \( t' \) it is possible to obtain an explicit expression for

\[
\Delta W = \int_0^{t'} \omega t' dt = f(\psi(Z)\psi(Z), \lambda, U_t, X)
\]

i.e. the proportionate change of the aggregate money wage rate over a period of length \( t' \).\(^{15}\) where \( X = [\alpha, \beta, \sigma, b, t'] \). Interest centers on the signs of \( \frac{\partial \Delta W}{\partial \lambda} \), \( \frac{\partial \Delta W}{\partial \phi} \), and \( \frac{\partial \Delta W}{\partial Y} \). First, it can be shown that

\[
\frac{\partial \Delta W}{\partial \lambda} = U^O t' \left[ \beta e^{-((\alpha+\beta)t')} + \beta^e (e^{\lambda t'} - e^{-(\sigma+\beta)t'}) \right],
\]

so that as long as \( \beta^e > 0 \) the effect of a growth in unionism on the aggregate money wage rate is positive.\(^{16}\) Second, it can be shown that

\[
\frac{\partial \Delta W}{\partial \phi} = \frac{\beta}{\alpha+\beta} t' - \frac{\beta}{(\alpha+\beta)^2} (1-e^{-(\alpha+\beta)t'}) - \frac{U^0}{\alpha+\beta} (e^{-(\alpha+\beta)t'} - e^{\lambda t'}),
\]

and

\[
\frac{\partial \Delta W}{\partial Y} = \frac{\alpha}{\alpha+\beta} t' + \frac{\beta}{(\alpha+\beta)^2} (1-e^{-(\alpha+\beta)t'}) + \frac{U^0}{\alpha+\beta} (e^{-(\alpha+\beta)t'} - e^{\lambda t'}).\]

\(^{15}\) The manipulations required to obtain \( \Delta W \) are straightforward but tedious, and the explicit relationship obtained is not very suggestive.

\(^{16}\) Lewis estimates \( 0.16 < D < 0.22 \) as an average over the period 1920-58 [23, p. 222]. He also estimates \( D \) at 5-year intervals over this period and his estimates range from .38 for 1930-34 to .02 for 1945-49.
Surprisingly, both of these expressions have ambiguous signs, although

\[
\frac{\partial \omega}{\partial \phi} + \frac{\partial \omega}{\partial \psi} = t' > 0, \text{ so that at least one of them must be positive. Intuitively,}
\]

increases in independent pressures on wage changes in the union (nonunion) sector need not lead to an increase in \( \Delta \omega \) over a time period of arbitrary length, even in this spillover model of wage determination, if the extent of unionism is declining (increasing) rapidly and if the spillover mechanism operates very slowly (i.e. if \( \alpha \) and \( \beta \) are very small). 17/ This suggests imposing as part of the spillover hypothesis the assumption that spillovers take place fairly rapidly. 18/ (15) and (16) then become

\[
(15a) \quad \frac{\partial \omega}{\partial \phi} \approx \frac{\beta}{\alpha + \beta} t' > 0
\]

\[
(16a) \quad \frac{\partial \omega}{\partial \psi} = \frac{\alpha}{\alpha + \beta} t' > 0
\]

17/ Note that in the long run (steady-state) wage inflation process increments in \( \phi(Z_u) \) and \( \psi(Z_n) \) must have positive effects on aggregate wage changes. This may be observed by taking \( t' \) to be very large in equations (15) and (16).

18/ This assumption seems consistent with what proponents of the spillover hypothesis have in mind. See Slichter [36], Pierson [31], and Eckstein and Wilson [12]. Note that (15) and (16) suggest that this assumption is essential for obtaining unambiguous predictions about the signs of \( \frac{\partial \omega}{\partial \phi} \) and \( \frac{\partial \omega}{\partial \psi} \) in a spillover model.
This implies that independent wage change forces will be relatively more important for the aggregate wage change the faster that the other sector catches up via the spillover mechanism.\footnote{19}

B. An Operational Wage Determination Model

The preceding discussion suggests that the rate of growth of union membership as well as the elements of the vectors $Z_U$ and $Z_N$ should have an effect on aggregate money wage changes. Since only the former has a clearly observable counterpart, further hypotheses relating the latter to observable variables must be specified.

1. It seems reasonable to suppose that the following set of variables reflect independent pressures on wage changes in the unionized sector:

   a. $UN$: unemployment as a percentage of the civilian labor force.

   b. $\Delta P$: the percentage rate of change of consumer prices

   c. $\Delta T$: the percentage rate of change of trade union membership

   d. $S$: an index of the degree of militancy of rank-and-file union members.

\footnote{19}{It is also interesting to note that
\frac{\partial \omega}{\partial U} = - B_0 (1-e^{-(\alpha+\beta-)t'}) + D_0 e^{\alpha t'} - e^{-(\alpha+\beta-)t'},
which has an ambiguous sign, regardless of the sizes of $\alpha$ and $\beta$, so that in this model the level of trade union membership does not necessarily have a positive effect on wage changes.
Now $\frac{\partial \delta}{\partial U^2} < 0$ and $\frac{\partial \delta}{\partial \Delta P} > 0$ for well-known reasons: excess demand pressures and union pressures to protect members' standards of living. Also $\frac{\partial \delta}{\partial \Delta T} > 0$ because union leaders will be the more militant the faster the union movement as a whole is growing. See Hines [18] and [19]. Finally, the ratio of the number of industrial strikes to union members is taken as an index of the degree to which union members, as distinct from their leaders, desire a high or low degree of union pushfulness, so that $\frac{\partial \delta}{\partial S} > 0$. Since $\frac{\partial \delta M}{\partial \delta} > 0$, it is predicted that UN will have a negative effect on $\Delta M$ and $\Delta P$ and $S$ will each have positive effects on $\Delta M$. Finally, the effect of $\Delta T$ on $\Delta M$ is twofold since

$$\frac{\partial \Delta M}{\partial \Delta T} = \frac{\partial \Delta M}{\partial \delta} \frac{\partial \delta}{\partial \Delta T} + \frac{\partial \Delta M}{\partial \delta} > 0.$$ 

The first effect on the right hand side reflects the effect of $\Delta T$ on union militancy while the second reflects the shift from the nonunion to the union sector of the weights in the weighted average aggregate wage.

