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Market Power and Wage Inflation

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The interactions between a firm's oligopoly power in its product markets and the wages paid to its employees have been discussed in both static and dynamic contexts. Most of the interest in the static aspects of this problem stems from the theory of countervailing power between large firms and multiplant unions which was proposed by Galbraith (1952). This theory suggests that unions will organize in industries where product market power exists, and that both imperfections will tend to raise wages and the quality of labor hired in the industries above what these would be in a competitive economy.[1] It does not, however, contain any implications for the behavior of wages over the business cycle, although some work has attempted to demonstrate that wage changes are larger in concentrated industries.[2] In this study we examine wage dynamics as they are affected by the firm's market power. We show that

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[1]Lewis (1963, pp. 177-78) uses industry data to estimate the effect of industrial concentration on wage levels, while Weiss (1966) estimates the same effect using Census data on the wages of individual workers.

[2]Levinson (1960) finds positive simple correlation coefficients of year-to-year wage changes and industrial concentration in manufacturing industries. Bowen (1960, p. 62) correlates concentration and wage changes in six sub-periods between 1947 and 1959. He finds no relation in three of these, but a significant positive relation in three others. Using a multivariate regression model in which concentration is an independent variable, Allen (1968) finds weak evidence that concentrated industries exhibit more rapid rates of wage inflation, while Eckstein and Wilson (1962) could not produce this result.

Perhaps the clearest statement of this view of inflation is in Cabinet Committee (1969, p. 138), "In brief, the discretionary power to demand and to grant wage increases in excess of average productivity gains can give an upward thrust to the price level [italics mine]." A more complex theory based on the ratchet effect on wages and prices of greater downward price rigidity in more concentrated industries is suggested by the Council of Economic Advisers (1969, p. 107). This formulation of the interaction between product and labor markets also implies that a strong antitrust policy will have a mitigating effect on inflation. It is tested in Section V of this paper.

market power does not have any long-run effect on the rate of change of money wages, but that it does have important effects on their time path.

I. Planning Firms and Wage Bargaining

The appropriate specification of the relation between wage inflation and the structure of product markets is suggested by the theory of the planning nature of firms with market power.[3] A large firm, the success of whose projects depends on being able to predict supply and product prices, will attempt to ensure the success of these products by making their prices less responsive to market forces. It does this by engaging in planning with a horizon of the same length as the expected lifetime of its new products. If it reduces the uncertainty inherent in unstable product and factor prices, its new products are less likely to show losses during their initial period on the market. In no field will the importance of such planning be greater than in the area of changes in wages paid to the firm's employees. During the ten-year period 1960-1969 the average manufacturing firm paid out 26 percent of its sales receipts in the form of compensation to its employees;[4] if attempts at controlling price

[3] Galbraith (1967, p. 37) states, "A firm cannot usefully foresee and schedule future action or prepare for contingencies if it does not know what its prices will be, what its sales will be, what its costs including labor and capital costs will be and what will be available at these costs. If the market is unreliable, it will not know these things. Hence it cannot plan . . . . Much of what the firm regards as planning consists in minimizing or getting rid of market influences."

[4] Computed from Survey of Current Business, July 1962, July 1966, and July 1970.

variation by firms with substantial market power exist anywhere, they should be observable in the behavior of wages paid by these firms. We should expect that the path of wages in these firms will not be influenced by the market forces which affect the course of wages in other firms for which industrial planning is not so important.

Consider two polar cases. The first is a firm which has enough market power and is of sufficient size to engage in planning behavior of the Galbraithian variety over a period of five to ten years, the lifetime of large new products it introduces to the market. If it is successful in its planning efforts, the short-run change in the money wages it pays will be:

(1) $\dot{W}(t) = \alpha_1 + \varepsilon_1(t)$,

where $\dot{W}$ is the rate of change of money wages, α_1 is some constant and ε_1 is a random error term.[5] In addition to not being affected by market forces, wage changes in the successful planning firm will deviate only slightly from the constant, α_1; i.e., the variance of the random error term ε_1; will be small relative to the variance of the disturbances affecting wage changes in firms which do not engage in planning.

[5] Reder (1962, pp. 293-294) suggests that there may exist a sluggishness of wage changes in response to changes in market forces. This notion is not empirically distinguishable from our hypothesis that firms with more discretionary power in the product market seek more constant rates of wage increases. George de Menil postulates that wage rigidity will arise in concentrated industries from the conjunction of firms' seeking a target rate of return to prevent entry and minimizing price variation to satisfy customers. With prices growing at a constant rate the firm can only ensure attainment of its target rate of return by also minimizing wage variability.

