PRICE AND WAGE CHANGE AND MARKET CONCENTRATION: UK. MANUFACTURING 1963-8*

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Introduction

One of the more interesting questions for inflation theorists is whether the aggregate rate of change of prices depends on market structure. The debate has a long history. Means (1935) first coined the term 'administered prices' during the Great Depression. Administered prices were defined as prices "set by administrative action and held constant for a period of time", whereas market prices were said to be "made in the market as the result of the interaction of buyers and sellers". Though there are difficulties in indentifying administered prices, since the line between setting a price, through "administered action" as opposed to the interaction of buyers and sellers is unclear, the phenomenon of administered prices has become associated with the firm having some degree of market power. The most useful empirical tests of the hypothesis have involved looking at the relationship between concentration (the measure of market power) and price change, for a sample of industries when allowance has been made for unit cost or demand change. (See eg. Cagan (1975), Dalton (1973), Lustgarten (1974), Philips (1971) and Weiss (1966)).

The empirical tests have been employed for a number of countries (with the notable exception of the UK which has receive d little attention) and are broadly favourable to the hypothesis that a significant characteristic of inflationary movements is that price increases in relation to costs tend to vary according to the degree of industry concentration[1].

Price change is in part a function of unit cost change so as well as the direct influence, a second manner by which market structure can influence the rate of inflation is via unit cost. In particular it has been suggested that monopolistic firms pay higher wages than competitive firms (See eg. Garbarino (1950), Gulbraith (1957) and

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1. While this evidence is consistent with the Means hypothesis it is also consistent with the converse hypothesis that prices are set in the market. Simple oligopoly theory (eg. kinked demand) predicts a variable response of price change to cost change given the trade cycle. (See Cowling and Cubbin (1972)).

moreover that "those workers enjoy the most rapidly rising wages who are employed in partially monopolistic or oligopolistic industries"[2]. (See eg. Schlesinger (1958), Garbarino 1950)). This hypothesis is not consistent with simple neo-classical profit maximising theory where the expectation is that the wage level will be lower under monopoly than perfect competition and further that wage increases will ceteris paribus be invariant to market structure. A number of reasons have been put forward to support the non-neoclassical view. Essentially these are that: (a) monopolistic firms balance the interest of employees, consumers and suppliers rather than maximise profits (See eg. Kaplan etal (1958), (b) the derived demand curve for labour of an oligopolistic or monopolistic industry is less elastic than the demand curve of a competative firm. This proposition derives from the famous Marshall/Hicks rules of derived demand (See Hicks (1963). As a consequence of the demand curve being less elastic it is asserted that unions will push for greater wage increase since the employment effect of such increases being granted is corresponding less than in the competitive case (See eg. Bronfenbrenner and Holzman (1963)). However, it is clear that even if the monopolists demand curve is less elastic than the competitors at a given level of employment, ultimately the demand curve will become elastic at some higher wage. Consequently it would appear that there is a limit to the wage increases which a union can obtain in the monopolistic industry given their wage - employment-trade off. (c) It is asserted that the product demand curve facing oligopolists is inelastic in the range at which prices are set. This is done by the firm as a method of regulating entry or for collusive reasons (Segal (1964)). It can also follow as a consequence of certain cost structures. (See Latham and Peel (1975)). Thus the demand curve for labour will be relatively more inelastic and once again the possibility of above average wages increases is possible. However as with proposition (b) ultimately price increases which follow as a result of wage increases will lead the firm to price in the elastic range of the demand curve and consequently limit the possibility of further wage increases.

Having teviewed briefly the evidence for a link between market structure and inflation the purpose of this paper is the examine the actual behaviour of price and wage for a sample of UK industries. Data limitations mean that we are committed to study the period 1963-68 using data at different levels of aggregation. (See data appendix). Our procedure with respect to the price equation is not original: we estimate the model suggested by Weiss (1966) and Phillips (1969) Dalton (1973). This is that price change is a function of a weighted / unweighted average of unit cost change, demand change and the concentration measure. Nevertheless the results are of some interest since as mentioned above no formal testing has been attempted on UK data.

Our analysis of the wage change equation is we hope more original.

2. An important alternative route which we do not discuss in this paper is via productivity (output per man). There is some evidence that productivity change is dependent on market structure. (See eg. Stigler (1956) Phillips (1956)).

The model of the Wage Equation:

Before the influence of concentration on wage change can be decided it is necessary as with the price equation to allow for alternative influences on wage change.

