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THE EFFECTS OF INCOMES POLICIES ON THE FREQUENCY AND SIZE OF WAGE CHANGES

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ABSTRACT

Along with house rents, wages have frequently been described as the "stickiest" prices in the economy, rarely adjusted more than once a year. Because of this stickiness (which arises from the transactions costs involved in changing wages), a distinction exists between the adjustment of wages and the size of that adjustment. This distinction has important implications for empirical investigations of the determinants of aggregate money wage changes because the equations fitted in these studies are almost invariably plagued with aggregation bias unless the non-synchronous pattern of wage settlements in different sectors of the economy is taken into account. This is a particularly relevant issue when evaluating the effectiveness of incomes policies since some policies have operated by postponing the implementation of new wage settlements (in which case they are directed towards the occurrence of the event) while other policies have taken the form of specifying a permissible ceiling on wage increases (in which case they are designed to affect the extent of occurrence of the event, but not its occurrence). One purpose of this paper is to reevaluate the effectiveness of incomes policies by making use of information from one industry both on the frequency of wage settlements and on the size of wage changes when a settlement takes place. Our empirical work leads us to conjecture whether the apparent "statistical significance" reported by researchers with respect to the performance of variables in models of <u>aggregate</u> wage changes reflects primarily the effects of these variables on the probability of wages being adjusted rather than the effects on the magnitude of wage changes conditional upon wages being adjusted.

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AND SIZE OF WAGE CHANGES

by

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I. Introduction

Theoretical models accounting for variations in the rate of wage or price inflation rarely take explicit account of the costs of altering money wages or prices[1] and yet these models are often evaluated in terms of their correspondence with data that presumably reflect the existence of such costs. At least until the more inflationary experiences of the 1970's, the normal situation in British and American manufacturing industry was for wage rates to be adjusted no more frequently than once a year even when there were indications of substantial inflationary pressures. The recognition of this point has induced economists to incorporate in some fashion the varying time pattern of wage settlements into their fitted models of wage inflation, but all have taken this pattern as exogenously determined.[2] By contrast, this paper considers the determinants of the probability of a new wage settlement in addition to the determinants of the size of the wage change conditional upon a new settlement taking place. This distinction between the occurrence of an event and the extent of the occurrence has been investigated in the

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literature on the demand for durable goods (e.g., Cragg [1971]), on labor force participation and hours of work (e.g., Heckman [1974]) and on the choice of transport modes (e.g., Domencich and McFadden [1979]), but with few exceptions it has been neglected in empirical work on wage inflation.[3]

Yet the distinction is an important one with immediate relevance to the evaluation of the effects of incomes policies on the course of wage inflation. The typical procedure in this literature is to regress some aggregate index of wage changes on dummy variables that indicate the presence or absence of incomes policies. But this procedure does not make full use of information about the <u>form</u> of these incomes policies. That is, incomes policies have sometimes taken the form of specifying a permissible ceiling on wage increases in which case they are designed to affect the extent of occurrence of the event, but not the occurrence of the event. On other occasions, incomes policies have operated by postponing the implementation of new wage settlements in which case they are directed towards the occurrence of the event. One purpose of this paper is to reevaluate the effectiveness of incomes policies by making use of information both on the frequency of wage settlements and on the size of wage changes when a settlement takes place.

Where the time pattern of wage settlements differs across sectors of the economy and where no account is taken of this non-synchronous pattern, then <u>aggregate</u> quarterly wage change equations will be plagued with aggregation bias. That is, under these circumstances, the coefficients of these aggregate wage change equations should not be expected to display any stability with respect to the addition or deletion of observations. This is by no means a novel point[4] yet its implications appear not to have been fully realized because the unproductive search for plausible and stable aggregate wage change equations continues unabated. A more appropriate strategy is to investigate the determinants of wage settlements and wage changes in individual sectors of the economy and this paper illustrates this strategy for the British coal industry. The task of extending the analysis to other industries will be addressed on another occasion.

