ECONOMETRIC MODELLING OF UK AGGREGATE INVESTMENT:

THE ROLE OF PROFITS AND UNCERTAINTY

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Abstract

This paper focuses on the determinants of aggregate investment spending in the UK for the industrial and commercial company (ICC) sector. It complements recent work by Cuthbertson and Gasparro (1995), who study an augmented Tobin's q model of investment in the manufacturing sector. Important focal points of our analysis are a role for real profits (internal funds), which allow firms to combat liquidity constraints when access to capital markets is not perfect, Chirinko (1987), and the impact of irreversibility and uncertainty in determining aggregate investment spending. Earlier work on manufacturing investment by Bean (1981a) developed a dynamic, error correction specification based on the flexible accelerator model. Following Cuthbertson and Gasparro we use multivariate cointegration techniques to discover a parsimonious dynamic model, which can explain the 1980s and early 1990s investment experience of the ICC sector. Our results show that a model based on investment and output alone does not cointegrate, and a short-run dynamic model of these variables suffers from heteroscedasticity. This may be consistent with the idea that increased (uncontrolled for) uncertainty has led to increased volatility in investment. The possibility that movements in the real price of gold reflect uncertainty in financial and other traded commodity markets is explored. Investigation of this more general model indicates that real profits and the real price of gold can enhance the explanation of investment spending by the ICC sector.

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ECONOMETRIC MODELLING OF UK AGGREGATE INVESTMENT: THE ROLE OF PROFITS AND UNCERTAINTY

1. Introduction

The modelling of investment expenditure, both at the microeconomic and aggregate level, is a topic that has occupied applied econometricians for a considerable number of years. Nevertheless in the 1980s and 1990s the modelling of aggregate investment expenditure has been a much less researched topic than, say, the modelling of aggregate consumer expenditure or the demand for money. This is somewhat surprising given the importance of physical capital accumulation to the process of economic growth, and the massive research effort on new growth theory. However there has been considerable recent theoretical interest in an approach which treats investment as a financial call option. This has led to a burgeoning literature on the effects of irreversibility and uncertainty on investment behaviour (Dixit and Pindyck, 1994). Some empirical work on this issue has been attempted, but is lagging behind, partly because of the difficulty of taking a theoretical innovation which essentially suggests that there is a greater range of inactivity before investment decisions may be triggered, and turning it into an empirical model of investment expenditure. Much of this empirical work has been concerned with the disaggregated level of the industry or, in particular, the firm. Such aggregate work as has been attempted has been generally concerned with the USA.[1] By and large this research has focused on the appropriate definition and specification of proxy

[1] See, for example, time series aggregate analysis performed by Ferderer (1993), Huizinga (1993), Goldberg (1993) and Episcopos (1995). Pindyck and Solimano (1993) present international aggregate cross-sectional analysis. Papers by Driver and Moreton (1991) and Price (1995, 1996) using UK aggregate data are an exception.

measures for uncertainty in order to assess the impact of increased uncertainty on the level or rate of investment activity.[2]

A rather older tradition in the empirical modelling of investment has been concerned with the potential role for profits in explaining investment activity. Early empirical work on investment, drawing from a more institutionalist or Keynesian methodological basis, pointed to the possibility that, in a world of less than perfect capital markets, firms may face liquidity constraints, and so current and past profitability can provide a cheaper and more readily accessible source of investment funds (Kuh 1963). Using flow-of-funds data, Bond and Jenkinson (1996) show that, over the period 1970-94, 93 per cent of all fixed investment was funded from internal sources in the UK, and 96 per cent in the US. Comparable figures for Germany and Japan are 80 per cent and 70 per cent respectively. If capital markets are perfect, and the tax treatment of different sources of finance are the same, then investment spending should never be limited by a shortage of internal finance. Yet a quarter of large British manufacturing companies, surveyed by the CBI, reported that their investment was constrained by the availability of internal finance. So the perfect capital market assumption is not supported by the economic facts, and real profits may be a useful proxy for potentially available internal funds.

