On Measuring the Impact of Wage-Price Controls:  
A Critical Appraisal

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* I wish to thank Professors Orley Ashenfelter and George de Menil for their helpful comments and suggestions in the early stages of this paper. The statistical analysis in Part II could not have been performed without Prof. de Menil who generously provided me with the data. I would also like to thank John Ortiz for his computational assistance, and the Industrial Relations Section at Princeton University for secretarial assistance. The paper was revised in the light of comments that I received at the Rochester Conference on Nov. 9-10, 1973. I alone am responsible for any errors in this paper.
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National governments have, for at least the last quarter century, assumed the responsibility for establishing policies that promote full employment, price stability, and economic growth. Monetary and fiscal policies have historically provided the necessary policy instruments. A policy that promotes one goal, say full employment, may often oppose another, say price stability. Following the now classic article by Phillips (1958), the "trade-off between inflation and unemployment" has been added to the jargon of the economics literature. These "trade-offs" were fully acknowledged by policy-makers well before Phillips. They were dissatisfied with traditional monetary and fiscal policies, and they proceeded to develop other policies that would hopefully help to control the macro behavior of the economy. These took the form of direct controls over wages and prices, (usually in wartime), or incomes policies that promulgated guidelines or enforced selective controls over certain sectors.

The fundamental question of whether controls and guidelines are desirable policy instruments is beyond the scope of the present paper. My objective here is far more modest, namely to evaluate the merits of several studies that have attempted to measure the impact of controls and guidelines in restraining the rate of wage and price inflation. These studies reveal four approaches to the measurement of the impact: (1) the case method, (2) tabular comparisons or the "before and after" approach, (3) the dummy variable method, and (4) the simulation approach.

In this paper, the term, "controls", will be used to stand for both formal wage and price controls (like those under phases I and IV of the Nixon administration), and informal guidelines, (like those under the
Kennedy-Johnson Guidestones). I fully recognize the heterogeneity of controls programs with respect to the size of the bureaucracy, the complexity of the rules and exceptions, the strength of the enforcement provisions, etc. However, it is convenient for this paper to lump all of these programs under the single label of "controls."

The studies that adopt the case method try to assess the impact of controls by examining particular episodes or cases. Thus, Sheahan (1967) described the circumstances surrounding the cases of (a) the steel price rollback in 1962, (b) copper prices in 1966, (c) hides where exports were banned to depress domestic prices, etc. Ulmon and Flanagan (1971) examined the ways in which temporary wage restraint was or was not achieved through incomes policies in several western European countries. These studies provide us with some useful insights, especially about the mechanics and politics of implementing a controls program. The cases are, however, not a random sample of economic activity, and quantitative estimates of what would have happened to wages and prices in the absence of controls cannot be easily gleaned from these studies.

The tabular comparison or "before and after" approach is exemplified by Bosworth (1972) who evaluated the impact of Nixon, Phase II. These studies assemble the pertinent data on the levels and rates of change in wages and prices that are supposed to be influenced by the controls. However, explanatory variables other than the presence of controls surely affect the rates of wage and price inflation. Hence, it is not possible to evaluate the partial impact of controls from a chart of the dependent variables alone.
The last two approaches — dummy variables and simulation — call for the specification of an economic model in which the rates of wage and price inflation are related to other economic variables. Historical time series data are then used to estimate the parameters of the model. I do not intend to provide a comprehensive survey of the voluminous literature on wage and price equations. I have instead selected four studies [by R.J. Gordon (1972), O. Ekstein and R. Brinner (1972), G.L. Perry (1970), R.G. Lipsey and J.M. Parkin (1970)] which, in my opinion, typify the kinds of models that are being used to quantify the overall impact of controls. The merits of these models as a basis for measuring the impact of controls can be appraised from two viewpoints; the econometric properties of the estimates generated by the model, and the theoretical foundations for the structural equations of the model. The plan of this paper is as follows:

In Part I, I briefly summarize the essential features of the four models and the theoretical rationale for their wage and price equations. The models differ in their econometric specification with respect to (a) measurement of the dependent variables, (b) the explanatory variables included in each equation, (c) the distributed lag structure for some of the variables, and (d) the functional form of the equations. The Appendix to this paper presents the detailed econometric model specification and least squares parameter estimates as reported in these four studies.

Part II is devoted to some econometric aspects. The discussion there does not question the theoretical structure of the models, and attention is directed to the statistical issues of measurement, tests for stability, and possible specification errors. The simulation method of comparing actual
predicted rates of wage and price inflation during a period of wage-price controls is then examined in Part III. It is argued there that the imposition of truly binding controls could (a) induce structural change, (b) affect the time paths of the so-called exogenous explanatory variables, and (c) lead to a discrepancy between the measured and true rates of price inflation. The salient conclusions of the paper are summarized in Part IV.

The paper tends to be negative because it is, after all, a critical appraisal. I regret that I cannot provide the reader with a demonstrably superior method for measuring the quantitative impact of wage-price controls, but I hope that the paper clarifies some of the theoretical and empirical pitfalls of the current methodology.

I

Four Models of Money Wage Rate and Price Level Determination

The models developed by Gordon (1972), Ekstein and Brinner (1972), Perry (1970), and Lipsey and Parkin (1970), [hereafter abbreviated G, EB, P, and LP], all purport to explain the time paths in relative rates of change of money wage rates and prices. According to these models, [excepting Perry who only estimated a wage equation], the inflationary process can be described by a system of two equations -- a wage equation and a price equation. Although the variables appearing in each equation differ from model to model, they all share a common theoretical rationale which I shall try to summarize in this section.

1.1 The Wage Equation

The relative time rate of change in hourly money wage rates, denoted by \( w \), can, according to these models, be decomposed into four additive
component changes which represent the relative adjustments in wage rates in response to (a) the secular growth in labor productivity or some other time related factor \( w^g \), (b) price inflation \( w^p \), (c) disequilibrium in labor markets \( w^d \), and (d) changing social security and personal tax rates \( w^t \).

Thus,

\[
w = w^g + w^p + w^d + w^t.
\]

For purposes of econometric estimation, counterparts to these relative continuous time rates of change must be developed from data for discrete time periods.

To explain the logic of these wage equations, let me begin with the third component of wage change, \( w^d \), which is the relative change in wages in response to disequilibrium in labor markets. The theoretical rationale here is the familiar one based on the Phillips curve which was originally articulated by Lipsey (1960). If \( D \) and \( S \) denote the demand and supply

\[1/\]

To facilitate typing and comparisons across models, I shall adopt my own notation which differs from that appearing in the four studies. Upper case letters like \( W \) and \( P \) will stand for level variables such as wage and price indexes at time \( t \). Lower case letters are used to denote relative time rates of change in continuous time.

\[
w = \frac{1}{w} \left( \frac{dW}{dt} \right) \quad \text{and} \quad p = \frac{1}{p} \left( \frac{dP}{dt} \right)
\]

The four component relative wage changes are defined in relation to the hourly wage rate \( W \). The differential in \( W \) can be decomposed into,

\[
dW = dW^g + dW^p + dW^d + dW^t
\]

The component relative wage changes are thus given by,

\[
w^g = \frac{1}{W} \left( \frac{dW^g}{dt} \right), \quad w^p = \frac{1}{W} \left( \frac{dW^p}{dt} \right), \quad \text{etc.}
\]
quantities of labor (appropriately measured) at time $t$, the fundamental behavioral postulate is a response function of the form.

$$w^d = f\left(\frac{D-S}{S}\right), \quad f(0) = 0, \quad f'(0) > 0.$$  

Most economists would agree about the sign of the derivative, $f'(0) > 0$, but there is considerable disagreement about whether there is a stable functional relationship $f$ between the relative excess demand for labor, $\left(\frac{D-S}{S}\right)$, and the relative time rate of adjustment in money wage rates $w^d$.

One cannot directly observe the relative excess demand for labor services. The four models use different proxy variables to approximate the unobservable relative excess demand. In the Perry wage equation, the disequilibrium component of wage change is linked to a linear combination of three variables: the inverse of a weighted unemployment rate $1/U^Y$, the dispersion of unemployment $D$, and the rate of change in secondary employment. R.J. Gordon uses a different combination of four variables; Perry's dispersion of unemployment $D$, disguised unemployment $U^d$, unemployment rate of hours $U^h$, and the gap between product and consumer price inflation rates $(p^X - p)$. In a steady state, the gap will be equal to zero. If we let zero subscripts denote the full employment values of the other variables, the Gordon wage equation invokes the following specification:

$$w^d_t = f\left(\frac{D-S}{S}\right)_t = \alpha_1 (D_t - D_0) + \alpha_2 (U^d_t - U^d_0) + \alpha_3 (U^h_t - U^h_0).$$

The component wage change, $w^d_t$, which adjusts wage rates to disequilibrium in the labor market is unobservable but can be inferred via regression techniques.
The controversy over whether the Phillips curve has shifted seems to confuse at least two meanings of the Phillips curve. On the one hand, the Phillips curve can be defined as a relationship between observable variables; e.g., the total wage rate change \( w \) and some function of the official unemployment rate \( U^0 \). On the other hand it could be defined as a relationship between conceptual, unobservable variables like the component wage change \( w^d \) and the relative excess demand for labor \( \left( \frac{D-S}{S} \right) \). Gordon and Perry argue that the Phillips curve has shifted due to a changing age/sex composition of the labor force, while Ekstein and Brinner contend that it has not shifted because prior studies (like G and P) did not impute enough of the total wage change \( w \) to price inflation. The controversy is mainly an empirical one over which set of proxy variables best measures the relative excess demand for labor services. All of these models must, however, assume that there is a stable relationship between \( w^d \) and the true relative excess demand for labor services \( \left( \frac{D-S}{S} \right) \).

Turn next to the component wage change, \( w^g \), attributed to the secular growth in labor productivity. In a steady state with stable prices, constant tax rates, and unemployment variables taking on values that correspond to zero excess demand, [denoted by zero subscripts in (1.2) above], the growth component of wage change is the sum of the intercept term \( \alpha_0 \) plus a linear combination of zero excess demand unemployment variables. Thus, in the Gordon model, we have,

\[
(1.3) \quad w^g_t = \alpha_0 + \alpha_1 D_0 + \alpha_2 U^d_t + \alpha_3 U^h_0 .
\]

The constancy of the right side of (1.3) implies that money wage rates will rise at a constant exponential rate presumably reflecting the long run
secular growth in labor productivity. The EB wage equation also included a variable to capture the effect of discrepancies between short and long run trends in labor productivity.

If the long run trend rate of growth in labor productivity has varied over the sample period, this fact should be incorporated into the econometric specification of the \( w^g \) component of wage change. In both the G and EB models, a variable representing the long run trend rate of growth in labor productivity, [which they claim, varied over the sample period], was explicitly included in their price equations but not in their wage equations. This is surely a specification error in the G and EB wage equations.

The inflation component of wage change \( w^p \) is intended to capture the adjustment in money wage rates in response to expected or actual inflation in consumer prices. In the P and LP models, \( w^p \) is assumed to be proportional to the one-quarter lagged rate of change in consumer prices. G and EB hypothesize more complicated relationships. The EB wage equation includes two variables; a four quarter moving average of actual rates of change in consumer prices, and the positive excess over 5 percent per year of the accumulated price inflation over a two year period. The rationale here is that for inflation rates below 5 percent, wages only adjust to the four quarter moving average which is their proxy for expected inflation. However, when inflation is high and sustained, wage change demands are accelerated to recapture all of the loss in real wage rates.

If we neglect the gap between product and consumer price inflation, \( (p^x - p) \), [which is rationalized as a partial measure of labor market disequilibrium], the inflation component of wage change in the Gordon model is given by,
(1.4) \[ w_t^P = \alpha_4 p_{t-1}^* = \alpha_4 \left[ a_1 p_{t-1} + a_2 p_{t-2} + \ldots + a_n p_{t-n} \right], \]

where \( p^* \) is the expected rate of price inflation defined as a polynomial distributed lag (with weights \( a_j \)) of lagged actual rates of price inflation, \( p_{t-j} \). The weights \( a_j \) were determined in two ways by minimizing (a) the residual variance of the wage equation or (b) an auxiliary interest rate equation. In the latter approach, the current and lagged rates of change in the money supply affect the \( a \)-weights thereby introducing the money supply as a determinant of the inflationary process in this indirect way. Although G agrees with EB that the relationship between \( w^P \) and price inflation could be non-linear, his data do not provide a clear verdict on this issue.

The final component \( w^T \) represents the adjustment of money wage rates to changes in personal and social security tax rates. It does not appear in the LP model, and in the others, it is assumed that gross hourly wage rates lag behind the changes in tax rates. To capture this component of wage change, the equation includes a variable which is a distributed lag in rates of change in the complements of tax rates.

To sum up, these models posit a wage equation in which the relative rate of change in gross hourly money wage rates can be decomposed into four additive components representing the adjustment of wages to (a) growth in labor productivity, (b) expected price inflation, (c) disequilibrium in the labor market, and (d) changing tax rates. The parameters are only constrained with respect to direction of change. Except for Gordon's parameter on the social security tax variable, the sample data are allowed to determine the parameter values.
The Price Equation

The theory behind the price equation is one of full-cost pricing. Prices in the long run are determined by unit costs and given an assumption of constant income shares, by standard unit labor costs. The derivation of the price equation from these assumptions is most clearly articulated in the LP model.\(^2\) G and EB also invoke this same assumption about the long run determinants of prices, but their price equations try to explain the short run fluctuations of prices about this long run relationship to standard unit labor costs. The spirit of these models is exemplified by the Gordon price equation which is of the form,

\[ p_t^x = \beta_0 + \beta_1 \left( \frac{w_t}{q_t} \right)^* + \beta_2 \left( \frac{q_t}{q_t} \right)^* + \beta_3 \left( \frac{c_t}{w_t} \right) + \beta_4 \left( \frac{u_t}{k_t} \right) + \text{error}, \]

where \( p_t^x \) is the one-quarter relative rate of change in product prices as measured by the GNP deflator for the private non-farm sector. In a steady state, the last three variables equal zero, and the rate of price inflation is determined solely by the rate of change in expected standard benef.

\(^2\) Lipsey and Parkin (1970), pp. 116-7, begin with the identity,

\[ P = WL + MT + CD + \tau, \]

where \( P \) = product price, \( L \) = labor input per unit of output (the reciprocal of \( Q \) = output per man-hour), \( T \) = import input per unit of output, \( CD \) = other material and capital costs per unit of output, \( \tau \) = unit profits, and \( W \) and \( M \) are the prices of labor and imports. They invoke three assumptions:
(a) \( CD = \mu (WL+MT) \), (b) \( \tau = \beta P \), and (c) \( T \) is a constant over time. Taking time derivatives and rearranging terms, they get,

\[ \frac{1}{P_t} \frac{dP}{dt} = \frac{1}{1+\theta} \left[ k_L \left( \frac{1}{L} \right) \left( \frac{dW}{dt} \right) + k_Q \left( \frac{1}{Q} \right) \left( \frac{dQ}{dt} \right) + k_M \left( \frac{1}{M} \right) \left( \frac{dM}{dt} \right) \right], \]

where \( k_L = WL/P \) and \( k_M = MT/P \) represent the labor and import shares of product price. The LP derivation implies that the coefficients of the relative rates of change in wage rates \( w \) and of labor productivity \( q \) should be numerically equal, but their data do not confirm this hypothesis.
unit labor costs, \((w/q')^x\) which is equal to the difference between expected wage inflation \(w^x\) and the trend rate of growth in labor productivity \(q'\). Gordon hypothesizes that \(\beta_0\) should equal zero, and with constant income shares, \(\beta_1\) should equal unity; his data do indeed confirm these hypotheses. The remaining three variables try to capture the short run price fluctuations about this long run norm due to (a) divergence between actual and trend labor productivity, (b) omission of some labor costs in \(w\), and (c) disequilibrium in the product market.

If actual labor productivity \(q\) climbs faster than the long run trend rate \(q'\), it allegedly exerts downward pressure on prices implying that \(\beta_2 < 0\). The variable, \((q/q')^x\) is a distributed lag on the ratio of actual to trend rates of growth in output per manhour.

Gordon uses a fixed weight index of straight-time wage rates \(W\) which omits some elements of labor compensation. The rate of change in the ratio of hourly total compensation to his wage index, \((c/w)\), is introduced to correct this omission. He argues that \(\beta_3\) should be positive because the omitted labor costs influence prices. This argument implies that the parameters should be constrained so that \(\beta_1 = \beta_3\), or "standard unit labor costs" should be measured by \((c/q')\). His results do not yield the equality, \(\beta_1 = \beta_3\), but this is partly due to the fact that \((c/w)\) is not entered as a distributed lag.