The second polar case is a firm whose wages are completely subject to the vagaries of market forces. It will have wage changes characterized by:

$$(2)\quad \dot{W}(t) = \alpha_2 + \beta_2 X(t) + \epsilon_2(t) \quad ,$$

where α_2 is a constant, β_2 is a vector of parameters, X is a vector of variables affecting the path of money wage changes and ϵ_2 is another random error term. The actual variables which make up the vector X are not particularly important for this theoretical discussion. The major point is that in this firm, unlike in the planning firm, wage changes are very strongly affected by market forces. The firm does not need to predict over a period of years what its wage bill will be, for its products are not dependent for their success on planned factor prices. Its wage rates will be affected by the economic conditions prevailing at the time when changes in wages are negotiated with the union bargaining for its employees or are granted unilaterally by the firm.

In reality no firm is characterized by the polar cases implied by equations (1) and (2). Even the firm for which planning is most important will to some extent be affected by market forces in its determination of changes in its money wage rates, and even firms with little market power will attempt to reduce uncertainty by exerting some control over the market forces affecting wage changes. For this reason we assume for simplicity that there is a linear relationship linking (1) and (2) along the continuum of a firm's ability to plan. If this assumption is valid we can take a weighted average of these equations to derive an equation characterizing wage changes in the firm which plans such that it has DIS_i amount of discretion over the extent to which market forces affect changes in wages. (DIS_i is some fraction between zero and one.) This average is:

$$\dot{W}(t) = [\alpha_1 DIS_i + \alpha_2(1-DIS_i)] + \beta_2(1-DIS_i)X(t) + \varepsilon_1(t)DIS_i + \varepsilon_2(t)(1-DIS_i)$$

or

$$(3)\quad \dot{W}(t) = \alpha_2 + DIS_i(\alpha_1-\alpha_2) + \beta_2(1-DIS_i)X(t) + \varepsilon_2(t) + DIS_i(\varepsilon_1(t) - \varepsilon_2(t)).$$

Divergence in wages between firms which have more or less monopoly power will occur if the overall effect of that power is not zero. Differentiating with respect to DIS in (3) and taking means over a sample period, we have:

$$\frac{\partial \dot{W}}{\partial DIS_i} = (\alpha_1-\alpha_2) + \beta_1\overline{X} \quad .$$

This derivative will be zero if there is no divergence in wages. Testing whether $\frac{\partial \dot{W}}{\partial DIS_i}$ equals zero provides a test of whether wages in firms with greater discretionary power rise more rapidly than wages elsewhere.

Some previous empirical work based on simple correlations has found that discretionary power has an effect on the average rate of wage increase. These results contradict the observation that wages in industries in which such power exists have not in the long run diverged from other wages. The results may be due to misspecifications of the estimating equations or to a failure to observe the sample over a period of sufficient duration. The more sophisticated specification of the effects of discretionary power on wage dynamics embodied in (3) should enable us to discover whether market power has any total effect on the rate of wage inflation in an economy, whether it merely produces short-run changes in the time path of wages or whether it has no effects whatsoever.

The theory we have outlined implies that the rate of inflation is not affected by product-market power except in the short run. Market power under this theory results in both the downward wage rigidity suggested by the Council of Economic Advisers (1969, p. 107) and upward wage rigidity as well. If the theory is empirically valid, vigorous antitrust policy will not lessen the rate of inflation but will only increase wage and price variability.[6] Rather than shifting the short-run Phillips curve downward to the left, it will merely pivot the curve to make it more vertical.[7]

II. Data and Estimation Techniques

The model embodied in (3) is estimated using data on the rate of change in wage rates negotiated by individual firms and the unions with which they bargain. Data on these firms covering the period January

[6] This view is contrary to that expressed by Samuelson and Solow (1960, p. 194).

[7] Lipsey and Parkin (1970) suggest that incomes policies will produce an opposite shift in the short-run Phillips curve.

