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Investment and Tobin's Q

Evidence from company panel data*

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A Q model of investment is estimated using data for an unbalanced panel of UK companies over the period 1975–86. Correlated firm-specific effects and the endogeneity of Q are allowed for using a Generalised Method of Moments estimator. In the calculation of Q we estimate the tax incentives available to individual companies. Q is found to be a significant factor in the explanation of company investment, although its effect is small and a careful treatment of the dynamic structure of Q models appears critical. In addition to Q, both cash flow and output variables are found to play an independent and significant role.

1. Introduction

In this paper we assess the extent to which Q models of investment provide an empirically fruitful framework for the analysis of individual firms'

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decisions to expand their capital stock. These models possess several appealing features which explain why they have become a popular theoretical and empirical setting for the analysis of investment behaviour. They can be derived from an explicit optimisation framework and rationalise why expectations about the future play a crucial role. The theoretical model is quite straightforward and relates the company investment rate to the ratio of the shadow value of capital and the unit price of investment goods. The shadow value is itself a forward-looking function of future expectations and is, as a result, unobservable. The empirical attraction of the Q model stems from a simple relationship between this ratio of shadow value to price, known as marginal Q, and the observable ratio of market valuation to replacement cost value of capital, known as average Q. To obtain this relationship, the additional assumptions of perfect competition, perfect capital markets, and linear homogeneity of the (gross) production and adjustment cost functions, are required. Under these assumptions unobservable expectations about the flow of profits generated by new investment are summarised conveniently by the observable average market value of capital [Hayashi (1982)]. However, despite these strong assumptions, average Q may still be felt to contain information concerning expectations relevant to investment decisions that is not otherwise available in econometric models that rely on output and user cost variables alone.

Q models have not been noticeably successful in accounting for the time series variation in aggregate investment [see, for example, von Furstenberg (1977), Summers (1981), Poterba and Summers (1983), and Poret and Torres (1989)]. Their explanatory power is low and serial correlation or dynamic structures including the lagged dependent variable are common. In addition, other variables reflecting liquidity constraints or the state of demand are often significant in the equations even though the standard formulation of Q models does not provide a satisfactory rationale for their inclusion. Previous results obtained using panel data are also mixed [see Hayashi and Inoue (1991), Fazzari, Hubbard, and Petersen (1988), Salinger and Summers (1983), and Chappell and Cheng (1982)]. Salinger and Summers find that the coefficient on Q in time series regressions for individual firms takes the expected sign in almost all cases but is statistically insignificant nearly half of the time. Hayashi and Inoue, in their study of a panel of Japanese firms, find a tax-adjusted Q variable to be a significant determinant of investment. However, a cash flow variable is also significant in some years when added to the model. Similarly, for a panel of US companies, Fazzari et al. find that cash flow has an important effect on investment in addition to Q variables.

In this paper we begin by testing a standard Q model of investment under different assumptions concerning the stochastic properties of Q at the firm level. The theory underlying the Q model implies reasonably strong restrictions on the stochastic properties of the model suggesting that careful treatment of the dynamic structure of the model and the choice of instruments used may be critical to the recovery of consistent estimates of the parameters. With the notable exception of Hayashi and Inoue (1991), previous microeconometric studies have adopted estimators which do not fully acknowledge these restrictions and as a result may not be consistent. These worries are confirmed in our empirical application on a panel of UK companies where the sensitivity to alternative stochastic assumptions is highlighted and in which we detect a more complex dynamic structure than is usually entertained. These important dynamic effects turn out to be consistent with average Q theory only under a strong autoregressive restriction on the error process which, perhaps surprisingly, does appear to be data-coherent. In the stochastic specification we allow for company-specific effects in the error term and for the endogeneity of regressors. This is achieved using a heteroskedasticity robust Generalised Method of Moments (GMM) estimator due to Hansen (1982) and White (1982). The GMM estimates are compared with those obtained by a number of other estimation methods the consistency of which relies on stronger assumptions about the nature of the error term and the exogeneity of Q. This comparison allows us to provide some assessment of the most appropriate estimation strategy for panel-data-based Q models.

In principle there are distinct advantages in exploiting data on individual firms. In the first place it allows the theory, developed in the context of a 'representative' firm, to be tested at the level at which it is formulated, so reducing econometric problems introduced by aggregation across firms. Aggregation problems may be important here since the standard Q model is specified in ratios and there are clear nonlinearities in the corporate tax system. Secondly, the estimates are obtained by using both the time-series and cross-sectional variation in the data. This should contribute to their precision and also allows consistent estimation in the presence of correlated company-specific effects. Finally, some variables can be measured more accurately at the individual firm level. This is certainly true for the market value of the firm and also for the effective factor prices it faces. In particular the widespread occurrence of tax exhaustion in the UK since the mid-1970s alters the effective price of investment goods, inclusive of tax incentives, according to the tax position in which the firm finds itself. In the calculation of Q we account for the tax incentives available to UK companies and we discuss how the value of these incentives changes when the firm has zero taxable profits, which is the case for a substantial fraction of UK firms during the estimation period.

In section 2 we briefly outline the main features of the theoretical model that provides the starting point for our empirical analysis. Although the theoretical issues are not new, we utilise this discussion to indicate the appropriate way to construct the firm-specific variables used in the empirical model and to assess precisely what stochastic restrictions are placed on the model by the optimising theory. Section 3 contains the empirical results and begins with a description of the data. Following from this some econometric issues are discussed and the results of our specification search for an appropriate Q model are presented. Having chosen what appears to provide the best description of the relationship between Q and the investment rate for our sample of companies, we then assess the degree to which other factors, in particular cash flow and output, contain additional explanatory power. Finally, in section 4 we conclude with a summary of our main findings.

2. Adjustment costs and investment: The Q model

The basic Q investment equation is derived from a standard model of a perfectly competitive firm that maximises the net wealth of existing shareholders when facing convex adjustment costs in changing its capital stock [see Summers (1981) and Hayashi (1982)]. The adjustment cost function, measured in output lost, is represented by $G(I, K, e)$, where I is investment, K is capital stock, and e summarises all unobservable stochastic factors that may influence a firm's adjustment costs. The gross production function, $F(K, X)$, is a function of capital and a vector of other factors, X, whose adjustment is assumed to be costless. Net output therefore equals $F - G$.

In specifying the optimisation problem for the firm we start from the usual capital market arbitrage condition:

$$\rho_t V_t = (1 - m_t)\theta_t D_t + (1 - z_t)(\mathrm{E}_t[V_{t+1}] - V_t - N_t), \tag{1}$$

where V_t is the market value of the firm's outstanding shares at the beginning of period t and $\mathrm{E}_t[V_{t+1}]$ is the conditional expectation of the market value at the beginning of period $t+1$, based on beginning of period t information. The parameter ρ_t is the required nominal rate of return on equity, m_t is the personal tax rate on dividends, z_t is the tax rate on capital gains, θ_t is the dividend received by the shareholder when the firm distributes one pound of retained earnings,[1] D_t is dividends paid, and N_t is new equity issued in period t. All payments are assumed to be made at the end of the period, but known at the beginning of the period.

