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This paper is circulated for discussion purposes only and its contents should be considered preliminary.

An Empirical Investigation into Income Distribution, Market Structure and Unionisation in UK Manufacturing 1980-1984.*

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ABSTRACT

This paper develops a Kaleckian income distribution model to assess the impact of various quantities on the share of wages in value added for the UK manufacturing sector. In particular it considers the empirical relationship between wage share, market structure, unionisation, and capacity effects. The sample is taken from three digit production industries over the period 1980-1984 and is analysed using sequential cross section and longitudinal techniques. The principle findings suggest that wage share and product market concentration display a consistent negative relationship, although the relationship is not necessarily linear as posited in earlier studies. Moreover, the findings indicate that union organisation ameliorates this monopolisation tendency and a positive relationship is established between unionisation and wage share. We also find that excess capacity negatively affects wage share in the sample period. The longitudinal results illustrate that industry specific effects are significant in explaining the income determination process. We conclude that omission of capacity or fixed effects in previous studies constitutes a specification bias caused by omitted variables.

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The theory of the distribution of the product of industry between wages and profits which is knocking about in current economic teaching consists of a number of propositions, each of which seems quite unexceptionable in itself, but none of which bears any relation to the rest.
Joan Robinson (1960).

Background and introduction.

Distributional issues occupy a peripheral state in current economic analysis. This reflects an overwhelming interest by neo-classical economists in questions of allocative efficiency, principally in a stylised competitive economy. However, these questions have not been ignored in the post-Keynesian school, which has a tradition stemming from the original foundations of Robinson (1933) and the investigations of Kalecki (1938). More recently, the question of the functional distribution and its determinants has been addressed by Cowling (1982), Henley (1986, 1987) and Reynolds (1984).

The purpose of this paper is to investigate empirically the sequential dynamics of the wage share determination process in UK manufacturing, focusing on the influence of market power in product and factor markets. Utilising the Cowling-Waterson (1976) result that market structure and firm interaction determine the excess of price over costs, a theoretical link is established between incomes accruing to factors and the oligopolistic pricing function. This, it will be recognised, is a development of the Kaleckian (1938) model.

The empirical analysis will focus on three distinct areas. First, the relation between market structure, conduct variables and wage share. The purpose is to ascertain whether the predictions of a Kaleckian distribution model, in particular that wage share is negatively related to the degree of concentration, are borne out in the sample period 1980-1984.[1] Whereas previous studies, in particular Cowling and Molho (1982) and Henley (1987), have considered only a linear concentration ratio when fitting a regression, this paper addresses the possibility that a critical concentration, or an asymptotic relationship, may exist.[2] In general we seek to identify if there exists

[1] This is pertinent since all previous analysis for the U.K. has made use of the Census of Production series SIC(1968). This study works at at three digit level of aggregation conceptually redefined under SIC(1980). As such there exists a ready impetus to assess whether previous known results are still valid. For a discussion of the differences in the organisation of production data, and conceptual redefinitions see Employment Gazette (1983) and CSO (1980)

evidence for the usually posited linear relationship between wage share and concentration at the industry level. Second, by using the longitudinal aspects of the data we can derive some conclusions concerning the sequential dynamics of wage share over the limited sample period. This contrasts to previous studies which have been purely cross-section, or comparative static, in approach. Section (2) outlines the particular way that this is achieved. Finally, we investigate the role of institutional or corporatist factors, particularly any ameliorating effect of organised labour, on the functional distribution in UK manufacturing in the early 1980's. To test this we exploit the model developed by Henley (1986) as applied for the UK.

1. The economics of wage share.

For empirical tractability we choose a distributional model that has its roots in the Cowling-Waterson (1976) oligopolistic pricing function. >From this short-run model distributional results are derived, defining the relation of industrial structure and collusion to macro distribution. This has the methodological advantage that the distribution question is linked to industrial composition and the strategic decisions of that industry, rather than to the conditions of factor input marginal productivity.

The model defines price determination in an industry k, where interdependence among N sellers is recognised and N is fixed. Inputs are purchased at given prices, and homogeneous output sold to price takers at a uniquely determined price P_k. Moreover the marginal cost function C_{ik}' is restricted to a constant up to capacity working, where subscripts allow for intraindustry variation. The firm's profit function is defined over traditional revenue and cost arguments.

$$\pi_{ik} = P_k X_{ik} - C_{ik} X_{ik} \tag{1}$$

A stationary value is obtained by maximising (1) with respect to own output yielding:

$$P_k - C_{ik}'(X_{ik}) = -X_{ik} P_k' \frac{dX_k}{dX_{ik}} \tag{2}$$

Defining the the conjectural variation as a linear composite : $1 + \sum_{i=j}^{N} \frac{dX_{jk}}{dX_{ik}}$. and the degree of

[2] Whether a critical concentration point exists has been discussed by White (1976), Geroski (1981) and Uri and Coate (1987) in the profits concentration literature. We apply this method to wage share in the next section, which addresses the case for supposed non-linearities.

intraindustry collusion as $\alpha_{ik} = \frac{dX_{ik}}{dX_{jk}} \cdot \frac{X_{jk}}{X_{ik}}$, the excess of price over cost by firm i in industry k is derived as:

$$P - C_{ik}'(X_{ik}) = -P_k'X_k\left\{\alpha_{ij} + (1-\alpha_{ij})\frac{X_{ik}}{X_k}\right\} \tag{3}$$

Defining the Lerner index for firms i in the k^{th} industry, L_{ik}, and MS_{ik} as the share of i^{th} output in market output, (3) implies the following Lerner condition:

$$L_{ik} = \eta_k^{-1}[\alpha_{ij}\ (1 - MS_{ik}) + MS_{ik}] \tag{4}$$

Where η_k is industry price elasticity of demand. Constraining the differential that defines collusion to be bounded in the unit interval intraindustry behaviour ranges from Cournot to cooperative activity. Cournot behaviour, $\alpha_{ij} = 0$, that is non-cooperation does not imply the competitive solution. Instead the Lerner condition is defined by firm market share and industry price elasticity of demand ie. $L_{ik} = \frac{MS_{ik}}{\eta_k}$, and if the product market is dominated by one seller the familiar monopoly solution results $L_{ik} = \eta_k^{-1}$.