The last variable, \((uf/k)\), is the rate of change in the ratio of unfilled orders to capacity in the manufacturing sector which, according to de Menil, is a proxy for disequilibrium in the product market.\(^3\) A fuller development of the rationale for this variable, based on a micro analogue to a macro price equation can be found in de Menil and Bhalla (1974)

\(^3\)A
response function similar to that for the $w^d$ component of wage change is presumed to apply in the product market. An increase in $(uf/k)$ signals arise in the relative excess demand for products leading to a price increase; i.e. $\beta_4 > 0$.

The price equation in the EB model is qualitatively similar and includes the same four explanatory variables differing only in their measurement and lag structures. The only substantive difference is that EB includes a variable measuring the gap between actual and standard unit labor costs. The detailed econometric specifications can be found in the Appendix.

1.3 Reduced Forms and the Phillips Curve

In order to direct attention to the basic structure of these two-equation inflation models, let me assume that (i) the trend rate of growth in output per manhour $q'$ is constant, (ii) total hourly compensation $C$ and the wage index $W$ change at the same rate, and (iii) the unemployment variables appearing in the wage equation can be described by a vector $U$.

Let $\lambda$ denote the lag operator so that $\lambda^j p_t = p_{t-j}$. Given our three assumptions, the wage and price equations can be written as follows:

$$(1.6a) \quad w_t = \lambda_0 + \lambda_1(\lambda)p_t + \lambda_2 U_t + \lambda_3(\lambda)\psi_t.$$  

$$(1.6b) \quad p^x_t = \beta_0 + \beta_1(\lambda)w_t + \beta_2(\lambda)q_t + \beta_3(uf/k)_t.$$  

$^4/$ In the Gordon model, $U$ would contain three elements, the dispersion of unemployment $D$, disguised unemployment $U^d$, and unemployment of hours $U^H$. In Perry, the reciprocal of a weighted unemployment rate $1/U^V$ and dispersion $D$ form the elements of $U$. 
where $A_1(\lambda), A_3(\lambda)$, etc. are polynomials in the lag operator which describe
the distributed lag structures. To close the model, we need an equation
linking the change in product prices $p^x_t$ to the change in consumer
prices $p_t$. It is assumed that the two move together, $p^x_t = p_t$. The
model thus contains two jointly dependent variables, $(w_t, p_t)$, and four
presumably exogenous, predetermined explanatory variables; $U_t$ is a vector
of unemployment measures and the rates of change in output per manhour $q_t$,
tax rates $\psi_t$, and the ratio of unfilled orders to capacity $(uf/k)_t$. This
system of simultaneous equations can be solved for the reduced-form equations
which will be of the form,

\[(1.7a) \quad w_t = \gamma_0 + \gamma_1 U_t + \gamma_2 \psi_t + \gamma_3 q_t + \gamma_4 (uf/k)_t.\]

\[(1.7b) \quad p_t = \delta_0 + \delta_1 U_t + \delta_2 \psi_t + \delta_3 q_t + \delta_4 (uf/k)_t.\]

The reduced-form parameters, $(\gamma_j, \delta_j)$, may involve rather complicated ratios
of polynomials in the lag operator $\lambda$. Two kinds of Phillips curves are
implicit in these reduced-form equations; one relating $w_t$ to the
unemployment variables $U_t$, and the other relating $p_t$ and $U_t$.

G and EB emphasize a Phillips curve defined in the $(p_t, U_t)$ space.
Since the official unemployment rate $U^0$ does not appear in his wage
equation, Gordon is forced to develop an auxiliary equation linking $U^0$
to his unemployment variables. The authors apparently derive this Phillips
curve by eliminating $w_t$; i.e. $(1.6a)$ is substituted into $(1.6b)$ to get,

\[(1.8) \quad p_t = B_0 + B_1(\lambda)[A_0 + A_1(\lambda)p_t + A_2 U_t + A_3(\lambda)\psi_t] + B_2(\lambda)q_t
+ B_3(uf/k)_t.\]
If $B_1(\lambda)$ and/or $A_1(\lambda)$ does not contain a constant term, equation (1.8) can be interpreted as a Phillips curve in which the current price change $p_t$ is a function of lagged price changes, $[B_1(\lambda)A_1(\lambda)]p_t$, and the other exogenous variables of the model. The short run Phillips curve is described by the partial relationship between $p_t$ and $U_t$ given the prior history of price inflation. By assuming that $p_t = p_{t-1} = \ldots = p_{t-n}$, [or equivalently, setting $\lambda = 1$], they derive the so-called long run Phillips curve.

The so-called short run trade-off between price inflation $p_t$ and current unemployment $U_t$ is determined by only the leading term in the polynomial in $\lambda$, $B_1(\lambda)A_2$. The position of the short-run Phillips curve is, however, dependent upon the prior history of price inflation which, in turn, depends on the prior time paths of the exogenous variables. The so-called long run Phillips curve exhibits less trade-off between price inflation and unemployment, [meaning a steeper slope in the $(p_t, U_t)$ plane], because this slope is obtained by summing over all of the coefficients for lagged wage changes in the polynomial $B_1(\lambda)$.

Two comments about the structure of these inflation models are in order here. First, the exogenous variables in the reduced-form equation are all real variables. Temporal variations in these real variables drive the inflation process in these models. Second, monetary and fiscal policies only affect inflation through their impact on the time paths of these four

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Let $B_1(\lambda)A_1(\lambda) = C(\lambda) = c_0 + c_1\lambda + c_2\lambda^2 + \ldots + c_n\lambda^n$. If $c_0 = 0$, then the current price change $p_t$ does not appear on the right side of (1.8). Further, $c_0$ will be equal to zero if either $B_1(\lambda)$ and/or $A_1(\lambda)$ does not contain a constant term.
real variables. Only the time paths of real variables matter, and the rates of wage and price inflation are, according to these models, invariant to the mix of monetary vs. fiscal policies.

II

Econometric Aspects of the Wage-Price Models

Suppose, for the moment, that we accept the theory behind these models and direct our attention to the problem of econometric estimation. The model is described by a system of two simultaneous equations, and this fact is acknowledged by all of the authors. EB stated that the parameters were also estimated by two-stage least squares, but they do not report these results because, according to them, the parameter estimates were not substantially different. The reduced-form equations for the EB model, [the counterparts to (1.7a) and (1.7b)] involve very complicated distributed lag structures, and one can only conjecture about whether EB estimated the correct reduced-form equations to obtain their 2SLS parameter estimates.

In this section, I propose to examine four aspects of the econometric methodology adopted in the four studies. In section 2.1, I argue that the high positive serial correlation in the residuals of wage equations using four-quarter rates of change, is the unavoidable outcome of the measurement of relative wage changes. A four-quarter wage change is inconsistent with the maintained assumption of serial independence for the true residuals. The meaning of stability in the wage or price equations is next examined in section 2.2. It is shown there that statistical stability of regression
coefficients can be compatible with economic instability when the parameter space is large, and the explanatory variables are highly colinear. The observable variables in a model should ideally be close empirical approximations to the conceptual variables in the underlying theoretical model. I contend in section 2.3 that in choosing among alternative observable variables or imposing a priori constraints on parameters, too much weight is attached to statistical criteria based on shaky data, (such as goodness of fit and plausible t-values), and not enough weight is given to the theory of a wage or price equation. Finally, in section 2.4, I suggest that the omission of data on (e) the proportion of workers getting wage changes in each quarter, (b) unemployment insurance programs, and (c) wages of public sector employees may have imparted possibly substantial specification errors in the parameter estimates.

2.1 Measurement of Relative Time Rates of Change

The dependent variable of a wage equation is the relative time rate of change in hourly wage rates which is conceptually measured by \( w \), the time derivative of the logarithm of hourly wages \( W \) at time \( t \).

\[
  w = \left( \frac{1}{W} \right) \left( \frac{dW}{dt} \right).
\]

An empirical counterpart to \( w \) must be developed because wage rates are only observed at discrete time intervals. Empirical studies have used one, two, and four-quarter relative rates of change calculated from average hourly wage rates \( W_t \) in quarter \( t \).
An interesting regularity is observed in the published wage equations. As one extends the interval over which the time rate of change is measured, the residuals of the fitted wage equations exhibit progressively larger positive serial correlation. Thus, when Gordon moves from one to two-quarter rates of change, the implicit serial correlation climbs from -.25 to +.33. EB and LP who both used four-quarter rates of wage change, reported serial correlations in the residuals of +.62 and +.39.\(^6\) It can be plausibly argued that this regularity is the result of the way in which rates of change are empirically measured.

Suppose that a correct specification of the model calls for a one quarter rate of change, \(w^1_t\).

\[(2.1) \quad w^1_t = c + \beta X^t + \epsilon^t,\]

where \(X^t\) is a vector of explanatory variables, and the random error \(\epsilon^t\) is independently, normally distributed with,

\[E[\epsilon^t^2] = \sigma^2, \quad E[\epsilon^t \epsilon^t_{-j}] = 0, \quad \text{(for all } j \neq 0).\]

If we had mistakenly used a two-quarter rate of change, \(w^2_t\), the model would have been given by,

\[(2.2) \quad w^2_t = A + B \bar{X}^t + u^t, \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad [\bar{X}^t = \frac{1}{2}(X^t + X^t_{-1})].\]

\(^6\) The first-order serial correlation coefficient \(r\) was calculated from the Durbin-Watson statistic \(DW\) via the following approximation:

\[DW = 2(1-r)\]

The DW statistics that were used for these calculations are reported in the Appendix. The "policy off" results were used for the LP wage equation.
By using the log approximation to measured rates of wage change, we can derive the relationship between the parameters of these two models. If hats are used to denote logarithms, the one- and two-quarter rates of wage change are defined as follows:

\[
\begin{align*}
\hat{w}_t^1 &= \frac{W_t - W_{t-1}}{W_{t-1}} \approx \hat{W}_t - \hat{W}_{t-1} \\
\hat{w}_t^2 &= \frac{W_t - W_{t-2}}{W_{t-2}} \approx \hat{W}_t - \hat{W}_{t-2}.
\end{align*}
\]

Hence, for small changes we have the following approximation:

\[
\hat{w}_t^2 \approx \hat{w}_t^1 + \hat{w}_{t-1}^1.
\]

It is apparent that \(A = 2\alpha\) and \(B = 2\beta\). More importantly, the random errors of the two models are related via,

\[(2, 3) \quad u_t = \epsilon_t + \epsilon_{t-1}.
\]

If \(\epsilon_t\) was serially independent, then \(u_t\) must be positively correlated with,

\[
E[u_t^2] = 2\sigma^2, \quad E[u_t u_{t-1}] = \sigma^2, \quad E[u_t u_{t-2}] = 0, \text{ etc.}
\]

It is evident that if we extend the interval for measuring relative rates of wage change from two to four quarters, the first-order serial correlation, \(r\), of the random errors in the wage equation must climb from \(r = +.5\) to \(r = +.75\).

Although \(P, EB,\) and \(LP\) all acknowledge that the \(t\)-statistics are not valid because of significant positive serial correlation in the residuals, they do not turn to generalized least squares to get maximum
likelihood estimators. The four quarter rates of change are rationalized on the ground that it increases the "signal to noise ratio". This presumes that the hourly wage rates $W_c$ contain unknown measurement errors. But, if the measurement errors were serially independent, they could not generate the pattern of larger serial correlation for longer intervals. Gordon found that one and two quarter rates of change yielded Durbin-Watson statistics that bracket two, and he accordingly reported results for regressions using both versions of the dependent variable. Although he erroneously suggests that serial correlation imparts bias in parameter estimates, his procedure of using shorter time intervals for measuring the rates of wage change is the preferred econometric method. Bowley and Milton (1973) applied generalized least squares to models using four quarter rates of wage change. If (2.1) is the correct model specification, the preferred method is to apply ordinary least squares to a model using one quarter rates of wage change.

2.2 Stability of Coefficients and the Guidepost Dummy

Every econometric model must tacitly assume that there is a stable structure for the economic process under investigation. In many studies, this is a maintained assumption that is never submitted to a statistical test. Gordon (1972) and Lipsey-Parkin (1970) have, however, questioned

---

7/ This misconception is implicit when he writes: "Hence, I exhibit estimates for both forms of the dependent variable and assume that the two estimated coefficients for each independent variable bracket the 'best' estimate." [Gordon (1972) p. 388]. The use of ordinary least squares when the errors are serially correlated, does not lead to bias in the point estimates of parameters but only affects the estimated covariance matrix of the parameter estimates.
the validity of this implicit stability assumption. Lipsey and Parkin employ a statistical technique due to Chow (1960) which uses an F-ratio to test for the stability of regression coefficients in two or more independent samples. There is, however, an important distinction that must be drawn between statistical and economic stability of the underlying structure. Two sets of parameter estimates, \((\mathbf{b}_1, \mathbf{b}_2)\), could, on the basis of a Chow test, be consistent with a common parameter vector \(\mathbf{\beta}\) implying a stable statistical structure. However, the point estimates of the parameters, \((\mathbf{b}_1, \mathbf{b}_2)\), could imply very different economic structures.

To demonstrate this distinction, I shall first briefly review the Chow test. Let \((\mathbf{Y}, \mathbf{X})\) denote the matrix of sample observations for the entire sample period of \(T\) observations. We can divide the sample into two parts, \((\mathbf{Y}_1, \mathbf{X}_1)\) for \(T_1\) observations and \((\mathbf{Y}_2, \mathbf{X}_2)\) for the remaining \(T_2 = T - T_1\) observations. One can then estimate two sets of parameter estimates and calculate the sum of squared residuals as follows:

\[
\begin{align*}
(2.5a) \quad \mathbf{y}_1 &= \mathbf{X}_1 \mathbf{\beta}_1 + \mathbf{\epsilon}_1, \quad e_1' e_1 = \mathbf{y}_1' \mathbf{y}_1 - \mathbf{y}_1' \mathbf{X}_1 \mathbf{b}_1. \\
(2.5b) \quad \mathbf{y}_2 &= \mathbf{X}_2 \mathbf{\beta}_2 + \mathbf{\epsilon}_2, \quad e_2' e_2 = \mathbf{y}_2' \mathbf{y}_2 - \mathbf{y}_2' \mathbf{X}_2 \mathbf{b}_2. \\
(2.5c) \quad \mathbf{y} &= \mathbf{X} \mathbf{\beta} + \mathbf{\epsilon}, \quad e' e = \mathbf{y}' \mathbf{y} - \mathbf{y}' \mathbf{X} \mathbf{b}.
\end{align*}
\]

If there are \(k\) elements in the parameter vectors, \((\mathbf{\beta}_1, \mathbf{\beta}_2, \mathbf{\beta})\), the null hypothesis that \(\mathbf{\beta}_1 = \mathbf{\beta}_2\), [implying stability of coefficients], can be tested by constructing the following F-statistic.

\[
(2.5d) \quad F = \frac{A}{B}, \quad A = \frac{e' e - e_1' e_1 - e_2' e_2}{k} \quad B = \frac{e_1' e_1 + e_2' e_2}{T - 2k}
\]
If $F$ exceeds the critical $F$-ratio, $F_Q$, for $n_1 = k$ and $n_2 = T-2k$ degrees of freedom, the sample data contradict the assumption of a stable structure. A less powerful test is available when $T_2 < k$; i.e. when the number of observations in one sample is less than the number of parameters. The power of the $F$ test to discriminate between economically significant parameter estimates appears to be quite weak when the parameter space is large, and the two samples are of vastly different sizes. To illustrate these points, I have attempted to replicate the Gordon wage equation. In Table 1, I present five sets of parameter estimates corresponding to five sample periods. The coefficients of the expected price inflation variable, $p^x_{t-1}$, and the gap between product and consumer prices, $(p^y-p)^x$, are the sums of lag coefficients involving OLS estimates for seven parameters in the polynomial distributed lags. Hence there $k = 12$ parameters estimated in each of these five regressions.

With the data and regression programs available to us, we were unable to replicate Gordon's results for the two samples beginning in 1954:1 and ending in 1966:4 and 1970:4. The least squares parameter estimates and summary statistics are reproduced below from Gordon's published results (pub1) and our replications (repl).

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8/ For this latter test, confer Johnston (1972), p. 207.