2. In addition to maintaining an equilibrium with the union wage rate, employers in the nonunion sector will adjust wages in response to the pressure of excess demand. Hence, following Lipsey [24], the major independent force on nonunion wage rates should be UN, with $\frac{\partial \omega}{\partial \text{UN}} < 0$, and hence

$$\frac{\partial \omega}{\partial \text{UN}} < 0.$$

---

20/ See Ashenfelter and Johnson [1] for evidence that the volume of strike activity is indicative of rank-and-file union wage pressures. Note also that the variable $S$ has the advantage that it reflects organized efforts to increase (or stop decreases in) wages even in periods when the unionized sector is small. Slichter [36] also draws attention to $S$ as an index of union pressure on wage levels.
Assuming a linear relationship in the relevant ranges of the variables, combining the above hypotheses about the determinants of $Z_u$ and $Z_n$, and adding a disturbance term yields the following estimating equation:

\[
\Delta W_t = \alpha_0 + \alpha_1 W_{nut} + \alpha_2 \Delta P_t + \alpha_3 \Delta T_t + \alpha_4 S_t + \epsilon_{1t},
\]

with the hypotheses $\alpha_1 < 0$ and $\alpha_2, \alpha_3, \alpha_4 > 0$.

II. Empirical Results

Equation (17) will be fitted to annual data on percentage changes in compensation (wages plus fringes) per man hour in U.S. manufacturing industries. In order to maintain continuity with the work of others, however, it is useful to proceed in two steps. First, (17) is estimated on the assumption that $\alpha_3 = \alpha_4 = 0$. This assumption is consistent with previous work on U.S. data and serves to introduce the problem of simultaneity in wage and price determination. Second, a three equation model including (17) is estimated.

A. Previous Empirical Results

Building on the work of Bodkin [5], Rees and Hamilton [32] present empirical results for the estimating equation

\[
\Delta W_t = \alpha_0 + \alpha_1 W_{nut} + \alpha_2 W_{nut-1} + \alpha_3 \Delta P_t + \alpha_4 \Theta + \epsilon_t,
\]

with hypotheses $\alpha_1 + \alpha_2 < 0$, $\alpha_3 > 0$, $\alpha_4 > 0$, where $\Theta = 0$ or 1 and $W_{nut}$ is a dummy variable taking the value unity for the operation of National Recovery Act codes in 1934 (which legislated a decline in weekly hours with no decline in weekly pay).\(^{21/}\) Their ordinary least squares (OLS) results are listed in

\(21/\)See the appendix for more precise definitions and data sources.
Table I as equations (18a) and (18c). Data limitations will require further work with equation (17) to span the period 1914-59, as opposed to the period 1901-57 used by Rees and Hamilton and Bodkin. For purposes of comparison, estimates for the former period are listed in the table as equations (18b) and (18d). As can be seen, changing the period of fit has little effect on the estimates.

Consider equation (18c) or (18d) where contemporaneous price changes appear as an independent variable. First, $\hat{\alpha}_1 + \hat{\alpha}_2 > 0$ so that the steady-state "Phillips curve" implied by either equation has a positive slope, contrary both to hypothesis and the evidence of other investigators. Second, $\hat{\alpha}_3 > 1$ so that an increase in prices leads to a larger increase in money wages, again contrary to the evidence of other investigators. Rees and Hamilton suggest that the cause of these surprising results may be simultaneous-equations bias. In particular, they conjecture that the OLS estimator of $\alpha_3$ in equation (18) may be biased upward because some of the "credit" for the effect of wage changes on price changes will have been given to the effect of price changes on wage changes.²²/ In order to avoid this difficulty Rees

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²²/ It is straightforward to show that for "large samples" their conjecture is correct. Assuming $\beta(> 0)$ to be the coefficient on $\Delta P$ in an equation determining $\Delta P$ the asymptotic bias in the OLS estimator $\hat{\alpha}_3$ of $\alpha_3$ in (18) is

$$\lim_{t \to \infty} \hat{\alpha}_3 - \alpha_3 = [\Delta (1-\alpha_3 \beta)]^{-1} \operatorname{var}(\text{UN}_t) \times$$

$$\operatorname{var}(\text{UN}_{t-1}) (1-\gamma \text{UN}_t \text{UN}_{t-1}) \times$$

$$\beta(\operatorname{var}(\epsilon_1) + \operatorname{cov}(\epsilon_1 \epsilon_2)),$$

where $\Delta(> 0)$ is the determinant of a positive definite matrix, and $\epsilon_2$ is the disturbance term in the price determination equation. Ignoring $\operatorname{cov}(\epsilon_1 \epsilon_2)$ for lack of information (although it seems likely to be positive) and noting that $(1-\alpha_3 \beta) > 0$ for stability of the system, it is clear that this bias is positive.
and Hamilton lag the price change variable and estimate a kind of reduced form wage equation, i.e. equation (18a) or (18b) in Table I. For this formulation \( \hat{c}_1 + \hat{c}_2 < 0 \), but not significantly so, and \( \hat{c}_3 < 1 \). Even this is achieved only at substantial expense, however, because \( R^2 \) drops from .83 to .55 and the standard error of estimate increases from 3.3 to 5.6 percentage points.

Clearly, the appropriate procedure for estimation of (18) requires treating \( \Delta P_t \) as endogenous and specifying those predetermined variables determining \( \Delta P_t \) but not \( \Delta W_t \). Accordingly, equation (18e) in Table I is the stage estimated two least squares (2SLS) version of (18) on the assumption that \( \Delta W_{t-1} \) and changes in average labor productivity (\( \Delta X_t \)) are additional predetermined variables.\(^{23}\) As can be seen from the table, \( \hat{c}_1 + \hat{c}_2 < 0 \), but not significantly so, \( \hat{c}_3 < 1 \), and there is little decline in explanatory power from (18b) to (18e). Although (18e) is certainly an improvement over (19b) it is clearly unsatisfactory. Since \( (\hat{c}_1 + \hat{c}_2) = 0 \), it essentially reads \( \Delta W = f(\Delta P) \).