The neo-classical wage determination hypothesis asserts that the rate of change of money wages is a function of the excess demand for labour and the expected rate of change of prices (see Hines (1971)).

$$\hat{w} = \lambda \frac{(D-S)}{S} + \hat{p}^e \tag{1}$$

where $\hat{w}$ is the rate of change of money wages per time period
D.S. are the demand and supply of labour
$\hat{p}^e$ is the expected rate of inflation
λ is an arbitary constant.

Such an equation, where excess demand is proxied by the unemployment rate, has provided a good explanation of money wage change for a number of different countries at both an aggregate and disaggregative level. (See eg. Gordon (1971), Turnovsky (1972), McCallum (1975) Holden and Peel (1976)).

There are a number of difficulties in estimating equation (1) on a "X" section basis at a disaggregative level. The interpretation of the rate of unemployment at a disaggregative level is unclear. In the UK the unemployed figure refers to the last occupation of the individual. He is assigned to the unemployment total for that industry. Such a person can be part of the effective labour supply to an number of alternative industries and hence determine ceteris paribus the path of money wages in those industries. Consequently at the level of aggregation of this study the possibility of serious mispecification can arise with the use of the unemployment rate as a proxy for excess demand. We suggest on alternative procedure. It is the case that for a given labour supply schedule the elasticity of labour demand is important in determining the path of the rate of change money wages. For example suppose that the demand for labour (D) is a function of the real wage in the following way.

$$D = w^{-a} \tag{2}$$

The elasticity of demand with respect to the real wage is given by a We wish to discover how money wages vary with a

For simplicity write,

$$\hat{w} = zw^{-a} \tag{3}$$

where z = a constant
Taking logs,

$$\log \hat{w} = \log z - a \log w \tag{4}$$

Totaly differentiating with respect to time

$$\frac{1}{w} \frac{d\hat{w}}{dt} = -\{a \hat{w} + \log w \frac{da}{dt}\} \tag{5}$$

Clearly for

$$\frac{da}{dt} > 0,$$

i.e. a more elastic demand curve

$$\frac{1}{w}\frac{d\hat{w}}{dt} < 0$$

In other words the more elastic the demand curve for labour *ceteris paribus* the lower the rate of change of money wages. The rules of derived demand for a perfectly competitive industry were derived by Marshall / Hicks and more as follows:

(a) the demand for anything is likely to be more elastic the more readily substitutes for that thing can be obtained

(b) the demand for anything is likely to be elastic the more elastic is the further thing which it contributes to produce

(c) the demand for anything is likely to be less elastic the less important is the part played by the cost of that thing in the total cost of some other thing, in the production of which it is employed provided that the elasticity of demand exceeds the elasticity of substitution[3].

Although these rules were derived explicitly for a competitive industry they have been applied within the framework of alternative market structures. For instance, as mentioned earlier, Bronfenbrenner and Holzman (1963) when considering the factors which might facilitate cost-push inflation state by implication that the more inelastic the demand for the product, the more inelastic the demand for labour in oligopolistic or monopolistic industries. Such an application of the Hicks rules to situations other than perfect cometition is in general invalid. This is shown in Latham and Peel (1976 (α), 1976 (b)) but can be simply seen from the formula for the elasticity of derived demand for factor (a) given by Yeung (1972) who derives the formula for the case of a monopolist facing a constant price elasticity of demand. For simplicity consider the formula when the elasticity of supply of the other factor is infinite. We obtain

$$a = \left(\sigma\frac{n-1}{n} + k\,(n-\sigma)\right)\left(\frac{n}{n-1}\right) \qquad (6)$$

where

σ is the elasticity of substitution

n is the price elasticity of demand

k is the share of wages

3. There is a fourth rule with respect to the elasticity of supply of the alternative factor. However, in this paper we assume the influence to be constant across the sample.

On differentiating to obtain the response of (a) to a change in the price elasticity of demand we obtain

$$\frac{\partial a}{\partial n} = \frac{k}{(n-1)^2} \{n(n-1) - (n-6)\} \tag{7}$$

Clearly if $6=1$

$$\frac{\partial a}{\partial n} = k > 0 \text{ which is the Hicks result}$$

But if $n(n-1) - (n-6) > 0$ (eg. if $n=1.\cdot 1$, $\sigma=0.2$)

$$\text{then } \frac{\partial a}{\partial a} < 0$$

In other words the response with respect to the price elasticity of demand is ambiguous[4]. The other results are as in the Hicks case. Thus, though the elasticity of demand for a factor will depend upon n, k, σ in all market structures the precise response can differ between structures. Since direct observations on σ and n are not available we make the following heroic assumptions in this paper. That σ is constant across the sample (for example all firms could be typified by a Cobb-Douglass technology, $\sigma=1$) and that the price elasticity of demand can be proxied by the ratio of advertising to sales. There are two rationale for this approach, neither being entirely satisfactory. First, the Dorfman-Steiner (1954) result shows a direct inverse proportionality between advertising / sales and the price elasticity of demand if the advertising elasticity of demand across the industries is not too disimilar. Secondly, it has been argued that a high advertising to sales ratio is indicative of entry barriers which ceteris paribus imply a lower price elasticity of demand. (See eg. Comanor and Wilson (1967). Though see Cowling (1975) for an alternative view).