9/ I am deeply indebted to Prof. George de Menil of Princeton University who provided me with the quarterly time series data for these regressions. Prof. de Menil also developed the programs to estimate the distributed lags. The polynomial distributed lags implicit in the Table 1 regressions are slightly different from those used by Prof. R.J. Gordon. This difference in the lag structures probably accounts for my inability exactly to replicate Gordon's results.
Comparison of Parameter Estimates for the Gordon Wage Equation
(one quarter rates of change: published vs. replications)

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>publ.</td>
<td>repl.</td>
</tr>
<tr>
<td>In Disp. Un. D</td>
<td>.0410</td>
<td>.0495</td>
</tr>
<tr>
<td>Disag. Un. Ud</td>
<td>-.602</td>
<td>-.6258</td>
</tr>
<tr>
<td>In Un. Hours Uh</td>
<td>-.422</td>
<td>-.3658</td>
</tr>
<tr>
<td>Price Infl. p*</td>
<td>.660</td>
<td>.5799</td>
</tr>
<tr>
<td>Ln Prod. Price Gap (p-x-p)*</td>
<td>.684</td>
<td>.7675</td>
</tr>
<tr>
<td>Pers. Tax ¥e</td>
<td>.115</td>
<td>.1080</td>
</tr>
<tr>
<td>ln σ²</td>
<td>.00182</td>
<td>.00184</td>
</tr>
<tr>
<td>DW</td>
<td>2.49</td>
<td>2.61</td>
</tr>
</tbody>
</table>

Although there are some disturbing discrepancies, (e.g. in the coefficient of the price inflation variable p*), I felt that the replications were close enough to Gordon's published results to warrant their use in tests for the stability of regression coefficients.

Gordon fitted his wage equation to three overlapping samples beginning in 1954:1 and ending in 1966:4, 1968:4, and 1970:4. An eyeball inspection of the parameter estimates for these three overlapping samples [presented in Gordon (1972), pp. 398-9] conveys an illusion of stability; compare, for example, (G1*) and (G2-a*) in the preceding table. He concluded, (apparently on the basis of a visual examination of parameter estimates) that the structure of his wage equation was stable over the entire sample period, 1954:1-1970:4. Gordon compares the estimates of β and β₁ in equations (2.5c) and (2.5a) above. Since (2.5a) contained
Table 1
Replications of the Gordon Wage Equation for Alternative Samples*

<table>
<thead>
<tr>
<th>Variable</th>
<th>(G1)</th>
<th>(G2-a)</th>
<th>(G2-b)</th>
<th>(G3-a)</th>
<th>(G3-b)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept $\alpha_0$</td>
<td>0.0054</td>
<td>0.0076</td>
<td>-0.0483</td>
<td>-0.0015</td>
<td>0.0062</td>
</tr>
<tr>
<td></td>
<td>(2.49)</td>
<td>(2.71)</td>
<td>(-1.65)</td>
<td>(-0.30)</td>
<td>(0.84)</td>
</tr>
<tr>
<td>Dispersion $D$</td>
<td>0.0495</td>
<td>0.0411</td>
<td>0.4086</td>
<td>0.1741</td>
<td>0.0381</td>
</tr>
<tr>
<td></td>
<td>(2.18)</td>
<td>(1.43)</td>
<td>(1.80)</td>
<td>(2.61)</td>
<td>(0.61)</td>
</tr>
<tr>
<td>Decrease in $U^d$</td>
<td>-0.6258</td>
<td>-0.8128</td>
<td>1.4209</td>
<td>-0.5095</td>
<td>-0.6358</td>
</tr>
<tr>
<td></td>
<td>(-3.43)</td>
<td>(-4.18)</td>
<td>(0.94)</td>
<td>(-1.40)</td>
<td>(-1.11)</td>
</tr>
<tr>
<td>Unemployment $U^h$</td>
<td>-0.3658</td>
<td>-0.3885</td>
<td>0.1647</td>
<td>-0.1140</td>
<td>-0.6304</td>
</tr>
<tr>
<td></td>
<td>(-1.39)</td>
<td>(-1.37)</td>
<td>(0.14)</td>
<td>(0.32)</td>
<td>(-1.18)</td>
</tr>
<tr>
<td>Price Infl. $p^*$</td>
<td>0.5799</td>
<td>0.2939</td>
<td>0.9740</td>
<td>0.0735</td>
<td>0.8402</td>
</tr>
<tr>
<td></td>
<td>(2.96)</td>
<td>(1.08)</td>
<td>(0.72)</td>
<td>(0.23)</td>
<td>(1.96)</td>
</tr>
<tr>
<td>Prod. Price Gap $(p^x-p)^*$</td>
<td>0.7675</td>
<td>0.3546</td>
<td>1.9314</td>
<td>0.0569</td>
<td>1.8310</td>
</tr>
<tr>
<td></td>
<td>(2.45)</td>
<td>(1.08)</td>
<td>(0.39)</td>
<td>(0.13)</td>
<td>(1.60)</td>
</tr>
<tr>
<td>Pers. Tax $\Psi^e$</td>
<td>0.1080</td>
<td>0.2366</td>
<td>0.0891</td>
<td>0.2702</td>
<td>0.0286</td>
</tr>
<tr>
<td></td>
<td>(1.85)</td>
<td>(3.25)</td>
<td>(0.42)</td>
<td>(1.76)</td>
<td>(0.37)</td>
</tr>
</tbody>
</table>

Summary statistics

<table>
<thead>
<tr>
<th></th>
<th>(G1)</th>
<th>(G2-a)</th>
<th>(G2-b)</th>
<th>(G3-a)</th>
<th>(G3-b)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$R^2$</td>
<td>0.7953</td>
<td>0.6468</td>
<td>0.8863</td>
<td>0.7148</td>
<td>0.8606</td>
</tr>
<tr>
<td>$\sigma_e$</td>
<td>0.00184</td>
<td>0.00173</td>
<td>0.00157</td>
<td>0.00172</td>
<td>0.00195</td>
</tr>
<tr>
<td>DW</td>
<td>2.61</td>
<td>2.70</td>
<td>2.49</td>
<td>2.75</td>
<td>2.93</td>
</tr>
<tr>
<td>No. of Obs.</td>
<td>68</td>
<td>52</td>
<td>16</td>
<td>32</td>
<td>36</td>
</tr>
</tbody>
</table>

*The dependent variable is the one quarter rate of change in the Gordon wage index. Coefficients of $p^*$ and $(p^x-p)^*$ are the sum of distributed lag coefficients. The $t$-statistics are shown in parentheses. Each equation contained $k = 12$ parameters.

**The periods corresponding to the five equations are as follows:

(G1) = 1954:1-1970:4
(G2-a) = 1954:1-1966:4
(G2-b) = 1967:1-1970:4
(G3-a) = 1954:1-1961:4
(G3-b) = 1962:1-1970:4
52 of the 68 observations in (2.5c), it is not surprising that the two parameter vectors look alike. In Table 1, column (G2-b), I present the least squares parameter estimates for \( \beta_2 \) for the smaller sample of \( T_2 = 16 \) observations, 1967:1-1970:4. The residual sum of squares from (G1), (G2-a), and (G2-b) produced an F-ratio of \( F = 1.656 \) which is well below the 5 and 1 percent critical values of \( F_{0.05} = 1.98 \) and \( F_{0.01} = 2.62 \) for \( n_1 = 12 \) and \( n_2 = 44 \) d.f. The hypothesis that \( \beta_1 = \beta_2 \) is also accepted when the entire sample is divided into approximately equal halves, columns (G3-a) and (G3-b) of Table 1. The F-ratio here was \( F = 0.927 \). The application of the Chow test thus conclusively demonstrates that the parameters of the Gordon wage equation satisfy the assumption of statistical stability.

What would the reader infer if he were only shown the parameter estimates for (G1) and (G2-b) instead of what he was shown, namely (G1) and (G2-a)? The Chow test would still have indicated no statistical difference in the vectors of parameter estimates. The parameter estimates now look very different. The coefficient of disguised unemployment \( U^d \), for example, was -0.63 for the entire sample (G1) and +1.42 for the sub-sample (G2-b). A similar reversal in the sign of the regression coefficient also occurs for the unemployment rate of hours \( U^h \). The magnitudes of the other parameter estimates change dramatically in the two independent samples, (G2-a) vs. (G2-b). The illusion of stability is created by comparing results for the entire sample period to those for a sub-sample which contains most of the observations that were used to fit the equation for the entire period. Although most of the parameter estimates of (G2-b) are insignificant (judged by t-tests), the entire equation was statistically significant at conventional
levels even if it has only $T_2 - k = 4$ d.f. A comparison of (G3-a) and (G3-b), [with $T_1 = 32$ and $T_2 = 36$ observations], reveals that even though the two parameter vectors could have come from the same population, the point estimates of parameters fluctuate wildly. With a large parameter space, ($k = 12$ in the Gordon wage equation) and colinearity among the explanatory variables, statistical stability of the regression coefficients may be consistent with an unacceptably large confidence ellipsoid that can accommodate wide variations in parameter estimates. What is important for public policy is not statistical stability alone, but rather tight confidence ellipsoids for statistically stable parameter estimates.

The method of dummy variables was used by Perry (1970) and Ekstein-Brinner (1972) to measure the impact of the Kennedy-Johnson guideposts on wage inflation.\textsuperscript{10} The appropriateness of this method depends on the validity of two assumptions: (a) stability of the slope parameters and (b) homoscedasticity. To test the validity of the former assumption, I re-estimated a simplified version of the Perry wage equation in which the rate of change in secondary employment (which was insignificant in Perry's preferred equation) was omitted. The dependent variable which was an overlapping four quarter rate of change in Perry's study was redefined in this replication to represent one quarter rates of change. In the first column,

\textsuperscript{10} Ekstein and Brinner included a guidepost dummy in their wage equation but not in their price equation. Presumably, guideposts only affected price movements through their impact on unit labor costs. This view of the Guideposts is contrary to the verbal analysis supplied by Sheahan (1967). Sheahan's descriptions of the cases involving the steel price rollback, copper prices, hides, etc. indicate that the Guideposts were to some extent, intended to have a direct effect on prices in addition to any indirect effects through changing unit labor costs.
(P1), of Table 2, I present the least squares estimates for my version of
the Perry wage equation including his Guideposts dummy variable \( G \).\(^{11}\) Five
sets of parameter estimates corresponding to the entire sample period,
years 1954:1-1970:4, and four sub-samples are then obtained for a wage equation
excluding the Guidepost dummy. It seems to be generally agreed that the
Guideposts were in effect from 1963:1 through 1966:4, but there is some
doubt about its efficacy at the start and end of the Guidepost experience.
The results shown in (P3-a) and (P3-b) assume the longer six-year period,
1962:1-1967:4 for the Guidepost era. My results in (P1) differ markedly
from those reported by Perry, [confer Appendix equation (A.1) below]. Part
of the differences is due to different sample periods and omission of the
secondary employment variable, but most is probably due to the definition
of the wage change variable. When Gordon tried to replicate Perry's wage
equation using one and two quarter rates of wage change, he also got results
that were very different from Perry's parameter estimates.

In my replication of the Perry wage equation, (P1) in Table 2, the
Guidepost dummy has the highest \( t \)-value of 6.95. The elimination of
\( G \) in (P2) thus leads to a much larger standard error \( \sigma_e \). However, the
issue at hand is not the standard error but rather the stability of
The \( G \) parameters in the periods of "guideposts off" vs. "Guideposts on"; i.e.

\(^{11}\) Perry's dummy variable, \( G \), is not the usual 0-1 dummy variable. In
the strong quarters of the Guidepost era, 1962:4-1966:4, \( G \) takes on the
value of one. However, it climbs up to one from a value of 0.25 in 1962:1, 0.50 in 1962:2, and 0.75 in 1962:3. In a similar manner, it linearly
retreats in 1967:1 and reaches a value of zero in 1967:4. The same ad hoc
Guidepost dummy variable was adopted by EB in their wage equation.
Table 2

Replications of the Perry Wage Equation for Alternative Samples*

<table>
<thead>
<tr>
<th>Variable</th>
<th>Equation**</th>
<th>(P1)</th>
<th>(P2)</th>
<th>(P3-a)</th>
<th>(P3-b)</th>
<th>(P4-a)</th>
<th>(P4-b)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept $\alpha_0$</td>
<td></td>
<td>0.0014</td>
<td>0.0028</td>
<td>0.0008</td>
<td>0.0003</td>
<td>0.0013</td>
<td>-0.0095</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.29)</td>
<td>(1.96)</td>
<td>(0.64)</td>
<td>(0.10)</td>
<td>(1.07)</td>
<td>(-0.99)</td>
</tr>
<tr>
<td>Perry Un. $1/U_Y$</td>
<td></td>
<td>0.0001</td>
<td>0.0006</td>
<td>0.0001</td>
<td>0.0006</td>
<td>0.0002</td>
<td>-0.0013</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.46)</td>
<td>(1.73)</td>
<td>(0.33)</td>
<td>(0.72)</td>
<td>(0.70)</td>
<td>(-0.69)</td>
</tr>
<tr>
<td>Dispersion $D$</td>
<td></td>
<td>0.1065</td>
<td>0.0277</td>
<td>0.1180</td>
<td>0.0417</td>
<td>0.0922</td>
<td>0.2619</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(4.31)</td>
<td>(0.95)</td>
<td>(3.92)</td>
<td>(0.56)</td>
<td>(3.13)</td>
<td>(1.26)</td>
</tr>
<tr>
<td>Price Infl. $P_{t-1}$</td>
<td></td>
<td>0.0922</td>
<td>0.2788</td>
<td>0.0613</td>
<td>0.1339</td>
<td>0.1343</td>
<td>0.2416</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.37)</td>
<td>(3.41)</td>
<td>(0.75)</td>
<td>(0.73)</td>
<td>(1.70)</td>
<td>(0.81)</td>
</tr>
<tr>
<td>Pers. Tax $\psi$</td>
<td></td>
<td>1.0008</td>
<td>0.9380</td>
<td>0.9803</td>
<td>0.9023</td>
<td>1.0161</td>
<td>1.4158</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(5.21)</td>
<td>(3.69)</td>
<td>(3.10)</td>
<td>(2.85)</td>
<td>(3.33)</td>
<td>(2.43)</td>
</tr>
<tr>
<td>Guidepost $G$</td>
<td></td>
<td>-0.0179</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-6.95)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Summary Statistics

- $R^2$:
  - (P1) 0.7601
  - (P2) 0.5733
  - (P3-a) 0.7534
  - (P3-b) 0.6414
  - (P4-a) 0.7142
  - (P4-b) 0.6663

- $\sigma_\epsilon$:
  - (P1) 0.00199
  - (P2) 0.00263
  - (P3-a) 0.00213
  - (P3-b) 0.00195
  - (P4-a) 0.00219
  - (P4-b) 0.00204

- DW:
  - 1.80
  - 1.21
  - 1.57
  - 1.77
  - 1.47
  - 2.44

- No. of Obs.:
  - 68
  - 68
  - 44
  - 24
  - 52
  - 16

*The dependent variable is the one quarter rate of change in hourly total compensation $c_t$. The price inflation variable is the lagged one quarter rate of change in the CPI. The t-statistics are shown in parentheses.

**The six equations correspond to the following sample periods:

- (P1) and (P2) = 1954:1-1970:4
- (P3-a) = entire period excluding 1962:1-1967:4
- (P3-b) = 1962:1-1967:4
- (P4-a) = entire period excluding 1963:1-1966:4
- (P4-b) = 1963:1-1966:4
We want to compare (P3-a) vs. (P3-b) and (P4-a) vs. (P4-b). The Chow test for these two comparisons produced F-ratios of $F = 8.582$ and $F = 7.003$ which exceed the 1 percent critical value of $F_{01} = 3.38$ for $n_1 = 5$ and $n_2 = 58$ degrees of freedom implying rejection of the hypothesis of stable coefficients. Thus, even if we accept the logic of the Perry wage equation, we must, in the light of the sample data, reject the dummy variable method for measuring the impact of Guideposts in restraining the rate of wage inflation. The smaller parameter space of the Perry wage equation, $[k = 5$ parameters as opposed to $k = 12$ in the Gordon wage equation] could possibly have increased the power of the F test. Time and data availability precluded an extension of the Chow test to the Ekelstein-Brinner wage equation. The use of the dummy variable method to quantify the effect of a non-quantifiable variable [such as the presence or absence of Guideposts] can only be justified on statistical grounds if an F test for the stability of regression coefficients implies acceptance of the null hypothesis that $\beta_1 = \beta_2$ in the two sample periods.