Condition (1) states that the return on equity given by the dividend yield and capital gain must equal the market return on comparable assets. The firm's objective is to maximise the wealth of existing shareholders. Solving (1)

[1]Under a classical system of corporation tax θ_t takes the constant value of unity. Under an imputation system, as in the UK, $\theta_t = (1 - c_t)^{-1}$ where c_t is the rate of imputation.

forward for V_t yields

$$V_t = \mathrm{E}_t \sum_{j=0}^{\infty} \beta_j (\gamma_{t+j} D_{t+j} - N_{t+j}), \tag{2}$$

where $\gamma_t = (1 - m_t)\theta_t/(1 - z_t)$ is the tax discrimination parameter that determines the relative tax advantage of dividends against retained earnings. The parameter β_j is the discount factor defined as

$$\beta_j = \prod_{i=0}^{j} (1 + r_{t+i})^{-1}, \qquad j = 0, 1, 2 \ldots, \tag{3}$$

where

$$r_t = \rho_t/(1 - z_t).$$

Dividends can be substituted out of (2) using the following definition of sources and uses of funds:

$$D_t + i_t(1 - \tau_t) B_t = R_t + B_{t+1} - B_t + N_t, \tag{4}$$

where i_t is the nominal rate of interest, B_t is the stock of (one-period) debt at the beginning of period t, and τ_t is the corporate tax rate. R_t is the firm's after-tax net revenue received at the end of the period, defined as

$$R_t = (1 - \tau_t)[p_t F(K_t, X_t) - p_t G(I_t, K_t, e_t) - w_t X_t] \\ - p_t^I (1 - u_t) I_t + \tilde{A}_t, \tag{5}$$

where p_t is the price of the firm's output, w_t is the nominal input price vector associated with X_t, p_t^I is the price of investment goods, u_t denotes first-year allowances, and $\tilde{A}_t$ the value of writing down allowances on past investments that can be claimed in period t.

Several papers analyse financial policy in the context of this model. For example, Poterba and Summers (1983) discuss the choice between retention and new equity finance by introducing nonnegativity constraints on dividend payments and new equity issues. Hayashi (1985) and Chirinko (1987) have extended the model to include optimal debt policy. We do not investigate these issues in this paper. Instead we assume that $\gamma = 1$ and that debt is given exogenously.[2]

[2] Note that the investment equation derived under this assumption is also the one that results from the assumptions that $\gamma < 1$, new equity is the marginal source of finance, and a constant proportion of new investment is financed by debt [Poterba and Summers (1983)]. The assumption that $\gamma = 1$ may be justified on the grounds of the possibility of tax exhaustion [Keen and Schiantarelli (1988)].

Denoting by λ_t^K the shadow price associated with the capital accumulation equation $K_t = (1-\delta)K_{t-1} + I_t$, where δ is the rate of depreciation, the first-order condition for investment is

$$(1-\tau_t)p_t\left(\frac{\partial G_t}{\partial I_t}\right) + (1-n_t)p_t^I = \lambda_t^K, \tag{6}$$

which sets the marginal cost associated with an additional unit of investment equal to its shadow price, where n_t is the expected present value of current and future investment allowances on a pound of investment expenditure in period t. In the empirical application we allow for the possibility that the firm may be tax-exhausted in current or future periods. Since in the UK losses may be carried forward indefinitely at nominal value, tax exhaustion leads to a postponement of tax effects, which we capture by discounting over the period of tax exhaustion.[3] The first-order conditions for optimal variable inputs are standard marginal productivity conditions.

The first-order condition for capital defines the Euler equation describing the evolution of λ_t^K, the shadow value of capital, according to

$$(1-\tau_t)p_t\left(\frac{\partial F_t}{\partial K_t} - \frac{\partial G_t}{\partial K_t}\right) + \mathrm{E}_t\left(\frac{\lambda_{t+1}^K(1-\delta)}{(1+r_{t+1})}\right) = \lambda_t^K. \tag{7}$$

This is the condition found in dynamic rational expectations models with capital or asset accumulation [see, for example, Hansen and Singleton (1982)]. It shows that λ_t^K is the expected present value of current and future marginal products of capital net of adjustment costs, and will in general depend on current adjustment cost shocks e_t. The Euler equation cannot be estimated directly since λ_t^K is not observed.

A rearrangement of (6) yields the equation

$$\frac{\partial G_t}{\partial I_t} = \left(\frac{\lambda_t^K}{(1-n_t)p_t^I} - 1\right)\frac{(1-n_t)p_t^I}{(1-\tau_t)p_t}, \tag{8}$$

which shows that investment depends on the ratio between the shadow value of a unit of new capital and its replacement cost, i.e., $\lambda_t^K/(1-n_t)p_t^I$. Typically this ratio is labelled marginal q. Note that when marginal q equals unity investment proceeds at a rate such that marginal adjustment costs are zero.

[3]This approach assumes that τ_t and n_t do not depend on current investment, which is not strictly correct in the presence of tax exhaustion with incomplete loss offset. A model which allows for this dependency is developed in Devereux, Keen, and Schiantarelli (1991).

Choosing a quadratic form for G which is homogeneous in K_t and I_t, for example $(\phi/2)[(I/K)_t - \alpha - e_t]^2 K_t$, eq. (8) can be written as

$$\left(\frac{I}{K}\right)_t = \alpha + \beta(q_t - 1)\frac{(1-n_t)p_t^I}{(1-\tau_t)p_t} + e_t, \tag{9}$$

where $\beta = 1/\phi$ and where α is a 'normal' rate of investment at which adjustment costs average zero. Notice that all expectations concerning the marginal product of capital are summarised in q_t through the shadow value λ_t^K. The only stochastic term involved in (9) represents the unobservable factors e_t in adjustment costs. These may contain firm-specific effects and time effects common to all firms, in addition to an idiosyncratic time-varying shock. Moreover, the latter may be serially correlated, giving rise to a dynamic specification of the investment equation characterised by common factor restrictions. This is discussed further below.[4]

Like the Euler equation, (9) is not empirically implementable since λ_t^K, and therefore q_t, is not directly observable. Nevertheless, under the assumption of linear homogeneity of $F(K, X)$ and $G(I, K)$ we may, following Hayashi (1982), write q_t for each firm as[5]

$$q_t = \frac{V_t - A_t + H_t}{(1-\delta)\hat{p}_t^I(1-n_t)K_{t-1}}, \tag{10}$$

where $\hat{p}_t^I = p_t^I/(1+r_t)$ is the discounted price of investment goods,

$$A_t = \sum_{j=0}^{\infty} \beta_j \tilde{A}_{t+j}, \tag{11}$$

$$H_t = \sum_{j=0}^{\infty} \beta_j \left[i_{t+j}(1-\tau_{t+j})B_{t+j} - \left(B_{t+j+1} - B_{t+j}\right)\right]. \tag{12}$$

$\tilde{A}_{t+j}$ is the expected $t+j$ value of the depreciation allowances on investment made before period t and A_t is therefore the expected present value of such tax savings. H_t is the expected present value of all cash flows associated with debt, including interest payments and the additional funds derived from the

[4] Note that more general dynamic structures are not permitted since if lagged I_t or K_t were arguments of G, for example, then future adjustment costs would be directly controllable by current investment decisions. In this case the variable q_t in the optimal plan (9) would not be sufficient to summarise future expectations and past decisions.