However, to derive distributional results, we can multiply the function in (4) by market share and sum across all firms i=1...N, rearrange, and define the degree of monopoly in industry k, μ_k as:

$$\frac{\pi_k + F_k}{R_k} = \mu_k = \frac{\alpha_k}{\eta_k} + \frac{(1-\alpha_k)}{\eta_k}.H_k \tag{5}$$

Where H_k defines the Herfindahl index. R_k is aggregate revenue in k, and F_k is overhead costs, comprising salaries, rent and depreciation. Deriving the first moment about the origin of μ_k for all industries defines the average degree of monopoly for a manufacturing economy.

Let the result be multiplied by sales revenue to value added, which defines the distributional result for profitability, π, plus overheads, F, in value added, Y. Since value added is composed of the wage bill, overhead costs and profits, it follows that the wage bill in value added is negatively related to the determinants of μ, ie.

$$\frac{W_k}{Y_k}\frac{R}{Y} - 1 \quad 3 \tag{6}$$

A similar result to this is found in Henley (1987)[4] We note that the oligopolistic pricing decision may be constrained by potential raiders, as with contestable markets. This is ruled out due to capacity barriers, which are either strategic, as in Spence (1977), or unplanned, as with Baran and Sweezy (1966). It is at least possible to define a credible threat to players in the game to sustain collusive behaviour.[5]

Since α_k lies in the unit interval, and if $\eta_k > 0$ and finite, equation (6) predicts a negative relation between wage share and the Herfindahl concentration measure. However the point here is that the relation need not be a priori linear. If we let $\frac{(1-\alpha_k)}{\eta_k} = \beta_k$ be the true value coefficient on concentration, and then estimate a linear form where $\beta_k = \beta^*$, then we induce an error into the analysis. If the true relation β_k is non-linear the error generated is $H_k \cdot (\beta_k - \beta)$ which is non-zero, unless β_k is identical in all industries. In some instances this is analogous to ignoring industry specific effects. Thus it is clear that estimating a linear equation, and repressing information conveyed in the unobservables α_k and η_k creates a bias if α_k and η_k are themselves functional to market structure. Viewing the relation from the side of what will enhance profitability and profit share, we may posit that the conjectural variation term will vary directly with concentration, since co-operation is more likely to be sustained in the presence of relative fewness. Likewise, the price elasticity of demand will vary inversely with concentration, since low values are associated with the possibility of extracting monopoly rents. In general β_k will vary directly with the Herfindahl index, which implies, ceteris paribus, that profit share position is enhanced, and in consequence a wage share position is worsened.

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[3] Function (6) has some noted properties. (i) AS $\eta \rightarrow \infty$, the quantity, $\frac{W}{Y} \rightarrow 0$, that is a price elasticity of demand commensurate with the Smithsonian paradigm produces the perverse result that wage share is zero. The paradigm requires a firms price elasticity with respect to demand be infinite, if Pareto optimality is to ensue. (ii) If $\alpha = 1$ the monopoly solution dominates for all H. This is ruled out theoretically. See either Cowling (1982) or in an international framework Cowling and Sugden (1987).

[4] The limits to functional redistribution are documented in Henley (1986). For a historical perspective see Bronfenbrenner(1971).

[5] Although Dixit (1980) demonstrates that Spence (1977) had investigated an imperfect equilibrium, Burlow et. al. (1985) illustrate instances where the Spence result holds. In general, Kreps et. al. (1982) show credible deterrence holds in a dynamic asymmetric game. The economics of strategic competition and commitment is reviewed in Vickers (1985).

To model the afore argument we draw on the critical concentration ratio hypothesis of Bain (1951), which is nested within this analysis. It specifies a target concentration level, which allows the coefficient β_k to take on particular values within some range of concentration. The wage share concentration relation may be specified such that if $H_k \leq H^*$ where H^* is a latent critical concentration measure, the estimating form is δH_k. This implies that $\beta_k = \delta$ for values of concentration less than H^*. For $H_k > H^*$, the estimating form is γH_k, which implies that $\beta_k = \gamma$ for values greater than the critical concentration measure. In this analysis we posit that a critical concentration might exist if some of the other factors that effect the wage share process produce their effect at some latent value of concentration. If the bargain over value-added is viewed as an asymmetric game, in which concentration is a decision variable under the ultimate control of oligopolists, the familiar monopoly solution may not dominate. Indeed there may exist a concentration point up to which oligopolists make distributional gains. However, beyond this latent point, the gains to concentration are outweighed by increased worker militancy, through channels including a slowdown in productivity growth, strike threats, and the withdrawal of goodwill. The impetus to such worker action being supernormal profit share. This line of argument suggests that an inverse asymptotic relation between wage share and concentration. Hence, a ready methodology may be the imposition of a general asymptotic function, or the use of linear splines, as an approximation to that function. In general we seek to identify whether the concentration-wage share relation is linear or otherwise.

A salient feature of the model is that production worker wage share is contemporaneously defined over structure and conduct variables, whilst not explicitly allowing a role for organised labour. This shortfall is modified, albeit heuristically, by considering the effects of Trades Unions on this structure. Hirsch and Addison (1986) document the orthodox effect that wage share gains depend on the technical conditions of substitutability of production techniques. It is argued that at best Unions will have a redistributive effect vis a vis other workers. The redistributive transformation between capital and labour is homogeneous to degree zero.[6] Within this Kaleckian

[6] Henley (1987) uses neoclassical production technology, and Eulers product exhaustion theorem applied to oligopolistic market structure to illustrate this distributional result. However the use of such auxiliary marginal productivity assumptions are neither necessary nor sufficient for the result to go through.

framework, the effect of Labour on profit and wage share is a direct application to price cost margins. Kalecki (1971) postulates that Union activity enhances wage share when the profit to wage ratio is high. The bargaining leverage is enjoyed to the extent that wage share demands are perceived compatible with high profits at ruling prices.

Initially, wage demands are fed into product price increases as firms try to maintain a target profit position. If this is to degree one, wage share gains are zero. With a Sweezian demand function, or variant, the factor to which prices follow costs is assumed less than unity. There will be distributional gains, on the assumption of a stable productivity regime. That is technology is constant, and resource combinations are held at their previous rates of utilisation. We test this assertion as a general case of institutional arrangements favouring workers.