1954 through June 1970 are used in our estimates; earlier data which are available are not used in order to avoid contaminating the estimates with possible residual effects of the wage controls of World War II as well as the direct effects induced by the wage stabilization policies of the Korean period. The equation is estimated over the entire sample of contracts available during the period of observation and over a subsample from which contracts embodying complete cost-of-living escalator clauses are deleted. In order to ensure that most of the manufacturing sector is covered by the firms in our sample we include only one large firm from each two-digit SIC industry. Since some of these industries have no firm which is significant and others are dominated by firms whose main products are in other industries, the sample contains only fourteen firms, each of which is listed in Table I along with the name of the two-digit industry in which much of its activity is concentrated.

The wage changes variable $\dot{W}$ is computed using the method discussed in Hamermesh (1970).[8] Each observation is the annual rate percentage change in the base wage paid by the firm compared to the wage negotiated in its previous contract. These bargain data are the appropriate ones to use because they account correctly for the timing of wage decisions and thus enable us to abstract from differences among industries which are the result of differences in the frequency of wage changes.

[8] The data are computed from the Bureau of Labor Statistics, Wage Chronologies and Current Wage Developments, and from the Bureau of National Affairs, Daily Labor Reports. An appendix listing these data is available upon request from the author.

Table I

Firms in the Sample and Their Major Industries

Firm	DIS 100	Major Industry
Armour	26.64	Food
American Tobacco	72.34	Tobacco
Dan River	32.58	Textiles
Weyerhaeuser	23.64	Lumber Products
Simmons	18.30	Furniture
International Paper	25.26	Paper Products
American Cyanamid	37.85	Chemicals
Sinclair	35.53	Petroleum Products
Firestone	55.83	Rubber Products
International Shoe	24.25	Leather
Pittsburgh Plate Glass	67.57	Stone, Clay and Glass
U.S. Steel	33.32	Primary Metals
General Electric	54.49	Electrical Equipment
General Motors	55.21	Transportation Equipment

They also reflect more directly than would earnings data the firm's efforts to control variations in factor prices. Earnings data vary depending on productivity and on decisions by plant managers and foremen to vary the skill mix of employment. These individuals are hardly likely to be concerned with the need for longer-run planning, but the higher executives involved in negotiating wages will be more aware of this need.

Only two variables, the rate of unemployment in the labor force and the change in the aggregate price level, have consistently been found to have an effect on wage changes in the United States, and these are the variables which comprise the vector $\dot{X}$ in our model. The change in the consumer price index, $\dot{C}$, is computed at an annual rate between contract dates, exactly as is the $\dot{W}$ variable. The unemployment rate variable U is the seasonally adjusted unemployment rate in the month when the wage negotiations are completed.

The major problem arises in the choice of a variable to represent DIS , the firm's discretionary power and desire to engage in controlling variations in factor prices. We need a measure of product-market concentration based not on an industry but rather on the specific product markets in which each firm operates. The ideal measure would be a weighted average of the percentages of shipments accounted for by the firm in each small industry in which it operates. Given the lack of these data the most appropriate measure is a weighted average of the 1963 four-digit SIC top-four concentration ratios, in which the weights indicate the extent of the firm's participation in each industry. The particular weights we use are the firm's plant equivalents in each four-digit industry as estimated by the 1961 Fortune plant-and-product survey.[9] While there may be substantial errors in these data as well as the usual conceptual problems in defining measures of market power, their use in computing DIS should reflect far more accurately than would the raw industrial concentration data the actual amount of market power available to each firm. Because the firms in our sample are fairly large we can furthermore be sure that they are among the top four at least in those four-digit industries which comprise the bulk of each firm's output.

Because our measure of discretionary power and the firm's desire to plan may not be accurate, instead of estimating (3) we initially estimate:

$$(4)\quad \dot{W}(t) = \alpha' + \alpha''DIS_i + \beta'\dot{C}(t) + \beta''\dot{C}(t)\cdot DIS_i + \gamma'U^{-1}(t) + \gamma''U^{-1}\cdot DIS_i + \delta GPOST + \varepsilon'(t),$$

where U is included in hyperbolic form following a substantial number of studies which have demonstrated that this form of the variable produces the best explanation of variations in changes in money wage rates. We have added a dummy variable GPOST for the years 1962-66 to account for the possible coincidence of lower than expected wage changes with the wage guideposts in effect in these years. We can test for the importance of DIS_i in explaining wage changes by testing the constraint:

$$(5)\quad \alpha'' = \beta'' = \gamma'' = 0.$$

[9] I am indebted to my colleague Charles Berry for making these data available to me.

If this constraint is not valid, we may conclude that DIS belongs in the equation in some form.