[5] Eq. (10) is derived by first multiplying (6) by I_t, (7) by K_t, and taking the difference between the resulting expressions. This gives a difference equation, the solution to which can be rearranged using the homogeneity assumptions to generate (10).

issue of new debt. This term is commonly proxied by the stock of debt at the beginning of the period.[6] The right-hand side of (10) is known as average or Tobin's q, and after substituting for q_t in (9) the observable regressor is usually referred to as tax-adjusted Q. Note, then, that the average Q model avoids the direct use of the Euler eq. (7) in estimation. However, this is achieved at the cost of strong technological and capital market assumptions.

3. Data, estimation, and results

3.1. The data

The principal data requirement for the estimation of the model developed above are cross-section and time-series data for tax-adjusted Q and the gross investment rate. Two sources for this data have been used: company accounting data, made available to us by Datastream International, and share price and related data from the London Share Price Database.

Datastream provides company accounting data for all UK quoted companies from 1968. At the time of writing this paper we had access to periods up to 1986. The London Share Price Database offers a similar coverage. We selected from Datastream companies whose main activity was manufacturing. We excluded companies which had changed the date of their accounting year end during this period, so that all sets of accounts used would cover a 12-month period. Most of these companies were then matched to data available from the London Share Price Database. We did not use data for the earliest three years, for reasons discussed below, concerning estimation of the replacement value of the capital stock. In addition, other data was sometimes incomplete (for example, lacking data on the capital stock), and so these (and either previous or subsequent) records for that company were eliminated. Companies that had made major acquisitions were also dropped. Excluding companies with less than ten records, we were left with an unbalanced sample of 532 companies with a number of records varying between ten and sixteen.

The detailed computation of the variables used is described in the data appendix. Company investment includes direct purchases of new fixed assets and those acquired through acquisitions. The firm's market value is an average for the three months prior to each accounting year. The book values of debt and net current assets are deducted to obtain the valuation of fixed

[6] Eq. (12) can be written:

$$H_t = B_t[1 + i_t(1-\tau_t)] - \sum_{j=1}^{\infty} \beta_j \left[r_{t+j} - i_{t+j}(1-\tau_{t+j}) \right] B_{t+j}.$$

The second term vanishes if $r_t = i_t(1-\tau_t)$ for all t. More probably, it is positive, which provides a rather arbitrary justification for ignoring the mark-up factor in the first term.

capital. The principal complexities involve estimating the replacement cost values of the capital stock and the tax parameters. The former are not available in the data, and must be estimated from historic cost data. The replacement cost value for 1968 was taken to be equal to the historic cost value at that date, and thereafter updated using the perpetual inventory method, allowing a different depreciation rate for plant/machinery and land/buildings. To reduce the impact of the starting assumption on our results we did not use the first three years of data in estimation. In fact, we have found that our results are not very sensitive to the particular measure of the capital stock used [see Bond and Devereux (1989)].

The tax parameters vary from company to company for three reasons. First, the value of the investment allowance in any period depends on the breakdown of investment by asset in that and previous periods. Second, because companies declare their accounting results at different points in the year and statutory tax rates may change, then companies may face a different average tax rate. Third, we take into account that the firm faces a nonlinear tax schedule as a result of tax exhaustion. Because of generous allowances within the tax system, taxable losses have been common in the UK even though accounting losses have been rare. Virtually all empirical studies have ignored this phenomenon, yet we estimate that a substantial proportion of our sample, varying between 26% and 40%, was tax-exhausted in each year. Only around 10% of the sample was never tax-exhausted.

Even studies that do consider tax exhaustion generally assume only that the tax rate in a period of tax exhaustion is zero. However, with carry forward provisions in the tax system, tax liabilities are in fact simply delayed until the firm ends its period of tax exhaustion and so they should be discounted over this additional period of time. In principle we therefore require, for each period, the expectation at that date of any future periods in which the company will be tax-exhausted (which depends on expectations of both future tax rates and future company behaviour). In this paper, we assume that the firm has rational expectations of periods of tax exhaustion and substitute realisations. In order to assess the periods in which firms actually have zero taxable profits we use a model of the UK corporation tax system which applies tax rules to company accounting data to estimate tax liabilities [see Devereux (1986)].[7] In principle this makes our estimates of the tax parameters τ_t, n_t, and A_t, and hence Q_t, dependent on future as well as current shocks to the investment process. We tested for the importance of this endogeneity by recalculating these parameters under the assumption of no tax exhaustion and using the recalculated values of Q_t to instrument our preferred measure. The results were very similar to those reported in the next section.

[7]Data are predicted beyond 1986 using a simple autoregressive model [see Devereux (1986)].

3.2. Estimation and results

The investment equation for firm i derived in section 2 can be summarised most conveniently for empirical work as

$$\left(\frac{I}{K}\right)_{it} = \alpha + \beta Q_{it} + e_{it}, \tag{13}$$

for $i = 1,2,\ldots,N$ and $t = 1,2,\ldots,T$, where

$$Q_{it} = (q_{it} - 1)\frac{(1 - n_{it})p^I_{it}}{(1 - \tau_{it})p_{it}}, \tag{14}$$

and q_{it} is defined as in eq. (10). We adopt the following component structure for the disturbances:

$$e_{it} = \alpha_t + \alpha_i + \nu_{it}. \tag{15}$$

As we noted in the theoretical discussion of section 2, the error term may contain company-specific effects α_i and time-specific effects α_t, as well as an idiosyncratic shock ν_{it}. Indeed, there is nothing in the theory that restricts ν_{it} to be an innovation. However, as we also noted, more general dynamic relationships between the investment rate and Q that are not generated by an autoregressive process for ν_{it} are not consistent with the standard theory. The firm-specific effect can be interpreted as a component of the 'normal' rate of investment at which the firm's adjustment costs are zero.

Our econometric analysis begins with an assessment of the appropriate stochastic assumptions on Q_{it} and the components of e_{it}. For consistent estimation of the parameter β in (13), the stochastic properties of Q_{it} are crucial. Asymptotic arguments will rest on limiting properties for large N; T will be considered finite throughout. If Q_{it} is uncorrelated with both α_i and ν_{is} for all s and t, and ν_{it} is uncorrelated across time, then the standard variance-components GLS estimator is appropriate. Q_{it} is, in this case, strictly exogenous. Alternatively, Q_{it} may be correlated with the fixed effect α_i but may still be uncorrelated with all ν_{is}, in which case Q_{it} is strictly exogenous with respect to ν_{is} alone and the within groups estimator would be appropriate but GLS becomes inconsistent.