II. Conceptual Framework

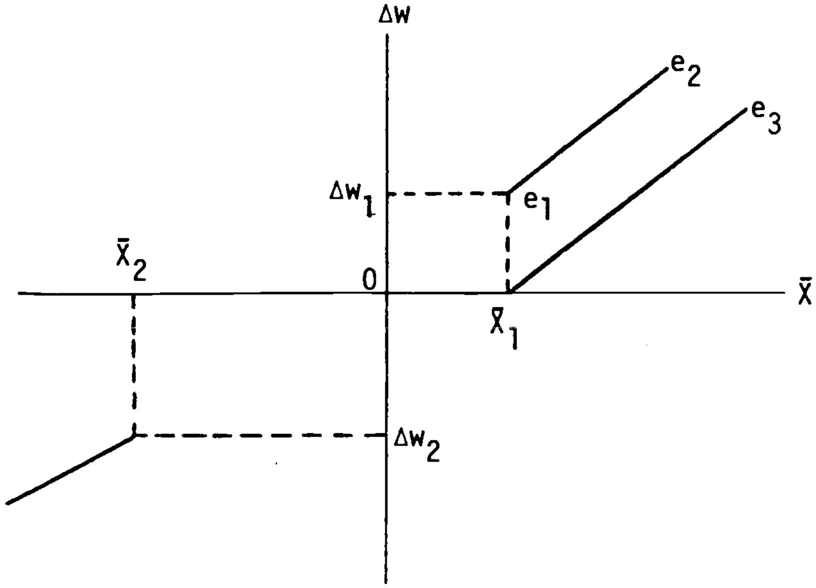
Collective bargaining agreements in the British industrial relations system are not legally binding contracts so that, unlike the unionized sector in the United States, the interval between wage settlements is not predetermined by the most recent agreement. On the contrary, negotiations toward a new wage settlement may be initiated in any quarter. If there were no costs to negotiating and implementing new wage agreements, then a new wage settlement would occur whenever the determinants of wage changes called for a new wage level. In fact, it has been argued that the transactions costs of negotiating and executing a new wage contract are by no means negligible nor should such costs be assumed to be unchanged over time. These costs include the prior determination by each party of the bargaining posture to be adopted, the resources consumed in conducting negotiations between the two parties, the expenses borne by each party in the event of a strike, and the administrative costs of changing payroll records. Rees' [1970] statement of the nature of these transactions costs in the labor market is particularly compelling:

> "Except in hyper-inflations, employers generally seem to behave as though the costs of changing wages were very substantial. When faced with labor shortages, they frequently lower hiring standards, raise expenditures on recruitment, and contract-out work before raising wages.... [E]very change in wages tends to raise questions of internal equity that can be very troublesome to solve. Even if it is decided to raise all wages uniformly, it must be decided whether this is to be by a uniform percentage or a uniform amount, and whether the increase is to extend all the way up the structure of wages and salaries or only part way. The problems are not unlike those faced by the United States Congress in separating problems of the level of income tax from issues of tax reform. Although we know very little about the exact nature of the costs of making wage changes, we can infer that they exist. Wages are, next to house rents, the stickiest general class of prices in the economy, seldom adjusted more frequently than once a year. This stickiness may be reinforced by unionism and collective bargaining, but it was present long before unions arrived" (p. 234).

In fact, during the period from the first quarter of 1948 (1948 I) to the last quarter of 1975 (1975 IV), the coal miners' nationally negotiated wage rates were changed in only 24 of the 112 quarters. Wage rates were not changed each year: five years witnessed no change in wage rates (in 1948, 1949, 1952, 1959, and 1972) and there was an occasion on which wages were adjusted in successive quarters and another occasion on which eleven quarters separated wage changes. Even during the inflationary 1970's (until 1975 IV), wage rates in coal mining were renegotiated on only five occasions. A stylized representation of the situation is given in Figure 1 where the vertical axis measures the proportionate change in wage rates (Δw) and where $\bar{X}$ on the