Having modified (3) by transforming it into (4), we can test whether firms which are most likely to engage in planning (for which DIS equals one) respond at all in the short run to changes in the market variables U^{-1} and $\dot{C}$. This can be done by imposing on (4) the constraints:

$$(6) \qquad \beta' = -\beta'' \text{ and } \gamma' = -\gamma'' \quad .$$

If this pair of constraints is valid we may conclude that, if our index showed that the top-four concentration ratio were one hundred in all industries a firm participates in, that firm would also not vary its rate of increase in money wages in response to short-run changes in labor-market conditions.

Using (4) we can also test whether DIS has any overall effect on the rates of change in wages. The appropriate statistic is the derivative:

$$(7) \quad \frac{\partial \dot{W}}{\partial DIS_i} = \alpha'' + \beta''\bar{\dot{C}} + \gamma'' \overline{U^{-1}} \quad ,$$

where we have taken expected values of the independent variables. Assuming that these variables are fixed in repeated samples, we can compute the variance of this derivative using the parameter variances and covariances and then test whether the derivative differs from zero. If it does not, we may conclude that a firm's market power has no effect on the average rate of growth of its employees' wage rates, and that an antitrust policy which entails reducing these firms' market power would not directly affect the long-run rate of wage inflation.

III. Empirical Results

Equation (4) alone and with constraints (5) and (6) imposed is estimated using the wage bargains data for the years 1954-1970. The weighted ordinary least squares estimates are presented in Table II.[10] Longer contracts are weighted more heavily so that the parameter estimates we derive are more comparable to those based on periodic observations of more aggregated data. Examining first the results for equation (4), we see that the signs on $\dot{C}$ and U^{-1} are positive, as is predicted by theory. The independent variables which we use account for over half of the variation in $\dot{W}$, a very high proportion considering that we are pooling cross-section and time-series data. The guidepost variable is negative and highly significant as in previous work in which only this variable, $\dot{C}$ and U^{-1} are included. An interaction term between this variable and DIS produced a t-value of only -.6 in another formulation of this equation. There is thus only weak evidence that the wage-price guideposts affected wage bargaining more in firms in which product market power is greater.

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[10] Ordinary least squares yields consistent estimates of (4). The covariance of DIS and the residual in (4) is:

$$E(DIS_i \cdot \varepsilon'(t)) = E\left\{DIS_i[\alpha_1-\alpha_2]\cdot[\varepsilon_2(t) + DIS_i(\varepsilon_1(t) - \varepsilon_2(t))]\right\}$$

from (3). Since $E(\varepsilon_2(t)) = E(\varepsilon_1(t) - \varepsilon_2(t)) = 0$, the covariance of DIS and the residual in (4) is zero whenever DIS is independent of the ε_1 and ε_2 (which was assumed in deriving (3)). A similar result can be demonstrated for the other terms in (4) involving DIS. What is true is that the residuals in (4) are heteroskedastic.

Table II

Estimates of (4) and (6), 1954-1970[a]

	All Contracts		Contracts without Escalators	
Equation	(4)	(6)	(4)	(6)
Variable				
Constant	-3.25	-2.18	-3.83*	-2.80*
	(-1.78)	(-1.83)	(-2.05)	(-2.25)
DIS	9.80*	6.81*	9.93*	6.95*
	(2.38)	(4.91)	(2.34)	(4.84)
$\dot{C}$	.77*		.33	
	(2.61)		(.99)	
$\dot{C}\cdot$DIS	-.78		-.39	
	(-1.16)		(-.53)	
U^{-1}	32.44*		39.72*	
	(3.65)		(4.22)	
$U^{-1}\cdot$DIS	-41.40*		-48.28*	
	(-2.04)		(-2.26)	
$\dot{C}$(1-DIS)		.76*		.30
		(4.26)		(1.46)
U^{-1}(1-DIS)		27.49*		35.21*
		(4.96)		(5.83)
GPOST	-1.38*	-1.39*	-1.29*	-1.29*
	(-4.39)	(-4.44)	(-4.01)	(-4.04)
R^2	.597	.595	.603	.601
N =	118	118	99	99
$\frac{\partial \dot{W}}{\partial DIS_i}$ =	-.16	-.23	-.87	-.94
	(-.21)	(-.21)	(-.88)	(-1.09)

[a]t-values in parentheses;

*Significant at the 5% level.