These estimators all assume the dynamic structure in (13) and (15) is correct. However, if we are unwilling to assume that Q_{it} is strictly exogenous with respect to ν_{is}, or wish to entertain the possibility of more general dynamic models including the lagged dependent variable, then both the within groups estimator and the GLS estimator are inconsistent. Correlation between Q_{it} and ν_{it} (and α_i) may well arise, for example through the dependence of λ^K_t on e_t in eq. (7). We therefore use an instrumental variable approach on a first-differenced (13) in which the instruments are weighted optimally so as to form a Generalised Method of Moments (GMM) estimator [see Hansen (1982)]. For example, if Q_{it} is endogenous and ν_{it} serially uncorrelated, then both $Q_{i,t-2}$, $(I/K)_{i,t-2}$, and further lags are valid instruments for the first-differenced equation for firm i in period t. Data from periods $3,\ldots,T$ can be used in estimation.[8] In fact, from the definition of the data used Q_{it} is measured at the beginning of the period, while, clearly, adjustment costs occur during the period. This raises the possibility that empirically Q_{it} can be treated as predetermined, which would allow the use of $Q_{i,t-1}$ as an additional instrument.

Omitting time effects for notational simplicity, the GMM estimator has the form

$$\hat{\beta} = (x'ZA_N Z'x)^{-1} x'ZA_N Z'y, \tag{16}$$

where x is the stacked vector of observations on ΔQ_{it} and y is the stacked vector of observations on $\Delta(I/K)_{it}$. The instrument matrix Z has the form illustrated in footnote 8. In the presence of general heteroskedasticity across both firms and time, the optimal choice for A_N is given by

$$A_N = \left(\frac{1}{N} \sum_{i=1}^{N} Z_i' \,\Delta\hat{\nu}_i \,\Delta\hat{\nu}_i' Z_i \right)^{-1}, \tag{17}$$

where the vectors $\Delta\hat{\nu}_i$ are consistent estimates of the first-differenced residuals for each firm [see White (1982)]. These are obtained from a preliminary consistent estimator of β, setting $A_N = (N^{-1}\sum_{i=1}^{N} Z_i' H Z_i)^{-1}$, where H is a matrix with twos on the leading diagonal, minus ones on the first off-diagonal, and zeros elsewhere. In all cases this first-stage estimation gave similar, though less well determined, coefficient estimates.

The use of endogenous variables dated $t-2$ (or predetermined variables dated $t-1$) as instruments is only valid if ν_{it} is serially uncorrelated, implying a first-order moving average error term in the differenced model. It is therefore important that we test for the presence of higher-order serial correlation, and for this purpose we employ the one degree of freedom test

[8] For each firm, the instrument set that exploits all available linear moment restrictions is (dropping firm subscripts):

$$Z_i = \begin{pmatrix} Q_1 & 0 & 0 & \cdots & 0 & \cdots & 0 & \vdots & (I/K)_1 & 0 & 0 & \cdots & 0 & \cdots & 0 \\ 0 & Q_1 & Q_2 & \cdots & 0 & \cdots & 0 & \vdots & 0 & (I/K)_1 & (I/K)_2 & \cdots & 0 & \cdots & 0 \\ \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \vdots & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot \\ \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \vdots & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot \\ 0 & 0 & 0 & \cdots & Q_1 & \cdots & Q_{T-2} & \vdots & 0 & 0 & 0 & \cdots & (I/K)_1 & \cdots & (I/K)_{T-2} \end{pmatrix}.$$

For firms with incomplete data the rows of Z_i corresponding to the missing equations are simply deleted, and missing values in the remaining rows are replaced by zeros. The matrix Z in (16) is obtained by stacking the Z_i for each company. In Monte Carlo simulations, Arellano and Bond (1991) show some considerable efficiency gains compared to simpler instrumental variables estimators.

(m_2) proposed by Arellano and Bond (1991). This exploits the property that $N^{-1/2}(\Delta\hat{\nu}'_{-2}\,\Delta\hat{\nu})$ is asymptotically normally distributed with known variance, where $\Delta\hat{\nu}$ is the stacked vector of residuals from the differenced model and $\Delta\hat{\nu}_{-2}$ is the conformable vector of residuals lagged twice. Essentially this tests for zero elements on the second off-diagonal of the estimated serial covariance matrix. This permits a test for second-order serial correlation to be computed without requiring an auxiliary regression – in which $\Delta\nu_{it}$ is assumed to be MA(2) – to be estimated. In this context we also report robust Sargan tests of the overidentifying restrictions that our estimator exploits.

In the absence of higher-order serial correlation, the GMM estimator provides consistent estimates of the parameters in equations like (13). This remains the case even when the lagged dependent variable and other endogenous regressors are introduced into the model, provided that a valid instrument set continues to be used. Note though that if Q is measured with error, which would induce negative correlation between the current shock ν_{it} and Q_{it}, then $Q_{i,t-1}$ could no longer be a valid instrument and its inclusion could be expected to induce a downward bias in the estimate of β. Although differencing may exacerbate measurement error bias in general, the first-difference estimates are robust to permanent measurement error and by omitting $Q_{i,t-1}$ from the set of instruments we are able to assess the importance of white noise measurement errors [Griliches and Hausman (1986)]. However, we will also present levels estimates for comparison.

Following Hayashi and Inoue (1991) we may also wish to use future values of Q in our instrument set Z for the first-differenced models. If ν_{it} is serially independent, then under efficient markets $Q_{i,t+1}$ can be shown to be independent of current and past innovations to the investment process. If valid, the inclusion of such future instruments, where available, can be expected to increase the efficiency of the GMM estimator. This discussion of the stochastic specification of the Q model emphasises the potential importance of assumptions on the exogeneity of Q for consistent estimation. In the context of the GMM estimator described above this will correspond to the appropriate timing of instruments and, as a result, will form an important aspect of our empirical investigations.

Turning to the empirical results themselves, in table 1 we produce some estimates of the basic Q model in first-differenced form. The sample contains 532 firms and 4739 observations. We maintain a common sample period 1975–86 across all the specifications and in all cases the heteroskedastic-consistent standard errors associated with the GMM estimator are reported in parentheses. Our estimation procedure was implemented on a microcomputer using GAUSS 1.49B and the DPD program [see Arellano and Bond (1988)]. In practice this limits the dimension of the instrument matrix we could use. The instrument matrix we use has the general form described in footnote 8 allowing the instruments used in each period to increase as we move through the panel and more observations become available. However,

Table 1

Some basic Q models in first differences.[a]