In addition to the the effects defined in (6) and the institutional labour effects, we consider the effects of macro fluctuations on wage share. To date the model contemporaneously defines wage share, and is made tractable by a stochastic process. However, Kalecki (1971 pp.124-137) provides a business cycle model where profit and investment interaction drive the dynamics of the system. Implicit is that if this changes the functional distribution, by changing profit share, then wage share responds to the dynamics of adjustment. To model this process we draw on two aspects of the monopoly capitalist literature, namely the existence of excess capacity and collusive (and paradoxically rivalrous) behaviour. The hypothesis is that the likely pricing response of oligopolistic firms, faced with a reduction in demand, is to collude in the face of common adversity. Sustained price chiseling is at best transitory. The import being that a rise in excess capacity, enhances the average degree of monopoly, which, ceteris paribus, implies deleterious effects on the residual wage share. Note that the result is sustainable, only to the extent that collusion is stable. Cowling and Sugden (1987) demonstrate the costs and benefits to collusion and acquiescing. In particular we might expect an initial decline in the degree of monopoly if the onset of slump condition are perceived as transitory. In which case a certain amount of destocking occurs. When uncertainty is resolved, and the extended nature of the slump recognised, the gains to acquiescing are strictly outweighed by the benefits of collusion. The general postulate is that in times of extended depression the degree of monopoly is enhanced, and in the absence of slippages there is a consequent wage share squeeze.

2. Empirical Framework

The estimating equations for empirical investigation into wage share-concentration and union sequential dynamics are given for the period 1980-1984. For expositional purposes the format of the equations will be applied to sequential cross section and longitudinal data. The rationale is that application of the model to a cross section series allows comparison with previous studies.[7] It suffers the drawback that macro-fluctuations cannot be controlled for. Also, issues concerning the interaction of market structure, conduct, and wage share in periods of macro fluctuation are best analysed using panel techniques. Moreover, utilising the longitudinal aspect allows industry specific effects to be considered.

To make the function in (6) operational we posit a linear form for the sequential cross section as:

$$WS^* = X_i'\beta + Z_i'\alpha + \varepsilon_i \tag{8}$$

Where WS^* is $n\times 1$ observations on wage share. X is $n\times(k-1)$ matrix of regressors.[8] Z is an $n\times 1$ vector of units and α is a scalar. The term β is a vector of unknown population parameters; Let ε_i be white noise. Moreover, assume the disturbance vector ε_i is independent and identically distributed (i.i.d.) and homoscedastic ensuring pairwise uncorrelation across the cross section sample space. Let ε have a multivariate distribution such that the above is summarised:

$$\varepsilon_i \;\; i.i.d.\,(0, \sigma^2 I_n)$$

Assuming the Gauss-Markov axioms are valid, estimates of the β vector are obtained by least squares, or equivalently maximising the likelihood under normality assumptions. This structure is applied to the sequential cross section. To investigate the time/capacity effects on the dynamics of wage share, panel data techniques are employed. The modelling strategy is a fixed

[7] An important caveat to this analysis, is that the data is constrained to five longitudinal observations beginning in 1980. This implicitly limits all conclusions drawn to be valid contingent on the end points of 1980 and 1984. For the behaviour of wage share over the long period see Henley (1986) and for the aggregate degree of monopoly Cowling (1982).

[8] The components of X are found in the Data Appendix, and their contribution to WS^* discussed at the end of this section.

effect (FE) structure[9], the parsimonious function of which is defined as:

$$WS^*_{it} = X_{kit}'\beta + Z_k'\alpha_i + \varepsilon_{it} \tag{9}$$

Where WS^*_{it} is wage share across i at t; X_{it} is a vector of exogenous variables. α_i is individual industry specific effects; β is a fixed homogeneous population parameter.

This formulation allows comparison with and without industry specific effects by appropriate restrictions on the vector (α). A consistent estimator for the FE model is provided by Least Squares Dummy Variables (LSDV). This method is applied instead of a first difference transform, despite the loss of degrees of freedom. The point is discussed in Hsiao (1985, 1986).[10]

Chowdhury and Nickell (1985) recognise that estimation of (9) generates biased coefficient estimates caused by correlation between individual effects (α_i) and the matrix X_{kit}. To avoid this, axiomatically assume $E(X_{kit}\alpha_i)=0$. The parsimonious model is estimated with and without industry specific effects. Ignoring such parameter heterogeneity in α, such that $\alpha_i = \alpha^*$, where α^* is 1×1, may induce inconsistent estimates. This is contingent on α_t being a relevant, but excluded, explanatory regressor.[11] In this case the problem is that

$$E(\beta_1) = \beta_1 + X'X^{-1}X_2'X_2^{-1}$$

For β_1 to satisfy the Gauss-Markov axioms either the excluded $\beta_2=0$, or $X'X^{-1}X_2=0$, or both. Consequently we examine this restriction. The parsimonious form is augmented by time specific effects, which are interpreted as capacity effects, to further investigate functional distribution predictions. The appropriate proxies for the forcing matrix X_{kit} are defined from the U.K. production series at a three digit level of aggregation. The dependent variable, WS^*, focuses on production worker wage bill in value added. This preserves the Kaleckian (1971) distinction between

[9] The relative merits and demerits of error component versus fixed effect models is surveyed by Mundlak (1978), and briefly by Johnston (1984). Mundlak argues that when using least squares dummy variables (LSDV) in a fixed effect (FE) model, inference is conditional on the particular intercept values inherent in the data at hand. Since no explicit assumption concerning the distribution is assumed, there is less likely to be specification error.

[10] LSDV can be computationally applied since industry specific effects are limited to N=59, across i.

[11] Witness Kmenta (1971), pp. 392-395; in particular the regression in partitioned form for a discussion of the general point.

wage and salary earners and their distributional status. The latter, Kalecki reasons, enter as an overhead element and are codetermined with profit share.