Given the above discussion our basic specification of the rate of change of money wages is given by

$$\hat{w} = \hat{w}\left(k, \frac{A}{S}, C, U\right) \tag{8}$$

where

k is the share of wages

A/S is the advertising to sales ratio

C is the five firm concentration ratio or to allow for a discrete influence proxied by two dimmy variable

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4. This is contrary to Yeung's that the rules are the same as in Hicks original presentation.

D_1 =0–33
D_2 = 34–66
U is the unemployment rate

The concentration measure is added to our formulation to test for the administered wage hypothesis and the unemployment rate is added to pick up any influence of excess demand not picked up by our proxy measures, though for reasons given above the measure has serious conceptual limitations. Since trade union membership figures as a proportion of the laboul force have recently been made available for a limited number of industries we also experimented with the addition of changes in the proportion of labour force unionized (DT). The rationale for this measure has been provided by Hines (1971). The expected apriori signs on the above variables are

$$\frac{\partial \hat{w}}{\partial k} < 0\,, \frac{\partial \hat{w}}{\partial A/S} \lesseqgtr 0\,, \frac{\partial \hat{w}}{\partial C} \geqslant 0, \frac{\partial \hat{w}}{\partial n} < 0, \frac{\partial \hat{w}}{\partial AT} > 0, \qquad (9)$$

The Price Equation:

Our data for this study is a sample of 42 Minimum List Heading industries (M.L.H.) for the period 1963-68, (see data appendix). We attempted to explain the percentage change in price over the five year period. Our basic specification follows that of Dalton (1973), Weiss (1966, 1971), Philips (1969) and Sellekaerts and Lesage (1973) and is

$$\Delta p = \alpha_0 + a_1 C + a_2 \Delta ULC + a_3 \Delta UMC + a_4 \Delta Q \qquad (10)$$

where Δp is the rate of change in price over the period
ΔULC is the rate of change of unit labour costs
ΔUMC is the rate of change of unit material costs
C is the five firm concentration ratio
ΔQ is rate of change of output (the proxy for demand)
a_0, a_1, a_2, a_3, a_4, are constants. See data appendix for definitions of rates of change.

We experimented by weighting the unit cost terms by their relative shares as recommended by Cagan (1975). However, for this data sample the results were inferior. We also experimented by including two dummy variables to allow for a discrete influence of concentration. These were for medium concentration (33–66) and high concentration (67–100). Finally, we experimented by allowing the influence of unit cost and output change to vary with the degree of market concentration. A sample of our more interesting results are reported in Table I. The general conclusion in which emerges from our results is that for the period 1963–68 though the regression are well determined ($\bar{R}^2 > 0.81$) market structure seems to have exerted a negligible direct influence on price determination. The only evidence of any impact is reported in regression (4) where changes in output show a significant negative influence on price

change in the medium and highly concentrated industries relative to the low concentration ones. If changes in output can be interpreted as a proxy for changes in demand then this result suggests that medium and highly concentrated industries respond to increases in demand by lower prices. This is consistent with unit cost pricing and declining average cost for higher output levels. The difficulty with this interpretation is that output change will reflect net movements in both demand and supply which may operate in different directions. However, bearing this important caveat in mind, it is the most readily available proxy for demand and has been used by other researchers. The influence of the continuous concentration measure is always insignificant. The point estimate is negative. Similar negative insignificant results are reported by Philips for Benelux (1958–68, (1969), Sellekaertes and Lesage for Candada) (1963–66 (1973), Dalton for the U.S.A. 1963–66 (1973) and Weiss for U.S.A. 1959–63, 1963–65 (1966). Analysis of the UK for a different time period may show a significant impact of market concentration on price but our results suggest that for the period 1963–68, there is little, (if any) evidence of a direct influence of concentration on price change for our sample of industries[5].

