A Note on Estimating the Determinants of Changes in Wages and Earnings

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The practical importance for public policy of the determinants of aggregate money wage changes has spawned an extensive empirical literature on this subject in the last decade.\(^1\) Since little of empirical significance seems to have been settled by this literature, it is perhaps not surprising that it has started to receive substantial formal and informal criticism. On the one hand, there has been criticism of the weak connection between economic theories and empirical analyses of the determinants of money wage inflation.\(^2\) On the other hand, there is now a sizeable body of criticism of the econometric methodology that has been used in most empirical studies.

A stylized description of most of this empirical work is as follows: A relationship is fitted to the data with great apparent statistical reliability, as discerned from the precision of estimated parameters and from overall goodness of fit statistics. At the same time, however, when these estimated relationships are used to forecast outside the sample period they tend to be unreliable.\(^3\) Moreover, tests of stability on the

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\(^{1}\) For a recent survey of this literature, see Goldstein [10].

\(^{2}\) See, for instance, the remarks by Oi [16] or Leijonhufvud [12].

\(^{3}\) Penceavel [18] reports the results of such an experiment with some British models, for example.
structure of fitted relationships within the sample period tend to show them to be highly unstable.\textsuperscript{4} One suspects on the basis of these latter results that the implied statistical reliability of these relationships is illusory and that the statistical precision reported in most analyses has lead to an overstatement of the quality of the empirical information that actually exists on the determinants of the wage inflation process.

In this note we report the results of an empirical investigation of quarterly aggregate money wage changes of British manual workers where we pay fairly careful attention to matters of econometric specification and measurement. In particular, we take up two issues of measurement over which there is some contention. First, we examine the relationship between actual earnings changes and changes in the basic weekly wage. Since weekly earnings include all actual money payments for work during a week, while the basic wage is essentially the "normal payment" for a work week of a "normal" number of hours, the former is more nearly a measure of labor costs in production than is the latter. Moreover, during the period 1950 to 1970, earnings per week have increased in Britain at about 6.2 percent per year while the basic weekly wage has increased at about 4.7 percent per year.\textsuperscript{5} Thus, the ratio of weekly earnings to the basic weekly wage has crept up by about 1.5 percent per year. At the same time, virtually all systematic study of the wage inflation process has used changes in the basic hourly wage (defined as the basic weekly wage divided by an index of normal weekly hours) as the variable to be studied, presumably because these data have been available each quarter while measures of weekly earnings

\textsuperscript{4} See Lipsey and Parkin [13], or Oi [16], for example.

\textsuperscript{5} These are calculated as the means of the semi-annual first differences in the logarithms of earnings and wages, respectively, and inflated to annual rates by multiplication by two.
have been available only twice each year. In the first part of this paper we study the legitimacy of this procedure and inquire as to the causes of the observed increase in the ratio of weekly earnings to weekly wages.

Second, we examine the appropriate formulation of the rate of change of the aggregate money wage to be used as the dependent variable. In one of the earliest systematic attempts to estimate the determinants of quarterly money wage changes, Dicks-Mireaux and Dow [6] observed that the fraction of workers receiving a wage adjustment varied substantially from quarter to quarter. Consequently, if there existed a stable relationship between (a) the size of the wage changes of those workers who received adjustments and (b) some set of explanatory variables, there might be no stable relationship between changes in an aggregate wage index and the same explanatory variables. To surmount this measurement problem while still maintaining the existence of a stable relationship between the wage changes of those workers receiving adjustments and the explanatory variables of interest, Dow and Dicks-Mireaux worked with the assumption that each worker received a wage adjustment approximately once every four quarters. This implies a simple and regular functional relationship between the proportion of workers receiving wage adjustments and calendar time, where the former is assumed to take on precisely four values in a regular sequence throughout the sample period. Under this assumption the four-quarter change in the aggregate wage index can be expressed as a function of the moving four-quarter averages of the explanatory variables with weights equal to the (known) proportion of workers receiving adjustments in a given calendar quarter. Of course, if the error terms in the initial underlying equation were serially uncorrelated, the error terms in the
moving average version of this equation would not be. Unfortunately, since the four necessary values for the proportion of workers receiving adjustments was unknown, Dicks-Mireaux and Dow added the further assumption that one-fourth of all workers received wage adjustments each quarter. Under these conditions, however, the underlying equation for wage adjustments in any given quarter is a mirror image of the equations for the other three quarters and the rationale for the use of the four-quarter change in the dependent variable, as opposed to the one-quarter change, must rest in the properties of the error terms in the one formulation as compared with the properties of the error terms in the other. Following on the heels of Dicks-Mireaux and Dow, Perry [19] adopted the use of the four-quarter change in the aggregate wage index as the dependent variable and virtually all subsequent work did likewise in spite of Sargan's [23] cautionary arguments. 6/

In this paper we report the results of tackling this question in a more direct manner by including explicitly in the estimation procedure the available information on the proportion of workers receiving wage adjustments in a given quarter. The advantage of this procedure is that it puts the basic estimating equations in a form where the various assertions about the properties of the error terms of these equations can be put to a test.

I. Earnings and Wages

In this section, we consider the relationship between, on the one hand, changes in the actual weekly earnings of manual workers and, on

6/ See also the same arguments by Black and Kelejian [3] and, more recently, by Rowley and Wilton [22] in a series of papers and the references therein.
the other hand, changes in the weekly wage rates of these workers as set down in collective bargaining agreements or as determined by Wages Councils.

In understanding this relationship and in structuring the problem, we build upon the evidence from previous industry studies which have argued persuasively that it is appropriate to distinguish those workers paid on time rates from those on some system of payments-by-results.\(^1\) Since separate data on these two groups of workers are not available for the period under investigation,\(^2\) we first examine the determinants of time-rate earnings, then the relationship between time-rate earnings and the earnings of workers on payment-by-results, and finally their combined effect on aggregate earnings.