B. The Estimated Model

From the above results it is clear that an adequate description of wage determination in U.S. manufacturing requires that at least \( \Delta W_t \) and \( \Delta P_t \)

\( ^{23} \)Limited information maximum likelihood (LIML) estimates have been obtained for all equations where 2SLS estimates are shown. As is well known, the former have the advantage of being invariant under changes in normalization. Consequently, the LIML estimates are also reported in the text whenever they differ substantially from the 2SLS results.

The LIML estimates of equation (15) are not reported because they are virtually identical to the 2SLS estimates.
Table I
Determinants of the Rate of Change of Money Wages

<table>
<thead>
<tr>
<th>Equation</th>
<th>Period Covered</th>
<th>Estimating Technique</th>
<th>Coefficients on:</th>
<th>Constant</th>
<th>UNt</th>
<th>UNt-1</th>
<th>ΔPt</th>
<th>ΔPt-1</th>
<th>N</th>
<th>-R²</th>
<th>D.-W.</th>
<th>SEE</th>
</tr>
</thead>
<tbody>
<tr>
<td>18a</td>
<td>1901-57</td>
<td>OLS</td>
<td></td>
<td>3.942</td>
<td>-1.410</td>
<td>1.320</td>
<td>0.665</td>
<td>16.831</td>
<td>0.554</td>
<td>1.629</td>
<td>N.A.</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(1.272)*</td>
<td>(0.237)*</td>
<td>(0.251)*</td>
<td>(0.128)*</td>
<td>5.647*</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>18b</td>
<td>1914-59</td>
<td>OLS</td>
<td></td>
<td>4.610</td>
<td>-1.412</td>
<td>1.293</td>
<td>0.648</td>
<td>16.823</td>
<td>0.545</td>
<td>1.60</td>
<td>5.57</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(1.601)*</td>
<td>(0.271)*</td>
<td>(0.284)*</td>
<td>(0.145)*</td>
<td>6.150*</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>18c</td>
<td>1901-57</td>
<td>OLS</td>
<td></td>
<td>2.204</td>
<td>-0.417</td>
<td>0.462</td>
<td>1.058</td>
<td>11.073</td>
<td>0.831</td>
<td>2.098</td>
<td>N.A.</td>
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</tr>
<tr>
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<td></td>
<td></td>
<td>(0.786)*</td>
<td>(0.166)*</td>
<td>(0.158)*</td>
<td>(0.084)*</td>
<td>3.504*</td>
<td></td>
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<td></td>
</tr>
<tr>
<td>18d</td>
<td>1914-59</td>
<td>OLS</td>
<td></td>
<td>2.490</td>
<td>-0.395</td>
<td>0.428</td>
<td>1.057</td>
<td>11.150</td>
<td>0.840</td>
<td>2.08</td>
<td>3.34</td>
<td></td>
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<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(0.962)*</td>
<td>(0.188)*</td>
<td>(0.175)*</td>
<td>(0.093)*</td>
<td>3.723*</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>18e</td>
<td>1914-59</td>
<td>2SLS</td>
<td></td>
<td>3.435</td>
<td>-0.560</td>
<td>0.523</td>
<td>0.894</td>
<td>12.022</td>
<td>0.827</td>
<td>1.93</td>
<td>3.58</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(1.131)*</td>
<td>(0.216)*</td>
<td>(0.190)*</td>
<td>(0.133)*</td>
<td>3.890*</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: Estimated standard errors are shown in parentheses below the estimated regression coefficients. The t distribution is used for two-tailed significance tests and the asterisk denotes significance at the 5% level. The coefficient of determination is corrected for degrees of freedom. D-W is the Durbin-Watson statistic, a measure of the first-order serial correlation of the estimated residuals. SEE is the standard error of estimate for the regression equation (and N.A. denotes that it is not available for equations (a) and (c)).
be considered jointly determined.  In addition, $\Delta T_t$ should be considered endogenous. The reasons for this may be briefly stated here; more detail is provided by Eines [17] and Ashenfelter and Fencavel [3]. The militancy of trade union leaders is likely to be greatest when they are gaining new members. Moreover, trade unions are likely to be gaining new members when prices are rising rapidly and workers are anxious to maintain their real wage without incurring the costs of mobility in the labor market. All of this suggests that $\Delta T$ will be determined in part by $\Delta P$ and thus should be treated as endogenous.

A summary statement of the model to be estimated and associated hypotheses is as follows:

\begin{align}
\Delta M_t &= \alpha_0 + \alpha_1 \Delta N_t + \alpha_2 \Delta P_t + \alpha_3 \Delta T_t + \alpha_4 S_t + \alpha_5 N_t + \epsilon_1 t \\
& \alpha_1 < 0; \quad \alpha_2, \alpha_3, \alpha_4, \alpha_5 > 0. \\
\Delta P_t &= \beta_0 + \beta_1 \Delta M_t + \beta_2 \Delta N_t + \beta_3 \Delta T_t + \epsilon_2 t \\
& \beta_1, \beta_2 > 0; \quad \beta_3 < 0. \\
\Delta T_t &= \gamma_0 + \gamma_1 \Delta P_t + \gamma_2 \Delta E_t + \gamma_3 \Delta T_t - 1 + \gamma_4 D_t + \gamma_5 U_t + \gamma_6 (T/E)_t - 1 + \epsilon_3 t \\
& \gamma_1, \gamma_2, \gamma_3, \gamma_4, \gamma_5 > 0; \quad \gamma_6 < 0.
\end{align}

---

Perry [29] does not find this to be the case with quarterly U.S. data for the post-war period, nor does Dicks-Nibereux [10] with annual U.K. data for the post-war period.