Lipsey and Parkin (1970) presented an even more devastating analytic attack on the inappropriateness of using dummy variables to measure the impact of guidelines and controls. According to LP, controls might affect the time rates of change in wages and prices in three ways: (a) by changing the values assumed by the explanatory variables, (b) by altering the coefficients of the wage and/or price equations, or (c) by replacing the previous equations with new relationships linking $w_t$ and $p_t$ to different explanatory variables. If the (a) path held, the dummy variable
would be insignificant implying that the controls had no effect. Under (b) and (c), the coefficient of the dummy variable could be positive or negative even if controls were effective depending on initial conditions, how the parameters were shifted, or the relationship between the new variables and the old explanatory variables in the wage/price equations. The dummy variable technique would be the analytically correct one in only one special case of (b); namely when the controls policy changed only one parameter, the intercept term in the wage or price equation. The slope parameters and the variables appearing in the wage and price equations must be the same in both periods of controls and non-controls.

2.3 Correspondence between Conceptual and Empirical Variables

Measurement is one of the more neglected aspects of econometrics. Index numbers and time series data, compiled and published by government and private bureaus are indiscriminately equated to the conceptual variables of a model. Perry and Gordon have been quite imaginative in constructing complex variables from the published series; details of some of these are described in the Appendix. Attention is directed here to (A) the Gordon fixed weight wage index, and (B) the relationship between unemployment measures and the relative excess demand for labor services.

(A) The Gordon Fixed Weight Wage Index: R.J. Gordon (1971) introduced his fixed weight wage index which incorporated adjustments for overtime in

\footnote{It should be pointed out that if the controls and guidelines only affected the values of the explanatory variables, the simulation approach would also fail to capture the effects of controls. In the models reviewed here, the explanatory variables include unemployment measures \( U \), tax rates \( \tau \), labor productivity \( q \), and the ratio of unfilled orders to capacity, \( uf/k \).}
manufacturing, inter-industry employment shifts, and fringe benefits. In the first adjustment, the hourly wage rate in manufacturing industries was measured by straight-time average hourly earnings of production workers net of overtime premiums. Wage rates in the other industries were measured by average hourly earnings including overtime. In the second adjustment, these hourly wage rates were combined into a fixed weight index \( W_1 \) using 1963 employment weights.\(^{13}\) The third adjustment was made by multiplying \( W_1 \) by the ratio of total employee compensation to wages and salaries as reported in the Commerce statistics. If we let \( C = \) hourly total compensation and \( W_2 = \) hourly wages and salaries as reported in the Commerce statistics, the Gordon wage index \( W \) is defined by,

\[
(2.6a) \quad W = W_1 \left( \frac{C}{W_2} \right),
\]

where upper case letters denote the levels of the wage indexes at time \( t \). The dependent variable of the wage equation is the relative time rate of change indicated by lower case letters.

\[
(2.6b) \quad w = \left( \frac{1}{W} \right) \left( \frac{dW}{dt} \right) = \left( \frac{1}{W_1} \right) \left( \frac{dW_1}{dt} \right) + \left( \frac{1}{C} \right) \left( \frac{dC}{dt} \right) - \left( \frac{1}{W_2} \right) \left( \frac{dW_2}{dt} \right) = w_1 + c - w_2.
\]

The definition in (2.6a) allows us to derive an equivalent expression for the rate of change in the Gordon wage index.

\[
w = w_1 + \alpha(f - w_2)
\]

\(^{13}\) For the period prior to 1964, Gordon calculated this fixed weight wage index \( W_1 \) from the disaggregated industry data on average hourly earnings. Since 1964, the Bureau of Labor Statistics has compiled and published this series; confer Handbook of Labor Statistics, 1972, Table 98, p. 221.
where \( \alpha = \frac{F}{C} \) is the ratio of fringe benefits to total compensation, and
\( f = \left( \frac{1}{F} \right) \frac{dF}{dt} \) is the relative rate of change in hourly fringe benefits
unadjusted for inter-industry employment shifts or overtime. It is reasonable to suppose that workers would forgo higher straight-time wage for more non-taxable fringe benefits. But the third adjustment which ignores inter-industry employment shifts seems to assume that the marginal rate of substitution between wages and fringes is a constant. In my opinion, the hourly total compensation \( C \) (the variable used by Perry) is to be preferred to the Gordon fixed weight wage index \( W \) as a closer approximation to the conceptual dependent variable of a wage equation.

Finally, the difference in the relative rates of change in hourly total compensation and the Gordon wage index, \( (c-w) \) is included as an explanatory variable in the G and EB price equations, on the ground that \( W \) excluded the labor costs of supervisory workers and overtime work. The unit labor cost variable in these price equation is straight-time wages.

The hourly total compensation \( C \) can be decomposed into four components: \( W_1 \) = hourly straight-time wages for a fixed industry mix, \( I \) = an hourly industry mix adjustment which can be positive or negative, \( J \) = the average hourly outlays for overtime and workers not included in the straight-time wage rate, and \( F \) = hourly fringe benefits. Thus, \( W_2 = W_1 + I + J \), and \( dC = dW_2 + dF \). Division by \( C \) yields,

\[
\frac{dC}{C} = \left( \frac{1}{W_2} \right) \frac{dW_2}{W_2} + \left( \frac{F}{C} \right) \left( \frac{dF}{F} \right) = (1-\alpha)w_2 + \alpha f
\]

If the expression on the right is substituted for \( c \), we get the equivalent expression. The Commerce statistics on fringe benefits (the difference between total compensation and wages and salaries, \( C-W_2 \)) excludes some fringe benefits such as paid holidays and vacations, sick leave privileges, etc. If we properly calculated fringe benefits "per hour worked" rather than "per hour paid", we find substantial differences in hourly fringe benefits across industries and establishment size categories. The larger firms in banking and finance, and in the unionized firms in mining, construction, transportation, and manufacturing pay much higher hourly fringe benefits. A shift in employment toward the "high fringe benefit" industries and firms would be accompanied by a rise in average hourly fringe benefits \( F \) for the entire private labor force even if there were no change in \( F \) in each industry and firm.
adjusted for inter-industry employment shifts. If prices are theoretically determined by all labor costs, the conceptually correct measure of unit labor costs should have been based on hourly total compensation per unit of output, \( c/q' \) instead of straight-time labor costs (inflated for fringe benefits) given by the Gordon index \( w/q' \). An alternative way of accomplishing this would have been to constrain the parameter of \( c/w \) in the price equation to be equal to the parameter of standard unit labor cost \( w/q' \) given that the same lag structure is imposed on both variables.

(B) Unemployment and Labor Market Disequilibrium: Lipsey (1960) developed a simple model linking the unemployment rate \( U \) to the relative excess demand for labor services, \( \frac{D-S}{S} \). Holt (1967) and others have extended the model by introducing the number of job vacancies, labor turnover variables, and labor force characteristics, etc. as determinants of the disequilibrium in some aggregate labor market. The LP and EB wage equations retain the original simple structure wherein disequilibrium in the labor market is measured by a single variable, the official unemployment rate \( U^0 \). Perry that in the late 1960's, the official unemployment rate failed to give an accurate reflection of tightness in the labor market because of the changing age/sex composition of the labor force. He proposes to measure disequilibrium in the labor market by a linear combination of two variables, (a) the reciprocal of an income-weighted unemployment rate \( 1/U^Y \) and (b) an index

15/ In the LP wage equation, \( U^0 \) is entered linearly contrary to Lipsey's rationale for using the reciprocal of \( U^0 \). The linear form was chosen to facilitate estimation and to direct attention to the shift in the \( (w, U) \) trade-off during periods of income policy. The EB wage equation used the familiar functional form of \( 1/U^0 \).
of the relative dispersion of unemployment rates $D$.\textsuperscript{16/} Gordon rejects the $1/U^Y$ variable on statistical grounds, accepts dispersion $D$, and adds two variables of his own, the disguised unemployment rate $U^d$ and an unemployment rate of hours $U^h$.

Perry begins with the obviously correct premise that all members of the labor force do not provide equivalent flows of labor services. However, the official unemployment rate $U^0$ attaches the same weight to all unemployed persons whether they be black teen-agers or middle-aged engineers. Perry argues that the downward pressure on an aggregate wage index will be greater when the pool of unemployed persons contains larger fractions of individual who, on average, earn higher hourly wages and work longer hours per year. He thus weights both the labor force and the pool of unemployed persons, (classified by age and sex but not color), by their average annual wage earnings relative to the wage earnings of prime age males, 35-44 years of age. Females and young males who typically experience higher unemployment rates and who are accounting for an increasingly larger share of the labor force, receive much smaller weights in the Perry unemployment rate $U^Y$. For a given age/sex mix, the cyclical fluctuations in $U^Y$ are smaller than those in $U^0$. Holding age/sex specific unemployment rates constant, the dispersion of $U^Y$ is less than that of $U^0$ when the age/sex mix is varied.\textsuperscript{17/}

\textsuperscript{16/}Perry also included the rate of change in secondary employment as a third labor market variable, but its coefficient was insignificant.\textsuperscript{17/} The formulas defining $U^Y$ and $D$ can be found in Appendix A.1. The propositions about the dispersion of $U^Y$ and $U^0$ are based on a numerical example presented in the Appendix.
link his rate to a theoretically correct measure of equivalent labor services in \( \frac{D-S}{S} \), the spirit of his approach is to be applauded.

Perry's dispersion index D is a weighted average of the absolute values of the relative divergence between age/sex specific unemployment rates \( U_i \) and his weighted unemployment rate \( U^X \). Lipsey (1960) argues that for any given aggregate unemployment rate, the rate of wage inflation would be higher, the larger is the dispersion (across labor markets) in unemployment rates. This rationale presumably prompted the Perry dispersion index. If age/sex is the criterion for identifying "separate" labor markets, one cannot use the implicit assumption in the Lipsey analysis that the unemployment rate \( \bar{U} \) which is consistent with zero excess demand is the same across markets. Loosely speaking, the natural rate of unemployment (which we temporarily equate to \( \bar{U} \)) is that rate which minimizes the sum of search costs for workers and firms given some unspecified, but presumably stable, distribution of stochastic demands for labor services. According to Hall (1972), the typical teen-ager may hold three or four jobs a year resulting in a very high turnover rate and frequent spells of unemployment between jobs. Part of this turnover probably reflects the youth's search for a lifetime occupation. A typist who is seeking employment is likely to incur lower search costs and shorter search time than one who now decides it is time to seek other, and hopefully more satisfying, employment in another occupation. The natural rate for prime age males who, for the most part, are set in their occupational affiliations, is likely to be far lower than that of teen-agers entering the labor force or females re-entering the labor market. A dispersion index which comes closer to the Lipsey rationale for this variable, is one which measures
the divergence between actual and natural rates, \((U_i - \bar{U}_i)\), where the natural rate \(\bar{U}_i\) varies across age/sex groups.

A major problem with the Gordon unemployment variables is that statistical criteria, such as goodness of fit and plausible t-values, determine the variables that appear in his preferred equation. Prior theoretical considerations carry little weight in the choice. Thus, in Gordon (1971), he defines a total unemployment rate of labor services, \(U^L\), as the sum of (a) the official unemployment rate \(U^0\), (b) a disguised unemployment rate of discouraged workers \(U^d\), and (c) an unemployment rate of hours \(U^h\).

\[ U^L = U^0 + U^d + U^h \]

The wage change component \(w^d\) due to disequilibrium in the labor market could be related to this total unemployment rate in either a nonlinear or linear fashion. The former, \([w^d = \beta(1/U^L)]\) is not attempted. A simple linear relation, \(w^d = A + B U^L\), implies a model specification in which the parameters of \(U^0\), \(U^d\), and \(U^h\), are constrained to be the same. Gordon, however, adopted the procedure of allowing the data to determine the parameters which resulted in the omission of \(U^0\) in his wage equation. This latter procedure tacitly implies a definition for the total unemployment rate as a weighted sum of the three components, namely, \(U^L = b_0 U^0 + b_1 U^d + b_2 U^h\) where the weights are determined by the sample data. Even in

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18/ It is shown in the Appendix that \(U^d\) can be interpreted as a nonlinear trend adjustment based on the reciprocal of the labor force participation rate, and that \(U^h\) is also a trend adjusted series for average weekly hours.
this context, the results are unacceptable since \( b_0 = 0 \) meaning that official unemployment contributes nothing to total unemployment.

The published wage equations appear to place too much reliance upon regression results based on shaky data, and not enough on the underlying conceptual variables of the theory. If an economist were asked to estimate a demand function for jelly, the theory of consumer choice demand indicates which variables ought to be included in the equation. If the sample data generated a positive or insignificant own price elasticity, the economist would probably question his data or econometric model.

2.4 Specification Errors

Econometric models almost always involve some specification errors of omitting pertinent variables. In many instances, the biases resulting from these omissions are negligible. In this section, I would like to remain within the spirit of these two-equation models, and direct attention to what I regard, as three possibly important omissions in the specification of the wage equation.

A. Proportion of Workers Getting Wage Changes: An individual's hourly money wage rate, \( W_{it} \), can differ from his wage in the preceding quarter, \( W_{it-1} \), for two reasons. First, he could have changed jobs, or second, the wage rate on his existing job could have been changed by his employer. The fraction of workers falling under the first category depends on promotion (demotion) rates within firms and turnover rates of labor between firms. The fraction in the second category is determined by the frequency and timing of wage change decisions. The importance of this factor was brought to my attention by O. Ashenfelter who together with
J. Pencavel is incorporating this variable in an empirical analysis of the Phillips curve for the United Kingdom; see Ashenfelter and Pencavel (1974).

Let \( \lambda_t \) denote the proportion of workers getting wage changes in the \( t \)-th quarter due either to job changes or wage rate changes by employers. If \( w_{1t} \) is the relative time rate of change in wages for workers getting wage changes, [meaning that \( w_{2t} = 0 \) for workers not getting wage changes], the relative time rate of change in the aggregate wage rate will be the weighted average of \( w_{1t} \) and \( w_{2t} = 0.19/\)

\[
(2.7) \quad w_t = \lambda_t w_{1t}
\]

It seems reasonable to suppose that the rate of wage change, \( w_{1t} \), for workers getting wage changes, should depend on some vector of explanatory variables, \( X_t \), representing unemployment measures, price inflation variables, etc. Thus, we have,

\[
(2.8a) \quad w_{1t} = \alpha + \beta X_t + \epsilon_t.
\]

If we substitute (2.8a) into (2.7), we get a wage equation in which the observed rate of change in an aggregate wage rate index, \( w_t \), is related to the proportion of workers getting wage changes, \( \lambda_t \), and a vector of explanatory variables \( X_t \) that affect the magnitude of wage changes when they occur.

\[19/\] Equation (2.7) tacitly assumes that the level of wages, \( W \), of workers getting wage changes is equal to the average wage rate \( \bar{W} \) of all workers. If the change group had a different average wage, \( k = (W_t / \bar{W}) \neq 1 \), equation (2.7) would have to be replaced by,

\[
w_t = k \lambda_t w_{1t}
\]
(2.8b) \[ w_t = \alpha \lambda_t + \lambda (\lambda_t X_t) + u_t, \]

where \( u_t = \lambda_t \epsilon_t \) is the random disturbance in the wage equation.

Ashenfelter and Pencavel (1974) found that the proportion of British workers getting wage rate changes by employers, varied widely from quarter to quarter. When they included \( \lambda_t \) and the interaction of \( \lambda_t \) and \( X_t \) in a Phillips-type wage equation, the explanatory power rose sharply.

It is my understanding that the people who forecast changes in aggregate money wages, do indeed, look at the timing of major wage contracts to determine the fraction of workers who will be getting wage changes. If the steel, auto, rubber, and chemical workers are all locked into three-year wage contracts, it is a simple matter to estimate their wage changes between contracts. However, if all of these unions were negotiating contracts in the third quarter of 1974, the change in the aggregate wage index for 1974:3 is far more likely to reflect the existing and anticipated labor market conditions at that time. Even in the non-union sector, workers are typically under explicit or implicit contracts that involve periodic, (possibly annual) reviews of hourly wage rates. Moreover, I would conjecture that in periods of rapidly changing excess demands for labor or unanticipated cost-of-living changes, the proportion of workers getting wage changes \( \lambda \), would rise. The omission of this variable, \( \lambda = \) the proportion of workers getting wage changes, not only enlarges the size of the unexplained residual variance, but also makes it impossible to estimate the parameters of the true wage change equation given by (2.8a).