Consider first the weekly earnings of a time-rate worker averaged over a given period of time \(t\), \(E^T_t\). If \(W_t\) is his basic weekly wage, \(C_t\) his actual hours worked, and \(H_t\) the number of hours worked before overtime rates apply (in other words, "normal" weekly hours), then the

\(^1\)Some of these studies are listed and their arguments summarized in R. H. Phelps Brown's [4] survey.

\(^2\)From the end of the Second World War until April 1961, the Ministry of Labour undertook enquiries every two years into the extent of payment-by-results methods. These data reveal that, from approximately the beginning of our estimating period (in October 1951 to be exact) to the last survey in April 1961, the fraction of wage-earners who received at least part of their pay according to some sort of incentive payments scheme changed imperceptibly from 32 to 33 percent. (These figures relate to manual workers employed in the same industries as those covered by the earnings data below, namely, manufacturing, construction, transport, public utilities, and mining other than coal.) The only comparable study since 1961 was conducted by the National Board for Prices and Incomes [15] in December 1967 from which they conclude ",.....there is no clear evidence of any overall movement to, or away from, payments-by-results since the Ministry of Labour's 1961 enquiry" (page 8). Of course, these figures are aggregated over different industries and different classes of workers and this stability in the percent of all workers on some form of payment-by-results conceals some interesting variation over time in particular sub-categories. See the discussion in Chapter 2 of [15].
relation

\[ E_t^n = \mathcal{W}_t + \sigma(\mathcal{W}_t/H_t)(C_t - H_t) = \mathcal{W}_t[1 + \sigma(C_t - H_t)/H_t] \]

holds as an identity where \( \sigma \) is the overtime premium ratio.\(^2/\) Hence, \( \sigma(\mathcal{W}_t/H_t) \) is the hourly rate for overtime hours and \( C_t - H_t \) measures the number of overtime hours worked. Since \( \sigma(C_t - H_t)/H_t \) is typically less than 0.20,\(^10/\) \( \ln[1 + \sigma(C_t - H_t)/H_t] \approx \sigma(C_t - H_t)/H_t \) and the first difference in the (natural) logarithm of (1) is

\[ \Delta \ln E_t^n = \Delta \ln \mathcal{W}_t + \sigma \Delta (C_t - H_t)/H_t \]

where we have assumed that the overtime premium ratio (\( \sigma \)) has not changed.\(^11/\)

According to equation (2), the proportionate change in the weekly earnings of time workers (\( \Delta \ln E_t^n \)) is the sum of two components: the first is the proportionate change in the basic weekly wage (\( \Delta \ln \mathcal{W}_t \)) and the second is the overtime premium ratio multiplied by the change in the ratio of overtime to normal hours. Alternatively, equation (2) says that the proportionate change in the ratio of weekly earnings to wages is \( \sigma \Delta (C_t - H_t)/H_t \).

\(^2/\)Of course, \( \sigma \) is not the same for all categories of overtime hours. For instance, overtime hours worked during a weekday are normally paid at time-and-a-quarter, time-and-one third, or time-and-one-half while for almost all Sunday work the rate is double time. (A recent listing of \( \sigma \) across different industries and at different hours as determined by collective bargaining agreements or Wages Councils is contained in Appendix II of [5]). Unfortunately, the data available to us do not distinguish between these different categories of overtime hours so we are obliged to work with a single value for \( \sigma \).

\(^10/\) Suppose overtime is worked at the rate of time-and-one-half (\( \sigma = 1.5 \)). At the aggregate level \( (C_t - H_t)/H_t \) never exceeds 0.10 in which case \( \sigma(C_t - H_t)/H_t \leq 0.15 \).

\(^11/\) Although much has been written concerning changes in the hours of overtime worked in the last 25 years, there has been very little work on the general movements of \( \sigma \) over this time so that it is difficult to determine the appropriateness of our assumption about zero change in \( \sigma \). (Issues of the Ministry of Labour's Time Rates of Wages and Hours of Labour in the earlier part of our estimating period do not contain the information on \( \sigma \) that the more recent issues have.) For example, consulting Whybrow's [24] comprehensive study of overtime work we could find no reference to this issue.
In fact, the ratio of overtime to normal hours for all workers has increased steadily over the postwar period in Britain from around zero in 1950 to approximately 0.10 in 1970. Applied to time-rate workers and assuming time-and-one-half for overtime ($\sigma = 1.5$), this alone implies from equation (2) that the ratio of weekly earnings to wages would increase by about 15 percent from 1950 to 1970 for this group.

Consider now the weekly earnings of payment-by-results workers, $E_t^R$. As a first approximation, for an individual working on some form of payment-by-results system, let $E_t^R = \Pi_t Q_t$ where $\Pi_t$ stands for his remuneration per unit of output produced (the "piece-rate") and $Q_t$ his weekly output.\(^{12}\)

The piece-rate, $\Pi$, will change over time and we must enquire how we should expect it to vary. Suppose that, at the time the piece-rate is set for period $t$, it is expected that a typical worker will be able to produce $Q_t^*$ units of output. Then $\Pi_t$ will be determined so that $\Pi_t Q_t^*$, expected future earnings, will be closely related to $W_t$ so that we may write the behavioral relation

\[
(3) \quad \Pi_t Q_t^* = kW_t
\]

where $k$ is a constant. The behavioral content of equation (3) may be rationalized in two (not necessarily exclusive) ways. On the one hand, we may conceive of a schedule relating the supply of workers to the payments-by-results sector relative to the time-rate sector as a function of the

\(^{12}\) The approximation here arises from our specification of a pure piece-rate system of remuneration whereas the pay of most workers in the payments-by-results sector is computed in a far more complex fashion. Marriott's [14] study may be consulted for a thorough description of the wide variety of different methods of wage payment. Although there will be aspects of labor market behavior in which it is appropriate to discriminate among these different incentive payments schemes, such distinctions cannot usefully be made with these data at the aggregate level. Of course, some finer distinctions among wage payments methods may, indeed, yield useful insights at a dis-aggregated level.
rewards in the one sector compared with the other. Equation (3) is then equivalent to the assertion that the relative supply of workers to payment-by-results employment is perfectly elastic at a relative return of $k$.  