Figure 1

Transactions Costs and Wage Changes

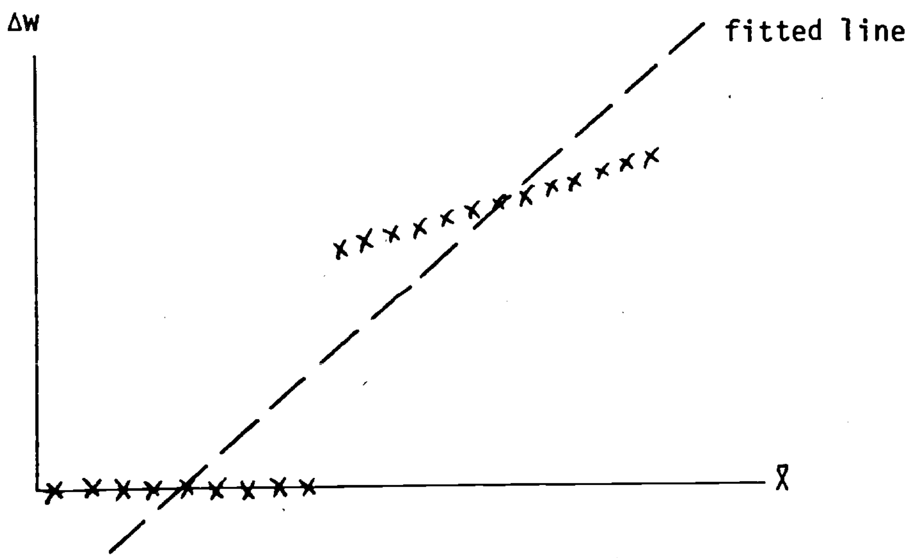


horizontal axis in some relevant exogenous variable that exerts a positive effect on Δw. The argument here is that, for values of $\bar{X}$ between $\bar{X}_1$ and $\bar{X}_2$, wage rates are completely unresponsive so that the relationship between $\bar{X}$ and nonzero values of Δw is only observed for values of $\bar{X}$ greater than $\bar{X}_1$ and less than $\bar{X}_2$. Of course, neither $\bar{X}_1$ nor $\bar{X}_2$ need be unchanged over all observations. The relationship between Δw and $\bar{X}$ for values of $\bar{X}$ greater than $\bar{X}_1$ may look like e_1e_2 in Figure 1 according to which at $\bar{X}_1$ the change in wage rates is a value (Δw_1) that just offsets the unit costs of making adjustments or it may look like $\bar{X}_1e_3$ according to which infinitessimal values of Δw are feasible.[5] The purported downward stickiness of wages implies that $\bar{X}$ has to assume much larger absolute values for any downward movement in wage rates to be observed (i.e., $0\bar{X}_2 > 0\bar{X}_1$) and/or that, for nonzero observed values of Δw, the partial derivative of Δw with respect to $\bar{X}$ is less in the south-west quadrant of the graph than in the north-east quadrant. The distinction between these two hypotheses in the literature on "downward wage rigidity" is rarely made, let alone tested.

The presence of these transactions costs calls for statistical methods that discriminate between (i) the effects of a variable on the probability that wages will change and (ii) conditional on wages changing, the effects of that same variable on the magnitude of the wage change. It is not appropriate to measure these second effects by applying conventional least-squares to the nonzero observations on Δw since, in this circumstance, the composition of the sample is being chosen on the basis of the values of the left-hand variable (that is, observations with zero values for Δw are being discarded) so that the conditional expectation of the error term is not zero. Nor is it appropriate to pool the limit (i.e., $\Delta w = 0$) observations with the nonlimit (i.e., $\Delta w \neq 0$) observations and estimate a least-squares regression over the entire sample. The gist of the bias that results in this event is illustrated by Figure 2 where the crosses indicate the observations describing the "true" relationship between an exogenous variable $\bar{X}$ and Δw and where the dashed line is the (biased) least-squares relationship estimated over all observations.[6]

Thus, consider the familiar situation in which in an attempt to reduce inflation the government appeals to the parties involved in collective bargaining to display "restraint" in their wage negotiations. A frequent research procedure to evaluate the effects of this policy of

Figure 2



wage restraint takes the form of a least-squares regression of an index of wage changes on a dummy variable that takes the value of unity when such a policy is in effect and of zero otherwise. In the presence of transactions costs in adjusting wages, this procedure does not provide an unbiased estimator of the effect of the incomes policy on the probability of wages being adjusted in any quarter nor of the effect of the incomes policy on the size of the wage change conditional on a wage settlement taking place. Moreover, the size and sign of the bias of the effect of the incomes policy depend upon the relationship between the dummy variable and all the right-hand side variables and, therefore, is likely to vary from study to study according to the particular regressors appearing in the regression and to the period of fit. Consequently, it comes as no surprise that the literature has reached little agreement on quantifying the impact of incomes policies on wage inflation.[7]

To address these sample selection problems, an estimation procedure proposed by Heckman [1979] has been applied: first a probit equation is estimated over the entire 112 quarters from 1948 I to 1975 IV to determine the probability of wages changing in the coal mining industry; then from this estimated probit equation the reciprocal of the Mills' ratio, M (the ratio of the ordinate of the estimated standard normal to the estimated tail area of the distribution), is constructed; and, finally, M so constructed is entered into a least-squares regression of Δw on a set of regressors estimated over the 24 quarters when wage rates actually increased.[8] Provided the equation disturbances are normally distributed, the ordinary least-squares estimator of the wage change equation that contains M is consistent, but it is not fully efficient since no account is taken of the complicated error structure.