(a) Unrestricted	(i)	(ii)	(iii)	(iv)	(v)	(vi)
ΔQ_{it}	0.0092 (0.0020)	0.0077 (0.0023)	0.0056 (0.0013)	0.0099 (0.0023)	0.0137 (0.0022)	0.0060 (0.0007)
$\Delta Q_{i,t-1}$	—	−0.0012 (0.0011)	−0.0013 (0.0007)	−0.0006 (0.0019)	−0.0004 (0.0022)	−0.0010 (0.0006)
$\Delta(I/K)_{i,t-1}$	—	0.2475 (0.0241)	0.2365 (0.0245)	0.2337 (0.0260)	0.2359 (0.0243)	0.2521 (0.0236)
m_2	−3.62	−0.66	−0.77	−0.80	−0.78	−0.64
z_1	21.1(1)	126.5(3)	117.6(3)	122.8(3)	170.5(3)	185.1(3)
z_2	267.1(12)	206.0(12)	227.3(12)	192.7(12)	175.9(12)	236.2(12)
Sargan	75.66(73)	85.72(75)	76.96(68)	82.50(71)	89.97(71)	92.24(75)
Stability	0.02(1)	3.84(3)	3.71(3)	1.38(3)	16.51(3)	7.68(3)
(b) Restricted	**(i)**	**(ii)**	**(ii)**	**(iv)**	**(v)**	**(vi)**
ΔQ_{it}	—	0.0076 (0.0023)	0.0056 (0.0010)	0.0097 (0.0023)	0.0144 (0.0022)	0.0059 (0.0007)
ρ	—	0.2448 (0.0238)	0.2365 (0.0242)	0.2378 (0.0256)	0.2389 (0.0243)	0.2520 (0.0236)
z_1	—	127.1(2)	134.3(2)	120.5(2)	166.8(2)	185.1(2)
Comfac	—	0.56(1)	0.00(1)	0.79(1)	1.97(1)	0.73(1)
Instruments	$Q_{t-2} \cdots Q_{t-8}$	$Q_{t-2} \cdots Q_{t-7}$ $(I/K)_{t-2}$	$Q_{t-1} \cdots Q_{t-5}$ $(I/K)_{t-2}$	$Q_{t-3} \cdots Q_{t-8}$ $(I/K)_{t-2}$	$Q_{t+4} \cdots Q_{t+1}$ Q_{t-3}, Q_{t-4} $(I/K)_{t-2}$	$Q_{t+3} \cdots q_{t-2}$ $(I/K)_{t-2}$

[a]1) Time dummies are included as regressors and instruments in all equations.

2) Asymptotic standard errors are reported in parentheses. Standard errors and test statistics are robust to general time-series and cross-section heteroskedasticity.

3) m_2 is a test for second-order serial correlation in the residuals, asymptotically distributed as N(0, 1) under the null of no serial correlation.

4) $z_1(k)$ is a Wald test of joint significance of the reported coefficients, asymptotically distributed as $\chi^2(k)$ under the null of no relationship.

5) $z_2(k)$ is a Wald test of joint significance of the time dummies.

6) The Sargan statistic is a test of the overidentifying restrictions, asymptotically distributed as $\chi^2(k)$ under the null.

7) The stability test is a Wald test of the hypothesis that the reported coefficients are common across the subperiods 1975–80 and 1981–86, asymptotically distributed as $\chi^2(k)$.

8) The Comfac statistic is a test of the common factor restriction that the dynamics are generated by an AR(1) disturbance, asymptotically distributed as $\chi^2(k)$.

as a result of the limitation noted above we retain for each cross-section only the most recent instruments. In the tables we report the set of instruments that were used where available.

Column (i) presents the results for a model that allows Q_{it} to be endogenous and also correlated with the fixed effects but assumes that ν_{it} is serially uncorrelated. Although the estimate of β is positive and significant, the m_2 statistic provides a signal of potential dynamic misspecification. Residual autocorrelation remained when lagged values of Q were added to the model, and also when lagged values of the investment rate were included in the instrument set. In the latter case the Sargan statistic decisively rejected the validity of these instruments. In column (ii) we therefore present a model under the same assumptions but including a further lag of both ΔQ_{it} and the change in the investment rate. Here we include $(I/K)_{i,t-2}$ among the instruments, which obliges us to use fewer lagged values of Q due to the limit on the total size of the instrument matrix. The statistical properties of this model are now much more acceptable, but a general model of this type is not consistent with the theoretical model expounded in section 2.

However, as we noted above, if the dynamic specification could be represented by persistence in the error terms entering model (13) above, then the empirical model would fit within the theoretical assumptions. For this to be the case here, ν_{it} in (15) would have to be AR(1) and the parameter estimates in column (ii) would have to satisfy the common factor restrictions for an AR(1) process. Specifically the coefficient on $\Delta Q_{i,t-1}$ must equal minus the product of the coefficients on ΔQ_{it} and $\Delta(I/K)_{i,t-1}$. This 'comfac' restriction is imposed on the unrestricted parameter estimates in the lower part of column (ii) by the minimum distance approach, with ρ being the estimate of the autoregressive coefficient.[9] The test statistic for these restrictions is then the minimised value of this criteria function which is asymptotically distributed as $\chi^2(k)$ under the null, where k is the number of restrictions. The results show that the data quite clearly do not reject the comfac restriction despite the significance of the individual coefficients. As a result, the theory behind the basic model does not appear to do too much violation to the data. We also find that these estimates pass a Wald test for the stability of the slope coefficients over the two halves of the sample period, split in 1981.[10]

[9] Letting $\pi = (\pi_1, \pi_2, \pi_3)'$ denote the unrestricted coefficients on Q_{it}, $Q_{i,t-1}$, and $(I/K)_{i,t-1}$ respectively, and $\theta = (\beta, \rho)'$ the restricted coefficients, with $\pi(\theta) = (\beta, -\beta\rho, \rho)'$ being the restriction, we choose θ to minimise $[g(\hat{\pi}) - g(\pi(\theta))]'\hat{\Omega}^{-1}[g(\hat{\pi}) - g(\pi(\theta))]$, where $\hat{\Omega} = (\partial g(\hat{\pi})/\partial\pi')\widehat{\text{var}}(\hat{\pi})(\partial g(\hat{\pi})/\partial\pi')'$ and $g(\pi) = (\pi_1, -\pi_2/\pi_3, \pi_3)'$ is chosen so as to make $g(\pi(\theta)) = (\beta, \beta, \rho)'$ linear in β and ρ.

[10] We test the significance of additional variables formed by interacting each of the original regressors with a dummy variable whose value is one for 1981–86 and zero otherwise. The Wald statistic is asymptotically distributed as $\chi^2(k)$, where k is the number of additional parameters.

Although column (ii) appears to be a reasonable specification for the microeconometric relationship between the investment rate and Q, we may wish to investigate some of the assumptions underlying these parameter estimates, in particular the importance of the endogeneity assumptions on Q_{it}. We first test for the possibility that Q_{it} is predetermined with respect to ν_{it}. By including $Q_{i,t-1}$ as an instrument, we investigate the possibility of biases due to correlation between $Q_{i,t-1}$ and the first-differenced error term $\Delta\nu_{it}$. Column (iii) shows that the estimate of β falls when $Q_{i,t-1}$ is included in the instrument set. This may suggest that measurement error in Q is leading to downward bias in the Q coefficients, and that this downward bias is more than offsetting any upward bias due to the simultaneous determination of $Q_{i,t-1}$ and $\nu_{i,t-1}$.