On the other hand, we may think of collective bargaining or Wages Councils establishing a value of $\Pi_t$ that preserves a "traditional" differential in pay packets between payment-by-results and time-rate workers. In this case, equation (3) is equivalent to the proposition that the "customary" wage differential is $100 \times (k-1)$ percent. In either case, substituting $\Pi_t = \kappa W_t / Q_t^s$ from equation (3) into the definition $E_t^R = \Pi_t Q_t$ and taking the first differences in logarithms gives

$$
\Delta \ln E_t^R = \Delta \ln W_t - \Delta \ln(Q_t / Q_t^s).
$$

According to equation (4), the proportionate change in the weekly earnings of payment-by-results workers ($\Delta \ln E_t^R$) will equal the sum of the proportionate change in the basic weekly wage plus the discrepancy between the change in actual and expected output per man.

---

13/ If the relative supply of payment-by-results workers were not perfectly elastic, then $k$ would be a function of the ratio of payment-by-results to time-rate workers. As long as this ratio remained constant, however, $k$ would also remain constant and there would be no operational significance to the assumption made in the text. In fact, as noted in footnote 8 above, the ratio of payment-by-results to time-rate employment among manual workers appears to have remained very stable over our estimating period. Of course, given our observations on relative employment, an equivalent operational proposition is not that the relative supply curve, but that the relative demand curve is horizontal at $k$. While this may be an appealing characterization within a particular industry, it strikes us as a less convincing description at a more aggregate level.

14/ On this interpretation, abundant evidence of a constant wage differential between payment-by-results and time-rate workers over the last 25 years is supplied by the minimum wage rates determined by collective bargaining or Wages Councils. For instance, the Ministry of Labour's Time Rates of Wages and Hours of Labour for October 1949 informs us that, for pieceworkers in the home grown timber trade, "piecework rates are to be such as to enable (cont'd. next page)
An index $E_t$ of the level of weekly earnings of all manual workers may be taken to be a geometric weighted average of the earnings of the time-rate and the payment-by-results workers so that $\Delta \ln E_t = R \Delta \ln E^R_t + (1 - R) \Delta \ln E^T_t$, where $R$ measures the fraction of all workers in the payment-by-results sector. Since it appears that at an aggregate level $R$ has changed imperceptibly in the two decades from 1950 to 1970, the proportionate change in $E_t$ may be written

$$\Delta \ln E_t = R \Delta \ln E^R_t + (1 - R) \Delta \ln E^T_t.$$  

The only unobservable component of either equations (2) or (4) is $\Delta \ln Q^u$. Suppose that these expectations follow a simple trend given by $Q^*_t = Q^*_0 e^{mt}$ in which case $\Delta \ln Q^*_t = m$, a constant. With this assumption, after the substitution of equations (2) and (4) in equation (5), we may write

$$\Delta \ln E_t = \Delta \ln W_t + \sigma(1 - R)\Delta[(C - H)/\bar{H}] + R \Delta \ln Q_t - Rm. \tag{5a}$$

Before we turn to the results of fitting equation (5a) to the data, it is useful to examine one possible objection to the assumptions underlying its development. In particular, (5a) concerns the relationship between weekly earnings and the weekly wage, yet it might be argued that interest centers more on the relationship between hourly earnings and the hourly wage. Correspondingly, it may be more appropriate to suppose that in equation (3) the piece-rate is determined so that the ratio of the earnings per hour of payment-by-results workers to that of time-rate workers is a constant, $k$. To allow for this specification, put the identity in equation (1) on a hourly basis by dividing through by $C_t$:

a worker of average ability to earn at least 25 percent above the appropriate time rates." The Department of Employment's equivalent publication in April 1973 [5] specifies the very same differential for these workers. This is just one of numerous examples that could be offered.
\[
\begin{align*}
(1a) \quad (E^T/C)_t &= (W/H)_t [(H/C)_t + \sigma (H/C)_t (C - H)_t / H] \\
&= (W/H)_t [1 + (\sigma - 1)(C - H)_t / C_t].
\end{align*}
\]

Since \((\sigma - 1)(C - H)/C\) will normally be a small number, we may replace equation (2) with
\[
(2a) \quad \Delta \ln(E^T/C)_t = \Delta \ln(W/H)_t + (\sigma - 1) \Delta [(C - H)/C]_t.
\]

Analogously, suppose that \(E_t\) is now set such that
\[
(3e) \quad \eta_t(Q^*/C_t) = \kappa(W_t/H_t),
\]

implying that the hourly earnings of payment-by-results workers are proportional to the normal hourly wage of time-rate workers. After the appropriate substitutions, the specification analogous to equation (5a) is
\[
(5b) \quad \Delta \ln(E/C)_t = \Delta \ln(W/H)_t + (1 - r)(\sigma - 1) \Delta [(C - H)/C]_t
\]
\[+ R \Delta \ln(Q/C)_t - m'R,
\]

where \(m'\) is derived from the assumption that \((Q^*/C)_t = (Q^*/C)_0 e^{m't}\).

The differences between equations (5a) and (5b) are obviously small.\(^{15}\)

Line 1 of Table 1 contains least squares estimates of the stochastic form of equation (5a) without the imposition of the constraint that the coefficient of \(\Delta \ln W_t\) equals unity. Precise definitions of variables and their sources are contained in the note to the Table. As others have found,\(^{16}\) the estimated coefficient of the variable \(\Delta \ln W_t\) is substantially

\(^{15}\)Note that in equation (3a) we have assumed that the hourly earnings of payment-by-results workers are proportional to normal hourly earnings \((W/H)\) and not to actual hourly earnings \(E^T/C\). If we were to assume the latter, equation (5b) would become \(\Delta \ln(E/C)_t = \Delta \ln(W/H)_t + (\sigma - 1) \Delta [(C - H)/C]_t
\]
\[+ R \Delta \ln(Q/C)_t - m'R.\] This differs from equation (5b) only in the interpretation of the coefficient on \(\Delta [(C - H)/C]_t\).