Comparing the results between the two samples, we see that there is substantially less response to changes in the cost of living in the second sample from which contracts including escalator clauses have been deleted. On the other hand, the response to the unemployment rate is much greater in this sample. The former result is not surprising, for an escalator clause assures that wage changes will respond completely to any change in the cost of living. When those contracts including this provision are deleted, we should expect that the average response will decrease. The latter result follows directly from the former, for holding other variables constant the positive correlation between $\dot{C}$ and U^{-1} produces an increase in the coefficient of U^{-1} when the coefficient of $\dot{C}$ falls.

Table III

F-Statistics of Imposing Constraints (5) and (6) on Equation (4)

	All Contracts	Contracts without Escalators
(5)	2.92* (3,111)	3.34* (3,92)
(6)	.31 (2,111)	.32 (2,92)

*Significant at the 5% level.

The hypothesis that wage changes are less variable in firms in which planning is likely to be more important, as indicated by a higher value of the DIS variable, is supported by the results in Table II. These firms apparently do respond less to changes in the market variables, especially to changes in the unemployment rate.[11] Moreover, the constant term involving DIS is positive and significant. Comparing equation (4) to an equation in which (5) is imposed, we find that the F-statistics listed in Table III for the inclusion of DIS in this equation indicate a significant difference between the explanatory power of the two equations. The extent of a firm's ability to control the market is important in describing the path of wages paid to its employees, and our evidence indicates that wages in firms with this type of power have time paths which are less affected by market forces than wages elsewhere in the economy.

The apparent decline in wage variability as market power increases is not merely an artifact produced by the possible effect of more frequent wage negotiations by firms with less market power. In our sample the rank correlation between DIS and the average length of contract is -.05, indicating that firms with more market power negotiate wages as frequently as other firms. The rigidity introduced by wage contracts of fixed duration is thus not the cause of our results.

Imposing constraint (6) on equation (4) is always justified according to our estimates. As Table III shows, the F-values of the tests of this constraint are not significantly different from zero. We may thus conclude that, if we had a firm in our sample in which the variable DIS took on

[11] Equation (4) was also estimated with the threshold specification of the effects of changes in the cost of living on wage changes (Hamermesh 1970). The interaction term of the cost of living with DIS is nearly zero for changes in the cost of living less than 2 percent per year, but is significant for changes above 2 percent per year.

a value of unity and if our linear weighting procedure used to derive (3) is correct, we would find that the response of changes in its wage rates to the market variables we have used would be zero. Not only is it essential to include interaction terms between DIS and other variables; our estimates show that market forces would not affect wage decisions in the short run in firms which were most likely to plan.

In none of the four estimated equations is $\frac{\partial \dot{W}}{\partial DIS_i}$ significantly different from zero. The average effect of DIS on changes in wage rates (the product of $\frac{\partial \dot{W}}{\partial DIS_i}$ and the average value of DIS) is -.35 percentage points per year in the second sample. Integrating over the sixteen-year sample observation period, we find that even the highest estimate of the effect of DIS on wage differentials is only -6 percent, and its variance is such that the true effect could very possibly be positive. These estimates provide some evidence that, as predicted by the theory of the supply of labor to firms and industries, wage divergence is not produced by interfirm differences in market power.

Using the estimates involving constraint (6), we can analyze what the average response of wage changes to market forces would be. Since the weighted average value of DIS is .40, we see that the average response of wage changes to a one percent rise in the cost of living is .46 percent in the first sample and .18 percent in the second. The effect of a one percent rise in the unemployment rate evaluated at the mean unemployment rate during the sample period is to lower the rate of wage inflation by .69 percent in the first sample and .89 percent in the second.

Both of these are similar to those results derived in a number of previous studies (Perry 1966).[12]

Although the ordinary least squares estimates of (4) are consistent, the derivation of this equation suggests that they may be inefficient because of heteroskedasticity in the residuals. Using (3) and assuming there is no correlation between $\varepsilon_1(t)$ and $\varepsilon_2(t)$, we have:

$$(8) \qquad \mathrm{Var}\ (\varepsilon') = \mathrm{Var}\ (\varepsilon_2) + \overline{\mathrm{DIS}}_i^2(\mathrm{Var}\ (\varepsilon_1) - \mathrm{Var}\ (\varepsilon_2)) \quad .$$