III. Empirical Results

This section reports an analysis of the determinants of the frequency of wage changes and of the magnitude of wage changes in the British coal industry from 1948 to 1975. The wage series used in the empirical analysis is that determined by the national agreements between the National Coal Board and the National Union of Mineworkers and it covers all manual workers in the coal mining industry. As is well known, the actual earnings of coal miners differed from these nationally negotiated rates, particularly so for piece-rate workers. Nevertheless, for most workers in this industry, the largest changes in the level of their earnings were associated with the re-negotiation of their wage rates by the national agreements. Naturally, as with other collective bargaining agreements, the re-negotiation of wage rates constituted only one class of issues settled by these national agreements. For instance, changes in the structure of pay accompanying the 1955 National Day Wage Agreement and the 1966 National Power Loading Agreement were inextricably linked with decisions concerning the overall level of wages. Therefore, by focusing on the determinants of changes in the level of wage rates, we necessarily provide an incomplete analysis of the frequency of wage settlements in this industry.[8]

A. The Frequency of Wage Settlements

The probit equation estimated to determine the conditional probability of a wage change taking place in the coal industry is as follows:

$$\text{(1)} \quad \text{prob.}\ (\Delta w_t > 0) = N[\alpha_0 + \alpha_1 \Delta p_t + \alpha_2 \Delta x_t + \alpha_3 \Delta c_t + \alpha_4 U_t + \alpha_5 D_t + \alpha_6 (U.D)_t + \sum_{i=1}^{3} \alpha_{7i} S_{it} + \sum_{i=1}^{3} \alpha_{8i} I_{it}]$$

where precise definitions of the variables are given in Table 1 and where N denotes the standard cumulative normal distribution function. This equation contains the variables most often used to account for wage changes at the aggregate level (namely, the percentage change in retail prices (Δp_t), the unemployment rate (U_t), and the incomes policy variables (the I_{it}'s)) and it has been augmented with two industry-

Table 1

Descriptive Statistics on and Definitions of Variables

(The current quarter is denoted by t while quarter $t - s$ indicates the period when wage rates were last adjusted. Standard deviations are given in parentheses beneath mean values.)

	All quarters	Quarters for which $\Delta w_t > 0$	Quarters for which $\Delta w_t = 0$
$\Delta w_t = 100 \cdot (w_t - w_{t-s})/w_t$ where w_t is an index of coal miners' basic weekly wage rates (including the value of allowances in kind)	1.631 (5.094)	7.613 (8.816)	0
frequency of $\Delta w_t > 0$	.214 (.412)	1.000	0
$\Delta p_t = 100 \cdot (p_t - p_{t-s})p_t$ where p_t is the retail price index	3.867 (3.353)	5.632 (3.805)	3.384 (3.070)
$\Delta x_t = 100 \cdot (x_t - x_{t-s})/x_t$ where x_t is the output of deep-mined coal per manshift	1.550 (5.032)	2.879 (4.727)	1.187 (5.077)
$\Delta c_t = 100 \cdot (c_t - c_{t-s})/c_t$ where c_t is an index of the selling price of coal	4.777 (5.526)	6.422 (5.808)	4.328 (5.394)
U_t = aggregate unemployment percentage	1.995 (.749)	1.923 (.676)	2.014 (.770)
$D_t = 0$ from 1948 I to 1966 III, $= 1$ from 1967 IV to 1975 IV	.339 (.476)	.333 (.482)	.341 (.477)
$(U.D)_t = U_t \cdot D_t$	.929 (1.363)	.890 (1.313)	.940 (1.384)
$S_{1t} = 1$ for the first quarter of the year, $= 0$ otherwise	.250 (.435)	.333 (.482)	.227 (.421)
$S_{2t} = 1$ for the second quarter of the year, $= 0$ otherwise	.250 (.435)	.292 (.464)	.239 (.429)
$S_{3t} = 1$ for the third quarter of the year, $= 0$ otherwise	.250 (.435)	.083 (.282)	.295 (.459)
$I_{1t} = 1$ in those quarters when the government implements a wage "freeze", $= 0$ otherwise	.205 (.406)	.083 (.282)	.239 (.429)
$I_{2t} = 1$ in those quarters when the government implements a "twelve month rule", $= 0$ otherwise	.205 (.406)	.292 (.464)	.182 (.388)
$I_{3t} = 1$ in those quarters when the government implements a "wage ceiling", $= 0$ otherwise	.134 (.342)	.167 (.381)	.125 (.333)