\(^{16}\)See the results in Gillion [7], Godley and Nordhaus [8] and Pesaran [20].
less than unity, casting some doubt on the validity of the arguments that underly the development of equation (5a). Indeed, if it were correct, this result would have the disturbing implication that the ratio of weekly earnings to the weekly wage declines by over 4 percent when the weekly wage increases by 10 percent. This can hardly be the case for an indefinite period without the ratio of earnings to wages falling below unity, which seems unlikely. A more satisfactory explanation for the small coefficient on $\Delta \ln W_t$ runs in terms of simultaneity bias. This bias may arise because the equilibrium relationship (3) between the weekly pay packet of time-rate workers and the weekly earnings of payment-by-results workers implies that weekly wages and earnings are determined jointly. That is, the equilibrium relationship (4) does not imply that $\Delta \ln W_t$ is exogenous and that $\Delta \ln E_t^R$ adjusts to changes in it. Indeed, it is often argued that the determination of time-rates follows on the heels of the setting of payment-by-results rates. Consequently, we may confuse the causation running from $\Delta \ln W_t$ to $\Delta \ln E_t$ with the causation running in the opposite direction.

One way to examine the nature of this difficulty is to recognize explicitly that earnings and wage changes may be jointly determined in a given half-yearly period. The results in line 2 of Table 1 represent an effort to do this by treating $\Delta \ln W_t$ as a right-hand endogenous variable in the estimation of equation (5a). The exogenous variables excluded from (5a) are taken to be the reciprocal of the unemployment rate and the rate of change of prices, two variables taken to be determinants of $\Delta \ln W_t$ in many studies. As can be seen from the table, the two-stage least squares (2SLS) estimate of the coefficient of $\Delta \ln W_t$ is now nearly unity and is certainly not significantly different from unity. On the other hand,
## Table 1

Estimates of Equations (5a) and (5b) using Semi-Annual Data from April 1950 to October 1969

<table>
<thead>
<tr>
<th>Line No.</th>
<th>Dependent Variable</th>
<th>Estimating technique</th>
<th>( \Delta \ln \bar{E}_t )</th>
<th>( \Delta \ln \bar{H}_t )</th>
<th>( \Delta \ln (C-H)/H_t )</th>
<th>( \Delta \ln Q_t )</th>
<th>( \Delta \ln (W/H)_t )</th>
<th>( \Delta \ln [(C-H)/C]_t )</th>
<th>( \Delta \ln (Q/C)_t )</th>
<th>( U_{\text{t}}^{-1} )</th>
<th>SEE</th>
<th>DW</th>
<th>( R^2 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>( \Delta \ln E_t )</td>
<td>OLS</td>
<td>-0.144</td>
<td>-0.578</td>
<td>1.074</td>
<td>-0.018</td>
<td>-0.604</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>2</td>
<td>( \Delta \ln E_t )</td>
<td>2SLS</td>
<td>0.046</td>
<td>1.017</td>
<td>1.139</td>
<td>-0.041</td>
<td>1.039</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>3</td>
<td>( \Delta \ln E_t )</td>
<td>2SLS</td>
<td>0.050</td>
<td>1.017</td>
<td>1.139</td>
<td>-0.041</td>
<td>1.039</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>4</td>
<td>( \Delta \ln (E/C)_t )</td>
<td>OLS</td>
<td>-0.141</td>
<td>-0.578</td>
<td>1.074</td>
<td>-0.018</td>
<td>-0.604</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>5</td>
<td>( \Delta \ln (E/C)_t )</td>
<td>2SLS</td>
<td>0.047</td>
<td>1.017</td>
<td>1.139</td>
<td>-0.041</td>
<td>1.039</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>6</td>
<td>( \Delta \ln (E/C)_t )</td>
<td>2SLS</td>
<td>0.051</td>
<td>1.017</td>
<td>1.139</td>
<td>-0.041</td>
<td>1.039</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>7</td>
<td>( \Delta \ln E_t )</td>
<td>2SLS</td>
<td>0.071</td>
<td>1.117</td>
<td>1.109</td>
<td>-0.057</td>
<td>1.039</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>8</td>
<td>( \Delta \ln (E/C)_t )</td>
<td>2SLS</td>
<td>0.076</td>
<td>1.117</td>
<td>1.109</td>
<td>-0.057</td>
<td>1.039</td>
<td>1.039</td>
<td>-0.0074</td>
<td>-0.074</td>
<td>(0.770)</td>
<td>(2.46)</td>
<td>(0.36)</td>
</tr>
</tbody>
</table>

The equations were estimated with 40 semi-annual observations on each variable from the first half of 1950 to the last half of 1969. The earnings and hours (E) data used in these estimates are derived from the surveys of manual workers in manufacturing, construction, transport (except for railway and London Transport workers), public utilities, and mining and quarrying (other than coal). These surveys were conducted in April and October of each year by the Ministry of Labour and, later, by the Department of Employment and the results are published in various issues of its \textit{Census}. The wage rate (W) and normal weekly hours (H) series pertain to manual workers in the same industries plus coal mining, the distributive trades, and agriculture. Q is the index of industrial production and is taken from the \textit{Monthly Digest of Statistics}. The unemployment rate measures the total number of registered unemployed as a percent of total employees and the series is published in various issues of the \textit{Ministry of Labour}.
Gazette or, more recently, the Department of Employment Gazette. For $U_t$ in these regressions, we have used the average of the unemployment rate in period $t$ and in period $t-1$. For any variable, $Z_t$, 
$\Delta \ln Z_t = \ln Z_t - \ln Z_{t-1}$. The standard error of estimate for the regression equation is given by $\text{S.E.}$ and the Durbin-Watson statistic by $DW$. In the 2SLS (two-stage least squares) estimates, $\Delta \ln W_t$ is treated as endogenous where excluded exogenous variables are the logarithmic change in retail prices (from the Monthly Digest of Statistics), the reciprocal of the unemployment rate, and $\Delta \ln W_t$. In lines 3 and 6, the equations are estimated by restricted least-squares where the constraint takes the simple form of the coefficient on $\Delta \ln W_t$ being restricted to the value of unity; this constrained value is indicated by the asterisk. All of these data may be obtained from the authors upon request.
as often happens in such cases, the residual variance associated with the
2SLS estimates in line 2 is much larger than that associated with the
conventional least squares estimates in line 1, so that a choice between
the two sets of estimates must depend partly on non-statistical criteria.
In this case the results in line 2 seem more plausible than those in line 1
in view of the long run implications for the ratio of earnings to wages
implied by the latter. Line 3 consequently contains the results of
imposing the constraint that the coefficient of $\Delta \ln Y_t$ equals unity in
equation (5a). The estimate of $\sigma (1 - R)$ from line 3, which is the
coefficient of $\Delta [(C - H)/H]_t$, is around 1.13. If we take $R = .33$,
as it was in 1961, then $\sigma$ is estimated at 1.7, which falls comfortably
within the range of 1.25 to 2.0 at which most overtime hours are rated.11/
The coefficient of productivity changes ($\Delta \ln Q_\tau$) has an unexpected sign,
though it is not significantly different from zero by conventional criteria.
This might result either because our measure of $Q$ covers all workers,
rather than payment-by-results workers alone as is called for by equation (4),
which would give rise to measurement errors or because $\Delta \ln (Q_t/Q_\tau^*) \approx 0.0$
throughout most of this period. In the latter case, the actual and expected
output of payment-by-results workers would be moving together and hence
it would not be possible to measure the effects of $\Delta \ln (Q_t/Q_\tau^*)$ on earnings.
This difficulty is probably made even more serious in view of our especially
naive assumption about the time path of $Q^*$.