The relationship between the growth of union membership and price increases has been recognized for some time. H. Davis [9] presented early evidence to support this hypothesis in the U.S., U.K., France and Germany. Note also that the argument in the text might also apply to the variable $S_t$. Some attempts to develop a structural equation for $S_t$ were made, but none of these could explain more than a small fraction of the variance in this variable and none contained even marginally significant coefficients on any of the endogenous variables noted in the text.
1. Equation (19) is simply a restatement of (17) with the dummy variable \( N_t \) added.\(^{26/}\)

2. Equation (20) is a simplified price determination equation which follows closely the work of Bodkin [5] and others. It is not designed to test between competing macroeconomic hypotheses of price determination and is surely consistent with most of them. Its explicit purpose in this model is to provide a reasonable framework for dealing with simultaneity biases.

3. Equation (21) is a simplified version of a more elaborate equation carried over from the work of Ashenfelter and Parentavel [3]. \( E_t \) is the level of employment in the "unionized" sectors of the economy (primarily

\(^{26/}\)Two points are worthy of note with respect to equation (19). First, comparing (19) with (18) it will be seen that the variable \( W_{t-1} \) has not been included in the former. The reason is because in a 2SLS preliminary test of (19) which included \( W_{t-1} \) its coefficient estimate dropped to .18 with an estimated standard error of .25. Leaving \( W_{t-1} \) out of (19) had very little effect on the other coefficient estimates. One implication of this result is that \( \Delta W \) does not have a significant effect on \( \Delta M_t \). Second, from the arguments of Section I it might be supposed that \( \Delta M_t - \Delta E_t \) should have been introduced into (19). The reason that this variable is not included in (19) is because in a 2SLS preliminary test which included \( \Delta E_t \) and \( \Delta T_t \) separately the estimated coefficient of the former was only .08 with an estimated standard error of .11. Leaving \( \Delta E_t \) out of (19) had very little effect on the other coefficient estimates. This result is consistent with a number of hypotheses, none of which can be tested directly with the data available. (a) In terms of the arguments of Section I it may be that

\[
\frac{\Delta M}{\Delta L} = 0 \quad \text{so that} \quad \frac{\Delta M}{\Delta L} = \frac{\Delta M}{\Delta L} \frac{\Delta L}{\Delta L},
\]

i.e. the effect of union growth on \( \Delta M \) operates solely through its effect on bargaining militancy and not through its effect on the weights in the weighted-average aggregate wage. (b) It may be that \( \Delta M_t - \Delta E_t \) and \( \Delta E_t \) should each enter (19) separately, the latter as a proxy for changes in the excess demand for labor. The above result might then be rationalized on the grounds that the coefficients of these two variables are nearly equal.
manufacturing), $D_t$ the percentage of membership in the House of Representatives which is affiliated with the National Democratic Party, and $U^D_t$ the unemployment rate in the preceding trough of the business cycle. Briefly, the $\Delta E_{t-1}$ reflect the trade cycle forces affecting union membership; $U^D_t$ and $D_t$ are proxies for the social and political forces involved in trade union growth; and the "saturation hypothesis" is measured by $(T/E)_{t-1}$.

Two preliminary issues remain. First, the period of fit for equations (19) - (21) is 1914-1959, and is prescribed by data limitations. It is a period, however, which is likely to be favorable to the hypotheses associated with (19). In particular, trade union membership was an especially small fraction of manufacturing employment prior to World War I. Second, since the system (19) - (21) requires simultaneous-equations estimation procedures it is unclear which of the several methods available should be used. As is well known, the small-sample properties of the available procedures are essentially unknown. Rather than choose among them, 2SLS, LIML, and three stage least squares (3SLS) estimates of (19) - (21) have been computed. Where the 2SLS and LIML estimates are not substantially different, only the 2SLS and 3SLS results are presented. Other things equal, presumably the 3SLS estimates are to be preferred.

Estimates of equations (19) - (21) are contained in Tables II, III, and IV. Each equation fits the data well and most of the estimated coefficients in these tables would be judged significantly different from zero on the conventional criteria.27/

27/ As is well known, some care must be exercised in judgments as to significance in models like this. In addition, it is worth making explicit exactly what pre-testing of these equations has been done. In addition to the tests noted in footnote 24 the following additional tests were made before settling (Continued on following page.)
1. The estimated versions of equation (19) lend substantial support to the hypotheses advanced in Section I. The estimated coefficients of both $\Delta T_t$ and $S_t$ are well over twice their estimated asymptotic standard errors regardless of the estimation procedure used. The same is true of the estimated coefficient of $\Delta W_t$, the unemployment rate. Neither result is very surprising in view of Figure 1, which contains a graph of $\Delta T$, $UN$, and $\Delta W$ as functions of time. One implication of these results is that market forces have had an important effect on $\Delta W_t$, but that a theory of institutional change is helpful for isolating these forces over any long historical period. Finally, the estimated effect of $\Delta P_t$ on $\Delta W_t$ varies widely among (19a), (19b), and (19c); from essentially zero in (19b) to about 0.5 in (19c). Presumably, the latter (3SLS) estimate is to be preferred.

2. Equation (20) provides a very tight fit to the data and each estimated coefficient is highly significant. Note that the coefficient of $\Delta X_t$ is significantly different from -1.0, in contrast to the conventional assumption used in computing the reduced form relationship between $\Delta P_t$ and $UN_t$ [29]. The derived reduced form estimate of the effect of $\Delta X_t$ on $\Delta P_t$ is obtained below, and it differs considerably from the structural estimate.