To capture conduct and structure variables, the expedient five firm concentration ratio [CONC] is adopted instead of the Herfindahl measure. To test whether there are non-linearities in the concentration-wage share relation we adopt three simple forms of the general model. First given that the model is specified in levels, we can impose the form $\beta \log CONC$, which imposes an asymptotic relation on the data. This notion is consistent with our prior reasoning that as concentration increases, it becomes increasingly difficult for the oligopolist group, however constituted, to capture further rents at the same rate. The implication, then, is that we expect β to be negative on the variable LOGCONC. Second, the inverse of concentration, INVCONC, is tested as a possible functional form. Again, as concentration increases, wage share will decline, but as ameliorating factors in the wage share determination process begin to act as an effective constraint on oligopolistic pricing, the rate of decline of wage share slows. We expect a positive coefficient. These two methods impose an asymptote on the relation between concentration and wage share in accord with our postulates, but it may be argued that this is necessarily a restrictive position to adopt. To avoid this problem of making a priori demands on the data, we invoke a third method to test for non-linearities in the relation. Namely, we adopt a simple switching regime to find a critical concentration.[12] If the notion that an asymptotic relation is valid then we expect the first segment [SPL1] to be negative, and the second [SPL2] to be negative, with the magnitude of the coefficient on the second linear spline, [SPL2], to be less than the former [SPL1]. This would constitute evidence for an approximation to an asymptotic relationship.[13]

The variable [ADV] is intended as a determinant of the inertia in the price elasticity of

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[12] It is, of course, possible to find a whole host of non-linearities in the concentration range, and use different methodologies to identify where the break occurs. For a summary see Geroski (1981). For the purposes here, we need only identify whether the relationship is linear or otherwise, hence a two way switching regime.

[13] An alternative test of interactions and nonlinearities would be to specify the model as a Cobb function, which is additively separable in logs, and then to impose the inverse of concentration on the resulting parsimonious form. Indeed, if we were to assume the unobservables were constant over time, then we could estimate a log-linear form in differences and this would get rid of the bias. This, in effect, is the Cowling-Waterson (1976) solution to the problem.

demand as discussed in Cowling and Molho (1982). Given that advertising to sales statistics have not been generated since 1968 we are forced to use the expedient variable costs of non-industrial services received. The properties and limitations of this proxy are given in Henley (1986). A Priori we restrict [CONC] and [ADV] negative.

We augment this structure by considering the role of corporatist or institutional factors that may affect the wage share process. This idea is introduced, and utilised, by Henley (1986). Such factors are defined in terms of institutional realities that may impinge, or enhance the probability of workers gaining a higher share of value added. In particular we consider the hypothesised ameliorating tendency of worker organisation on the wage share/concentration relation. Axiomatically, we constrain the coverage measure to be exogenous to the bargaining process to satisfy stochastic process requirements for desirable estimators. The relative merits of this judgement are dealt with in Geroski, Hamlin and Knight (1982). However, if Union power is latent it is reasonable to assume it more readily approaches exogeneity, than an activity measure which is clearly endogenous. The coverage measure [COVER], following Kalecki's (1971) arguments, is assumed to positively affect wage share.

In addition we proxy other corporatist or institutional variables as relevant controls in the wage share process. The number of manual employees to the total in an industry [MANUAL] reflect the degree of potential corporate control that can be exercised by workers. The realisation of such potential over the work process inhibits the growth in concentration and profit share, and positively affects the wage share position. Similarly, average plant size [PLANT] may at first be thought to positively affect wage share. As the number of employees per plant increases we would expect the logic of collective action to become known, and in the absence of decreasing returns to organising labour, a positive coefficient emerges. However, if wage gains are at the expense of productivity, through the collective action of the employers, then this affect may be nullified. The sign is ambiguous.

A variable capturing the degree of capital intensity [CAP] is included to remove the highly concentrated-high capital intensity equivalence problem. It follows that resulting concentration wage share effects represent oligopolistic rent seeking and not technological imperatives. In the sequential cross-section a growth measure [GROW] is added, and interpreted as an estimate of capacity/macro fluctuations experienced in the industry at a given year. In the longitudinal results time dummies [TD81 to TD84] perform this function, and are measured from the base year 1980. We interpret the coefficients in a specific way, as an indication of capacity effects, or macroeconomic activity, on wage share from a given point. In consequence if 1984 is a "bad" time compared to 1980 a negative estimate is expected and so on.[14]

3. Sequential Cross Section Results.

The least squares results for the maintained wage share equation are presented in table 1 for the period 1980-1984. A cursory analysis suggests the results are consistent with our priors, but we have to limit this conclusion especially given that the concentration term enters only in a linear form. We find the linear term, [CONC], is always and everywhere correctly signed but is only significant in the last two years at the 5% level. The variable [ADV] which is proxying the extent to which oligopolists can create demand inertia, and hence influence the magnitude of the price elasticity of demand, is always negative as expected. Furthermore, it is insignificant at the 5% level only in the last year. It follows that our proxies for the structure and conduct elements of the distributional model are only established in the case of the conduct variable.

The augmented controls are of interest. The control variable [GROW] displays erratic behaviour, being significant in two, and negative in three of the years. It suggests that this variable is identifying time/capacity dynamic effects. The two strongly negative years, 1980 and 1981, are consistent with a rise in the degree of monopoly, that occurred over the sample period,[15] placing downward pressure on the position of wage share. Post 1982 may have seen some institutional catching up effects, with growth now positively affecting the wage bill in value

[14] Ideally we would have preferred to include regressors to proxy spare capacity at the industry level. The obvious candidate would have been the CBI spare capacity variable, but this was not available for SIC (1980) until 1983. Macro measures, say from the Wharton School, would play exactly the same role as the time dummies when enter into the reduced form model.

[15] Evidence for the course of the aggregate degree of monopoly over the sample period is given later in the text.

added. It is clear that the process is not adequately modelled in the sequential cross section. We note that capital intensity [CAP] has an insignificant effect on wage share, a result consistent with Cowling and Molho (1982).

The corporatist or institutional variables, [PLANT], [COVER] and [MANUAL] indicate an array of results in accord with our priors. The coverage measure [COVER] is positively correlated everywhere, and significant in four of the years. This gives a fairly robust indication that union presence, in support of the Kaleckian hypothesis, is an important statistical determinant of wage share over the sample period.

This result also holds for the manual worker strength variable [MANUAL]. As the ratio approaches unity we observe highly significant positive effects in all years. The concentration of manual workers, en masse, exerts a potential control over the work process that out strips the gains that union membership affords. The two (distinct) channels to wage share gains are therefore drawn. The average plant measure [PLANT], which was signed in an ambiguous way, turns out to be positive in all years but significant at 5% only in the first two. The results suggest that with average plant size increasing, the ability of labour to capture rents, is enhanced.