17/ Even so, it should be observed that our measures of $C$ and $H$ cover
all workers, though the arguments leading to equation (5a) imply that
they should be measured for time-rate workers only. Hence our estimate
of the aggregate value of $\sigma$ might be biased by measurement error.
Lines 4, 5, and 6 in Table 1 are the estimates of the stochastic form of equation (5b) that correspond to the estimates of equation (5a) in lines 1, 2, and 3. As can be seen from the table, the former are nearly identical to the latter after due account is taken of the relationship between the parameters in these two equations. Finally, in lines (7) and (8) we report the results of adding the reciprocal of the unemployment rate to equations (5a) and (5b) to test for any effect of measured unemployment on earnings changes that might operate independently of wage changes. The results of adding $U_t^{-1}$ to these equations were independent of the estimation procedure and so we supply only the 2SLS results. As can be seen from the table, the unemployment ratio has neither a statistically significant nor quantitatively important independent effect on earnings.

In short, while earnings per week increased at about 6.2 percent per year over the period 1950 through 1970, about 75 percent [=(4.7/6.2)100] of this can be accounted for by a stable equilibrium relationship with the growth in the basic wage rate. In addition, the ratio of overtime hours to normal weekly hours has been growing by about .5 percent per year,\textsuperscript{18} contributing another 10 percent [= 100(1.13) .5/6.2] to the explanation of the growth in average annual earnings over this period. The remaining per annum growth in average earnings is unexplained.

II. The Frequency of Wage Settlements

Suppose that in a given quarter of a year we divide manual workers into two groups: those who receive adjustments to their wages in that quarter and those who do not. If we write $\lambda_t$ for the fraction of all workers receiving adjustments, it is curious, of course, that this growth in overtime has come about almost entirely because normal hours of work have dropped steadily over the 1950 to 1960 period, while actual hours of work have remained constant. This invites an explanation for the increase in overtime hours in terms of a conscious effort to manipulate average hourly earnings, but we do not pursue this issue.
workers who receive adjustments in a given quarter \( t \), \( (W/H)^A_t \) for the hourly wage of those workers who receive wage adjustments, \( (W/H)^N_t \) for the hourly wage of those workers who do not receive adjustments, and take the aggregate hourly wage \( (W/H)_t \) to be a geometric weighted average of \( (W/H)^A_t \) and \( (W/H)^N_t \), then the approximate value of \( \ln(W/H)_t \) is 
\[
\ln(W/H)_t = \lambda_t \ln(W/H)^A_t + (1 - \lambda_t) \ln(W/H)^N_t. \tag{22}
\]
Since \( \Delta \ln(W/H)^N_t = 0 \) by construction, the change in \( \ln(W/H)_t \) is
\[
\Delta \ln(W/H)_t = \lambda_t \Delta \ln(W/H)^A_t. \tag{6}
\]

According to (6) the proportionate change in the aggregate hourly wage index is the product of the fraction of workers receiving wage adjustments in a quarter and the proportionate change in the wage that these workers receive. Since the value of \( \lambda_t \) is often known in advance of the period \( t \) and there is often informed speculation on the plausible value of \( \Delta \ln(W/H)^A_t \) in advance of the period \( t \), it is our impression that practical forecasts of the change in the aggregate wage index are often made with a relationship such as (6). 20/

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This expression is an approximation because the weights \( \lambda_t \) should be the share of each group's wages in the total wage bill, not the fraction that each group is in total employment. The quality of the approximation obviously depends on the extent to which \( \lambda_t [(W/H)^A/(W/H)_t] \) differs from \( \lambda_t \) or, alternatively, the extent to which \( (W/H)^A_t \) differs from \( (W/H)_t \). Obviously, it would be useful to explore this issue if the data were available.

20/ See Artis [1], for example.
How most workers have normally had their wage rates reviewed or re-negotiated on a regular basis, but not so frequently as once each quarter. Presumably the frequency with which wage rates are reviewed or negotiated is determined by the real personnel and other costs associated with these activities. Although very little is known about the nature and determinants of these costs, they apparently result in what are sometimes substantial periods between wage adjustments. As a consequence, most theoretical analyses of the determinants of wage changes have been taken to refer to the determinants of $\Delta \ln(W/H)^A$, the wage changes of those workers whose wages are being reviewed or re-negotiated, while presumably assuming that the timing of the adjustment was exogenously fixed.\footnote{For an example of this for the case of the nonunion firm, consider the analysis by Phelps \cite{Phel72} of the firm's decision regarding wage offers. For an example of this for union-management negotiations, consider the analysis by Ashenfelter and Johnson \cite{AshenfelterJohnson72} or Pencavel \cite{Pencavel72} of the firm's decision to take a strike or offer a wage increase acceptable to the union rank and file. In view of the importance of variations in $\lambda_t$ in explaining the variation in aggregate wage changes, as we shall see below, it might be useful to address the determinants of the timing of wage negotiations directly in such models.}