3. As might be expected, equation (21) does not fit the data as well as (19) or (20). The estimated coefficients for (21) are very similar to those reported elsewhere [3], although the standard errors reported here

(Continued from previous page.) on the estimated model presented in Tables II-IV: (a) A variable measuring the percentage change in the number of work stoppages per year was substituted for $S_t$. This variable was significant by conventional standards and its inclusion had little effect on the other estimates, although its coefficient was smaller with respect to its standard error than was that of $S_t$. (b) The percentage change in farm prices was added as an exogenous variable in (20). Although positive, its estimated coefficient was not significantly different from zero on the conventional criteria and so it was dropped from (20).
### Table II
Determinants of the Rate of Change of Money Wages, 1914-1959

<table>
<thead>
<tr>
<th>Equation</th>
<th>Estimating Technique</th>
<th>Constant</th>
<th>UNt</th>
<th>△Pt</th>
<th>N</th>
<th>△Tt</th>
<th>St</th>
<th>R²</th>
<th>D.-W.</th>
<th>SEE</th>
</tr>
</thead>
<tbody>
<tr>
<td>(19a)</td>
<td>2SLS</td>
<td>3.456*</td>
<td>-0.503*</td>
<td>0.335</td>
<td>16.384*</td>
<td>0.436*</td>
<td>0.747*</td>
<td>0.801</td>
<td>2.28</td>
<td>3.719</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.405)</td>
<td>(0.182)</td>
<td>(0.255)</td>
<td>(4.213)</td>
<td>(0.151)</td>
<td>(0.302)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(19b)</td>
<td>LIML</td>
<td>4.129*</td>
<td>-0.663*</td>
<td>0.083</td>
<td>17.544*</td>
<td>0.577*</td>
<td>0.894*</td>
<td>0.742</td>
<td>2.26</td>
<td>4.234</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.600)</td>
<td>(0.207)</td>
<td>(0.290)</td>
<td>(4.796)</td>
<td>(0.172)</td>
<td>(0.344)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(19c)</td>
<td>3SLS</td>
<td>3.824*</td>
<td>-0.451*</td>
<td>0.487*</td>
<td>11.509*</td>
<td>0.407*</td>
<td>0.525*</td>
<td>0.809</td>
<td>2.18</td>
<td>3.630</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.125)</td>
<td>(0.142)</td>
<td>(0.207)</td>
<td>(3.257)</td>
<td>(0.121)</td>
<td>(0.228)</td>
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<td></td>
<td></td>
</tr>
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</table>

[See notes to Table I]
Table III
Determinants of the Rate of Change of Prices, 1914-1959

<table>
<thead>
<tr>
<th>Equation</th>
<th>Estimating Technique</th>
<th>Constant</th>
<th>$\triangle W_t$</th>
<th>$\triangle W_{t-1}$</th>
<th>$\triangle X_t$</th>
<th>$R^2$</th>
<th>D.-W.</th>
<th>SEE</th>
</tr>
</thead>
<tbody>
<tr>
<td>(20a)</td>
<td>2SLS</td>
<td>-1.253*</td>
<td>0.720*</td>
<td>0.174*</td>
<td>-0.485*</td>
<td>0.871</td>
<td>1.73</td>
<td>2.337</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.482)</td>
<td>(0.057)</td>
<td>(0.048)</td>
<td>(0.098)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(20b)</td>
<td>3SLS</td>
<td>-1.403*</td>
<td>0.730*</td>
<td>0.154*</td>
<td>-0.404*</td>
<td>0.870</td>
<td>1.63</td>
<td>2.347</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.450)</td>
<td>(0.040)</td>
<td>(0.051)</td>
<td>(0.087)</td>
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<td></td>
</tr>
</tbody>
</table>

[See notes to Table I]
Table IV
Determinants of the Rate of Growth in Union Membership, 1914-1959

<table>
<thead>
<tr>
<th>Equation</th>
<th>Estimating Technique</th>
<th>Constant</th>
<th>$\Delta P_t$</th>
<th>$\Delta E_t$</th>
<th>$\Delta E_{t-1}$</th>
<th>$D_t$</th>
<th>$U^P_t$</th>
<th>$(T/E)_{t-1}$</th>
<th>$R^2$</th>
<th>D.-W.</th>
<th>SEF</th>
</tr>
</thead>
<tbody>
<tr>
<td>(21a)</td>
<td>2SLS</td>
<td>-13.418*</td>
<td>0.659*</td>
<td>0.201</td>
<td>0.018</td>
<td>0.255*</td>
<td>0.298*</td>
<td>-0.037</td>
<td>0.672</td>
<td>1.61</td>
<td>4.849</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(3.900)</td>
<td>(0.195)</td>
<td>(0.128)</td>
<td>(0.119)</td>
<td>(0.084)</td>
<td>(0.146)</td>
<td>(0.041)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(21b)</td>
<td>3SLS</td>
<td>-13.483*</td>
<td>0.562*</td>
<td>0.283*</td>
<td>0.036</td>
<td>0.262*</td>
<td>0.237</td>
<td>-0.025</td>
<td>0.663</td>
<td>1.55</td>
<td>4.919</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(3.329)</td>
<td>(0.171)</td>
<td>(0.108)</td>
<td>(0.099)</td>
<td>(0.070)</td>
<td>(0.122)</td>
<td>(0.034)</td>
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</tr>
</tbody>
</table>

[See notes to Table I]
Figure 1
The Rate of Change of Union Membership ($\Delta T$)
The Rate of Change of Money Wages ($\Delta M$) and
The Civilian Unemployment Rate: 1914-1959.
are somewhat larger because of the smaller sample size used. Note that $\Delta T_t$ is positively correlated with $U^p_t$, so that it cannot be interpreted as a proxy for excess demand in labor markets. In fact, the simple correlation coefficient between $\Delta T_t$ and $UH_t$ is +.18. Finally, $\Delta P_t$ has a strong effect on $\Delta T_t$, so that trade union pressure on wages may be interpreted as partially the result of attempts by organized laborers to maintain real wages.

III. Further Results and Implications

In this Section additional evidence is examined for its implications concerning the effect of trade unions on wage inflation. First, Hines' [17] model of wage determination in the U.K. is examined for its post-sample predictive ability as compared with two well-known alternative models. An important reservation concerning the stability of Hines' estimated model is also noted. Second, the restricted and unrestricted reduced forms of equations (19)-(21) are investigated and estimates of the historical tradeoffs between $\Delta W_t$, $\Delta P_t$ and the exogenous variables are obtained.