B. Unemployment Insurance Programs: The extent of coverage under unemployment insurance programs and the size of benefits should, in principle,
affect the duration and rate of unemployment. Greater coverage and higher unemployment benefits (relative to weekly wages) should lengthen the optimum search period thereby raising the "natural rate" of unemployment. Schmidt (1973) found that the mean duration of unemployment was considerably longer for male workers who received unemployment compensation. Covered unemployment (by State and Federal unemployment insurance programs) climbed from 57.6 to 72.0 percent of the civilian labor force over the period 1954 to 1970. That the unemployment rate of covered workers is below the rate for all workers, (rows 2-a and 2-b of Table 3), seems to contradict the thesis advanced here. However, when a worker exhausts his unemployment benefits, (usually 26 weeks), he no longer is counted among the "insured unemployed" and is shifted to the uninsured unemployed pool. Unemployment benefit payments as a percentage of either gross weekly earnings, row 5-a, or of spendable weekly earning for a worker with three dependents, row 5-b have climbed at rather modest annual compound growth rates of 0.34 and 0.66 percent a year over the period, 1960-70. The data shown in Table 3 reveal that coverage under unemployment insurance programs has substantially increased, and benefit payments have risen moderately. The omission of this factor may not be serious in an aggregate wage equation, but it may be important in explaining the level and composition of unemployment.

C. Wages in the Public Sector: All of the published wage equations for the U.S deal with the rate of change in money wages in the "private nonfarm sector". Two reasons have been advanced to justify the choice
Table 3

**Insured Unemployment and Average Weekly Benefits**
*(selected years, 1954-70)*

<table>
<thead>
<tr>
<th>Item</th>
<th>1954</th>
<th>1960</th>
<th>1965</th>
<th>1970</th>
<th>Mean&lt;sup&gt;a&lt;/sup&gt;</th>
<th>Trend&lt;sup&gt;b&lt;/sup&gt;</th>
<th>Growth Rate 1960-70</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Insurance Employment as percent of Civ. labor force</strong></td>
<td>57.6</td>
<td>66.5</td>
<td>69.3</td>
<td>72.0</td>
<td>67.7</td>
<td>0.83</td>
<td>0.80</td>
</tr>
<tr>
<td><strong>Unemployment rates</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>2-a. Civ. labor force</td>
<td>5.55</td>
<td>5.53</td>
<td>4.52</td>
<td>4.94</td>
<td>4.91</td>
<td>-0.08</td>
<td>-1.12</td>
</tr>
<tr>
<td>2-b. Insured Emp.</td>
<td>5.60</td>
<td>4.47</td>
<td>2.81</td>
<td>3.48</td>
<td>3.82</td>
<td>-0.18</td>
<td>-2.47</td>
</tr>
<tr>
<td>Avg. Weekly Unemployment Comp.</td>
<td>24.93</td>
<td>32.87</td>
<td>37.19</td>
<td>50.34</td>
<td>35.10</td>
<td>1.42</td>
<td>4.35</td>
</tr>
<tr>
<td><strong>Average Weekly Earnings</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>4-a. Gross</td>
<td>64.52</td>
<td>80.67</td>
<td>95.06</td>
<td>119.46</td>
<td>88.04</td>
<td>3.19</td>
<td>4.01</td>
</tr>
<tr>
<td>4-b. Spendable 3 dependents</td>
<td>60.85</td>
<td>72.96</td>
<td>86.30</td>
<td>104.61</td>
<td>79.42</td>
<td>2.56</td>
<td>3.57</td>
</tr>
<tr>
<td><strong>Unemployment Comp. as percent of Avg. Weekly Earnings</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>5-a. Gross</td>
<td>38.64</td>
<td>40.75</td>
<td>39.12</td>
<td>42.14</td>
<td>39.72</td>
<td>0.17</td>
<td>0.34</td>
</tr>
<tr>
<td>5-b. Spendable</td>
<td>40.97</td>
<td>45.05</td>
<td>43.09</td>
<td>48.12</td>
<td>43.92</td>
<td>0.37</td>
<td>0.66</td>
</tr>
</tbody>
</table>


**Means are based on annual data, 1954-70. Average weekly unemployment compensation costs and spendable earnings of production workers are in current dollars.**

**The figures here are the least squares slope of a linear trend line.**
of this wage rate series: (1) wages in the private nonfarm sector are
determined by market forces, and (ii) the private nonfarm sector generates
a majority of the GNP.\textsuperscript{20}

The public sector is demanding an increasingly larger share of the
nation's labor force. Civilian employees of Federal, State and Local
governments accounted for 13.4 percent of total civilian employment in
1960 and 16.6 percent in 1970. In Table 4, I have assembled data on
the time path of wages for selected industries and groups of workers.
The last two columns present the annual compound growth rates measured
between the end-points of the periods, 1964-70 and 1967-70.

The time paths of the fixed weight wage indexes varied slightly
across industries with Construction exhibiting the most rapid growth in
the last four years, Panel B of Table 4 compares the indexes for four
measures of wages of production workers in all manufacturing industries.
The differences among these four measures are seen to be quite small.

The most striking comparison is that between panel C, [which shows
salary indexes for three groups of public sector employees] and either

\textsuperscript{20} According to Table C-10, The Economic Report of the President, 1972
p. 205, the percentage of GNP generated by the "private nonfarm sector"
was as follows for selected years:

<table>
<thead>
<tr>
<th>Year</th>
<th>Percent</th>
</tr>
</thead>
<tbody>
<tr>
<td>1954</td>
<td>82.3</td>
</tr>
<tr>
<td>1960</td>
<td>83.4</td>
</tr>
<tr>
<td>1965</td>
<td>85.0</td>
</tr>
<tr>
<td>1968</td>
<td>85.4</td>
</tr>
<tr>
<td>1970</td>
<td>85.3</td>
</tr>
</tbody>
</table>

Part of the GNP generated by the private sector is, however, the result
of government purchases of goods and services.
The rates were converted to percentages for this table.

\[ M_{70} = (1 + \frac{g}{10})^6 \]

Growth rates at the solution to the following equation:

\[ \frac{M_4}{M^{64}} \]

and \( M^{70} \) denote the wage indexes in 1964 and 1970. The

\[ \frac{g}{10} \]

appearing on pp. 221-1, 222-6, and 262-3.

Source: Handbook of Labor Statistics, (1972), selected tables

<table>
<thead>
<tr>
<th>Year</th>
<th>Wage Index</th>
<th>Wage Index</th>
<th>Wage Index</th>
<th>Wage Index</th>
<th>Wage Index</th>
<th>Wage Index</th>
</tr>
</thead>
<tbody>
<tr>
<td>1964</td>
<td>7.06</td>
<td>128.0</td>
<td>117.0</td>
<td>89.40</td>
<td>0.90</td>
<td>0.90</td>
</tr>
<tr>
<td>1965</td>
<td>7.77</td>
<td>122.5</td>
<td>114.0</td>
<td>93.50</td>
<td>0.90</td>
<td>0.90</td>
</tr>
<tr>
<td>1966</td>
<td>6.42</td>
<td>110.0</td>
<td>107.0</td>
<td>99.00</td>
<td>0.90</td>
<td>0.90</td>
</tr>
<tr>
<td>1967</td>
<td>5.19</td>
<td>106.5</td>
<td>101.0</td>
<td>95.7 0</td>
<td>0.90</td>
<td>0.90</td>
</tr>
<tr>
<td>1968</td>
<td>4.46</td>
<td>106.6</td>
<td>101.0</td>
<td>93.6 0</td>
<td>0.90</td>
<td>0.90</td>
</tr>
<tr>
<td>1969</td>
<td>4.77</td>
<td>106.7</td>
<td>101.0</td>
<td>93.6 0</td>
<td>0.90</td>
<td>0.90</td>
</tr>
<tr>
<td>1970</td>
<td>5.94</td>
<td>106.6</td>
<td>101.0</td>
<td>93.6 0</td>
<td>0.90</td>
<td>0.90</td>
</tr>
</tbody>
</table>

**Table 4**

Indexes of hourly wage rates and salaries, 1964-70
panels A or B. Over the period, 1967-70, the average salaries of Federal
civil servants climbed at an annual compound rate of 9.1 percent compared
to an annual rate of increase of only 6.6 percent in the hourly wage rates
for the total private nonfarm sector. The maximum salaries of firemen
and policemen (in cities of more than 100,000 persons) rose at an annual
rate of 8.6 percent, and the average salaries of all teachers by 7.0
percent. The period, 1967-70, was one of rapid wage inflation. The data
of Table 4 conclusively demonstrate that over this period, wages in the
public sector, (which makes up one-sixth of the national labor market)
rose relative to wages in the private sector.

Most of the recent developments in the wage equation literature
were prompted by the inability of earlier wage equations to track the wage
explosion of 1967-70. The Perry dispersion index and the Ekstein-Brinner
threshold price variables would both have been highly correlated with a
dummy variables for the last three or four years of the sample period,
1967-70. These variables have replaced some of the explanatory variables
in earlier wage equations, (such as corporate profits) because their
inclusion enables the wage equation to track the recent historical
experience. The data of Table 4 suggest that the time path of public
sector wages closely follows the time paths of unemployment dispersion and
the EB threshold price variable. The omission of public sector wages from
a wage equation for the private nonfarm sector may have imparted signif-
icant specification biases not only in earlier wage equations but also
in the-current P, G, and EB models. The rationale for its inclusion is
incomplete because I have not developed a theory of wage determination
for the public sector. However, when wages rise in the public sector, competition for labor will compel employers to raise their wages in order to retain qualified workers.

One could point to other omitted variables such as the conspicuous absence of any monetary or fiscal policy variables, but I shall defer discussion of these omissions to the concluding section.

III

Measurement by Simulation

Measurement by simulation is a familiar technique that was employed by Lipsey-Parkin (1970) to estimate the quantitative impact of British incomes policy. Wage and price equations are estimated from sample data for periods when there were no controls. The values that were actually observed for the explanatory variables during the "controls on" periods are substituted into the fitted equations to obtain forecasts of the precipitated rates of change in wages and prices, \( \hat{w}_t, \hat{p}_t \) that would have prevailed if there had been no controls. The residuals between actual and predicted rates of change, \( e_t = (w_t - \hat{w}_t) \), are then interpreted as the "best estimates" of the quantitative impact of controls in restraining the rates of wage and price inflation. In the G and EB models, the simulations are somewhat more complicated for two reasons: (a) assumptions must be invoked about the trend rates of growth in labor productivity, potential labor force, and potential hours of work in order to construct the values for some of the explanatory variables, and (b) initial conditions [meaning the past histories of wages, prices, tax rates, and labor productivity], now affect the calculations because of the distributed lags for
some variables in the model. These simulations that utilize data on the actual values of explanatory variables in the "controls on" period, will be referred to as ex post simulations.

Simulations of another kind which I shall call ex ante simulations are also developed by G and EB to assess the implications of alternative macro-economic policies. Values of the explanatory variables of the model are generated to conform to alternative policy scenarios. Gordon (1971) Appendix B, developed a system of eleven equations which generate the values of the explanatory variables for his ex ante simulations. An examination of this system of eleven equations reveals that all of the explanatory variables of the model are determined by only one exogenous variable, X = the level of real GNP at time t. High and low unemployment policy scenarios can thus be traced back to alternative assumptions about the time path of real GNP. 21/ I shall not discuss these ex ante simulations but shall instead, evaluate the merits of the ex post simulations as a technique for measuring the quantitative impact of wage-price controls.

If the assumptions of the econometric model are fully met, the ex post simulation approach is surely the correct method for measuring the impact of wage-price controls. The key assumptions that must be met include the following:

(a-1) The explanatory variables of the model are truly predetermined, exogenous variables.

21/ Gordon (1972) argues on p. 385 that the level of nominal GNP is exogenous and determined by current and past monetary and fiscal policies. His model is supposed to explain the division of nominal GNP between real output and price level changes. However, as I read his description of the simulations, this position is forgotten, and real GNP is the driving force in simulating the time paths of wage and price inflation.
(a-2) The structure of the model including the functional form of equations, control variables appearing in the equations, and structural parameters are shown stable and appropriately estimated.

(a-3) The random disturbances of the model are orthogonal to the explanatory variables and satisfy the usual assumptions of normality and serial independence.

The quality and quantity of sample data are unfortunately inadequate to perform powerful statistical tests on these assumptions. The validity of these assumptions must thus be evaluated on the bases of theoretical considerations, judgements, and circumstantial evidence.

3.1 Stability of the Structural Equations

Let me turn first to the assumption of a stable structure. Assume for the moment that the explanatory variables of the model are truly exogenous with respect to the wage-price inflation process. The explanatory variables in the G and EB models include various unemployment variables and the rates of change in tax rates $\Psi$, the ratio of current to trend labor productivity ($q/q'$), and the ratio of unfilled orders to capacity ($uf/k$). At least two considerations suggest that the structural equations are likely to shift in anticipation of or in response to the imposition of truly binding controls. In the simulation approach, the quantitative impact is measured by the sum (or average) of the residuals between actual and predicted rates of wage-price inflation during those periods in which the controls were imposed. Let me denote this sum (or average) by $e_t = (w_t - \hat{w}_t)$ where $t$ represents the period of effective controls. If workers and firms anticipate the imposition of wage-price
controls, it is reasonable to expect that they will try to establish above equilibrium wages and prices so that the controls will have less impact on them. Hence, for any vector of explanatory variables, \( X_{t-j} \), in periods preceding the imposition of controls, the actual rates of wage and price inflation would be larger if the controls program were anticipated.

Consider next, the behavior of wages and prices after the controls are lifted. The structure of the wage and price equations tacitly assume pseudo-equilibrium paths of adjustment in wages and prices in response to the disequilibrating forces described by the unpredictable time paths of the exogenous explanatory variables. When controls are truly binding, workers and firms are unable to achieve these pseudo-equilibrium (and presumably optimal) adjustment paths. The immediate short run response of wage and price changes in the post-controls period, is unlikely to conform to the earlier structural equations in which wages and prices adjusted in a stable fashion to changes in the explanatory variables \( X_{t+j} \). The argument is even stronger for the distributed lag structures built into these models. Suppose that the rates of consumer price changes were held equal to zero for a period of three years as the result of a truly binding controls program. If prices begin to climb when the controls are ended, is it reasonable to expect that workers will still use the same long distributed lag of prior actual consumer price inflation rates in calculating the expected rate of price inflation \( p^* \)? I suspect not. Anticipation of and reaction to wage-price controls can thus be expected to alter the structure of the wage/price equations in the periods preceding
and following the effective controls period. Assumption (a-2) is thus violated, and a correct assessment of the full impact of wage-price controls calls for simulations over a longer comparison period. We must estimate not only the divergence between actual and predicted rates of wage inflation during the controls period, \( e_t = (\hat{w}_t - \hat{\hat{w}}_t) \) but also the discrepancies in the pre- and post-controls periods, \( e_{t-j} = (\hat{w}_{t-j} - \hat{\hat{w}}_{t-j}) \) and \( e_{t+j} = (\hat{w}_{t+j} - \hat{\hat{w}}_{t+j}) \). Lipsey and Parkin (1970) partially allowed for these structural changes by adopting a very broad definition of the "policy on" quarters; when in doubt, the quarter was included in the "incomes policy on" sample.

Second, a policy of selective controls could alter the structure of the wage/price equations. An extreme scenario serves to illustrate this point. In the price equation, the change in product prices, \( p^X \), is assumed to be linearly related to standard unit labor costs \( (w/q')^* \). This formation tacitly assumes a relatively stable output mix for the economy. Suppose that binding controls were imposed on only one sector, say Construction. Wages and prices in this sector are fixed at below market equilibrium values. As a consequence, labor and other resources are shifted to other sectors resulting in a very different output mix for the private nonfarm sector. If the relationship between price changes and changes in unit labor costs in Construction differed markedly from the relationship in other sectors, the parameters of an aggregate price equation would be altered by such a policy of selective controls. Under the Nixon plan, wage and price controls are only applied to big unions and large firms. If binding, these selective controls would tend to divert
resources toward small firms and non-unionized sectors. In the FRB-MIT model, separate wage equations were estimated for the union and non-union sectors, and these exhibited very different structures. The Nixon policy of selective controls would thus imply a structural change in an aggregate wage equation applicable to all workers. By simply substituting the observed value of $X_t$ into the fitted wage/price equations, we do not get an unbiased estimate of the expected rate of wage and price inflation in the absence of wage-price controls when the controls policy itself has altered the structural equation.