Let us suppose, then, that we proceed by assuming that $\Delta \ln(W/H)^A$ depends on some vector of variables $X_t$ suggested by an appropriate theoretical argument and that we write

\begin{equation}
\Delta \ln(W/H)^A_t = a + \beta X_t + \varepsilon_t
\end{equation}

for a linear approximation of this relationship, where $a$ and $\beta$ are parameters and $\varepsilon_t$ is a disturbance term. If one wishes to explain the aggregate change in wages, $\Delta \ln(W/H)_t$, equation (7) will not do, however,
until it is substituted into (6) to obtain

\[(6a) \quad \Delta \ln(W/H)_t = a \lambda_t + b \lambda_t X_t + \nu_t \]

where \( \nu_t = \lambda_t \epsilon_t \). Notice that if \( \text{var}(\epsilon_t) \) is a constant in (7), then the variance of the disturbance term in (6a) is not constant, because \( \text{var}(\nu_t) = \lambda_t^2 \text{var}(\epsilon_t) \) and depends systematically on \( \lambda_t \). We return to this issue below.

In equation (6a) \( \lambda_t \) enters multiplicatively on the right hand side. Apparently to overcome the lack of data on \( \lambda_t \), most authors have assumed that every worker receives a wage adjustment once a year. Under this assumption \( \sum_{i=0}^{3} \lambda_{t-i} = 1.0 \) for all values of \( t \); that is, the sum of the fraction of workers receiving wage adjustments over any four-quarter period equals unity. If we now sum both sides of (6a) over any four-quarter period we have

\[(6b) \quad \ln(W/H)_t - \ln(W/H)_{t-4} = a + b \sum_{i=0}^{3} \lambda_{t-i} X_{t-i} + \sum_{i=0}^{3} \nu_{t-i} . \]

In (6b) the four-quarter change in the aggregate wage is a function of the weighted four-quarter average of the \( X \)'s, with weights equal to the fraction of workers receiving settlements in each of the preceding four quarters. This is not quite the end of the story, however, because the \( \lambda_t \) that form the weights on the right hand side of (6b) are still unknown. To overcome this final difficulty, it is typically assumed that \( \lambda_t = 1/4 \) for all periods. Under this assumption (6b) becomes

\[(6c) \quad \ln(W/H)_t - \ln(W/H)_{t-4} = a + b \sum_{i=0}^{3} \frac{1}{4} X_{t-i} + \sum_{i=0}^{3} \nu_{t-i} \]

\[= a + \frac{1}{4} \beta \sum_{i=0}^{3} \epsilon_{t-i} + \xi_t . \]
As several authors have pointed out, under the assumption that $\lambda_t = 1/4$ which leads to (6c), there is no reason why (6a) could not have been fitted directly. Indeed (6a) and (6c) are now observationally equivalent except that the disturbance term $z_t = \sum_{i=0}^{3} v_{t-i}$ in (6c) will be serially correlated if the disturbance term $v_t$ in (6a) is serially independent, and vice versa. The covariance between $z_t$ and $z_{t-j}$ is simply $(4-j)\text{var}(v_t)$ for $j = 0, \ldots, 3$ and zero otherwise, so long as the $v_t$ are serially independent. This has prompted Rowley and Wilton [22] to fit the four-quarter change equation (6c) by generalized least squares using this disturbance covariance matrix, although it is difficult to see why so complicated a method should be used when under the same assumptions the direct fitting of (6a) is entirely appropriate.

In this paper we take a different approach to these problems and attempt to estimate equation (6a) directly by introducing measures of $\lambda_t$ into the estimation process. In this way we can put hypotheses on the constancy of the $\lambda_t$ and the properties of the disturbances $v_t = \lambda_t z_t$ to direct tests.

To begin with, we calculate the mean value of the $\lambda_t$ in our sample period of the second quarter of 1951 through to the first quarter of 1970. This gives $\bar{\lambda} = .27$, which is not much different from the $\lambda_t = .25$ that has been assumed in many studies using quarterly data. On the other hand, the standard deviation of the $\lambda_t$ over this period is .14, indicating that the $\lambda_t$ can hardly be considered constant at their mean value. This suggests that fitting equation (6a) under the assumption of constancy

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22/ See Sargan [23] and Black and Kelejian [3].
in the $\lambda_t$ is not likely to produce very satisfactory results and we turn next to what happens when we do this.

Line (1) of Table 2 contains the least squares results of fitting equation (6a) assuming $\lambda_t = 1/4$ for all observations, and using a simple average of the reciprocal of the aggregate unemployment rate in the current and preceding quarters as well as the lagged four-quarter proportionate change in consumer prices as elements of the explanatory vector $X_t$. These are the two variables most often used to explain movements in aggregate money wage changes: the reciprocal of the unemployment rate is normally interpreted as a measure of the excess demand for labor and the rate of change of prices is introduced to allow for real rather than nominal wage changes to respond to the excess demand for labor. Although we do not discuss their theoretical rationale here, both of these variables raise serious problems of measurement in the context of equation (6a). For the elements of $X_t$ in that equation are meant to refer to the conditions surrounding the wage determination of the workers who are in the process of receiving wage adjustments, that is, the conditions surrounding the workers who are counted in the numerator of the ratio of $\lambda_t$ in a given quarter. This implies that the correct measure of the excess demand for labor to use in equation (6a) in quarter $t$ would be a measure of the excess demand for labor in the markets for workers who are in the process of a wage review or re-negotiation. Only if we may think of the aggregate economy as a single labor market would the aggregate unemployment rate be an appropriate measure of the excess demand for these workers. Since the consumer price level is presumably a uniform measure of the cost of living faced by a worker, it might be thought that this
variable would not be susceptible to this measurement problem. In fact, this is not the case because the change in the price level should be measured for workers receiving wage adjustments in period \( t \) by the change over the period since this group of workers last received a wage adjustment. Now under the assumption that \( \lambda_t = 1/4 \), workers receiving wage adjustments in period \( t \) must have last received wage adjustments in period \( t-4 \), so that \( \ln P_t - \ln P_{t-4} \) is the proportionate change in prices since their last wage adjustment. This is the rationale for using the four-quarter change in the price level as an explanatory variable in line 1 of Table 2, though we have allowed for delayed adjustments by lagging it one period. On the other hand, under the more realistic assumption that \( \lambda_t \) varies over time, this procedure may introduce substantial error of measurement. Both of these possible sources of measurement error could seriously bias the estimators of the coefficients on line 1.