In the presence of measurement error neither $Q_{i,t-1}$ nor $Q_{i,t-2}$ would be valid instruments. They are therefore excluded from the instrument set in column (iv). The effect is a large increase in the coefficient on Q_{it} compared with column (iii), and a smaller increase compared with column (ii), as would be expected in the presence of measurement error. The standard error of the Q coefficient in column (iv) is only marginally greater than in column (ii), which suggests that there is surprisingly little efficiency loss by excluding $Q_{i,t-2}$ from the instrument set. However, both (ii) and (iv) show a loss in efficiency compared to (iii) due to the absence of $Q_{i,t-1}$ from the instrument set. All three models continue to satisfy the common factor restriction for an AR(1) process and the Wald test for parameter stability. We further considered the possibilities that the measurement error in Q is MA(1) rather than white noise and that there may be measurement error in the investment rate. Omitting $Q_{i,t-3}$ from the instrument set had very little effect on the coefficient estimates but further reduced their precision. Using $(I/K)_{i,t-3}$ in place of $(I/K)_{i,t-2}$ actually reduced the (unrestricted) coefficient on $\Delta(I/K)_{i,t-3}$ to 0.18 (standard error = 0.06), with little change in the remaining coefficients. Neither experiment leads us to reject the specification in column (iv).

In column (v) we examine the possibility that the conditions hold under which future instruments may be valid, by including future values of Q in the instrument set, beginning with $Q_{i,t+1}$ (as well as lagged values, beginning with $Q_{i,t-3}$). We continue to exclude Q_{it}, $Q_{i,t-1}$, and $Q_{i,t-2}$ from the instrument set on the grounds of measurement error in Q. As predicted by the theory, the presence of serial correlation in ν_{it} induces upward bias when using future-dated instruments, leading to a large increase in the estimate of β compared to column (iv). In this case, the hypothesis of stability of the slope coefficients across the two subperiods is also rejected.

Finally, in column (vi) we test the effect of assuming strict exogeneity of Q_{it} with respect to ν_{is}, as would be required for the consistency of GLS and within groups estimators, by adding Q_{it}, $Q_{i,t-1}$, and $Q_{i,t-2}$ to the instrument set in (v). The estimate of β now falls to below that found in column (iv),

Table 2
Alternative estimators of the levels investment model.[a]

(a) Unrestricted	(i)	(ii)	(iii)	(iv)	(v)
	OLS	OLS	Within groups	GLS	Corrected GLS
Q_{it}	0.0099 (0.0010)	0.0077 (0.0010)	0.0081 (0.0011)	0.0076 (0.0011)	0.0077 (0.0010)
$Q_{i,t-1}$	—	−0.0025 (0.0009)	−0.0002 (0.0010)	−0.0016 (0.0009)	−0.0016 (0.0009)
$(I/K)_{i,t-1}$	—	0.4387 (0.0266)	0.1957 (0.0293)	0.3373 (0.0275)	0.3225 (0.0297)
R^2	0.086	0.26	—	—	—
m_1	13.06	−1.90	7.86	−0.78	0.01
m_2	11.45	1.77	2.15	−0.42	−0.34
z_1	104.86(1)	388.6(3)	143.4(3)	240.8(3)	205.2(3)
z_2	112.93(11)	108.2(11)	134.0(11)	110.6(11)	31.4(11)
z_3	45.93(8)	18.8(8)	—	13.1(8)	12.6(8)
z_4	—	—	—	—	30.6(12)
(b) Restricted	**(i)**	**(ii)**	**(iii)**	**(iv)**	**(v)**
Q_{it}	—	0.0076 (0.0010)	0.0081 (0.0011)	0.0075 (0.0010)	0.0076 (0.0010)
ρ	—	0.4413 (0.0265)	0.1974 (0.0292)	0.3409 (0.0274)	0.3248 (0.0296)
z_1	—	361.7(2)	140.6(2)	228.7(2)	195.5(2)
Comfac		1.11	1.44	1.19	1.03

[a]1) Industry dummies are included in columns (i), (ii), (iv), and (v).
2) m_1 is a test for first-order serial correlation in the residuals, asymptotically distributed as N(0, 1) under the null of no serial correlation.
3) $z_3(k)$ is a Wald test of the joint significance of the industry dummies.
4) $z_4(k)$ is a Wald test of the joint significance of the time-varying coefficients on the initial condition added to each cross-section in column (v), as explained in the text.
See also notes to table 1.

suggesting that the simultaneity bias apparent in column (v) is dominated by the downward bias due to measurement error. Whilst both effects appear to be important there is the expected tendency for some offsetting to occur when strict exogeneity is imposed. Nevertheless our preference from these results would be for the specification in column (iv). This allows for an AR(1) disturbance and measurement errors in Q. It shows a reasonably precise and relatively large estimate of β in a theoretically consistent and stable model.

In table 2, we continue to examine the endogeneity of Q_{it}, by estimating the model in levels. A levels model would provide efficient parameter estimates if Q_{it} was uncorrelated with the fixed effect α_i. Moreover, biases resulting from white noise measurement errors can be shown to be smaller in this case. However, since the lagged dependent variable is correlated with the fixed effect by construction, we need to allow for this correlation if we are to reproduce the levels equivalent of estimators in table 1.

In the first column the OLS estimates equivalent to the first column of table 1 are presented. They show an estimate for the Q coefficient which is perhaps surprisingly in line with our preferred estimate in column (iv) of that table. However, there is strong evidence of serial correlation suggesting the presence of firm-specific effects. In column (ii) of table 2 we provide estimates of the more general dynamic specification. Here we find evidence that the coefficient on the lagged dependent variable in this level specification is biased upwards using OLS showing evidence of firm-specific effects (despite the inclusion of industry dummies). As would also be expected, the within groups estimates in column (iii) produce a downward bias on this coefficient. However, in each of these columns the coefficient on Q_{it} is virtually unaffected, suggesting that Q may be only weakly correlated with the fixed effect. If we could assume that Q_{it} was strictly exogenous with respect to ν_{is} and uncorrelated with the fixed effects, then GLS would be a consistent estimator. The standard variance-components GLS transformation sweeps out all of the correlation between the lagged dependent variable and the fixed effect except for the correlation with the initial observation of the dependent variable [see Hsiao (1986), for example]. However, adding this initial value to the GLS regression for each period removes this correlation and renders the estimator consistent.[11] As expected the estimate of ρ drops significantly compared to OLS and the correction reduces this coefficient still further. Nevertheless the estimate remains uncomfortably above those found in table 1 which casts doubt on the consistency of GLS in this application. However, a noteworthy feature of these dynamic levels models is the robustness of the common factor restriction.

Although we have found our measure of average Q to be a significant determinant of company investment, we know from section 2 that strong restrictions on technology, adjustment costs, competition, and stock market efficiency are required for average Q to be a sufficient statistic for marginal q. For this reason alone we might expect other variables such as cash flow to contain independent explanatory power. Moreover, if some firms are constrained from raising as much external finance as they would like due to capital market imperfections, then the availability of internal finance could limit their investment. We investigate this possibility by adding terms in net cash flow as a proportion of end-of-period capital stock [denoted (C/K)] to

[11]This assumes that the stochastic process for the initial condition has a constant mean. In more general cases the residual from an initial regression should be included in place of the initial observation [see Blundell and Smith (1991)]. We obtained very similar results when using the more general correction.