Since the structure variable, [CONC], failed to be statistically significant as a linear term, we now address the non-linear possibilities that it may take on. The inverse measure of concentration, [INVCONC], enters the parsimonious model in table 2.[16] We can note immediately that in all equations the measure is correctly signed, a positive coefficient implying an inverse relation between concentration and wage share, and significant in all years. In all cases the Log Likelihood value is enhanced with this alternative measure. This is evidence that for the given sample period, and the model specification, a non-linear concentration term outperforms its linear counterpart. Moreover, it suggests that information contained in the coefficient estimate on the concentration measure, about the price elasticity of demand and the conjectural variation term, varies

[16] For brevity we do not report the results of the log-linear, or the linear spline, version here. The conclusions drawn from augmenting the model by alternative configurations of concentration non linearities are not altered. For the linear spline case a grid search that maximised the likelihood function identified the suitable split. However, we were unable to maintain the assertion that the relationship was an inverse asymptotic one; only that non linearities were preferred to a simple linear version. This was based on a likelihood ratio test with values in the range 4.52 to 12.09 over the sample, where the critical value at 5% is 3.84.

across industries, and its exclusion induces bias. This implies that previous studies, that have not entertained the possibility of interactive non linear effects, have mis-specified the functional form. We note that the results found previously on the other regressors, that attempt to model the process of wage share determination, are not qualitatively affected by the introduction of the new proxy for concentration effects. These documented results are, of course, contingent on model adequacy to which we turn. The general significance of the equations as measured by the R^2 is always greater than that in the equation where a linear measure was used, the values lying in the bound 0.53 to 0.71. We also note that the functional form of the model passed a Ramsey RESET test.[17]

We also note that the model was checked for second order adequacy, ie whether heteroscedasticity was present in the model. A generalised LM test is used to check that the sample second moment about the mean is constant. A Goldfeld-Quandt test required certain data manipulation, and the power of the test is constrained by the optimal choice of the central observations. The Breusch-Pagan (BP) test is performed and fails at a critical value of 14.06 in three cases under the null of a spherical variance-covariance matrix. In consequence all reported t-ratios are transformed and are White (1980) consistent.[18]

4. Longitudinal Results

The results of the parsimonious pooled model are presented in table (3).[19] Initial inspection

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[17] We ascertained whether the comparison of the sample moments accord with the assumed conditional moments. First order adequacy is evaluated with a Ramsey RESET test. The maintained model is augmented with an optimal power series of retrieved fitted values, and the test illustrates whether further augmentation by other preferred regressors is necessary. For a discussion of these ideas see A. Pagan (1984). The test statistic is either an asymptotic Lagrange multiplier, distributed $\chi^2(k)$, or an exact F distribution with critical values here of 2.87 and 4.24 at the 5% and 1% levels. The model passes the test at 5% in years 1981 to 1983, and at 1% in 1981 and 1984.

[18] Alternative estimation procedures were experimented with. In particular the models were estimated by the Messer and White (1984) three step Instrumental Variable technique. The result which produced a Heteroskedastic consistent covariance matrix, which did not alter the qualitative results produced here.

[19] We tested for the adequacy of pooling the results from the sequential cross section. A test of homogeneity of slope and intercepts in the FE model over n industries and t time periods is carried out. The consequences for LSDV are in Johnson (1986) and Hsiao (1986). We employed an analysis of covariance test defined as a variant of Hsiao (1986) which yielded a calculated value 3.12, with a critical value at 1% of 1.69. On strict statistical grounds we reject the null of homogeneity of slope and constants, since it lies just outside the critical region. However, we still pool on the grounds of examining the importance of capacity effects, and realising that this result might indicate the importance of differential industry fixed effects.

of column (1) generates results consistent with the Kaleckian postulates, when industry and time effects are excluded. Considering first the structure and conduct variables [CONC] and [ADV], it is noted that both are correctly signed and significant at the 5% level. The predicted coefficient on the linear concentration term, [CONC], implies that a 10% increase in the concentration measure reduces wage share by 1%, as evaluated from mean values.[20] A general monopolising tendency has deleterious effects on production worker share of value added, supporting our central hypothesis concerning structure and conduct variables.

Turning to the corporatist, or institutional variables, we find that our priors are, in large part correct, and significant. As with the sequential cross section the potential ameliorating effect of organised labour, on the general monopolisation tendency, is borne out in the pooled analysis. The predicted coefficient means that a 10% increase in coverage, results in enhancing wage share by 2.6%. The magnitude implies that in an environment where the socio-legal framework does not benefit organised labour, we can expect a shift from workers share of the rent to oligopolists and those of a largely managerial or salariat status. We also find average plant size significant, as in the sequential cross section, and the story associated with the other corporatist variables remain unaltered. We note the overall significance of the regression. A reported adjusted R^2 of 0.52 is qualitatively similar to the sequential cross section, and has an improved exact F test of 47.44. The reported Breusch-Pagan (BP) statistic of 97.64 indicates a problem of heteroscedasticity, which was removed by using a White consistent estimator.

Columns (2) to (4) augment the model with three versions of the concentration measure to assess whether they perform better than the linear term. Column (2) shows that the inverse concentration, [INVCONC], is correctly signed and significant; as does column (3) with the logarithmic version. These two formulations are superior on the basis of a Likelihood Ratio test.[21] Once again this is suggestive that the precise relationship between concentration and wage share is not necessarily linear.

[20] This result is contrasts to Henley (1987) who finds the magnitude of the coefficient in the range 9% to 12% for US manufacturing three digit cross section in 1972, based on a logarithmic functional form.

Column (4) introduces the idea of the critical concentration measure to approximate the hypothesised inverse relation between wage share and concentration. Using the Quandt (1958) procedure the critical split was found that maximised the log likelihood function. The first linear spline [SPL1] is negative and highly significant, implying that up to the critical point concentration negatively affects wage share. Beyond this point any increase in concentration leads to higher wage share. This paradoxical result may come about if the bargain over value added favours the workers in highly concentrated industries. It is clear that the relation is not necessarily a simple linear one, and that the inter and intra-industrial effects complicate the precise nature of the relation. To this extent controlling for fixed effects may be necessary. Whilst it may not be very illuminating to draw conclusions on the exact nature of the "true" functional relation from this section, it can be seen that a simple linear form is not borne out by the data.

Table (4) augments table (3) by including time specific dummies. Column (1) indicates that concentration is still negatively related to wage share, and its estimate does not vary from that presented in table (3). This is true for all the other regressors in the model, whose contribution has not qualitatively altered. The model suffers heteroscedasticity problems, which were removed with a White estimation technique. We note the general increase in the Likelihood value to 364, which on the basis of a Likelihood Ratio test accepted, the augmented model with time dummies.