3.2 Shadow Prices and the Measurement of Price Inflation

Up to now, I have assumed (a-1) that the explanatory variables are truly exogenous. Suppose further that the method of least squares yields sufficiently accurate parameter estimates, and that the model is correctly specified for the non-controls period. One might think that under these ideal conditions, the simulation approach would provide us with an accurate estimate of the quantitative impact of controls in retarding the real rate of inflation. However, when wage-price controls are truly binding, they constitute an additional disequilibrating force

---

It is acknowledged by all that the model is described by a system of two simultaneous equations, but they all estimate the parameters by ordinary least squares. Lipsey and Parkin (1970) fully recognize not only this point but also the fact that explanatory variables like the unemployment rate are not exogenous to wage behavior in the labor market. They argue that in the absence of a full macro-model for the entire economy, it is impossible to avoid the simultaneous equation bias. Ekstein-Brinner claim that they also estimated the parameters by two-stage least squares, but I question whether they used the correct reduced-form equations (which involve complicated distributed lags) to obtain $\hat{\pi}_t$ and $\hat{p}_t^X$ for the 2SLS estimators. None of these authors tries to examine the possible simultaneous equation biases.
which is reflected in our measurement of the dependent variable. The observed (actual) rates of wage and price inflation during the controls period are very different from the real inflation rates.

Let \( p_t \) denote the relative rate of change in an aggregate price index during the controls period. It can be expressed as a weighted average of the relative rates of change in controlled prices, \( p^c \), and uncontrolled prices \( p^u \).

\[
(3.1) \quad p_t = kp^c_t + (1-k)p^u_t, \]

where \( k \) is approximately equal to the weight given to the controlled goods in an aggregate price index. With effective and binding wage-price \( p^c \) is a policy determined variable which is presumably set at a rate below what would have prevailed in the absence of controls. More importantly, binding price controls must be accompanied by rationing when the controlled prices \( P^c \) are below market equilibrium prices. Outputs of the controlled goods will be determined by the controlled prices at \( \bar{x}^c \) along the appropriate product supply curves. Tobin and Houthakker (1951) have shown that under rationing, the demand functions for the uncontrolled goods,

\[23/ \]

The level of an aggregate price index, \( \bar{P} \), is some weighted average of indexes for controlled prices \( \bar{P}^c \) and uncontrolled prices \( \bar{P}^u \).

\[ \bar{P} = k\bar{P}^c + (1-k)\bar{P}^u. \]

If we have a fixed weight price index, differentiation with respect to time yields,

\[
\frac{d\bar{P}}{dt} = \frac{1}{\bar{P}} \left( \frac{d\bar{P}^c}{dt} \right) = k \left( \frac{1}{\bar{P}^c} \left( \frac{d\bar{P}^c}{dt} \right) \right) + (1-k) \left( \frac{1}{\bar{P}^u} \left( \frac{d\bar{P}^u}{dt} \right) \right)
\]

The index numbers can always be defined so that initially, \( \bar{P} = P^c = P^u \), and we get the expression in (3.1).
(and hence their equilibrium prices \( P^u \) and rates of change in equilibrium prices, \( p^u \)) will depend on the rationed quantities \( X^c \). Two implications pertinent to measuring the impact of wage-price controls, can be drawn from the Tobin-Houthakker analysis of rationing.

First, there exists a vector of shadow prices, \( P^s \) such that the rationed quantities, \( X^c \), would have constituted the preferred, utility-maximizing amounts of those goods. From the viewpoint of consumers, the shadow prices of controlled goods, \( P^s \), and the market prices of uncontrolled goods, \( P^u \), form the pertinent arguments of their indirect utility functions.

If we are trying to measure the impact of price inflation on the welfare of consumers, the real rate of price inflation should be measured by a weighted average of the rates of change in shadow prices, \( p^s \), and uncontrolled prices, \( p^u \).

\[
(3.2) \quad p^r_t = kp^s_t + (1-k)p^u_t.
\]

A correct measure of the quantitative impact of controls is thus given by the residual, \( e^r_t = (p^r_t - \hat{p}_t) \). At the outset of binding wage-price controls, \( p^r_t > p_t \) since with binding controls, shadow prices \( P^s \) will exceed the controlled prices \( P^c \). According to the simulation approach, one would conclude that controls had a significant impact in retarding price inflation if the residual, \( e_t = (p_t - \hat{p}_t) \) is large and negative.

However, from the viewpoint of consumer welfare, a correct measure of the impact is given by the difference between the real rate of price inflation, \( p^r_t \), and the predicted rate \( \hat{p}_t \). Hence, even under ideal conditions, (meaning that the explanatory variables are exogenous, and the model is
correctly specified and estimated), the simulation approach does not give us an unbiased estimate of the real impact because the controlled prices differ from the shadow prices.

Second, demands for the uncontrolled goods (with money incomes constant) could rise or fall depending on whether the controlled goods are substitutable or complementary with the uncontrolled goods. If substitution dominates, (as it is likely to do), the uncontrolled price $P^u$ will rise faster than they would have risen in the absence of controls. Indeed, one can construct exceptional cases in which the aggregate price index (including the below market equilibrium controlled prices), would climb when controls are imposed. It is presumably this sort of case that is implicit in the Friedman (1966) argument on how controls might have no effect on the time path of an aggregate price index.

3.3 Labor Productivity and Inflation

Turn next to the question of whether the explanatory variables in these models can be regarded as exogenous with respect to either the wage-price inflation process or the imposition of controls. For the purposes of simulation, the two key explanatory variables are the unemployment rate $U$ and labor productivity measured by output per manhour $Q^p$. If one accepts the basic postulate of a wage equation, [that there is a stable relationship between the wage change component $\omega_d$ and the excess demand for labor services], the prior history of wage change should have a feedback effect on the unemployment rate. In these

---

24/ In the Gordon model, the simulations also require data on $D$, $U^d$, $U^h$, and $(u^d/k)$ as well as assumptions about the trend rates of growth in labor productivity $q'$, potential labor force, and potential hours.
models, monetary and fiscal policies only indirectly affect the rates of wage and price inflation through their impact on the time path of the unemployment rate and the other explanatory variables. That the imposition of controls would have no effect on unemployment rates, seems to me to be an untenable assumption. If the actual time path of unemployment rates is itself influenced by wage-price controls, the forecasts generated by the simulations do not give us estimates of the predicted rates of wage and price changes that would have prevailed in the absence of controls.

Two opposing views have been advanced about how wage-price controls will affect the time path of labor productivity, \( q = \left( \frac{1}{Q} \right) \left( \frac{dO}{dt} \right) \). Sheahan (1967) and Solow (1966) argue that controls over wages and prices can lead to greater real output growth if the controlled sectors are monopolized or oligopolized. The rationale here rests on a micro-economic analogue wherein controls can achieve what anti-trust has failed to accomplish, namely restrain monopoly by imposing effective price ceilings which, in turn, induces greater output by the monopoly. Tobin (1972) rejects this thesis and argues that monopoly can only affect the level of prices and not the time rate of change in prices. The alternative thesis of competitive markets argues that binding wage-price controls must be accompanied by shortages and decreases in real output and employment. Recall that in the price equation, the rate of price inflation, \( p^X \), was inversely related to the rate of change in the ratio of actual to trend labor productivity \( (q/q') \). Thus, if labor productivity is retarded by controls, the smaller value for \( (q/q') \) during the controls period will, (when substituted in the fitted price equation) increase the predicted rate of price inflation
\( \hat{p}_t \). As a result, the residual, \( e_t = (p_t - \hat{p}_t) \), becomes more negative implying a larger quantitative impact of controls. To the extent that the imposition of controls is responsible for a slower rate of growth in labor productivity, the explanatory variables are not exogenous, and the simulation approach will overstate the quantitative impact of controls in retarding the rate of price inflation.

The preceding discussion fails to come to grips with what I regard as the principal weakness of the simulation approach, namely that the models they use, are simply mis-specified. Wage changes are linked to price inflation, and vice versa, without ever analyzing the fundamental forces driving the inflationary process. One could probably get very nice empirical relationships in which the mutton price is linked to the wool price and vice versa, but until we introduce the supply price of the sheep, the explanation of mutton and wool price movements is incomplete. In the reduced-form equations of these models, the rates of change in money wages and prices depend only on real variables which are assumed to be exogenous. However, somewhere in the background, the time paths of these real explanatory variables are presumably influenced by monetary and fiscal policies.

IV

Concluding Remarks

The problem of measuring the quantitative impact of wage-price controls and guidelines is a difficult one at both the conceptual and empirical levels. Judged by a market test of publishability, the four
models examined here, appear to represent what a substantial part of the profession regards as the best available solution to this problem. In this paper, I have presented my critical appraisal of the merits of these models as a basis for measuring the impact of wage and price controls. If, for the moment, we accept the theoretical structure of these two-equation models, what conclusions can we draw from this appraisal?

1. Rejection of the dummy variable method: A majority of the studies that used a dummy variable to measure the impact of the Kennedy-Johnson Guideposts, 1962-67, supported the conclusion that the Guideposts restrained the rates of wage and price inflation. Sheahan (1967) reported that these studies indicated a "downward deflection" in the rate of wage inflation of between 0.8 and 1.6 percent a year. The t-statistic for the coefficient of the guidepost dummy variable was indeed significant in both the P and EB wage equations. However, Gordon using different explanatory variables, found no significant effect for this dummy variable. The dummy variable method is a statistically valid procedure if and only if the sample data support the assumption of equal slope coefficients in periods with and without the guideposts. My replication of the Perry wage equation, (using a one quarter rate of wage change) showed that we must reject the assumption of equal slopes. The theoretical arguments advanced by Lipsay and Parkin (1970) provide us with an even stronger reason for rejecting this technique. This conclusion holds even when one accepts the theory behind a wage equation.

2. One vs. four quarter rates of change: A moving sum (or average) of independent random variables will exhibit positive serial correlation. This theorem explain the empirical regularity of high positive serial
correlation in the residuals of fitted wage equations using four quarter, overlapping rates of wage change. An implication of the analysis in section 2.1 is that the use of four quarter rates of wage change involves a specification error in the assumed probability distribution of the random disturbances. On statistical grounds, the Gordon procedure of using one quarter rates of wage change, is to be preferred even though it may yield lower values for the t-statistics and $R^2$.

3. Stability of coefficients: Any acceptable wage or price equation must satisfy the condition that it is stable; otherwise, it cannot be used as a reliable basis for prediction. We must, however, draw a distinction between economic and statistical stability. Gordon concluded from an examination of three sets of parameter estimates, (from overlapping samples), that the parameters of his wage equation were stable. The Chow test for statistical stability of the regression coefficients, [based on a partitioning of the entire sample], was employed by Lipsey-Parkin who found instability between samples for "policy on" and "policy off" periods. The one quarter version of the Perry wage equation, [Table 2 above] failed to pass the Chow test for stability. However, when I replicated the Gordon wage equation, the sample data confirmed the hypothesis of stable coefficients; namely the estimated parameter vectors for two independent samples were not significantly different. An example serves to illustrate the importance of my distinction between economic and statistical stability. A coefficient of 0.6 for the price inflation variable in a wage equation has vastly different policy implications than a coefficient of 1.3. However,
the two parameter estimates, (0.6 vs. 1.3), could easily lie within a
95 percent confidence ellipsoid for the true parameter vector. My
intuition suggests that the power of the Chow F test to discriminate
economically meaningful differences in vectors of structural parameters,
gets progressively weaker as the parameter space is expanded. The number
of structural parameters in the LP, G, and P wage equations were respectively
4, 12, and 5. The wide differences in the point estimates of parameters
shown in Table 1 suggest that the confidence ellipsoid for the 12 parameters
in the Gordon wage equation is likely to be very large.

4. Omitted Variables: The discussion in section 2.4 suggests that the
wage equations in these models may have been mis-specified because they
omitted at least two pertinent explanatory variables. The evidence on
unemployment insurance programs in section 2.4 was inconclusive. On the
bases of logic and arithmetic, the rate of change in an aggregate wage
index is obviously related to the proportion of workers getting wage
changes during any quarter. It is impossible for wages to change in quarter
t if no workers got wage changes during that quarter, \( \lambda_t = 0 \). The
rationale for including wages in the public sector in an equation explaining
wage changes in the private nonfarm sector, rests on two arguments. First,
the empirical evidence of Table 4 reveals that the time path of wages in
the public sector is very different from that in the private nonfarm sector;
the ratio of public to private wage rates has climbed over the period 1967-70.
Second, if public and private sectors compete for labor, (as they surely
do), wage movements in one sector must affect wages in the other.
5. The flexible nature of the underlying theory: The correspondence between the underlying theory of a wage or price equation and the empirical variables appearing in the econometric models, appears to be quite weak. Goodness of fit and plausible t-values based on shaky time series data, appear to be more important than adherence to an accepted theory of wage and price dynamics in determining just which variables should, or should not, be included in the final wage and price equations. An example of the flexible character of the theory is provided by the corporate profits variable. The profits variable was popular in the wage equations prior to 1968, but when profits failed to track the wage explosion of 1967-70, it was quickly discarded, and the recent wage equations do not even mention "ability to pay" and "liquidity" as determinants of short run wage changes.

6. Structural changes due to wage-price controls: A maintained implicit assumption of these models is that there are stable pseudo-equilibrium paths of wage and price adjustments to disequilibrating shocks described by the time paths of the unpredictable (but presumably exogenous) explanatory variables. There are compelling reasons to expect that workers and firms will not follow the same pseudo-equilibrium adjustment paths, (which prevailed in the absence of controls) in the period just preceding and following the imposition of binding wage-price controls. Their responses in anticipation of or in reaction to controls will manifest themselves in structural changes in the wage and price equations in the periods adjacent to the controls experience. In the light of these considerations, the
discrepancies between actual and predicted rates of wage/price inflation must be estimated via the simulation approach for a comparison period extending ahead of and beyond the period of controls in order to get an estimate of the full impact of controls in retarding the rates of wage and price inflation.

7. **Effect of controls on labor productivity:** The discretionary price and wage thesis argues that the imposition of selective controls on the monopolized sectors of the economy will lead to larger rates of growth in real output and employment when wage and price inflation rates are suppressed in these sectors. The opposing competitive markets thesis contends that if controls are truly binding, they must create shortages and retard the growth rates in real output and employment. If the latter thesis holds, the actual rate of change in labor productivity becomes an endogenous variable that is, in part, determined by the presence of wage-price controls. According to the price equations in these models, a slower growth rate in labor productivity implies a higher predicted rate of price inflation. Thus, treating labor productivity as a truly exogenous variable, (which is incorrect in the competitive markets thesis), imparts an upward bias in the predicted rates of price inflation during the period of controls. It is unclear whether similar or opposing biases are introduced by the other so-called exogenous variables such as the unemployment rate, tax rates, and the ratio of unfilled orders to capacity.

8. **Observed vs. implicit real rates of price inflation:** In the simulation approach, the "actual" rate of price inflation, \( p_t \), is equated to the observed price inflation rate as measured by official government price
Indexes which are weighted averages of movements in controlled and uncontrolled prices. Stigler and Kindahl (1970) and others have documented some of the biases in these published price indexes. In section 3.2, I introduced the concept of a real rate of price inflation, \( p^R_t \), which was defined as a weighted average of the rates of change in uncontrolled prices, \( p^u_t \) and shadow prices of the controlled goods, \( p^s_t \). The welfare of consumers can be measured by their indirect utility functions, \( U = U(P, M) \), where \( P \) is a vector of product prices and \( M \) is money income. When wage-price controls are truly binding, the shadow prices for the controlled (rationed) goods and market equilibrium prices for the uncontrolled goods make up the correct price vector in each consumer's indirect utility function which measures his utility under controls. The implicit real rate of inflation, \( p^R_t \), thus gives us the correct measure of price inflation under controls in terms of the welfare position of consumers.