Note to Table 1: The data on Δw_t, Δp_t, and U_t are taken from issues of the Department of Employment Gazette (or formerly the Ministry of Labour Gazette). Information on x_t is taken from issues of the Ministry of Technology Digest of Energy Statistics (or formerly Ministry of Fuel and Power Statistical Digest) and the data on c_t are from the Monthly Digest of Statistics. There are 88 quarterly observations for the sub-sample defined by $\Delta w_t = 0$ and 24 quarterly observations when $\Delta w_t > 0$.

specific variables, the percentage change in coal prices (Δc_t) and the percentage change in coal output per man-shift (Δx_t). All percentage change variables have been measured between the current quarter t and the quarter when wages were last altered. Given the absence of a reliable series on the expected rate of price inflation, we have followed the procedure of expressing the expected rate by a distributed lag function of the actual rate of price inflation (Δp). Seasonal dummy variables (the S_{it}'s) have been included to remove any systematic seasonal pattern in wage settlement dates while the presence of the dichotomous variable D_t [10] and of the interaction between the unemployment rate and D_t permits the effect of movements in unemployment on wage settlements to differ before and after the apparent shift in the unemployment-vacancy relationship in the mid-1960's.[11]

Inferences concerning the effects of incomes policies on the course of wage inflation are drawn from dichotomous variables indicating whether or not an incomes policy is in effect. The use of dichotomous variables to measure the effects of incomes policies is customary and it is also customary to lament the inadequacies of the procedure. This conventional practice is modified here by the use of three dichotomous variables that distinguish the effects of different forms of incomes policies: the variable I_1 identifies quarters when a wage "freeze" was in effect and ostensibly all wage increases were prohibited; the variable I_2 identifies quarters when a "twelve month rule" was instituted under which the time between wage settlements was mandated to be no less than one year; and the variable I_3 identifies quarters when wage ceilings were in effect, that is, quarters when the government declared wage increases to be no greater than a certain specified level. If the wage freeze achieves its stated purpose, then I_1 will reduce the probability of wages changing, while I_2 will operate in a similar fashion unless the period between wage negotiations would have been no less than twelve months in the absence of this incomes policy. A policy of imposing ceilings on wage increases is designed to affect the magnitude of the wage change and not the frequency of wage settlements.

The appendix to Fallick and Elliott [1981] was used to determine which incomes policies were operative in each quarter. The variable I_1 takes the value of unity from 1948 I to 1951 II, from 1961 III to 1962 I, from 1966 III to 1967 II, and from 1972 IV to 1973 I; I_2 takes the value of unity from 1967 III to 1970 IV and from 1973 II to 1975 II; and I_3 takes the value of unity from 1956 I to 1956 IV, from 1962 II to 1963 I, from 1965 II to 1966 II, and from 1975 III to 1975 IV. Of course, the identification of particular types of incomes policy is fraught with difficulties and it is important to determine whether the results are sensitive to the classification of the quarters with particular incomes policies. For this reason, the consequences of a second classification of incomes policies were examined and the results from this are reported below.

Estimates of the probit index parameters of equation (1) are given in column (i) of Table 2 with estimated partial derivatives of the probability of wage change function evaluated at sample means given in column (ii).[12] Thus the greater the increase in retail prices (Δp)

since the previous wage settlement, the higher the probability of wages changing in the current quarter. In particular, at sample means, a small increase in Δp above its average value raises the probability of a new wage agreement by 6.6% and this estimate is highly significant by conventional criteria. Increases in labor productivity raise the probability of a new wage settlement although this effect would not be judged as significantly greater than zero by the usual standards. By these same criteria, no significant association exists between movements in coal prices and in the frequency of wage settlements. On a conventional likelihood ratio test, the hypothesis that the probability of a wage increase is independent of movements in unemployment (that is, a joint test of α_4 and α_6) cannot be rejected at conventional levels of significance. Other things equal, seasonal variations in wage settlements appear to be of negligable importance. Of the three incomes policy dummy variables only the coefficient on that pertaining to the wage freeze, I_1, appears to be significantly less than zero: the results suggest that, at the mean values of the variables, the imposition of a wage freeze reduces the probability of a wage settlement in coal mining by almost 24%.