Briefly considering columns (2) to (4) we note that the non linearities tell the same story as before, each being correctly signed and significant. The inclusion of time specific effects have not altered our conclusions. Addressing column (1) again we find that the time and capacity effects negatively affect wage share over the sample period.[22] Moreover, if we consider the magnitude of the coefficients we see that they are monotonically increasing their (negative) effect between 1980 and 1984. This conclusion is valid both for the magnitude of the coefficient, and for the intensity of the effect as captured by the asymptotic t ratios. Thus if these time dummies are capturing changes in macro-economic activity, and the sample period is associated with

[21] The test statistic is defined as $-2log(\hat{\theta}_0 - \hat{\theta}_1)$, where $\hat{\theta}_1$ is an estimator under the null. See Engle (1983) for a discussion of the asymptotic theory.

[22] Conventional interpretation is of all effects occurring in t unexplained by the conditional model. Our analysis is the not unreasonable assumption that these are primarily macroeconomic in origin.

slump conditions, this is evidence for the proposition that there is a wage share squeeze in extended depression. However, the negativity of the coefficients alone does not constitute direct evidence, since we cannot distinguish between capacity and other effects. If, there is a tendency for a wage share squeeze in the period, we would also expect a rise in the aggregate degree of monopoly.

Table (5) matches events that occurred in the sample period with the time dummy coefficient estimates, obtained from table (4) column 1. It illustrates that over the sample period, instead of an erosion of a monopolisation tendency, as measured in the aggregate degree of monopoly, there has been an increase. This increase, which came about after an initial decline associated with uncertainty about the extended nature of the slump, is concurrent with sustained excess capacity as captured in the unemployment variable. The persistence of labour excess capacity, as a proxy for general excess capacity, supports the proposition that the degree of collusion, captured by μ , is enhanced in a slump. It is seen that as the labour excess capacity measure, UNEMP, becomes persistent, the negative effect of the coefficient estimates on wage share becomes monotonically greater in magnitude and significance. If we use an alternative proxy for excess capacity, namely electricity consumption in the UK manufacturing sector, CU, we find the proposition still holds. The capital utilisation measure falls post 1979, and marginally fluctuates, but does not reach its pre 1980 utilisation rate. Again, the existence of excess capacity, coupled with the time dummy estimates and the degree of monopoly, lends support to the notion that wage share is squeezed under slump conditions.

5. Fixed Effects Results.

The results of the fixed effect model are presented in table (6). The inclusion of fixed effects markedly adds to the value of the likelihood, rising to 604.6, from the previous value without fixed effects of 364. The inclusion of industry specifics is accepted on the strength of a highly significant Likelihood Ratio test. The individual fixed effects are not reported. Column 1 gives the analogous model to table (4) column 1 but augmenting by industry specific factors. There are some noteworthy features of the model. First, the variable proxying the degree of control over the price elasticity of demand that an oligopolist can exert, [ADV], is now insignificant but correctly signed. The structure variable, [CONC], retains its significant negative effect. This result implies that even when factors unique to an industry are taken into consideration, market structure effects still exert a deleterious influence on the wage share position. The corporatist variables, in particular COVER, display the qualitatively same story as. If we consider the time dummies, these too, are qualitatively similar and display a negative monotonic effect to wage share. >From this we conclude the result that concentration is negatively related to wage share, and that institutional factors positively affect wage share in the sample period. Industry specific effects are important explanatory regressors in the wage share determination process,[23] and only detract from the significance of the conduct variable. In consequence, exclusion of these effects represents omitted regressor bias, in previous studies. We note, however, that column (2) to (4) illustrate that whilst this negative relation between concentration and wage share is established, it is not necessarily linear. The concentration term in columns (2) and (3) perform less well, on the basis of their Likelihood values, than does the linear term CONC, although the signs and general significance of each coefficient estimate, are established. However, the linear spline does perform marginally better in column (4), on the basis of a Likelihood Ratio test, and we conclude that there is tentative evidence for non-linearities in the wage share relation when industry factors are taken into account. Of interest are the signs of the estimates. The first spline [SPL1] is negative, but insignificant, suggesting that when industry effects are controlled for, the concentration effect up to that latent point is not important in explaining the wage share determination process. However, beyond the critical point, [SPL2] indicates that the deleterious effects on wage share are coming through the highly concentrated industries. The estimated relationship is not an inverse one, given the slope estimate of the second spline, hence we cannot claim support for an inverse relationship when fixed effects are controlled for. The coefficient estimate implies that a 10% increase in concentration leads to a 0.5% reduction in the share of production worker value added. The precise nature, however, of this relation needs further investigation, since we are only

[23] Under the null of irrelevance of inclusion these would be simultaneously zero. This is rejected by the data.

identifying in the aggregate estimation, a complex story concerning particular industry concentration-wage share effects.

Conclusions

Theoretical and empirical distributional considerations remain an enigma. The focus of this paper provided support for various Kaleckian postulates. First wage share consistently bears a negative relation to concentration, although it is not necessarily linear. We have found limited support for an inverse relation, especially in the sequential cross section results, but this conclusion needs to be tempered. When the model is pooled we only find consistent support for non-linearities, and the precise nature of the relationship, although always negative, remains unresolved. This negative relation holds despite the complexities with fixed effects. Second, union, and other institutional arrangements, ameliorate the monopolisation process. This result holds in the pooled results, inclusive of fixed effects. We also found in the longitudinal results that industry specific factors were important in the wage share determination process. The result implies that previous studies are prone to a certain specification bias, caused by omitted variables. The conduct variable [ADV] performed less well when fixed effects are taken into account. Finally, dynamic time and capacity effects proved important in explaining wage share over the sample period. The critical assumption here is that we are considering a restricted sample, and we cannot generalise this result to the behaviour of wage share over the longer period. However, it is clear that capacity effects along with industry specific factors and product market structure are important in the income determination process in the early 1980's. Moreover, it is established that there is a tendency for unionisation to offset a monopolisation process, and enhance wage share, in accord with a Kaleckian distribution theory.