If the planning firm is successful in its efforts to control variations in the price of its labor, we should expect the second term in (8) to be negative. If this hypothesis is valid, the average residual in (4) will have a smaller absolute value in those firms which are more likely to engage in planning, i.e., where DIS_i is large. Tests for this type of heteroskedasticity were made by regressing the absolute values of the residuals on DIS_i. In all cases this latter variable explained less than .1 percent of the variance in the absolute values of the residuals. The result implies that our ordinary least squares estimates are producing efficient parameter estimates, at least as far as this source of difficulty is concerned. It suggests, however, that random disturbances have as much

[12]Equation (4) and constraint (6) were also estimated over a sample of industry data using four-quarter changes in wage rates for the years 1957-1969. While these data are less appropriate for estimating an equation derived from the planning hypothesis than the firm data we have used, they corroborate to some extent our results. The coefficient on the term $U^{-1}\cdot\mathrm{DIS}$ is negative and highly significant, but the coefficient on the interaction term between DIS and the cost of living is positive but insignificantly different from zero. Because these data are composed of changes in the skill mix of employment as well as changes in productivity through wage incentive payment schemes, the failure to observe the correct sign on this second interaction term does not conflict with our evidence using the more appropriate firm data.

effect on wage changes in firms likely to engage in planning as they do in other firms. While planning firms are able to reduce the effects of market forces on wage changes, they are apparently unable to reduce random variations in the time paths of their wage rates.

IV. The Effects of Differential Union Organization

The estimates obtained in Section III may arise from the positive correlation between the degree of product market power and the extent of union organization. In our sample the rank correlation between DIS_i and UNZ_i, the extent of collective bargaining coverage in the two-digit industry where most of the firm's output is located, is +.65. Rosen (1969) has presented some evidence that the extent of unionization is positively correlated with the union relative wage effect in unionized plants. Because their wage is thus substantially above the competitive level, unionized firms in heavily unionized industries will not need to raise wages very much to attract labor when product demand increases. Similarly, when product demand falls the greater power of unions in firms in highly unionized industries will enable them to be more capable of maintaining at least some wage increases than unions elsewhere. For both of these reasons we should thus expect UNZ_i to increase the stickiness of wages. This effect is not caused by the institution of fixed-duration wage contracts in our estimates, but rather by the behavior of firms and unions in industries among which union power varies.

We use as a measure of UNZ_i the 1958 Bureau of Labor Statistics data in Lewis (1963, p. 274). To account for the effect of differential unionization we could add the following variables to (4): UNZ_i, $\dot{C}(t)\cdot UNZ_i$ and $U^{-1}(t)\cdot UNZ_i$. These terms would correspond to the three terms containing DIS_i which are already included in (4). Because of the high degree of collinearity between DIS and UNZ this procedure was not successful. It produced estimates which changed drastically when the nineteen contracts containing escalator clauses were dropped. To circumvent this difficulty we form the variable:

$$(9) \quad COMB_i = \omega DIS_i + (1-\omega)UNZ_i \quad .$$

Both the unconstrained version of (4) and the form with constraint (6) imposed are then estimated using COMB with the values of ω = 0, .25, .5, .75 and 1 in each place where DIS appears. Searching over this grid is equivalent to a stepwise maximization of the likelihood function; the maximum likelihood estimate of ω is that value for which the sum of squared residuals in (4) is minimized.

The results of using this combined form of both product-market power and the extent of unionization are shown in Table IV. In all four estimated equations $\hat{\omega} = .5$. In the unconstrained version of (4) $\hat{\omega}$ is significantly different from both zero and one, while in the constrained version of (4) the parameter estimate is substantially greater than its standard error. These results indicate that greater values of both DIS and UNZ decrease the variability of wages. Each of these factors has an independent effect on interindustry differences in short-run wage dynamics.

Table IV

Estimates of (4) and (6), 1954-1970, using COMB[a]