A number of variations on this equation specification yielded results that differed in no substantive sense from those reported in column (i) of Table 2. For instance, instrumental variables were found for the rate of change of coal prices,[14] but treating Δc_t as endogenous in this way did not materially affect the results. The same was true when the unemployment rate was measured as an average over four quarters instead of its one-quarter value used in the estimates in column (i).

Table 2

Estimates of the Determinants of the Probability of Wages Changing

(estimated standard errors in parentheses)

	(i) probit index	(ii) partial derivatives	(iii) linear probability	(iv) probit index	(v) partial derivatives
constant	-1.077 (.900)		.224 (.213)	-.717 (.927)	
Δp_t	.285 (.085)	.066	.069 (.018)	.226 (.073)	.055
Δx_t	.033 (.036)	.008	.008 (.009)	.026 (.034)	.006
Δc_t	-.007 (.045)	-.002	-.001 (.010)	.004 (.042)	.009
U_t	-.301 (.465)	-.070	-.099 (.113)	-.421 (.521)	-.103
D_t	3.280 (2.772)	.763	.746 (.465)	.967 (1.852)	.237
$(U.D)_t$	-1.266 (.947)	-.295	-.269 (.177)	-.473 (.730)	-.116
S_{1t}	.360 (.422)	.084	.084 (.102)	.256 (.401)	.063
S_{2t}	-.005 (.494)	-.001	-.001 (.116)	.170 (.474)	.042
S_{3t}	-.373 (.559)	-.087	-.061 (.120)	-.407 (.540)	-.100
I_{1t}	-1.754 (.609)	-.408	-.370 (.116)		
I_{2t}	-.540 (.969)	-.126	-.131 (.171)		
I_{3t}	.206 (.498)	.048	.078 (.114)		
I_{4t}				-.825 (.451)	-.202
I_{5t}				.050 (.714)	.012
ℓnL	-42.95			-47.27	

Notes: The logarithm of the likelihood function at its maximum is given by ℓnL.

Some researchers (e.g., Henry [1981]) have allowed for the possibility of an acceleration in the rate of wage inflation in the period following the relaxation of an incomes policy and this "catch-up" feature was investigated in these data by constructing dummy variables for those periods when an incomes policy was dropped and was not replaced by another form of incomes policy. In fact, the probability of a wage settlement exhibited no significant association with this "catch-up" variable.

Finally, equation (1) was modified to allow for a different classification of incomes policies, a classification that turns on whether the incomes policies have been "voluntary" on the one hand or "statutory" (or "compulsory") on the other hand. The dichotomous variable I_{4t} takes the value of unity in those quarters in which a "voluntary" incomes policy was in operation and of zero otherwise while the variable I_{5t} takes the value of unity in those quarters in which a "statutory" or "compulsory" incomes policy was in effect and of zero otherwise. The categorization of incomes policies into these two forms follows Parkin and Sumner ([1978] Table 1.1): I_{4t} takes the value of unity from 1948 I to 1951 II, from 1956 I to 1956 IV, from 1961 III to 1963 I, from 1965 II to 1966 II, and from 1974 II to 1975 III; and I_{5t} takes the value of unity from 1966 III to 1970 IV, from 1972 IV to 1974 I, and in 1975 IV. The estimates listed in column (iv) of Table 2 are the maximum likelihood results from fitting equation (1) after replacing the three incomes policy variables I_1, I_2, and I_3 with the two variables I_4 and I_5. The corresponding estimated partial derivatives of the probability of wage change function evaluated at sample means is given in column (v). The probit estimated coefficient on I_5 is clearly insignificantly different from zero by conventional criteria while that on I_4 is significantly less than zero and it suggests that, at the mean values of the variables, a voluntary incomes policy reduces the probability of a wage settlement in coal mining by almost 18%. In other words, these results given no support to the often-stated belief that a statutory incomes policy is more effective in reducing the frequency of wage settlements than a voluntary incomes policy.