A. Hines' Model of Wage Determination

Hines' preferred three-equation model of wage determination bears a strong resemblance to equations (19) - (21) above. A summary statement of his model is

\[
\begin{align*}
\Delta W_t &= f_1(\Delta T_t, \Delta P_t; T_t, \Delta P_{t-1}, X_{1t}) \\
\Delta P_t &= f_2(\Delta W_t; X_{2t}) \\
\Delta T_t &= f_3(\Delta P_t; T_{t-1}, X_{3t})
\end{align*}
\]  

(22)

where the disturbance terms have been suppressed; the $f_i$ are linear in their arguments; the $X_{it}$ are vectors of exogenous variables; and $\Delta T_t$ is defined as the arithmetic change in the percentage of the labor force unionized.
Hines' estimated version of (22) provides a very tight fit to the data and $\Delta T_t$ was found to be a highly significant explanatory variable. It seems plausible, therefore, to compare the post-sample predictive ability of (22) with some other well-known U.K. models of wage determination as a further test of the importance of trade unions in the wage inflation process.

Unfortunately, Hines' estimated version of (22) cannot be used for this purpose. Although it does not appear to be widely known, the 2SLS estimates of (22) imply that this system is explosive. In particular, it is well known that the reduced form of a structural model such as (22) is a system of three simultaneous difference equations and that the solution for each of the endogenous variables depends upon the characteristic roots ($\lambda_i$) of that part of the reduced form coefficient matrix relating to the lagged endogenous variables. Necessary and sufficient conditions for stability of such a system are that $|\lambda_i| < 1$ for all $i$. In the case of (22), there are two non-zero

---

\textsuperscript{20} Write (22) in the conventional form $AY_t = B_1 Y_{t-1} + B_2 X_t + \varepsilon_t$, where $A$ and $B_1$ are the $(3 \times 3)$ coefficient matrices of the endogenous and lagged endogenous variables; $B_2$ is the $(3 \times 3)$ coefficient matrix of the exogenous variables; $Y_t$, $Y_{t-1}$, and $X_t$ are the appropriately dimensioned observation vectors; and $\varepsilon_t$ a $t'(3 \times 1)$ vector of disturbance terms. Then the $\lambda_i$ are the characteristic roots of $A^{-1}B_1 = \pi_i$. Further, if $|\lambda_i| > 1$ for some $i$ then the final form of a model like (22) is undefined and since the sample moments of the predetermined variables need not converge to their population moments even the large-sample properties of simultaneous equation estimators applied to (22) are unknown. See, for example, Goldberger [14].
roots and these are readily estimated to be $\lambda_1 = 7.55$ and $\lambda_2 = .80$, so that this estimated system is explosive.\(^{29/}\) Under these conditions it is not clear how Hines' estimated version of (22) should be interpreted and it seems unlikely that the estimated reduced form of (22) would provide meaningful predictions.\(^{30/}\)

An alternative scheme is to discard Hines' complete system and work with the single (OLS) equation determining $\Delta W_t$. Some results on the post-sample predictions of this latter equation as compared with two other well known studies—R. G. Lipsey's [24] and L. A. Dicks-Hireaux's [10]—are contained in Table V.\(^{31/}\)

\(^{29/}\) It is possible, following Theil and Boot [29], to compute the estimated asymptotic standard error of $|\lambda_1|$ and hence the appropriate large sample test statistic for the hypothesis $|\lambda_1| < 1$. With $\lambda_1$ as large as 7.55 this did not seem necessary. It should be noted in passing that the estimated system (19) - (21) of Section II above does not suffer from this difficulty. On the contrary, it is highly damped, with $\hat{\lambda}_1 = .22$ and $\hat{\lambda}_2 = -.02$ from the 3SLS estimates.

\(^{30/}\) Indeed, Hines' 2SLS estimates of (22) were solved for the reduced form and the derived predictions were ridiculous.

\(^{31/}\) Hines' predictions are obtained from the OLS estimated equation

$$\Delta W_t = .7445 + 1.511\Delta T_t + .0639 T_{ct} + .6199\Delta P_t -.0409 \Delta P_{t-1} -.1243UN_t,$$

which was applied to the 1921-1961 peacetime period. Lipsey's predictions are obtained from the OLS estimated equation

$$\Delta W_t = .74 + .43UN^{-1}_t + 11.18UN^{-4}_t + .03\Delta UN_t + .62\Delta P_t,$$

which was applied to the 1923-57 peacetime period. Finally Dicks-Hireaux's predictions are obtained from the OLS estimated equation

$$\Delta W_t = 3.72 + .39\Delta P_t + .14\Delta P_{t-1} + 2.44\Delta D_{t-1/4},$$

which was applied to the 1946-1959 period and where $D_t$ is an index of the demand for labor. Note that the definitions of the variables in these British studies differ both from those in the analysis of American data in this paper and from one another.
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Root Mean Square Error (RMSE)</td>
<td>.396</td>
<td>1.138</td>
<td>1.498</td>
<td>1.918</td>
<td>1.716</td>
</tr>
<tr>
<td>Theil's Inequality Coefficient</td>
<td>.079</td>
<td>.221</td>
<td>.312</td>
<td>.350</td>
<td>.307</td>
</tr>
</tbody>
</table>
Recall that both Lipsey's and Dicks-Mireaux's models contain variables measuring $U_{it}$ and $\Delta p_{t}$, while only Vines' contains these and $\Delta T_{t}$. Both the root mean square prediction error (RMSE) and Theil's (revised) inequality coefficient (U) are presented in Table V. The results of this comparison are striking. The RMSE from Vines' equation is about one-third of what it is for Lipsey's and about one-fifth of what it is for Dicks-Mireaux's equation. This evidence clearly supports the hypothesis that trade unions have had an impact on British wage determination and may be interpreted as further evidence in support of the hypotheses advanced in Section I.