	All Contracts		Contracts without Escalators	
Equation	(4)	(6)	(4)	(6)
Variable				
Constant	-5.18*	-2.30	-6.76*	-3.13*
	(-2.01)	(-1.74)	(-2.53)	(-2.27)
COMB	13.51*	6.53*	15.93*	6.98*
	(2.40)	(3.93)	(2.70)	(4.09)
$\dot{C}$	1.12*		.27	
	(2.46)		(.49)	
$\dot{C}\cdot$COMB	-1.50		-.22	
	(-1.53)		(-.19)	
U^{-1}	40.13*		55.37*	
	(3.13)		(3.94)	
$U^{-1}\cdot$COMB	-55.90*		-80.15*	
	(-2.00)		(-2.63)	
$\dot{C}$(1-COMB)		.83*		.31
		(4.29)		(1.39)
U^{-1}(1-COMB)		29.21*		38.29*
		(4.90)		(5.91)
GPOST	-1.46*	-1.43*	-1.38*	-1.36*
	(-4.71)	(-4.62)	(-4.37)	(-4.28)
R^2	.608	.600	.616	.605
N=	118	118	99	99
ω	.5*	.5	.5*	.5
	(2.16)	(1.33)	(2.53)	(1.93)
$\frac{\partial \dot{W}}{\partial COMB_i}$ =	-.72	-1.00	-1.29	-1.59
	(-.67)	(-.88)	(-.96)	(-1.35)

[a]t-values in parentheses;

*Significant at the 5% level.

The other results in Section III are corroborated when we estimate (4) using both DIS and UNZ . The interaction term with the unemployment rate is always significantly different from zero, and the cost-of-living interaction is negative but not significant. As the F-values in Table V demonstrate, the inclusion of the three terms involving COMB adds significantly to the explanatory power of the equation. Furthermore, the tests of constraint (6) suggest that wages would not vary about their trend in a firm with very great monopoly power in a heavily unionized industry.

Table V

F-Statistics of Imposing Constraints (5) and (6) on Equation (4) Estimated Using COMB

	All Contracts	Contracts without Escalators
(5)	4.06* (3,111)	4.42* (3,92)
(6)	1.22 (2,111)	1.31 (2,92)

*Significant at the 5% level.

V. Of Ratchets and Spillovers

The discussion in the previous sections focused on the question of whether discretionary power in the product and labor markets produces larger changes in wage rates or affects merely the time path of wage changes. In this section we analyze in detail the more subtle question of whether wage increases in firms where this power exists produce an indirect effect on wages in firms where this power is absent. (We denote indirect effects due to product market power as ratchets and those due to labor market power as spillovers.) Even in competitive labor markets wage increases in one firm will affect wages in other firms through the mechanism of the adjustment of labor supply to changes in relative wages among firms. What the ratchet-spillover hypothesis must imply is that there are certain firms or industries whose wage increases provide the spur to later wage increases in other firms or industries.

Previous attempts to specify this mechanism have ignored the problem of timing and simply included a variable representing wage changes in a key industry or sector as an additional determinant of wage changes in other sectors.[13] Our use of data explicitly linked to the timing of wage decisions enables us to circumvent this problem. We construct variables representing wage changes in firms where the most substantial discretionary power exists. If the ratchet-spillover hypothesis is correct, wage changes at any point in time in firms where this power is absent will, *ceteris paribus*, be positively related to some function of the wage increases

[13]Eckstein and Wilson (1962, pp. 394-396) find a significant effect on wage changes in non-key industries of wage changes over the same period in industries in what they designate as the key group. McGuire and Rapping (1968) include variables representing the change in wages in the automobile and steel industries in equations explaining variations in wages in other industries. Both the wages in these two key industries and those in the other industries are calculated over the same time period, but there is only weak evidence that auto and steel wages have independent positive effects on wages in other industries.

in key firms which occurred since the last time wages in the non-key firm were adjusted. We thus form variables of the sort:

$$SPILL_i(t) = \begin{cases} 0 & \text{if the i'th firm is among the top five in the sample in its value of DIS (UNZ);} \\ WKEY & \text{if the i'th firm is not among this group, and where WKEY is the average wage increase in the top five firms between time } t \text{ and the last time the i'th firm negotiated a wage change.} \end{cases}$$

Separate variables are calculated for DIS and UNZ in order to test the hypothesis both for ratchet effects on wages stemming from firms with the greatest amount of product market power and for spillover effects from firms in industries which are most heavily unionized.[14] The cut-off point of five top firms in the ranking of DIS and UNZ is completely arbitrary, but variations around this cut-off produced no qualitative changes in our results.

The variable DISSPILL is entered as an additional variable in equation (4'), and the variable UNZSPILL is entered into an equation like (4') but with UNZ appearing in each place where DIS appears in (4'). The results of this set of estimates are presented in Table VI. We do not present estimates of $\frac{\partial \dot{W}}{\partial DIS_i}$ or $\frac{\partial \dot{W}}{\partial UNZ_i}$ because of the difficulty of

[14] The top five firms in the sample ranked by the value of DIS are American Tobacco, Pittsburgh Plate Glass, Firestone, General Motors and General Electric. The top five in the ranking by percent of workers unionized in the firm's major industry are U.S. Steel, Sinclair, Pittsburgh Plate Glass Firestone, and General Motors.