B. The Size of Wage Changes

Consider now the second stage equation accounting for the size of the wage change in coal mining conditional upon a new settlement being negotiated. This condition is satisfied in only 24 of the 112 quarters from 1948 I to 1975 IV so a relatively small number of observations are available to account for the magnitude of wage increases. The least-squares fitted equation is

$$\text{(2)} \qquad \Delta w_t = \beta_0 + \beta_1 \Delta p_t + \beta_2 \Delta x_t + \beta_3 \Delta c_t + \beta_4 U_t + \beta_5 D_t + \beta_6 (U.D)_t + \sum_{i=1}^{3} \beta_{7i} S_{it} + \sum_{i=1}^{3} \beta_{8i} I_{it} + \gamma M_t + \varepsilon_t$$

where M_t is the inverse of the Mills' ratio formed from the probit estimates in column (i) of Table 2, where ε_t is a stochastic disturbance term, and where the other variables are defined in Table 1. The estimates of equation (2) are contained in column (i) of Table 3 and it is evident that, with so few degrees of freedom, little can be said with confidence

Table 3

Estimates of the Determinants of the Magnitude of Wage Increases

(estimated standard errors in parentheses)

	column (i)	column (ii)	column (iii)	column (iv)
constant	-.300 (.279)	.011 (.080)	.162 (.371)	.007 (.056)
Δp_t	4.978 (3.617)	.824 (.530)	-1.130 (4.615)	.818 (.326)
Δx_t	-.056 (.532)	-.593 (.267)	-.806 (.543)	-.593 (.197)
Δc_t	.168 (.411)	-.039 (.376)	.042 (.281)	.052 (.270)
U_t	-.079 (.070)	-.011 (.039)	.022 (.105)	-.020 (.032)
D_t	.868 (1.252)	-.433 (.567)	.474 (.348)	.547 (.291)
$(U.D.)_t$	-.241 (.379)	.159 (.159)	-.061 (.128)	-.099 (.087)
S_{1t}	.090 (.048)	.051 (.035)	.036 (.051)	.054 (.025)
S_{2t}	.026 (.050)	.008 (.048)	.034 (.036)	.043 (.029)
S_{3t}	-.014 (.068)	.030 (.057)	.098 (.119)	.051 (.039)
I_{1t}	-.289 (.242)	-.014 (.047)		
I_{2t}	-.234 (.369)	.131 (.195)		
I_{3t}	.025 (.042)	.014 (.041)		
I_{4t}			.069 (.176)	-.004 (.029)
I_{5t}			-.221 (.101)	-.211 (.095)
M_t	.215 (.186)		-.130 (.308)	
σ	.022	.026	.019	.019
R^2	.874	.857	.895	.893

Notes: The standard error of the regression is given by σ.

concerning the determinants of the magnitude of wage increases. Two of the incomes policy dummy variables have negative coefficients, but the confidence intervals span a wide area in terms of their economic implications. The consequences of ignoring the censoring of the left-hand variable are given by the estimates in column (ii) of Table 3 which omit the Mills' ratio variable, M. The point estimates change noticeably although, once again, the standard errors accompanying these coefficients do not permit confident inferences. Columns (iii) and (iv) of Table 3 replace the three incomes policy dummy variables I_1, I_2, and I_3 with I_4 (denoting a "voluntary" incomes policy) and I_5 (denoting a "statutory" or "compulsory" incomes policy). The Mills' ratio variable, M, in column (iii) is formed from the probit estimates in column (iv) of Table 2. Once again, it is difficult to derive confident inferences although on a conventional t-test the coefficient on the statutory incomes policy variable, I_5, does appear to be significantly less than zero. Even if variables such as the seasonal dummy variables, coal price changes, and productivity changes are arbitrarily excluded from these least-squares equations or if coal price changes are treated as an endogenous variable, the resulting parameter estimates on the remaining variables do not differ appreciably from those given in Table 3. Surely these results must be interpreted as being very discouraging with respect to our knowledge of the determinants of the magnitude of wage changes in the coal industry.

IV. Conclusions

Since the transactions costs of adjusting wages have been largely neglected in the literature on wage inflation, insufficient attention has been directed to the distinction between the frequency of wage settlements on the one hand and the magnitude of wage changes negotiated by these settlements on the other hand. Government policy-makers have been acutely aware of this distinction, however, and have directed their incomes policies at one time towards reducing the occurrence of wage settlements and at other times towards reducing the size of wage increases. This paper has presented estimates of the effectiveness of these incomes policies both on the frequency of wage settlements and on the size of wage increases in the coal mining industry. An analysis of a single industry avoids the aggregation bias that plagues much of the research on wage inflation. Reasonably precise estimates were derived of the effect of various variables on the frequency of wage settlements, but no such precise estimates were obtained for the size of wage changes. This leads naturally to the conjecture that the apparent "statistical significance" reported by researchers with respect to the performance of variables in models of aggregate wage changes reflects primarily the effects of these variables on the probability of wages being adjusted rather than the effects on the magnitude of wage changes conditional upon wages being adjusted. As for the implications of the incomes policy variables, a wage freeze appears to have a large and significant negative effect on the probability of the occurrence of a new wage settlement. The effects of a voluntary and a statutory incomes policy are different: the frequency of wage settlements appears to be affected more by a voluntary incomes policy while the magnitude of wage changes seems to be responsive to a statutory incomes policy.