B. Reduced Forms of the Basic Model

Table VI contains two sets of estimates of the reduced form equations from (19) - (21) which explain $\Delta U_{t}$ and $\Delta p_{t}$. The first are unrestricted reduced form estimates (OLSRE), i.e. those obtained by direct application of OLS to the exogenous variables, and the second are the "solved" 3SLS estimates. Since it is of some interest to test hypotheses concerning the reduced form parameters, estimated (asymptotic) standard errors of the estimated coefficients are enclosed in parentheses beneath them. In addition, Table VI

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32/ The RMSE is simply $\left[ \frac{1}{n} \sum (P_{i} - \hat{P}_{i})^2 \right]^{1/2}$, where the $P_{i}$ are the predicted and the $\hat{P}_{i}$ the actual values. Clearly accurate predictions give rise to a small RMSE. U is defined as $\text{RMSE} / (\Sigma_{i}^{2}/n)^{1/2}$ so that when $U = 0$ the predictions are perfectly accurate and when $U = 1$ the prediction procedure leads to the same RMSE as naive no-change extrapolation. This inequality coefficient differs from that used by Theil in [37] and his preference for the measure applied here is expressed in [38], especially Chapter II.

33/ The procedure for calculating the standard errors of the reduced form coefficients from a structural model is worked out by Goldberger, Nagar, and Odeh [15]. The standard error of estimate (SEE) for the solved 3SLS structure is estimated by inserting the 3SLS estimates of $A$ and $\Sigma$ into $\Sigma_{e} = A^{-1} \Sigma (A^{-1})'$, where $\Sigma = E(\epsilon_{t} \epsilon_{t}')$ using the notation of footnote 29. The SEE is then $(\Sigma_{e})^{1/2}$. 
Table VI
Estimated Reduced Form Equations Determining
the Percentage Rates of Change of Money, Wages and Prices

<table>
<thead>
<tr>
<th>Endogenous Variable</th>
<th>Estimating Technique</th>
<th>Constant</th>
<th>Nt-1</th>
<th>St</th>
<th>Xt-1</th>
<th>Xe</th>
<th>Xe-1</th>
<th>Ut</th>
<th>(U/e) E-1</th>
<th>D</th>
<th>R²</th>
<th>D-W</th>
<th>SEE</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta N_t$</td>
<td>OLSRF</td>
<td>-6.424</td>
<td>.382</td>
<td>1.373*</td>
<td>.167</td>
<td>-.109</td>
<td>.373*</td>
<td>.161</td>
<td>-.130</td>
<td>.056</td>
<td>.123</td>
<td>16.607*</td>
<td>.76</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(4.099)</td>
<td>(.216)</td>
<td>(.369)</td>
<td>(.129)</td>
<td>(.225)</td>
<td>(.122)</td>
<td>(.139)</td>
<td>(.135)</td>
<td>(.052)</td>
<td>(.093)</td>
<td>(.50)</td>
<td></td>
</tr>
<tr>
<td>$\Delta Y_t$</td>
<td>Solv SLS</td>
<td>-5.597</td>
<td>-.944*</td>
<td>1.099*</td>
<td>.231</td>
<td>-.606</td>
<td>.241</td>
<td>.031</td>
<td>.202</td>
<td>-.021</td>
<td>.223*</td>
<td>24.104*</td>
<td>----</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(5.058)</td>
<td>(.347)</td>
<td>(.906)</td>
<td>(.161)</td>
<td>(.357)</td>
<td>(.128)</td>
<td>(.084)</td>
<td>(.137)</td>
<td>(.031)</td>
<td>(.108)</td>
<td>(.913)</td>
<td></td>
</tr>
<tr>
<td>$\Delta P_t$</td>
<td>OLSRF</td>
<td>-3.002</td>
<td>-.328*</td>
<td>.792*</td>
<td>.385*</td>
<td>-.600*</td>
<td>.004*</td>
<td>-.032</td>
<td>-.096</td>
<td>.032</td>
<td>.058</td>
<td>9.569*</td>
<td>.818</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(2.704)</td>
<td>(.147)</td>
<td>(.251)</td>
<td>(.089)</td>
<td>(.194)</td>
<td>(.083)</td>
<td>(.095)</td>
<td>(.106)</td>
<td>(.035)</td>
<td>(.063)</td>
<td>(3.831)</td>
<td></td>
</tr>
<tr>
<td>$\Delta P_t$</td>
<td>Solv SLS</td>
<td>-5.486</td>
<td>-.689*</td>
<td>.803*</td>
<td>.323*</td>
<td>-.866*</td>
<td>.176</td>
<td>.022</td>
<td>.147</td>
<td>-.016</td>
<td>.163*</td>
<td>17.396*</td>
<td>----</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(3.774)</td>
<td>(.265)</td>
<td>(.377)</td>
<td>(.121)</td>
<td>(.299)</td>
<td>(.095)</td>
<td>(.062)</td>
<td>(.101)</td>
<td>(.023)</td>
<td>(.081)</td>
<td>(6.941)</td>
<td></td>
</tr>
</tbody>
</table>

$R^2 = .68$
$c^2 = .82$
contains Hooper's trace correlation coefficient squared ($\tau^2$) as estimated for the OLSRF and the largest canonical correlation coefficient squared ($\gamma^2$). See Hooper [20]. $\gamma^2$ is analogous to the conventional $R^2$ and may be interpreted as the fraction of the variance (in a linear combination) of the endogenous variables explained by (a linear combination of) the exogenous variables.

1. Following the methodology suggested by Christ [8], it is useful to compare the OLSRF and solved 3SLS estimates of the reduced form for major discrepancies. On the whole, these are not large. Whenever a reduced form coefficient may be judged significantly different from zero the differences between the two estimates of it are typically small. For example, the coefficient of $S_{et}$ is significant in each equation and differs little on account of estimation procedure. One important exception is the case of the unemployment rate. In particular, solved 3SLS estimate of the implied "Phillip's Curve" is much steeper than the OLSRF estimate. Finally, since the equation determining $\Delta W_t$ contains $\Delta X_t$ and $\Delta W_{t-1}$ it may be interpreted as a test of Kuh's productivity-distributed-lag (PDL) model [21]. The PDL model predicts positive coefficients on $\Delta X_t$ and $\Delta W_{t-1}$. Note that the coefficient of $\Delta X_t$ has the wrong sign for the PDL interpretation, although neither the coefficient of $\Delta X_t$ or $\Delta W_{t-1}$ is significant.