Table 3

Additional variables in the differenced investment model.[a]

(a) Unrestricted	(i)	(ii)
ΔQ_{it}	0.0070 (0.0023)	0.0063 (0.0026)
$\Delta Q_{i,t-1}$	−0.0018 (0.0019)	−0.0036 (0.0020)
$\Delta(C/K)_{i,t-1}$	0.1880 (0.0229)	0.1931 (0.0244)
$\Delta(C/K)_{i,t-2}$	0.0140 (0.0168)	0.0201 (0.0176)
$\Delta(Y/K)_{it}$	—	−0.0131 (0.0023)
$\Delta(Y/K)_{i,t-1}$	—	0.0073 (0.0023)
$\Delta(I/K)_{i,t-1}$	0.2001 (0.0222)	0.2072 (0.0211)
m_2	−0.86	−0.87
z_1	186.7(5)	231.8(7)
z_2	118.2(12)	137.5(12)
Sargan	83.19(67)	82.04(65)
Stability	15.9(5)	112.4(7)
Comfac	8.35(2)	18.63(3)
Instruments	$Q_{t-3} \cdots Q_{t-5}$ $(C/K)_{t-2}, (C/K)_{t-3}$ $(I/K)_{t-2}$	Q_{t-3}, Q_{t-4} $(C/K)_{t-2}, (C/K)_{t-3}$ $(Y/K)_{t-2}, (I/K)_{t-2}$

[a]See notes to table 1.

our preferred (unrestricted) first-differenced GMM model. Our measure of net cash flow is obtained by adding back accounting depreciation to the book value of post-tax profits.

These results are reported in column (i) of table 3 and confirm that lagged cash flow does contain additional explanatory power, although Q_{it} remains statistically significant. Note that in estimation $(C/K)_{it}$ is assumed to be correlated with e_{it}. As a result $\Delta(C/K)_{i,t-1}$ is treated as endogenous and the earliest instrument we use is $(C/K)_{i,t-2}$. The current cash flow ratio was found to be insignificant and dropped from the specification, although current cash flow was highly significant when treated as exogenous. The hypothesis of parameter stability across the subperiods before and after 1981 is rejected here. Interestingly the coefficient on cash flow is significantly higher, and that on Q_{it} lower, for the earlier period which includes the severe recession after 1979 in the UK. The common factor restriction is also mildly rejected by these estimates.

Having found a role for cash flow in addition to average Q, column (ii) of table 3 investigates the importance of adding terms in the ratio of output to end-of-period capital stock (Y/K). Perhaps surprisingly, although the coefficients on output are strongly significant, those on average Q and cash flow remain close to those in column (i) and retain their significance. The comfac restriction and the hypothesis of stability across time are again rejected, with the coefficients on output appearing to be particularly unstable. The negative coefficient on contemporaneous output could reflect monopolistic product markets, in which case an additional term reflecting the present value of future monopoly rents should be deducted from the numerator in (10) [see Schiantarelli and Georgoutsos (1990)]. Certainly the coefficients on the two output terms, conditional on average Q and cash flow, do not appear to support an accelerator effect.

4. Conclusions

The aim of this paper has been to investigate the importance of Tobin's Q in the determination of investment decisions at the company level. We have approached this question by estimating a standard Q model on disaggregated panel data for 532 UK manufacturing companies over the period 1975 to 1986. The results have highlighted the sensitivity of parameter estimates in such models to the choice of dynamic specification, exogeneity assumptions and measurement errors in Q. Whilst Q was found to be a significant determinant of investment, its coefficient was small. Since the numerator of Q relates to the stock market value of the company, it appears that the investment rate is relatively unresponsive, at least in the short run, to variations in equity values. Indeed, the short-run elasticity of the gross investment rate to the equity market value indicates that a 10% rise in the equity market value would be associated with an immediate rise in the investment rate of only 2.5%.

To derive the theoretical relationship between the investment rate and average Q that is used in estimation, we are required to make a number of strong assumptions on technology and adjustment costs as well as on the efficiency of the stock market. However, the theory does allow for some dynamic extensions to the standard investment–Q relationship, although an unrestricted distributed lag formulation is ruled out. We find that dynamic generalisations are an important factor in our derivation of a data-coherent specification, but we also find that the autoregressive restrictions on the pattern of dynamics suggested by the theory are acceptable. In our estimated models we allow for individual firm-specific effects which, with the inclusion of lagged dependent variables implied by the dynamic generalisations, requires a careful choice of estimation technique. Since our average Q variable is also allowed to be endogenous and possibly correlated with the firm-specific effects, the initial estimator we choose to work with is of a Generalised Method of Moments type in which past variables are utilised as instruments. Estimates from this specification are then compared with those that make stronger assumptions on the exogeneity of Q in order to assess the sensitivity of our results. Further investigations of time stability and instrument validity were also considered and our preferred specification was one that allowed Q to be endogenous and correlated with the firm-specific effects. Indeed, the Q coefficient was found to be quite sensitive to misspecifications in these assumptions.

The structural restrictions placed on the average Q model suggest that this measure of Q may not completely reflect all the determinants of investment decisions in the way the theory would predict that marginal Q should. As a result we investigated the influence of two additional factors. The first, cash flow, has been suggested by a number of authors as an important influence on investment. Our results confirm this suggestion but again point to the importance of correctly dealing with the endogeneity of cash flow and its dynamic specification. Moreover, we find a continuing role for average Q. This conclusion is unaffected by the inclusion of our second factor, output, which although significant did not eliminate the importance of either Q or cash flow. In addition we argue that its negative sign is suggestive of monopoly effects rather than an accelerator model.

Data appendix

This appendix describes the calculation of the principal variables used in the estimation, and provides some summary statistics. There are two data sources: company accounting records from Datastream and share price and related data from the London Share Price Database. A sample of 532 companies was selected whose main activity (allocated by sales) was in the UK manufacturing sector, which were available from both data sources and which had continuous data for at least ten years. These companies were allocated to nine subsectors of manufacturing according to their main product as in table 4, and the structure of the sample by number of observations per company is given in table 5. Some descriptive statistics of the variables used in estimation over the period 1975–86 are shown in table 6.

The variables were constructed as follows (square brackets refer to the codes of Datastream items):

Investment $(p^I I)$: Total new fixed assets [435].

Cash Flow (C): Provision for depreciation of fixed assets [136] plus profit after tax, interest and preference dividends [182].

Total Sales (Y): Total sales [104].

Replacement cost of the capital stock $(p^I K)$: For most of the years of data, accounting rules did not require replacement cost to be declared.

Table 4

Group	Industry	SIC classes	Number
1.	Metals and metal goods	21, 22, 31	38
2.	Other minerals and mineral products	23, 24	25
3.	Chemicals and man made fibres	25, 26	30
4.	Mechanical engineering	32, 33	141
5.	Electrical and instrument engineering	34, 37	54
6.	Motor vehicles and parts, other transport equipment	35, 36	29
7.	Food, drink and tobacco	41, 42	73
8.	Textiles, clothing, leather and footwear	43, 44, 45	80
9.	Other[a]	46, 47, 48, 49	62

[a]'Other' includes 'timber and wooden furniture', 'paper and publishing', 'rubber and plastics', and 'other manufacturing'.