Footnotes

1 Among the rare exceptions are Barro [1972] and Sheshinski and Weiss [1977].

2 For instance, Rowley and Wilton [1973] address the temporal aggregation problem when an equal number of settlements each period are assumed to take place. Or Ashenfelter and Pencavel [1975] and Smith and Wilton [1978] augment conventional equations of quarterly wage inflation with a variable measuring the fraction of workers receiving new wage settlements.

3 The main exceptions include work by Coutts, Tarling, and Wilkinson [1976] and Elliott and Dean [1978].

4 It is recognized in Dicks-Mireaux and Dow's [1959] seminal paper and in Black and Kelejian [1972]. Evidence on aggregation bias is contained in Rowley and Wilton [1974]. Work with individual contract data is presented in Christofides, Swindinsky, and Wilton [1980], Hamermesh [1970], and Riddell [1979].

5 This second specification is Rosett's [1959] model of friction.

6 In fact, in the British coal mining industry, from 1948 to 1975 nationally negotiated wage rates never fell so all the nonlimit observations on Δw were positive values.

7 For instance, Parkin [1978] finds "no evidence at all either from the studies of the determinants of inflation or from more direct studies that wage and price controls materially affect the rate of inflation" (p. 49) while Henry [1981] claims, "Here the outstanding feature is the significant effects for incomes policy generally observed, <u>whatever the basic model of inflation used</u>" (p. 42, emphasis in original). Blackaby's [1978] survey of incomes policy concludes, "In sum, if we suppose that over this period there was a variety of forces tending to produce a rising trend in the increase in money earnings, then the experience of the late 1960's and early 1970's is consonant with the view that the two incomes policies during that period were successful, so long as they were in force, in holding back this process" (p. 392).

8 Thus, if $\hat{Q}$ is the predicted value of the index from a fitted probit on the determinants of a wage change taking place and if N is the standard cumulative normal distribution function, then $M_t = N'(Q_t)/N(\hat{Q}_t)$. See Heckman [1979] for further details.

9 The adjustment of wage rates by decisions reached at the national agreements was clearly prompted in the 1950's by movements in pay differentials between time-rate workers and piece-rate workers.

However, a continuous, quarterly, series on the earnings of these two types of workers is not available. More information on industrial relations in the coal industry is contained in McCormick [1979].

10 The date from which D_t takes the value of unity (the fourth quarter of 1966) is associated with the introduction of earnings-related supplements to unemployment compensation.

11 The national unemployment rate was used in these equations in preference to unemployment in the coal industry for several reasons. First, voluntary separations from the industry were "considerably affected by fluctuations in the general level of demand for labor" (McCormick [1979], p. 106) which is often indexed by the aggregate unemployment rate. Second, when a contraction in the coal industry's labor force was required after 1957, this was accomplished primarily through early retirements and through quits for other jobs and these individuals frequently escaped registration as unemployed. Third, it is by no means clear how to interpret movements in an industry's unemployment rate when labor supply curves to an industry are highly elastic.

12 The effect of a variable X on the probability is derived by multiplying the coefficient of the index, Q, by the normal density N'(Q) at the evaluation point; that is, ∂ prob./$\partial X = (dN/dQ)(\partial Q/\partial X) = N'(Q)(\partial Q/\partial X)$. At the sample mean values of the variables, the standard normal density is .228 for the estimates in column (i). As is evident from a comparison of the entries in column (ii) with those in column (iii), the ordinary least-squares estimates and the probit estimates of equation (1) imply very similar partial derivatives of the probability of wage change function.

13 In other estimates, the incomes policy dummy variable I_1 was interacted with Δp, Δx, U, D, and (U.D) to test the hypothesis that the whole relationship between the probability of wages changing and these variables alters when a "wage freeze" is in effect. A conventional likelihood ratio test fails to reject the null hypothesis.

14 The instrumental variables consisted of the percentage change in the price of gas and electricity, the percentage change in the price of construction inputs, and the percentage change in an index of consumer total expenditure. The source for these three variables in the Monthly Digest of Statistics.

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