All of these eight conclusions are predicated on the supposition that we accept as true, the underlying theoretical structure for these two-equation inflation models. In these four models, monetary and fiscal policies are relegated to positions as background variables that are ignored in the search for "stable" relationships. If there is a close and temporally stable relationship between the monetary and fiscal variable on the one hand, and the explanatory variables of these models on the other, it would be statistically impossible to identify which set of variables constitutes the true causal forces generating wage and price inflation.
I doubt if such a close relationship exists. Even if one found such a stable relationship, the logic of the underlying theory of inflation should, from a methodological viewpoint, dictate which set of variables properly belongs in these dynamic wage and price equations.25/

In closing, I would like to offer two, perhaps gratuitous, remarks. First, the defenders of wage-price controls and guidelines apparently accept the methodology set forth in the four studies examined in this paper, namely, measure the quantitative impact of controls in restraining the rates of wage and price inflation either by the coefficient of a dummy variable or by the accumulated sum of residuals in a simulation experiment. The opponents point to other impacts of binding wage-price controls such as the misallocation of resources due to shortages, the adjustments costs when controls are anticipated or ended, etc. I have attempted in this paper, to present my critical appraisal of the existing methodology for measuring the impact of wage-price controls. I regret that I cannot supply the reader with a demonstrably better model and estimation method to achieve the objective of measuring the full impact of controls. My second remarks is a sad commentary on the state of economic science, at least as it pertains to the economic theory of inflation and unemployment. The empirical evidence assembled by either the defenders or opponents of wage-price controls,

25/ An analogue from demand theory illustrates this point. The per capita consumption of gasoline is closely related to per capita vehicle ownership. One may choose to use this relationship if prediction of gasoline consumption is his objective. However, if we want to find the causal forces affecting gasoline consumption, our theory tells us that vehicle ownership must be treated as an endogenous variable.
is unlikely to have a significant effect in varying the numbers in the two camps. My personal evaluation of the merits of the two-equation inflation models in the LP, G, and EB studies is obvious from this paper. The views articulated by Houthakker (1972) on the place of wage-price controls as an instrument of macro-economic policy, come closest to my opinions on the subject. The forces that fundamentally drive the inflation process are our monetary and fiscal policies, and these models fail to come to grips with this issue.
APPENDIX

This appendix presents the least squares estimates for the preferred wage and price equations in the four published studies that were reviewed in this paper. I also try to summarize the way in which Perry and Gordon constructed their unemployment variables. Reference to the published studies reveals that the authors experimented with several alternative sets of explanatory variables. I have not attempted to review these experiments; the interested reader is referred to Gordon (1972) who has compared the parameter estimates for the P, G, and EB wage equations using different sets of unemployment variables, price inflation variables, etc. It should be repeated that the notation here is my own and differs from that appearing in the published studies.

A.1 The Perry Wage Equation

The dependent variable in the Perry wage equation is the four quarter, overlapping relative rate of change in the total hourly compensation of employees in the private nonfarm sector \( c_t \). In his preferred equation, [confer Perry (1970), eq. (3), Table 2, p. 425], \( c_t \) is related to six explanatory variables: \( p_{t-1} \) = the lagged one quarter rate of change in the Consumer Price Index, \( (1/U^Y_t) \) = the reciprocal of the Perry weighted unemployment rate, \( D_t \) = the Perry index for the dispersion of unemployment, \( G_t \) = the Kennedy-Johnson guidepost dummy variable, \( \Psi_t \) = a tax variable and \( f_t \) = the rate of change in the employment of secondary workers. The parameter estimates reported by Perry are shown in equation (A.1) where the numbers in parentheses denote the t-values, \( t_j = \hat{\alpha}_j / \sigma_{\alpha_j} \).
(A.1) \[ c_t = -0.0034 + 0.3453p_{t-1} + 0.0705(1/u^Y_t) + 0.0796D_t \]
\[ (-0.40) \quad (2.45) \quad (2.59) \quad (2.07) \]
\[ -0.0078G_t + 1.4025\psi_t + 0.0593f_t \]
\[ (-3.35) \quad (1.13) \quad (1.13) \]

[period, 1953:1-1968:4, \( R^2 = 0.777 \), \( \sigma_e = 0.0065 \), \( \rho = 0.669 \)]

When the dependent variable \( c_t \) is measured by a one quarter rate of wage change, there is a sharp drop in the explanatory power of the Perry wage equation; compare (A.1) above with column (P1) of Table 2.\(^{26}\) The definition for the Guidepost dummy variable \( G_t \) has already been described in footnote 11 above. The two interesting features of the Perry equation are his weighted unemployment rate \( U^Y \) and the dispersion index \( D_t \).

Perry argued that during the latter part of his sample period, movements in the official unemployment rate \( U^0 \) did not accurately reflect changes in the relative excess demand for equivalent labor services due to a rapidly changing age/sex composition of the labor force. According to Perry, a better measure of tightness in the labor market is provided by his weighted unemployment rate \( U^Y \). The pool of unemployed persons and the labor force is divided into age and sex groups where \( N_i \) is the number of unemployed persons, and \( L_i \) is the labor force in the \( i \)-th age/sex group. Let \( U_i = N_i/L_i \) is the age/sex specific unemployment rate. The official unemployment rate \( U^0 \) and the Perry weighted unemployment rate \( U^Y \) are simply two different variable-weighted averages of these age/sex

\(^{26}\) My replication also differed in two additional respects. First, the sample period was 1954:1-1970:4 instead of 1953:1-1968:4. Second, I omitted the last explanatory variable \( f_t \) which is not statistically significant in the Perry equation.
specific unemployment rates. The official rate which attaches the same weight to all unemployed persons can be written,

$$(A.2a) \quad U^0 = \frac{N}{L} = \frac{\sum N_i}{\sum L_i} = \frac{L_i}{L} U_i = \sum a_i U_i$$

where $L = \Sigma L_i$ = the aggregate number in the labor force. The weights $a_i$ represent the proportions of persons in the different age/sex groups. For any given official unemployment rate $U^0$, Perry argued that the downward pressure on an aggregate wage index would be greater, the larger is the fraction of unemployed persons who, on average, had earned higher hourly wage rates $W_i$ and worked longer hours per year $H_i$. The weight $y_i$ that Perry gives to unemployed persons $N_i$ and the labor force $L_i$ in the i-th group is given by,

$$y_i = \frac{W_i H_i}{W_0 H_0}$$

where the base group denoted by the zero subscript is taken to be the group of 35-44 year-old male workers. It will be noticed that $W_i H_i = y_i$, is simply the average annual wage earnings of persons in the i-th age/sex category. The Perry unemployment rate can thus be interpreted as an income-weighted average of the age/sex specific unemployment rates.

$$(A.2b) \quad U^Y = \frac{N^Y}{L^Y} = \frac{\sum y_i N_i}{\sum y_i L_i} = \frac{V_i L_i}{V_i L} U_i = \sum a_i U_i$$

A simple numerical example illustrates how these two rates behave when we vary the age/sex specific unemployment rates $U_i$, or the fractions $a_i$ falling into the different groups. Suppose that there are only two groups; $1 =$ prime age males, and $2 =$ secondary workers. Assume that the
annual wage earnings of secondary workers is only 0.4 times that of prime age males: \( y_2 = \frac{Y_2}{Y_1} = 0.4 \). Since group 1 is the base group, \( y_1 = 1 \).

It is further assumed that in periods of low, middle, and high unemployment, the group specific unemployment rates are: \( U_1 = 2.7, 3.0, 3.3 \) percent and \( U_2 = 8, 10, 12 \) percent. This pattern conforms to the cyclical one of widening relative differentials during a downswing. The ratio of prime age males to the total labor force, \( a_1 = \frac{L_1}{L} \), was allowed to vary from \( a_1 = 0.85 \) to \( a_1'' = 0.75 \). The aggregate official and weighted unemployment rates that correspond to these assumed conditions were calculated using equations (A.2a) and (A.2b).

<table>
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<tr>
<th>Hypothetical Official ( U^0 ) and Perry ( U^Y ) Unemployment Rates</th>
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<td>Rate (percent)</td>
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</tr>
<tr>
<td>Official Unemployment Rates ( U^0 )</td>
</tr>
<tr>
<td>Weight ( a_1 = 0.85 ) ( a_1' = 0.80 ) ( a_1'' = 0.75 ) ( a_1 = 0.85 ) ( a_1' = 0.80 ) ( a_1'' = 0.75 )</td>
</tr>
<tr>
<td>Low</td>
</tr>
<tr>
<td>Middle</td>
</tr>
<tr>
<td>High</td>
</tr>
<tr>
<td>Perry Unemployment Rate ( U^Y )</td>
</tr>
<tr>
<td>Weight ( s_1 = 0.9341 ) ( s_1' = 0.9091 ) ( s_1'' = 0.8824 ) ( s_1 = 0.9341 ) ( s_1' = 0.9091 ) ( s_1'' = 0.8824 )</td>
</tr>
<tr>
<td>Low</td>
</tr>
<tr>
<td>Middle</td>
</tr>
<tr>
<td>High</td>
</tr>
</tbody>
</table>
This exercise suggests that the dispersion of the Perry rate is smaller than that of the official rate \( U^0 \). When the fraction of secondary workers, \( a_2 = 1-a_1 \), climbs and the age/sex specific unemployment rates fall, [compare the high (i) and low (iii) figures], the official rate shows a relatively smaller reduction. This is apparently what happened in the 1967-70 period.

The Perry dispersion index, \( D \), is a weighted average of the relative deviations (in absolute values) of the age/sex unemployment rates \( U_i \) from the Perry weighted unemployment rate \( U^Y \) where the weights \( w_i \) are identical to those used for the unemployment rate as defined in (a.2b) above. More precisely, we have,

\[
D = \sum w_i \frac{U_i - U^Y}{U^Y}
\]  

(A.3)

In Table 5 below, I present the annual averages for (a) the one quarter rate of change in hourly total compensation \( c_t \) expressed at annual rates, (b) the official unemployment rate \( U^0 \), (c) the Perry weighted unemployment rate \( U^Y \) and (d) the Perry dispersion index \( D \) for the 17 years included in the sample period, 1954-70. The percentage decline in the official unemployment rate \( U^0 \) from 1967 to 1968 was roughly 5 percent, while that for the Perry rate \( U^Y \) was nearly 12 percent. Thus, in the period, 1967-69, the Perry rate implied a considerably tighter labor market relative to the official unemployment rate. In the table below, I have calculated the means and standard deviations of the quarterly rates of wage changes and the quarterly values (divided by 4) for the inverse of the Perry weighted unemployment rate, \((1/U^Y)\), and the Perry dispersion
index for three sub-samples and for the entire period, 1954:1-1966:4

<table>
<thead>
<tr>
<th>Period</th>
<th>1/U^Y</th>
<th>D</th>
<th>c</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1.186)</td>
<td>(0.900)</td>
<td>(0.290)</td>
</tr>
<tr>
<td></td>
<td>(1.453)</td>
<td>(1.058)</td>
<td>(0.301)</td>
</tr>
<tr>
<td></td>
<td>(1.511)</td>
<td>(0.749)</td>
<td>(0.232)</td>
</tr>
<tr>
<td></td>
<td>(1.419)</td>
<td>(1.677)</td>
<td>(0.293)</td>
</tr>
<tr>
<td></td>
<td>(1.802)</td>
<td>(2.301)</td>
<td>(0.388)</td>
</tr>
</tbody>
</table>

*Standard deviations are shown in parenthesis.

The last four years of the sample period, 1967-70, experienced the highest rate of wage inflation, and it is precisely over these four years that \((1/U^Y)\) and \(D\) show their greatest increases. If one defines a dummy variable that is equal to one in 1967:1-1970:4 and zero otherwise, the simple correlation of the wage change variable, \(c_t\), with this dummy was \(+.697\).

The same correlation (with this dummy variable) for the inverse of the Perry weighted unemployment rate was \(+.609\), and with the dispersion index, \(+.758\). However, over the period 1966-68, the correlation of \(c_t\) with \((1/U^Y)\) is considerably higher than that between \(c_t\) and \(D_t\). It is thus not surprising
Table 5

Wage Changes and Unemployment Variables for the Perry Wage Equation$^a$/

<table>
<thead>
<tr>
<th>Year</th>
<th>$c_t$</th>
<th>$U_0^t$</th>
<th>$U^y_t$</th>
<th>$D_t$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1954</td>
<td>3.03</td>
<td>5.5</td>
<td>4.56</td>
<td>25.34</td>
</tr>
<tr>
<td>1955</td>
<td>4.39</td>
<td>4.4</td>
<td>3.54</td>
<td>29.61</td>
</tr>
<tr>
<td>1956</td>
<td>5.38</td>
<td>4.1</td>
<td>3.30</td>
<td>32.22</td>
</tr>
<tr>
<td>1957</td>
<td>5.25</td>
<td>4.3</td>
<td>3.42</td>
<td>30.57</td>
</tr>
<tr>
<td>1958</td>
<td>3.73</td>
<td>6.8</td>
<td>5.76</td>
<td>23.11</td>
</tr>
<tr>
<td>1959</td>
<td>3.95</td>
<td>5.5</td>
<td>4.64</td>
<td>27.11</td>
</tr>
<tr>
<td>1960</td>
<td>3.80</td>
<td>5.5</td>
<td>4.52</td>
<td>27.44</td>
</tr>
<tr>
<td>1961</td>
<td>5.51</td>
<td>6.7</td>
<td>5.52</td>
<td>26.46</td>
</tr>
<tr>
<td>1962</td>
<td>3.40</td>
<td>5.5</td>
<td>4.45</td>
<td>29.97</td>
</tr>
<tr>
<td>1963</td>
<td>3.24</td>
<td>5.7</td>
<td>4.40</td>
<td>34.41</td>
</tr>
<tr>
<td>1964</td>
<td>2.79</td>
<td>5.2</td>
<td>3.84</td>
<td>39.24</td>
</tr>
<tr>
<td>1965</td>
<td>3.88</td>
<td>4.5</td>
<td>3.31</td>
<td>41.55</td>
</tr>
<tr>
<td>1966</td>
<td>4.94</td>
<td>3.8</td>
<td>2.67</td>
<td>44.28</td>
</tr>
<tr>
<td>1967</td>
<td>5.20</td>
<td>3.8</td>
<td>2.85</td>
<td>45.00</td>
</tr>
<tr>
<td>1968</td>
<td>6.89</td>
<td>3.6</td>
<td>2.50</td>
<td>50.13</td>
</tr>
<tr>
<td>1969</td>
<td>6.77</td>
<td>3.5</td>
<td>2.50</td>
<td>50.94</td>
</tr>
<tr>
<td>1970</td>
<td>7.05</td>
<td>4.9</td>
<td>3.74</td>
<td>45.87</td>
</tr>
</tbody>
</table>

| Mean | 4.54 | 4.90 | 3.85 | 35.49 |
| Std. Dev. | 1.37 | 1.02 | 0.99 | 9.20 |

$^a$/Key to symbols:

$c_t$ = average of the four quarterly rates of change in hourly total compensation for the private nonfarm sector expressed at annual rates.

$U_0^t$ = the official unemployment rate of the civilian labor force, 16 and older.

$U^y_t$ = the Perry weighted unemployment rate obtained by first taking the average for the four quarterly observations of $(1/U^y)$ and then taking its reciprocal as a percentage.

$D_t$ = the Perry index of the dispersion of unemployment.
that Gordon who used the longer sample period, 1954-1970, found that the
dispersion index provided a better fit to the wage change variable.

A.2 The Gordon Model

In his preferred wage equation, [Gordon (1972), Column 6, Table 3,
pp. 398-9], the relative rate of change in his fixed weight wage index,
$w_t$, (inflated for fringe benefits), is related to seven explanatory variables.
However, the coefficient on employer's contributions for social security
taxes, $\psi^s$, was constrained to equal minus one, and I have moved it to
the left side of the equation. The results shown below in equation (A.4)
correspond to Gordon's one quarter version for the dependent variable; the
two quarter version is not appreciably different.

(A.4) \[ w_t + \psi^s_t = \alpha_0 + 0.0410D_t - 0.602U^d_t - 0.422U^h_t + 0.660p^*_t - 1 \]
     \[ (1.91) \]
     \[ (-3.42) \]
     \[ (-1.65) \]
     \[ (3.71) \]
     \[ + 0.684(p^x_p)^*_t - 1 + 0.115\psi^a_t \]
     \[ (2.51) \]
     \[ (2.04) \]

[period, 1954:1-1970:4. $\sigma_e = .00182$. DW = 2.49]

where the numbers in parentheses are the t-values. The two novel features
of the Gordon wage equation are the absence of any aggregate unemployment
rate, [it is replaced by three other unemployment variables; the Perry
dispersion index $D_t$, the disguised unemployment rate of discouraged workers
$U^d_t$, and the unemployment rate of hours $U^h_t$], and the inclusion of a
short distributed lag of the gaps between the rates of change in product
and consumer prices, $(p^x_p)^*$. [An asterisk indicates that the variable
is included as a distributed lag.] The former feature has already been
discussed in section 2.3B above.
The justification for the variable, \((p^X - p)\), is described in the following excerpt:

"In a period of excess commodity demand, when firms raise product prices relative to labor costs, demand for labor increases, and they are willing to pay a higher wage. The same result occurs even when firms do not take the initiative in raising wages, when union leaders respond to a firm's increased profits by demanding higher wage increases than they would have otherwise. In the wage equations presented below, this 'marginal revenue product demand effect' is tested by including as an additional explanatory variable, the difference between the rate of increase in product prices measured by the NPD and consumer prices measured by the PCD." [Gordon (1971) p. 112]

This variable presumably measures the vertical gap between labor's marginal revenue product, MRP, and hourly wages \(W\). The unemployment variables in the equation presumably measure the horizontal gap, \(\frac{D-S}{S}\). If the elasticities of the labor demand and supply curves are stable, information about the size of one gap implies the size of the other and the two variables are, in theory, redundant. More importantly, Gordon's rationale for this variable suggests that firms and union leaders raise money wage rates in response to either anticipated or realized profits. In the empirical Phillips curve literature of the early 1960's, profits frequently appeared as an important explanatory variable. The failure of profits to track the wage explosion of 1967-70, (and possibly its unstable coefficient) has made it an unacceptable variable in the recent wage equations.