Table VI

Estimates of (4), 1954-1970, with Inclusion of Spillover Variables

	DISSPILL		UNZSPILL	
	All Contracts	Contracts without Escalators	All Contracts	Contracts without Escalators
Variable				
Constant	-1.06 (-.52)	-2.25 (-1.08)	-4.25* (-4.31)	-7.36* (-2.19)
DIS	6.24 (1.43)	7.57* (1.71)		
UNZ			10.28 (1.58)	15.73* (2.30)
$\dot{C}$	.84 (2.89)	.35 (1.06)	1.27* (2.14)	-.01 (-.02)
$\dot{C}\cdot$DIS	-.93 (-1.40)	-.42 (-.58)		
$\dot{C}\cdot$UNZ			-1.68 (-1.42)	.37 (.27)
U^{-1}	32.52* (3.72)	40.76* (4.37)	32.22* (2.02)	57.59* (3.26)
$U^{-1}\cdot$DIS	-39.93* (-2.00)	-49.34* (-2.33)		
$U^{-1}\cdot$UNZ			-36.00 (-1.11)	-78.62* (-2.21)
DISSPILL	-.30* (-2.24)	-.24 (-1.67)		
UNZSPILL			.08 (.75)	.05 (.45)
GPOST	-1.63* (-4.96)	-1.53* (-4.37)	-1.42* (-4.01)	-1.37* (-3.64)
R^2	.614	.615	.600	.600
N	118	99	118	99

[a] t-values in parentheses

*Significant at the 5% level.

attaching any meaning to these derivatives in equations in which a spillover variable based on concentration or unionization is included. There is no qualitative change in the results we obtained earlier on the DIS variable and on its interactions with the unemployment and price change variables. Moreover, the equations containing the unionization variable show precisely the same pattern of negative coefficients on the interaction terms and a positive one on the UNZ variable itself.

The ratchet-spillover variables never produce a significant positive additional push on wage increases in firms which are among those with less discretionary power or are in industries which are relatively weakly unionized. Indeed, the sign on the ratchet variable in the equations containing DIS is negative, although its effect is such that the rate of wage increases is decreased only by .3 percent per year in non-key firms. The expected positive effect does appear in the equations containing UNZ, but the coefficient of the spillover variable is not significant, and the addition to wage increases in the non-key firms is only .08 percent per year.

There are a number of problems with this admittedly weak test of the ratchet-spillover hypothesis. All of our data are based on large firms which are unionized, although these firms do differ substantially in the degree of competition they face in their product markets and in the amount of non-union labor which is potentially available to them. There is the further empirical difficulty that U.S. Steel is not included among the five firms in the key group used in the calculation of DISSPILL.

This neglect is the result of the DIS variable's basis on four-firm concentration ratios, which in the steel industry are quite low even though concentration measures based on more firms might produce a substantially different ranking for steel. We can conclude that in our weak test, which is nonetheless a more correct specification than that in previous studies, there is no evidence of the existence of ratchet or spillover effects emanating steadily from one group of firms or industries.

VI. Conclusions

These results demonstrate that wages react significantly less to market variables in those firms which have substantial market power and which are in heavily unionized industries. They thus show that it is incorrect to include industrial concentration alone as a variable explaining changes in wage rates. Instead, its inclusion must be coupled with interaction terms between it and other variables which cause changes in money wage rates. They also indicate that, while industrial concentration has no overall effects on wage inflation, attempts to slow wage inflation by pursuing policies to raise unemployment become less successful as firms' monopoly power increases. All of this suggests that antitrust divesture policy will not affect the overall rate of inflation in an economy but will only tilt the short-run Phillips curve toward the vertical. Finally, our weak tests of the ratchet-spillover hypothesis imply that there is no indirect effect of market power on wage increases.

Our results might be taken as some tentative empirical evidence in favor of the Galbraithian proposition about the behavior of large firms engaging in industrial planning. While other theories about monopoly behavior could be used to derive the relationship we have tested, the planning hypothesis has at least not been refuted in our tests. Our estimates indicate the usefulfess of attempts to test this hypothesis using data describing other areas of a firm's endeavor.

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