As can be seen from Table 2, the fraction of the variance in aggregate wage changes explained by the reciprocal of the unemployment rate and price changes when we assume \( \lambda_t = 1/4 \) is very small, namely, about .09. On the other hand, the residuals of the equation in line 1 of Table 2 show no serial correlation whatever with a measured Durbin-Watson statistic of 2. This situation differs considerably from what most researchers have reported when the four-quarter change in the aggregate wage rate is used as dependent variable. According to equation (6a) the regression coefficients in line 1 are estimates of \( 1/4 \alpha \) and \( 1/4 \beta \), where \( \alpha \) and \( \beta \) are the underlying parameters in equation (7). Accordingly, the estimate of \( \alpha \) from line 1 of Table 2 is .016 (.018) and the estimates of the \( \beta \)'s
Table 2

Estimates of Equation (6a) using Quarterly Data 1951 II - 1970 I

| Line No. | Dependent Variable | Constant | $U^{-1}_t$ | $\Delta \ln (P_{t-1}/P_{t-5})$ | $\lambda_t$ | $\lambda_t U^{-1}_t$ | $\lambda_t \Delta \ln (P_{t-1}/P_{t-5})$ | $D_t$ | SEE | DW | $R^2$
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<td>.975 ($10^{-2}$)</td>
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<td>$\Delta \ln (W/H)_t$</td>
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<td>.720</td>
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<td>3</td>
<td>$\Delta \ln (W/H)_t$</td>
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<td>.00323</td>
<td>.218</td>
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<td></td>
<td></td>
<td></td>
<td>.541 ($10^{-2}$)</td>
<td>1.76</td>
<td>.725</td>
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<td></td>
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<td>.00139</td>
<td>.0157</td>
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</tr>
<tr>
<td>4</td>
<td>$[\Delta \ln (W/H)_t]/\lambda_t$</td>
<td>.0349</td>
<td>.00323</td>
<td>.218</td>
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<td>.550 ($10^{-2}$)</td>
<td>1.72</td>
<td>.703</td>
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<td>.0101</td>
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<td></td>
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</tbody>
</table>

There are 76 quarterly observations underlying these ordinary least-squares regression estimates spanning the period from the second quarter of 1951 to the first quarter of 1970. All these data are published in issues of the Ministry of Labour Gazette or the Department of Employment Gazette except for observations on $D_t$ which are constructed from the published series on the vacancy rate in the National Institute Economic Review. A quarterly time series on the numbers of manual workers to use in the denominator of $\lambda_t$ is not published over this two decade period. However, at approximately the mid-point of this period (namely in 1961), the Ministry of Labour estimated that the wage series (N) related to something like 13-1/4 million manual workers (see Gazette, Jan. 1961). This would have represented 0.556 ($\approx 13,250/23,935$) of total civil employment. We assumed that this ratio of manual to total employees obtained throughout the period so that the denominator $\lambda_t$ is 0.556 of total civil employment in each quarter. As in Table 1, $U_{t-2}$ in these regressions is an average of the measured unemployment rate in quarter $t$ and in quarter $t-1$. In lines 4 and 5, the $R^2$ is reported in terms of $\Delta \ln (W/H)_t$ as dependent variable in order to make the $R^2$ in these lines comparable with the $R^2$ in the preceding lines. This is achieved by multiplying the predicted values in lines 4 and 5 by $\lambda_t$ and then finding the fraction of the variance in $\Delta \ln (W/H)_t$ that is removed by these predictions.
attached to $u_t^{-1}$ and price changes are $0.041 (0.029)$ and $0.384 (0.192)$ respectively, with estimated standard errors in parentheses. The former implies that the percentage rate of change of the hourly wage of those workers receiving adjustments will be about 1 percentage point higher for each 1 percentage point that the unemployment rate declines, where this partial derivative has been evaluated at an unemployment percentage of 2.0. Likewise, the coefficient on price changes implies that the wage changes of those workers receiving adjustments increases by about $\lambda$ of the increase in consumer prices that has taken place since these workers' last wage adjustment. On the other hand, neither of these effects is estimated with much precision and the effect of the unemployment rate would not be judged significantly different from zero by conventional criteria.

In line 2 of Table 2 we report the results of fitting equation (6a) by least squares without the assumption that $\lambda_t$ is a constant. This calls merely for multiplying the explanatory variables used in line 1 by $\lambda_t$, and entering these products as explanatory variables. According to the reasoning underlying this approach the appropriate equation is estimated without a constant term and we have followed this practice in line 2. Since the dependent variables are the same in line 1 and line 2 it is perfectly correct to compare the fraction of explained variance in the one against the other. As can be seen from the table, the standard error of the residuals in line 2 is nearly one-half as large as in line 1 and the fraction of explained variance increases to .72. In order to test the hypothesis that the constant term should be excluded from line 2, and also to force the mean of the residuals for this equation to equal zero so that the $R^2$ and Durbin-Watson statistics in lines 1 and 2 are fully comparable,
we report the results in line 3 of adding a constant term to our estimated version of equation (6a). Not only is the constant term quantitatively small and not significantly different from zero in line 3, but the other estimated coefficients are nearly identical to those in line 2. As can also be seen from the table, the estimate of α is larger and the estimates of the β's are smaller in line 2 than in line 1, though these parameters are estimated with enough precision in line 2 to be judged significantly different from zero on the conventional criteria.