The disguised unemployment rate \(U^d\) which is defined as the labor reserve \(R\) divided by the civilian labor force including the Armed Services \(L\), \([U^d = \frac{R}{L}]\), can be expressed as a nonlinear transformation of actual labor force participation rates. The labor reserve \(R\) is the difference between the potential \(L'\) and actual \(L\) labor forces. Let \(X_1 = \)
the population of prime age (25-44 year-old) male workers, and \( X_2 \) = the population of all other (secondary) persons. According to Gordon (1971), p. 152, the potential labor force of prime age males follows a negative trend, while that of secondary workers exhibits a positive trend plus an adjustment for the ratio of employment to population \( (E_2/X_2) \).

\[(A.5a) \quad L'_1 = X_1[a_1 - b_1 t] \]

\[(A.5b) \quad L'_2 = X_2[a_2 + b_2 t + c_2(E_2/X_2)] \]

He further assumes, perhaps implausibly, that \( (E_2/X_2) = .95 \) at full employment. Let \( \lambda_1 = L_1/X_1 \) and \( \lambda_2 = L_2/X_2 \) denote the actual labor force participation rates so that \( X_j = L_j/\lambda_j \). If these relationships are substituted into the definition of the disguised unemployment rate, [where \( L' = L'_1 + L'_2 \)], we get,

\[(A.5c) \quad U^d = \frac{L' - L}{L} = \pi\left[\frac{a_1 - b_1 t}{\lambda_1}\right] + (1-\pi)\left[\frac{a_2 + b_2 t + c_2(E_2/X_2)}{\lambda_2}\right] - 1 . \]

where \( \pi = L_1/L \) is the ratio of the labor force of prime age males to the total actual labor force. Movements in \( U^d \) are thus driven by the cyclical fluctuations in labor force participation rates, [especially of secondary workers \( \lambda_2 \)] and the secular decline in \( \pi \).

The unemployment rate of hours \( U^h \) is one minus the ratio of actual \( H_t \) to potential \( H^*_t \) average weekly hours in the \( t \)-th quarter.

\[ U^h_t = 1 - (H_t/H^*_t) \]

The potential hours \( H^*_t \) are calculated from a broken trend line connecting
the actual average weekly hours in the peak quarters in 1950:4, 1955:4, 1956:1 and 1968:3. The choice of an appropriate trend line appears to be somewhat arbitrary, especially in the last four or five years of the sample period.

In his preferred price equation, [Gordon (1972), column 3, Table 4, p. 407], the one quarter rate of change in product prices, \( p^x_t \), is related to four explanatory variables; \((w/q')_t^* = \) a distributed lag on prior rates of change in standard unit labor costs, \((q/q')_t^* = \) a distributed lag on prior rates of change in the ratio of actual to trend labor productivity, \((c/w)_t = \) the change in the ratio of hourly total compensation to the Gordon wage index, and \((uf/k)_t = \) the change in the ratio of unfilled orders to capacity. The least squares results are shown below where the coefficients of the lag variables are the sums of coefficients in the distributed lag structures.

\[
(A.6) \quad p^x_t = \beta_0 + 0.964(w/q')_t^* - 0.238(q/q')_t^* + 0.537(c/w)_t \\
(6.42) \quad (-1.92) \quad (2.88) \\
+ 0.023(uf/k)_t \\
(2.96)
\]

[period, 1954:2-1970:4 \( \sigma = .00212, \) DW = 2.47]

The trend rate of growth in labor productivity \( q' \) was obtained from a broken trend line connecting the actual rates of change in labor productivity in the peak quarters of 1950:4, 1955:2, and 1966:1. For the purposes of simulation, \( q' \) is a constant. Gordon argues that the price equation should
be consistent with an assumption of constant income shares which, in turn, implies a coefficient of unity for standard labor costs \((w/q')\). The results are consistent with this assumption.

The inclusion of the difference in rates of change in hourly total compensation and the Gordon wage index, \((c/w)\), was justified on the grounds that the latter did not include the labor costs of supervisory personnel and overtime work in manufacturing. As I have argued in section 2.3A above, the logic of a price equation suggests that standard unit labor costs should be measured by \((c/q')\) which implies that in the price equation, \(\beta_1\) and \(\beta_3\) should be equal. The results shown in (A.6) above are contrary to this constraint, but in fairness, the \((c/w)\) variable was not included as a distributed lag variable.

The ratio of unfilled orders to capacity \((uf/k)\) was included to measure the impact of disequilibrium in product markets on price movements. This variable appears to work better than other proxies such as the rate of change in finished goods inventories. Unfilled orders are, however, mainly found in the heavy durable goods industries, and a majority of the product markets in the private nonfarm sector rely on finished goods inventories to accommodate short run fluctuations in product demands. Is the observed statistical association between \(p^X\) and \((uf/k)\) spurious? A test can be constructed to shed light on this question. The aggregate price index \(P^X\) can be disaggregated into an index for those industries which report data on unfilled orders, \(P^{xa}\), and those that do not \(P^{xb}\). Separate price equations in which the same \((uf/k)\) variable is included in each, can be estimated for the rates of price change \(p^{xa}\) and \(p^{xb}\). The
statistical significance and stability of the coefficient for \( (uf/k) \) should be greater for the former \( p^x \) equation. If it were not, I would conclude that \( (uf/k) \) is simply a cyclically volatile variable which just happened to be correlated with \( p^x \), and we should question this aspect of the theory of a price equation.

A.3 The Ekstein-Brinner Model

In the EB wage equation, the four quarter percentage rate of change in the Gordon wage index, \( w_t \), is related to two price inflation variables, the reciprocal of the official unemployment rate \( 1/U^0 \), the Guidepost dummy variable \( G \), and a tax rate variable \( \psi \). The parameter estimates for their preferred wage equation, [Ekstein and Brinner (1972), p. 4] are shown in equation (A.7).

\[
(A.7) \quad w_t = 1.257 + 0.496p_t + 0.248[\Sigma p_t - 5]^a + 11.171(1/U^0)_t \\
- 0.705G_t + 0.022\psi_t \\
(5.90) \quad (7.31) \quad (4.02) \quad (10.07) \\
(6.91) \quad (3.22)
\]

[period, 1955:1-1970:4, \( R^2 = .935 \), \( \sigma_e = .329 \), DW = 0.769]

The \( t \)-values are presented in parentheses. The guidepost dummy \( G_t \) is the same as that used by Perry, and the tax variable is an arbitrary moving average over four quarters of the rate of change in the ratio of gross wages to wages net of taxes. I shall limit my remarks to the two price inflation variables.

The price inflation variable is a weighted four quarter moving average of the four quarter rate of change in the personal consumption deflator.
\[ p_t = 0.4p_t + 0.3p_{t-1} + 0.2p_{t-2} + 0.1p_{t-3} \]

By imposing arbitrary fixed weights, BE assume away the issue of how workers formulate expectations about price inflation. The procedure does, however, eliminate the need to estimate more parameters in a distributed lag model. The novel variable in the BE equation is \( \Sigma p_{t-5}^a \) which is defined as the positive excess of the accumulated inflation over two years over and above 5 percent (or 2.5 percent per annum). More precisely,

\[ \Sigma p_{t-5}^a = [(p_t + p_{t-1} + \ldots + p_{t-7}) - 5], \text{ if positive, zero otherwise.} \]

The argument here is that when price inflation is below 2.5 percent a year, only part of the higher cost of living (the parameter of \( \Sigma p_t \) implies about half) is passed on through higher wage rates; hence workers during a mild inflation suffer a loss in real wages. However, when inflation exceeds 2.5 percent a year, workers become aware of the loss in purchasing power and bargain for higher wages. As a consequence, with rapid price inflation, each percentage point rise in \( P \) is reflected in an approximately 1 percent rise in hour wages. Gordon (1972) agrees that in principle, the partial relation between the price inflation component of wage change, \( w^P \) and expected price inflation \( p^* \) could be nonlinear. However, his data did not yield a clearcut verdict on this hypothesis. Except for possibly one or two quarters, virtually all of the non-zero observations for \( \Sigma p_{t-5}^a \) occur in the last four years of the sample period. My hunch is that if they had included a dummy variable equal to one for 1967:1-1970:4 and zero otherwise alongside this threshold price variable, its significance would have sharply declined.
The Ekstein and Brinner price equation is qualitatively similar to that of Gordon but differs in three important respects. First, instead of allowing the data to determine the appropriate distributed lags, they again impose an arbitrary fixed weight moving average on both wages and labor productivity. Second, they introduce two variables describing the difference between current and moving average values for wages and labor productivity to capture the very short run effects. Third, the unfilled orders to capacity variable is included as a positive or zero variable similar to their threshold price variable; i.e. it implies a ratchet effect. The preferred price equation [Ekstein-Brinner (1972), p. 4] is given by,

\[
\begin{align*}
\rho^X_t &= 0.007 + 1.000(\overline{w}_{t-1} \cdot 0.65) + 0.351(w_t - \overline{w}_{t-1}) - 0.199(q_{t-1} \cdot 0.65) \\
& \quad - 0.089(q_t - \overline{q}_{t-1}) - 0.517(\overline{w}_t - \overline{c}_t) + 0.120(uf/k)^a_t \\
& \quad (-2.29) \quad (-3.21) \quad (2.30)
\end{align*}
\]

[period, 1955:1-1971:4, $R^2 = .707$, $\sigma_e = .223$, DW = 1.954]

where $\overline{w}_t = .4w_t + .3w_{t-1} + .2w_{t-2} + .1w_{t-3}$, and $\overline{q}_t$ and $\overline{c}_t$ employ the same (.4, .3, .2, .1) weights for labor productivity and hourly total compensation. The subtraction of -.65 in the first and third explanatory variables is based on the assumed long run trend rate of growth in labor productivity of +0.65 percent per quarter.

The coefficient of unity for $(w_{t-1} - .65)$ again implies that price changes in "the long run" fully reflect all changes in standard unit labor costs. The equation, however, also identifies a separate "short run effect" wherein prices in the very short run have an elasticity of +.35.
with respect to deviations between the current rate of wage change and the expected wage change, \((w_t - \overline{w}_{t-1})\). The distributed lag structure that is implicit in the EB price equation can be explicitly expressed by a polynomial in the lag operator \(\lambda\) as shown in (1.6b) above.\(^{27}\)

\[
B_1(\lambda)w_t = [.1442 + .1045\lambda + .0934\lambda^2 + .0781\lambda^3 + .0649\lambda^4]w_t.
\]

In a similar fashion, the EB price equation identifies separate "long run" and "short run" effects of changes in labor productivity on prices. The polynomial distributed lag, \(B_2(\lambda)q_t\) is also described by a fourth-degree polynomial in \(\lambda\). Given these long distributed lags in the structural equations, the reduced form equations [which must be estimated for two-stage least squares] are likely to be very complicated. Further, we must solve for these reduced-form equations in order to derive the partial effects of a change in any one exogenous variable, (like the current rate of change in labor productivity) on the two jointly dependent variables.

Finally, the EB price equation involves ratchet effect in the variable, \((uf/k)^a_t\) which is defined as follows:

\[
(uf/k)^a_t = [(\frac{UF}{K})_t - (\frac{UF}{K})_{t-1}], \text{ if positive and zero otherwise.}
\]

---

\(^{27}\) In (1.6b), we wrote the structural equation as,

\[
p^x_t = B_0 + B_1(\lambda)w_t + B_2(\lambda)q_t + B_3(uf/k)_t.
\]

From (A.8) above, if we collect terms involving wage rate changes, we have,

\[
p^x_t = 1,000\overline{w}_t + .351(w_t - \overline{w}_{t-1}) - .517(\overline{w}_t - \overline{c}_t),
\]

where \(\overline{w}_t = .4w_t + .3w_{t-1} + .2w_{t-2} + .1w_{t-3}\). By temporarily neglecting \(\overline{c}_t\) (which includes \(w_t\)) I get the polynomial shown in the text.
When the ratio of unfilled orders to capacity rises, it is accompanied by an increase in product prices, but when \((UF/K)\) falls, it does not put downward pressure on prices. The nonlinearity is justified on the ground that it provided a better fit to the sample data.

A.4 The Lipsey-Parkin Model

The LP model is a comparatively simple one involving no distributed lags in the rates of change in wages and prices. The dependent variable of the wage equation is a centered four-quarter rate of change in hourly wage rates \(w_t\), which is related to three explanatory variables: the unemployment rate \(U\), the lagged rate of change in consumer prices \(p_{t-1}\), and the rate of change in union membership \(n_t\).

\[
(A.9a) \quad w_t = \alpha_0 + \alpha_1 U_t + \alpha_2 p_{t-1} + \alpha_3 n_t + \epsilon_t.
\]

The rationale for \(n_t\) is that when union membership is increasing, it reflects greater aggressiveness on the part of union leaders which, in turn, adds upward pressure to wage changes.

The LP price equation is also a simple one in which the one quarter rate of change in consumer prices \(p_t\) is related to the rates of change in wage rates, \(w_t\), lagged import prices \(m_{t-1}\), and labor productivity \(q_t\).

\[
(A.9b) \quad p_t = \beta_0 + \beta_1 w_t + \beta_2 m_{t-1} + \beta_3 q_t + \epsilon_t.
\]

Lipsey and Parkin convincingly argue that the structural wage and price equations are likely to be very different in periods when market forces determine the rates of wage and price change than in periods when these market forces are constrained by incomes policy. They submit this theoretical hypothesis to a statistical test by estimating the two
structural equations, [(A.9a) and (A.9b)] for two independent samples. The first sample included only observations for periods (quarters) when there were no incomes policy, while the second was limited to observations for the "policy on" quarters. An application of the Chow test revealed that the hypothesis of structural stability (between "policy off" and "policy on" periods) could be rejected. In Table 6 below, I show the least squares parameter estimates for these three samples.
Table 6

Least Squares Parameter Estimates for the Lipsey-Parkin Model*
(t-values in parentheses)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Entire Period</th>
<th>Policy Off</th>
<th>Policy On</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>A. The Wage Equation</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>constant</td>
<td>4.147</td>
<td>6.672</td>
<td>3.919</td>
</tr>
<tr>
<td>(4.26)</td>
<td>(5.79)</td>
<td>(2.27)</td>
<td></td>
</tr>
<tr>
<td>Unemployment, $U_t$</td>
<td>-0.891</td>
<td>-2.372</td>
<td>-0.404</td>
</tr>
<tr>
<td>(-1.77)</td>
<td>(-3.64)</td>
<td>(-0.56)</td>
<td></td>
</tr>
<tr>
<td>Price inflation, $P_t$</td>
<td>0.482</td>
<td>0.457</td>
<td>0.227</td>
</tr>
<tr>
<td>(5.76)</td>
<td>(6.25)</td>
<td>(0.93)</td>
<td></td>
</tr>
<tr>
<td>Union membership, $u_t$</td>
<td>3.315</td>
<td>0.136</td>
<td>3.764</td>
</tr>
<tr>
<td>(2.09)</td>
<td>(0.07)</td>
<td>(1.61)</td>
<td></td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.616</td>
<td>0.856</td>
<td>0.138</td>
</tr>
<tr>
<td>DW</td>
<td>0.742</td>
<td>1.231</td>
<td>0.724</td>
</tr>
</tbody>
</table>

| **B. The Price Equation**               |               |            |           |
| Constant                                | 1.374         | -0.140     | 3.874     |
| (2.51)                                  | (-0.16)       | (5.65)     |           |
| Wage change, $w_t$                      | 0.562         | 0.851      | 0.014     |
| (5.53)                                  | (5.52)        | (0.10)     |           |
| Import price, $m_{t-1}$                 | 0.085         | 0.073      | 0.001     |
| (4.60)                                  | (2.93)        | (0.04)     |           |
| Labor productivity, $q_t$               | -0.145        | -0.092     | -0.198    |
| (-3.48)                                 | (-1.90)       | (-2.68)    |           |
| $R^2$                                   | 0.697         | 0.843      | 0.241     |
| DW                                      | 0.946         | 1.274      | 1.088     |

REFERENCES