The Wage Equations

A wider range of data is available for tests of the influence of concentration on wage change for a variety of different periods and levels of aggregation. We are able to examine this relationship for a sample of 45 MLH industries between 1962-3, 1967-8 and 1963-68. Further, since data on trade-union memberships are available at the Standard Industrial Classification level (S.I.C.), we were able to obtain consistent data for twelve industries for 1963 and also for 1968. As twelve observations is regarded as too small a sample the data are pooled for the two years to give 24 observations in all. (See data appendix for fuller sources). The theory developed in the earlier section implies that some of the coefficients can vary between different market structures. As with the price equations we experimented by including slope dummies on the various coefficients and also allowed for the discrete influence of market concentration. We also experimented with a variety of linear and non-linear formulations. There was little to choose between them and consequently we report the linear specification. The unemployment rate variable is entered non-linearly (ie. 1/U) as the apriori theory (see eg. Lipsey (1960) suggests. The wage variable is weekly earnings.

No significant correlations were obtained for the period 1963-8 so we do not report any results. The difficulty could be in the timing of the independent variable which is unclear since the wage change variable refers to a five year period. Ho-

5. Data limitations preclude investigation of our sample of industry for other years. If the administered price hypothesis is thought of as a cyclical phenomena, that is the mechanism by which demand increases are translated into price increases in different market structures with less (see Galbraith (1957)) then investigation of earlier periods would be additionally desirable. In this context the choice of terminal years in our study is dictated by data availability rather than phases of the cycle though the output change variable should hopefully capture some of this influence.

wever, neither simple averages nor the use of the 63 or 68 data for the independent variables revealed any significant correlations. The slope / intercept dummies were always insignificant and consequently we do not report these results.

A sample of our more interesting results are reported in Table II. It should be noted that since we assume that price expectations are constant across industries this variable does not appear in the regressions.

The most significant results were obtained for the S.I.C. sample of industries where we are able to explain over seventy eight percent of the variation in money wages. The share of wages is significantly negative. This result accords with our a priori expectation. Quite interestingly the share of advertising excerts a significant negative influence on wage change. This result is not inconsistent with the theory. Concentration whether allowed for as a continuous or a discrete variable is insignificant. It is also worth noting that changes in the percentages of the labour force unionized the usual proxy for ilitancy was insignificant in all formulations.

The results for the MLH industries are mixed. As with S.I.C. industries the share of wages and advertising enter with a negative sign. However, only the share of wages is significant at normal levels of significance. Unlike the S.I.C. industries case concentration does have a significant impact for 1962/3 but not 1967/8. In fact the 1967/8 sample is somewhat of an enigma. No variables are significant except for the unemployment rate which has the incorrect a priori sign.

Conclusion

The purpose of this paper has been to examine the impact of market power proxied by market concentration on price determination in the UK. Our empirical procedure was to relate the rate of change of prices to market concentration having allowed for unit cost and demand changes. Our results were broadly consistent with the hypothesis that for our sample period there was little evidence of any direct impact of concentration on prices. Our results did show, as was to be expected, that wage change exerts a significant positive impact on price change. Consequently we examined the relationship between wage change and concentration having made allowance for differences in excess demand by proxying determinants of the elasticity of derived demand for labour. Broadly interpreted our results again suggested a negligible influence of concentration on wage change. The results offered more support for the hypothesis that the elasticity of derived demand for labour was important in explaining wage movement. In particular, the significant impact of the share of wages and advertising is of interest. To conclude our results are generally consistent with the neo-classical hypothesis and suggest that in the UK market structure has no lasting impact on price or wage change.

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DATA APPENDIX

The Price Equation MLH data (42) observations

MLH Orders 216, 217, 231, 240, (261/1-263) (272-273), 275, 276, 311/2, 312/1-313/3, 368, 411, 412, 413, 415, 416, 417, 419, 421, (431-432) 441, 442, (462/1-462/2) 443, (461/1-461/2, (463/3) (463/1-463/2) 469/1-169/2, (472-473, 481, 491/2, 492.

The price data obtained from the *Board of Trade Journal* and figures made available by the *Board of Trade.*

Concentration: Census of Production 1963, 1968 Table 44

Sales: Census of Production 1963, 1968 Table 1. Wage and Salary and Material cost – Census of Production 1963, 1968 Table 3.

$$\text{Percentages were calculated as } \frac{X1968 - X1963}{X1963}$$

where X refers to the level of the variable

Output change was calculated as the difference between sales change and price change over the five year period.