This is hardly the end of the story, however, because we have still to examine the possible heteroscedasticity in the error terms of the estimated equation in line 2. Recall that the error term is $v_t = \lambda_t e_t$ in equation (6a). If $\text{var}(e_t) = \sigma^2 e$, a constant, in equation (7) then $\text{var}(v_t) = \sigma^2 e \lambda_t^2$ and depends systematically on $\lambda_t$. This suggests that it is necessary to test the hypothesis that $\text{var}(v_t)$ is a constant.

The necessary calculations for testing this hypothesis using a straightforward procedure devised by Goldfeld and Quandt [9] are contained in Table 3. To proceed we divide the 76 observations underlying the results in line 2 of Table 2 into two disjoint sets: those corresponding to the cases where $\lambda_t$ is above its median value in the sample and those corresponding to the cases where $\lambda_t$ is below its median value in the sample. The mean value of $\lambda_t$ and the sum of squares of $\lambda_t$ for these two sets of observations are contained in the first two rows of Table 3. Row 3 contains the sum of squared errors from fitting equation (6a) to the two separate sets of observations. Under the null hypothesis that $\text{var}(v_t)$ equals a constant, the ratio of the expectations of these sums of squared errors will be equal to unity. As can be seen from the table, the ratio of the sums of squares in columns 1 and 2 of row 3 is 3.42, which leads to a firm rejection of the hypothesis that $\text{var}(v_t)$ is a constant.
There is also a straightforward way to test the alternative hypothesis that \( \text{var}(\varepsilon_t) \) is a constant. Suppose we divide both sides of equation (6a) by \( \lambda_t \) to get

\[
(6d) \quad \Delta \ln(H/H_t) / \lambda_t = a + \beta X_t + \varepsilon_t .
\]

If \( \text{var}(\varepsilon_t) \) is a constant then (6d) is the appropriate form in which to estimate equation (6a). Line 4 of Table 2 presents the results of fitting equation (6d) to the data. Likewise, the appropriate F-ratio for a test of the null hypothesis that \( \text{var}(\varepsilon_t) \) in the fitted equation is homoskedastic is contained in row 4 of Table 3. As can be seen from this table there is no evidence for rejecting the formulation that \( \text{var}(\varepsilon_t) \) is a constant, with a calculated F-ratio of 1.37 that is well below the tabulated critical value of 1.75. We conclude, therefore, that equation (6d) is the appropriate stochastic specification and that the results in line 4 of Table 2 are to be preferred over the others in this table.

Unfortunately, the results in line 4 of Table 2 are not very encouraging with respect to our knowledge of the economic determinants of aggregate wage changes. The estimated coefficient on the reciprocal of the unemployment rate is now only about one-half its standard error and even the estimated coefficient on price changes would not be judged significantly different from zero on the conventional criterion. The same is true when Dicks-Mireaux and Dow's measure of excess demand replaces the unemployment rate in the equation, as can be seen from line 5 of Table 2. In effect, the underlying econometric specification of the determinants of aggregate wage changes has inexorably led us to specify a dependent variable that is the ratio of the aggregate wage change to the fraction of workers whose wages are reviewed or re-negotiated in a
Table 3
Tests of Homoscedasticity for Equation (6a)

<table>
<thead>
<tr>
<th>Row No.</th>
<th>Large Values of $\lambda_t$</th>
<th>Small Values of $\lambda_t$</th>
<th>Calculated F-ratio $^g/$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>Mean Value of $\lambda_t$</td>
<td>.371</td>
<td>.170</td>
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<tr>
<td>2</td>
<td>$\Sigma \lambda_t^2$</td>
<td>5.83</td>
<td>1.22</td>
</tr>
<tr>
<td>3</td>
<td>Sum of squared errors</td>
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<tr>
<td></td>
<td>with $\Delta n(W/H)_t$ as</td>
<td></td>
<td></td>
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<tr>
<td></td>
<td>Dependent Variable</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>.164 ($10^{-2}$)</td>
<td>.048 ($10^{-2}$)</td>
</tr>
<tr>
<td>4</td>
<td>Sum of Squared Errors</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>with $[\Delta n(W/H)_t]/\lambda_t$</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>.0142</td>
<td>.0195</td>
</tr>
</tbody>
</table>

Notes: $^g/$The tabulated critical level for the F-ratio at the .05 level is approximately 1.75 for 35 degrees of freedom (in numerator and denominator)
given quarter. As it turns out, the conventional economic variables are not of much practical use in accounting for movements in this dependent variable. Surely this is the accurate way to assess the quality of the theory and measurement that underlies our current understanding of the determinants of aggregate wage changes. 23/

III. Concluding Remarks

In this note we have examined two issues of contention in the estimation of the determinants of money wage and earnings changes for manual workers in the U.K. On the one hand, we have found that, even after allowing for the influence of payment-by-results workers, the major causes of the growth in earnings per week over the postwar period have been (a) the upward movement in wages and (b) the upward movement in the excess of actual hours worked over normal hours of work and the attendant growth in the fraction of total hours paid at premium rates. We can only speculate on the causes of the increase in overtime hours and further analysis is clearly required. Still, it is clear from our results that once we abstract from any changes in overtime hours, explanations of the determinants of wage changes will simultaneously provide the bulk of the explanation of the determinants of earnings changes. This result presumably rationalizes the interest that has been historically centered on empirical analyses of the determinants of wage changes.

23/ It is perhaps worth observing that, like Gillion [7] and Johnston and Timbrell [11], we have found no evidence that \( \lambda_t \) is itself related in any simple manner to the unemployment rate or the rate of change of prices.
On the other hand, in paying close attention to the stochastic specification underlying an equation designed to explain wage changes, we have been inexorably led by a series of tests to specify a form for the dependent variable that bears only an imprecise relation to the conventional variables that have been found in the past to influence wage changes. This leads us to suspect that much of the "statistical significance" and "goodness of fit" that has been reported in past studies of quarterly wage inflation is a mirage and results primarily from models with substantial errors in stochastic structure. True advances in the quality of our empirical understanding of the causes of wage inflation are more likely to result if researchers address themselves to these difficulties.
REFERENCES