TABLE 1

Sequential Cross Section Wage Share	1980	1981	1982	1983	1984
Constant	0.375	0.403	0.068	0.178	-0.074
	(3.02)	(3.25)	(0.982)	(1.51)	(0.673)
CONC/100	-0.089	-0.061	-0.090	-0.134	-0.109
	(1.79)	(1.42)	(1.72)	(2.60)	(2.33)
ADV	-0.895	-1.04	-1.10	-1.21	-0.670
	(2.08)	(4.20)	(2.77)	(2.07)	(1.48)
COVER	0.118	0.178	0.181	0.167	0.118
	(1.70)	(2.77)	(2.58)	(3.24)	(2.63)
GROW	-0.399	-0.402	0.005	-0.209	0.019
	(4.71)	(5.20)	(1.97)	(1.68)	(0.23)
CAP	-0.049	-0.213	-0.287	-0.004	-0.523
	(0.297)	(1.19)	(1.01)	(0.390)	(2.79)
PLANT	0.119	0.128	0.068	0.043	0.112
	(2.80)	(2.83)	(0.985)	(0.56)	(1.82)
MANUAL	0.641	0.513	0.647	0.599	0.670
	(6.01)	(6.71)	(9.32)	(8.64)	(9.31)
$\bar{R}^2$	0.71	0.62	0.53	0.58	0.64
F	21.34	14.90	10.16	12.51	16.02
Log L	80.41	86.45	75.07	83.54	88.85
BP	17.01	20.13	26.72	14.50	7.36
N	59	59	59	59	59
μ	0.424	0.405	0.401	0.374	0.375

Notes.

1. Dependent variable wage share defined as production wage bill to value added in each time period.

2. Absolute values of t-ratios based on White heteroscedastic standard errors are reported in parentheses.

3. BP is a Breusch-Pagan (1983) statistic for heteroscedasticity which is distributed as $\chi^2(k)$ where k is the number of regressors in each equation.

4. μ is defined as the mean of the dependent variable.

5. Log L is the log of the Likelihood function in each of the the successive cross sections.

6. The estimates for this table are based on the maintained model with the concentration term entering in a linear level form.

TABLE 2

Sequential Cross Section Wage Share					
	1980	1981	1982	1983	1984
Constant	0.248	0.313	-0.141	0.063	-0.152
	(1.94)	(2.58)	(-1.92)	(0.540)	(-1.36)
INVCONC	1.37	1.06	1.02	1.36	1.14
	(3.43)	(3.59)	(4.13)	(4.58)	(4.99)
ADV	-0.886	-1.02	-1.102	-1.159	-0.648
	(2.11)	(4.56)	(2.78)	(2.11)	(1.52)
COVER	0.141	0.204	0.190	0.186	0.135
	(1.97)	(3.09)	(2.45)	(3.81)	(3.10)
GROW	-0.389	-0.402	0.006	-0.194	0.018
	(4.78)	(5.60)	(2.11)	(1.65)	(0.21)
CAP	-0.005	-0.138	-0.242	-0.004	-0.514
	(0.031)	(0.762)	(0.831)	(0.380)	(2.84)
PLANT	0.103	0.115	0.034	0.009	0.070
	(3.14)	(3.33)	(0.59)	(-0.151)	(1.36)
MANUAL	0.663	0.518	0.641	0.581	0.652
	(6.33)	(6.72)	(9.40)	(9.14)	(9.58)
$\bar{R}^2$	0.72	0.65	0.54	0.61	0.66
F	23.18	16.37	10.67	14.48	17.77
Log L	82.26	88.3	75.92	86.33	90.98
BP	19.86	17.96	26.36	14.42	7.20
N	59	59	59	59	59
μ	0.424	0.405	0.401	0.374	0.375

Notes.

1. Dependent variable Wage Share defined as production wage bill to value added in each time period.

2. Absolute values of t-ratios based on White heteroscedastic standard errors are reported in parentheses.

3. BP is a Breusch-Pagan (1983) statistic for heteroscedasticity which is distributed as $\chi^2(k)$ where k is the number of regressors in each equation.

4. μ is defined as the mean of the dependent variable.

5. Log L is the log of the likelihood function in each of the cross section periods.

6. The estimates given in table 2 are based on the maintained model. The concentration term enters in an inverse form.

TABLE 3

Longitudinal Wage Share				
	WS	WS	WS	WS
Constant	-0.126	-0.210	-0.026	-0.087
	(3.39)	(6.02)	(0.536)	(2.58)
CONC/100	-0.088			
	(2.99)			
INVCONC		1.156		
		(6.93)		
LOGCONC/10			-0.423	
			(5.04)	
SPL1/100				-0.226
				(7.14)
SPL2/100				0.227
				(2.63)
ADV	-0.677	-0.667	-0.661	-0.626
	(2.10)	(2.10)	(2.03)	(2.32)
COVER	0.149	0.168	0.167	0.145
	(4.81)	(5.10)	(5.23)	(4.76)
GROW/100	-0.485	-0.464	-0.387	-0.613
	(0.342)	(0.332)	(0.274)	(0.460)
CAP/10	-0.217	-0.132	-0.157	-0.062
	(0.879)	(0.591)	(0.710)	(0.330)
PLANT/10	0.832	0.614	0.823	-0.197
	(2.48)	(2.32)	(3.00)	(0.405)
MANUAL	0.691	0.686	0.691	0.695
	(18.69)	(19.14)	(19.19)	(19.30)
$\bar{R}^2$	0.52	0.54	0.54	0.57
F	47.44	51.48	50.39	50.39
Log L	357.517	364.106	362.363	372.496
BP	97.64	79.51	96.57	85.07
N	295	295	295	295

Notes.

1. Absolute values of t-ratios based on White heteroscedastic standard errors are reported in parentheses.

2. BP is a Breusch-Pagan (1983) statistic for heteroscedasticity which is distributed as $\chi^2(k)$ where k is the number of regressors in each equation.