2. The reduced form equations for $\Delta P_t$ and $\Delta P_t$ in Table VI may be solved recursively to obtain the historical steady-state relationship between $\Delta P_t$ and the exogenous variables.34/ Substituting the 1950-1960 mean values of

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34/ In the notation of footnote 26, this matrix of coefficients is obtained as $\bar{\pi} = (I - \pi_1)^{-1} \pi_2$ where $\pi_2 = \lambda^{-1} E'$. 


all the exogenous variables except $U_{Nt}$ and $\Delta X_t$ into this relationship gives the family of Phillips curves defined by

$$\Delta P = 10.624 - 1.092UN - 1.104\Delta X,$$

using the solved 3SLS estimates of the reduced form. With $\Delta X = 3\%$ and $\overline{UN} = 4\%$ the estimate from (23) of $\Delta P$ is about $3\%$ per annum which does not seem unreasonable. For reasons noted in the next section, equation (23) should not be taken very seriously, but it is interesting to compare it with the results of other investigators. First, Bodkin [5] estimated the following with 2SLS for the same period:

$$\Delta P = 2.10 - .11\overline{UN}.$$

(24) is much flatter than (23) and implies virtually no tradeoff between $\Delta P$ and $\overline{UN}$. For example, (24) implies $\Delta P = 0.0$ with $\overline{UN}$ at about $19\%$, while (23) implies $\Delta P = 0.0$ with $\overline{UN}$ at about $7\%$ (on Bodkin's assumption that $\Delta X = 2.5\%$).

Equation (23) seems more reasonable than (24). Second, $\frac{\Delta P}{\overline{UN}} = -1.09$ from (23) may be compared with Perry's estimate [29] of $\frac{\Delta P}{\overline{UN}}$ for the post-war period. The latter ranges from $-0.65$ for $\overline{UN} = 6.0\%$ to $-1.45$ for $\overline{UN} = 4.0\%$ and appears consistent with the estimate from (23).

IV. Conclusions and Limitations

A number of formal hypotheses about the way in which trade unions may affect the aggregate rate of change of money wages have been advanced. These hypotheses appear to be consistent with U.S. data over the period 1914-59 and with British data over a similar period. In order to arrive at these results, however, a number of assumptions are required. Consequently, it seems appropriate to note their limitations:

1. In order to test hypotheses about the affects of trade unions on money wages it is necessary to specify a set of proxy variables measuring trade
union strength and activity. Aside from the fact that alternative proxies can always be specified, there is the further problem that these variables probably do not maintain an invariant relationship with the underlying forces they are presumed to measure. This implies that the parameters of equations like (19) - (21) are probably not stable through time.

2. Although the evidence supports the hypothesis that trade unions have been a causal factor in the determination of aggregate money wages, this does not imply that the long-run historical course of money wages would have been different without the presence of trade unions. In particular, the evidence in Section II may be interpreted to imply that the pressure of trade unions on wages is in part a response to the downward pressure on the rate of change of real wages brought about by price increases. On this interpretation trade unions take advantage of outward shifts in labor demand schedules before these shifts can result in the excess demand forces which would eventually pull money wages up. From this point of view trade unions make markets more rather than less perfect.

3. Finally, it should be clear from the standard errors of estimate of the $\Delta P_t$ and $\Delta W_t$ reduced form equations presented in Table VI that the estimated versions of equations (19) - (21) do not have any specific implications for policy in the post-War period. In particular, the estimates of equations (19) - (21) or their reduced forms are not useful tools for either policy or actual forecasts of $\Delta W_t$ and $\Delta P_t$ in the post-War period. To see this note that from Table VI the lower bounds on the asymptotic standard errors of forecasts
from the solved reduced forms \( \Delta M_t \) and \( \Delta P_t \) are estimated as 7.7 and 6.1 percentage points respectively.\(^{34/}\) Given that the standard deviations of \( \Delta M_t \) and \( \Delta P_t \) are on the order of only 3.5% and 4.3% for the post-war period, the estimated reduced form forecasts from equations (19) - (21) would have associated confidence intervals too wide for any practical use.

\(^{34/}\) The asymptotic forecast variance-covariance matrix is \( \phi = \Lambda + A^{-1} \Sigma (A^{-1})^T \), where \( \Lambda \) has positive diagonal elements which depend on the values of the exogenous variables. Since the square roots of the estimated diagonal elements of \( A^{-1} \Sigma (A^{-1})^T \) are listed as SEE in Table VI these serve as the required lower bounds.
References


References (Continued)


References (Continued)


Appendix

Definitions

\[ \Delta W_t = \frac{W_t - W_{t-1}}{W_{t-1}} \cdot 0.100 \] = the percentage rate of change of money wages

\[ \Delta P_t = \frac{P_t - P_{t-1}}{P_{t-1}} \cdot 0.100 \] = the percentage rate of change of the consumer price index

\[ T_t = \frac{1}{2} (T_t^* + T_{t+1}^*) \] = the level of trade union membership where \( T_t^* \) is the membership of unions given by L. Troy in the reference cited below


\[ \Delta T_t = \frac{T_t - T_{t-1}}{T_{t-1}} \cdot 0.100 \] = the percentage rate of change in union membership

\[ \Delta X_t = \frac{X_t - X_{t-1}}{X_{t-1}} \cdot 0.100 \] = the percentage rate of change in average productivity

\[ U_t \] = unemployment as a percentage of the civilian labor force

\[ \Delta E_t = \frac{E_t - E_{t-1}}{E_{t-1}} \cdot 0.100 \] = the percentage rate of change of employment in the "unionizable" sectors of the economy, identified as manufacturing, mining, construction, and transport and utilities

\[ \frac{T_t}{E_t} \cdot 0.100 \] = trade union membership as a percentage of "unionizable" employment

\[ S_t \] = the number of work stoppages

\[ D_t \] = the percentage of the major party membership in the House of Representatives affiliated to the National Democratic Party

\[ u_t^P \] = unemployment as a percentage of the civilian labor force measured at the pit of the preceding recession where the dating of the trough is taken from the NBER's reference cycles.
Sources


Democratic Party members: U. S. Bureau of the Census, Ibid., Table Y139-140, p. 691.