Table 5

Number of records per company	Number of companies
10	67
11	153
12	41
13	37
14	67
15	69
16	98

Table 6

	Mean	Std. dev.	Minimum	Maximum
I/K	0.090	0.068	0.001	0.785
Q	−0.978	1.473	−9.906	9.363
C/K	0.114	0.096	−0.689	1.856
Y/K	2.196	1.824	0.016	35.201

Hence we have access only to historic cost valuations of the capital stock, separately for plant and machinery and for buildings. Since investment is not split by asset, the split in gross fixed asset data is used to estimate the split in investment:

$$p_t^I I_t^P = p_t^I I_t \left[\frac{(GFP_t - GFP_{t-1})}{(GFP_t - GFP_{t-1} + GFB_t - GFB_{t-1})} \right], \tag{A.1}$$

$$p_t^I I_t^B = p_t^I I_t - p_t^I I_t^P, \tag{A.2}$$

where $p_t^I I_t^P$ and $p_t^I I_t^B$ denote the estimated value of investment in plant and machinery and buildings, respectively, and GFP_t and GFB_t denote gross historic cost values of plant and machinery [328] and buildings [327], respectively.

The replacement cost value of capital is calculated from the perpetual inventory formula:

$$p_{t+1}^I K_{t+1}^i = p_t^I K_t^i (1-\delta^i)\left(p_{t+1}^I / p_t^I\right) + p_{t+1}^I I_{t+1}^i, \tag{A.3}$$

for $i = P, B$. Values of δ of 8.19% for plant and machinery and 2.5% for buildings are used, taken from estimates in King and Fullerton (1984) for UK manufacturing industry. To obtain starting values for the perpetual inventory method we assume equality of replacement cost and historic cost valuations of the capital stock in the first year of data, usually 1968.

The replacement cost valuation of total fixed capital assets is

$$p_t^I K_t = p_t^I K_t^P + p_t^I K_t^B. \tag{A.4}$$

Share valuation (V): Using monthly observations (last trading day) on share prices for each company from the London Share Price Database, we calculate an average price for the three months preceding the accounting year. This average price is then multiplied by the total number of issued shares outstanding at the beginning of the year.

Debt and other assets (B): Book value of total loan capital with repayment due in excess of one year [321] plus the book value of other long-term liabilities (total deferred tax [312], other provisions [313], and minority interests [315]), less the book value of other assets (total intangibles [344], total investments [356], current assets [390], and other assets [359]).

Price indices: The output prices (p) are implicit value-added price deflators for the nine subsectors of manufacturing defined above, constructed from current price GDP and constant price GDP figures published by industry in various Blue Books. The price of investment goods (p^I) is an implicit price deflator for gross fixed investment by manufacturing industry, using data in the Economic Trends Annual Supplement (1988).

Tax parameters: These were estimated allowing for tax exhaustion. The existence of loss carry forward provisions implies that any liabilities or allowances are postponed until the company is once more in a tax-paying position. If the company does not pay any tax until period $t+n$, then the effective corporate tax rate, τ_t^*, is

$$\tau_t^* = \beta_n \tau_{t+n}, \tag{A.5}$$

with

$$\beta_n = \prod_{i=1}^{n} (1 + r_{t+i})^{-1}, \qquad n = 1, 2, \ldots,$$

where τ_{t+i} is the average statutory tax rate during period $t+i$ and β is the discount factor between periods t and $t+i$, using as a discount rate the rate on British government consols. We ignore the expectations operator for ease of exposition. In practice we substitute realised periods of tax exhaustion [estimated from the model described in Devereux (1986)] and the discount rate. However, we assume that all changes to statutory tax rates were unforeseen, except for those announced as part of the UK 1984 corporation tax reforms. Clearly, for $n = 0$ we revert to the case with no tax exhaustion.

To assess investment allowances we consider two types of asset, plant and machinery, and industrial buildings. Commercial buildings receive no allowance, and can therefore be ignored. Following calculations in Devereux (1986), we assume that industrial buildings constitute 65% of the total value of buildings.

Plant and machinery receive a first-year allowance, α^P, and in subsequent years a depreciation allowance, δ^P, on a reducing balance basis. Tax exhaustion can be allowed for by using τ^* rather than τ in the formula. The present value of tax allowances on a unit investment in plant and machinery is then

$$n_t^P = \tau_t^* \alpha_t^P + \sum_{j=1}^{\infty} \beta_j \tau_{t+j}^* d^P(t,j), \tag{A.6}$$

where $d^P(t,j)$ is the depreciation allowance in year $t+j$, given in the UK by

$$d^P(t,j) = \delta_{t+j}^P \left[\prod_{i=1}^{j-1} \left(1 - \delta_{t+i}^P\right) \right] \left(1 - \alpha_t^P\right), \tag{A.7}$$

where $\prod_{i=1}^{0} x_{t+i}$ is defined to take the value unity.

Industrial buildings receive a straight line depreciation allowance, δ^B, which remains fixed over the lifetime of the asset, together with an additional allowance in the year of investment, α^B. Hence, a building receives an allowance for a fixed number of years, T_t, defined from

$$\alpha_t^B + T_t \delta_t^B = 1. \tag{A.8}$$

Following a similar approach to that for plant and machinery, the present value of tax allowances on a unit investment in industrial buildings is

$$n_t^B = \tau_t^* \alpha_t^B + \sum_{j=0}^{T_t - 1} \beta_j \tau_{t+j}^* \delta_t^B. \tag{A.9}$$

The present value of tax allowances on a unit of new investment made in year t is then calculated as

$$n_t = \frac{n_t^P p_t^I I_t^P + 0.65 n_t^B p_t^I I_t^B}{p_t^I I_t}. \tag{A.10}$$

The present value in year t of tax allowances on investment in plant and machinery made in year $t-s$ $(s>0)$ is

$$A_{t-s,t}^P = p_{t-s}^I I_{t-s}^P \sum_{j=0}^{\infty} \beta_j \tau_{t+j}^* d^P(t-s, s+j), \tag{A.11}$$

and the present value of remaining allowances on all investments in plant and machinery made before period t is

$$A_t^P = \sum_{s=1}^{\infty} A_{t-s,t}^P = \sum_{j=0}^{\infty} \beta_j \tau_{t+j}^* \left(\sum_{s=1}^{\infty} d^P(t-s, s+j) p_{t-s}^I I_{t-s}^P \right). \tag{A.12}$$

Calculation of the present value of remaining allowances on an investment in industrial buildings made in period $t-s$ is analogous to (A.11):

$$A_{t-s,t}^B = 0.65 p_{t-s}^I I_{t-s}^B \sum_{j=0}^{T_{t-s} - s - 1} \beta_j \tau_{t+j}^* \delta_{t-s}^B. \tag{A.13}$$

The present value of remaining tax allowances on all investments in industrial buildings made before period t is then

$$A_t^B = \sum_{s=1}^{\infty} A_{t-s,t}^B, \tag{A.14}$$

although only periods for which $s < T_{t-s}$ need to be explicitly considered. The present value of all remaining tax allowances on investments made before period t is then simply $A_t = A_t^P + A_t^B$.

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