TABLE 4

Longitudinal Wage Share with Time Effects				
	WS	WS	WS	WS
Constant	-0.112	-0.197	-0.009	-0.068
	(3.18)	(5.90)	(0.212)	(2.08)
CONC/100	-0.088			
	(3.12)			
INVCONC		1.18		
		(7.68)		
LOGCONC/10			-0.429	
			(5.37)	
SPL1/100				-0.233
				(7.58)
SPL2/100				0.243
				(2.87)
ADV	-0.630	-0.614	-0.610	-0.567
	(2.09)	(2.08)	(2.02)	(2.33)
COVER	0.156	0.177	0.175	0.153
	(5.15)	(5.46)	(5.60)	(5.22)
GROW/100	-0.428	-0.379	-0.314	-0.537
	(0.329)	(0.298)	(0.243)	(0.449)
CAP/10	-0.143	-0.054	-0.081	-0.021
	(0.570)	(0.239)	(0.364)	(0.113)
PLANT/10	0.738	0.525	0.735	-0.354
	(2.28)	(2.17)	(2.87)	(0.752)
MANUAL	0.686	0.681	0.686	0.690
	(18.87)	(19.29)	(19.35)	(19.38)
TD81/10	-0.076	-0.086	-0.080	-0.106
	(0.527)	(0.615)	(0.563)	(0.766)
TD82/10	-0.091	-0.112	-0.102	-0.133
	(0.631)	(0.796)	(0.720)	(0.964)
TD83/10	-0.352	-0.370	-0.360	-0.397
	(2.43)	(2.63)	(2.53)	(2.86)
TD84/10	-0.363	-0.391	-0.377	-0.418
	(2.58)	(2.87)	(2.74)	(3.11)
$\bar{R}^2$	0.54	0.56	0.55	0.58
F	32.26	35.32	34.42	36.05
BP	110.97	94.97	111.87	96.49
Log L	364.000	371.580	369.389	381.277
N	295	295	295	295

Notes.

1. Absolute values of t-ratios based on White heteroscedastic standard errors are reported in parentheses.

2. BP is a Breusch-Pagan (1983) statistic for heteroscedasticity which is distributed as $\chi^2(k)$ where k is the number of regressors in each equation.

TABLE 5

Monopoly, Capacity and Time Effects

YEAR	UNEMP	CU	MONOP	TD/10
1979	5.1	110.01	0.622	
1980	6.6	100.63	0.595	
1981	9.9	98.19	0.601	-0.076
1982	11.4	99.73	0.616	-0.091
1983	12.6	99.84	0.635	-0.352*
1984	13.0	101.64	0.631	-0.363*

(i) n^* indicates that the time dummy coefficient is significant at the critical 5% level.

(ii) MONOP: Defines the degree of monopoly over the sample period. Disaggregated figures are available by industry on request. The degree of monopoly is defined as the ratio of value added minus the costs of operatives to value added, and is derived from the production series. The measure goes some way to alleviating the problem of possible changes in the degree of vertical intergration. (see Cowling 1983). Post 1979 measures are provided from the production series under SIC(1980).

(iii) UNEMP - Unemployment(%), constructed as a capacity measure of the degree of potential capacity in the national economy. The series is constructed from the National Institute Economic Review.

(iv) CU : Capacity utilisation measure based on taking mean values of the Central Electricity Generating Board quarterly electricity consumption data for the manufacturing sector of England and Wales. The merits of the measure are delt with in Heathfield and Allison (1987)

(v) TD/10 - Time Dummy coefficient estimate from longitudinal work. Intended as a capacity effect on wage share measured relative to the 1980 base.

TABLE 6

Longitudinal Wage Share with Time and Fixed Effects				
	WS	WS	WS	WS
Constant	-0.242	-0.179	-0.048	-0.218
	(1.31)	(1.09)	(0.357)	(1.26)
CONC/100	-0.269			
	(2.48)			
INVCONC		0.693		
		(2.05)		
LOGCONC/10			-0.549	
			(2.16)	
SPL1/100				-0.049
				(0.655)
SPL2/100				-0.482
				(2.62)
ADV/10	-0.371	-0.224	-0.297	-0.242
	(0.457)	(0.240)	(0.334)	(0.294)
COVER	0.885	0.601	0.719	0.831
	(2.85)	(2.65)	(2.71)	(2.92)
GROW/100	-0.169	-0.130	-0.148	-0.174
	(0.301)	(0.218)	(0.249)	(0.332)
CAP/10	-0.203	-0.194	-0.199	-0.192
	(1.51)	(1.44)	(1.46)	(1.47)
PLANT/10	0.874	0.449	0.558	0.121
	(0.716)	(0.315)	(0.406)	(1.07)
MANUAL	0.355	0.327	0.338	0.316
	(2.46)	(2.42)	(2.46)	(2.26)
TD81/10	-0.149	-0.165	-0.162	-0.138
	(2.13)	(2.17)	(2.15)	(2.13)
TD82/10	-0.181	-0.188	-0.190	-0.156
	(2.38)	(2.34)	(2.35)	(2.28)
TD83/10	-0.467	-0.460	-0.465	-0.449
	(5.15)	(4.99)	(4.99)	(5.43)
TD84/10	-0.495	-0.468	-0.483	-0.463
	(3.81)	(3.70)	(3.69)	(3.94)
$\bar{R}^2$	0.88	0.88	0.88	0.88
F	34.30	32.71	33.10	34.80
Log L	604.624	598.255	599.843	609.102
N	295	295	295	295

Notes

1. Absolute values of t-ratios based on White heteroscedastic standard errors are reported in parentheses.

2. BP is a Breusch-Pagan (1983) statistic for heteroscedasticity which is distributed as $\chi^2(k)$ where k is the number of regressors in each equation.

3. Coefficient estimates for industry specific effects are not reported within the table.

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DATA APPENDIX.

Descriptions and Means of Explanatory Variables.		
Variable	**Description**	**Mean**
WS	Production worker wage bill in value added, defined from census of production, various issues, deflated by relevant producer price index.	0.396
CONC	Five firm concentration ratio by sales, defined from the census of production.	40.38
ADV	The ratio of the costs of non-industrial services received to sales, defined from the census of production series.	0.005
PLANT	The ratio of total employment in an industry to the number of establishments.	0.116
COVER	Proportion of male manual employees covered by union agreement. Source: New Earnings Survey (NES 1985) table 190 coverage data by industry for male manuals.	0.685
MANUAL	The ratio of total operative employment to total employment in an industry.	0.703
GROW	Defined as the ratio of sales in year t minus sales in t-1 to sales in t-1, deflated by industry prices.	0.008
CAP	Capital intensity control, defined as the ratio of net capital expenditure to value added, deflated by industry prices.	0.103
INVCONC	Defined as the inverse of CONC.	
LOGCONC	Defined as the natural log of CONC.	
TD81 to TD84	Time dummies taking a value one for each respective year.	