Wage Equation MLH data (45) observations

MLH Orders (211 + 219) (212 + 213) 214, (215 + 221 + 229) 216, 217 218, 231, 232 + 239, 240, 261, 262, 263, 272, 273, 275, 276, 297, 311, 312, 313, 321, 322, 323, 361 (363, 367) 368, 369, 370, 382, 383, (391-9) 411-414, 417, 419, 431, 432, 433, 442, 443, 445, 450, (461-464) 491, 472, 473, 481, 491, (492 + 496)

Weekly earnings: British Labour Statistics – Historical Abstract Table 131

Concentration: Census of Production 1963, 1968 Table 44

Share of Wages: Census of Production 1963, 1968 Table 1

Advertising: Census of Production 1963, 1968 Table 4

Percentage Uemployment: Ministry of Labour Gazette Feb 1964, Feb 1968.

Percentages were calculated $\frac{X1963-X1962}{X1962}$ or $\frac{X1968-X1967}{X1967}$

SIC Orders 3, 4, 5, 6, 7, 9, 10, 11, 12, 13, 14, 15.

Weekly earnings: – British Labour Statistics – Historical Abstract Table 131

Concentration Data – Census of Production 1963, 1968, Table 44. Weights used are sales to aggregate MLH to SIC order.

Advertising: Census of Production 1963, 1968, Table 4

Share of Wages: Census of Production 1963, 1968 Table 1

Percentage Unemployement: British Labour Statistics – Historical Abstract Table 131

Trade Union Membership form The Degree of Unionization 1948-1968 by B. Burkitt and D. Bowers (1975) *Bulletin of Economic Research*

Table I The price equation 1963-1968

Dependent Variable		Constant	C_{63}	ΔULC	ΔUMC	ΔQ	CmΔQ	CHΔQ	Cm	CH	R^2
(1)	Δp	0.074 (1.9145)		0.1306 (2.1126)	0.4038 (10.0575)	−0.0458 (1.1952)					0.8145
(2)	Δp	0.0680 (2.0313)	−0.0002 (−0.7799) (C63+C68)	0.1294 (2.0817)	0.4051 (10.0288)	−0.0450 (−1.1688)					0.8125
(3)	Δp	0.0668 (1.9978)	−0.001 (−0.8405)	0.1304 (2.1012)	0.4050 (10.0413)	−0.0446 (−1.1594)					0.8130
(4)	Δp	0.0122 (1.6189)		0.1574 (2.5250)	0.4267 (10.5715)	−0.0095 (−0.2337)	−0.0334 (−2.1467)	−0.0319 (−2.2001)			0.8301
(5)	Δp	0.0792 (2.0585)		0.1559 (2.4224)	0.4174 (10.2241)	−0.0334 (0.8487)			−0.0336 (−1.6164)	−0.0341 (−1.7371)	0.8203
(6)	Δp	0.0684 (4.1254)		0.2017 (5.7891)	0.4238 (10.6045)				−0.0381 (−1.9079)	−0.0359 (−1.8469)	0.8217

t values in parenthesis
R^2 = corrected coeffecient of determination

Table II The wage equation

	Dependent Variable		Constant	C63/68	R	A/S	1/U	Cm	CH	ΔT	R^2
(7)	MLH	1962/3 ΔW	0.0659 (3.3245)	0.0003 (2.2180)	−0.0007 (−2.5122)	−0.0257 (−0.3223)	−0.00003 (−1.1851)				0.3058
(8)	MLH	1967/8	0.1093	−0.0001	−0.0007	−0.0227	−0.0005				
(8)		ΔW	(3.9601)	(−0.6299)	(−0.2036)	(−0.1686)	(−2.5855)				0.0641
(9)	MLH	1962/3	0.0793		−0.0008	−0.3668	−0.00003	0.0083	0.0142		0.2563
(9)		ΔW	(3.4243)		(−2.8831)	(−0.4445)	(−1.0169)	(1.0412)	(1.6075)		
	SIC	1962/3+ 1967/8	0.1335	0.00006	−0.0012	−0.0008	0.0002			−0.0007	0.7878
(10)		ΔW	(4.8723)	(0.0675)	(−2.7273)	(− 2.1738)	(2.3942)			(−0.8863)	
(11)		ΔW	0.1349 (5.4615)		−0.0012 (−2.8263)	−0.0010 (− 2.8283)	0.0002 (2.3400)	−0.0045 (−1.1955)	0.0037 (0.8067)		0.8322

t values in parethesis

R^2 = corrected coefficient of determination

Regressions (10) and (11) were run with a dummy variable taking the value 1 in 1963 0 in 1968. This was significant but is not reported for space reasons.