In a recent study of UK manufacturing industry, Cuthbertson and Gasparro (1995) embrace the theoretical framework of dynamic optimisation that underlies the popular q model. By augmenting this model to include the agency costs of debt and periodic demand constraints, they derive a model whereby $I = F(q, Y, G)$: q is effectively average q, Y is

[2] In a more direct approach, Pindyck and Solimano (1993) attempt to proxy the threshold level at which investment is triggered.

output, and G is capital gearing. This model is investigated using cointegration and ECM techniques, and with the aid of a number of dummy variables, the empirical model performs satisfactorily. One contentious issue in much macroeconometrics is the extent to which the theory and the empirical model are linked. Dynamic optimisation theory generates non-linear, structural relationships in the form of Euler equations. However, empirical macroeconomic research in the UK rarely estimates such structural relationships, mainly due to the difficulty of obtaining results consistent with the underlying theory.

Chirinko (1987), in a study of US aggregate investment, emphasises the importance of linking the financial and real sectors within the framework of dynamic optimisation. However, he notes the empirical performance of structural q models has been disappointing for three reasons. First, lagged variables are found to be significant determinants of investment, contrary to the theory (c.f. the Hall (1978) rational expectations permanent income hypothesis consumption function debate). Second, q itself is found to have weak explanatory power both within sample and out-of-sample. Third, the implied structural parameters are not believable, especially with respect to the adjustment of the capital stock to unexpected shocks. He shows that basic q theory can be misleading and requires augmentation to include endogenous financial policy, which is consistent with the augmented q framework of Cuthbertson and Gasparro. Chirinko amends the basic q analysis to account for firms participating in a number of financial markets, and estimates a non-linear, first-order structural condition which includes debt variables. The disappointment is that his more general model does not work very well empirically, and he proposes the *ad hoc* addition of internal funds and macroeconomic demand conditions, proxied by the change in real GDP. As such it is perhaps not too surprising that Cuthbertson and Gasparro chose not to estimate the structural Euler equations underlying their theory model.

Moreover, it might be argued that the representative firm, dynamic optimisation approach invites a microeconometric response. In this vein, Bond and Meghir (1994) adopt a discrete time representation of the firm's dynamic optimisation problem, and augment the q approach by taking account of imperfect competition and quadratic adjustment costs. This yields an Euler equation which includes output and cash flow. A further extension of their model to incorporate financial decisions augments the Euler equation with a debt variable. Other, sophisticated and disaggregated empirical analyses, such as those by Fazzari, Hubbard and Petersen (1988), Hayashi and Inoue (1991), and Blundell, Bond, Devereux and Schiantarelli (1992), support a role for real profit flows by revealing excess sensitivity of corporate investment to internal finance. Aggregate analysis of the relationship between investment and q has not been particularly successful for the UK (Oulton 1981, Poterba and Summers 1983), though Cuthbertson and Gasparro (1995), for the UK manufacturing sector, is an exception. There have been, however, no recent attempts to explain the time series of aggregate investment in terms of aggregate profit flows.[3]

The recent theoretical emphasis on the role of uncertainty is also sympathetic to a renewed interest in profits, given the role of current profitability as a signal of potential future earnings. Q theory relies on stock market sentiment as an indicator of the market value of capital to provide an accurate, sufficient statistic for investment. While profitability may be more backward looking, it may provide a more accurate reflection of the expectations of those responsible for fixed investment decisions. The ability of profits to explain and forecast

[3] Anderson (1981) constructs an aggregate "cash-flow" variable and uses this, in a distributed lag formulation, to explain investment. However, as Bean (1981b) points out, his approach raises stock-flow problems in the availability of finance.

investment over and above information contained in other variables is, in practice, an empirical issue.

In this paper we focus on the industrial and commercial corporate (ICC) sector of the UK economy. During the 1980s and 1990s rates of growth in gross fixed capital formation in this sector have been highly volatile and much more so than GDP. As Figure 1 shows, annualised growth in real investment peaked in the first quarter of 1985 at 45 per cent but had fallen to minus 13 per cent by the second quarter of 1991. Even by the standards of the volatility in personal sector spending in the last fifteen years, ICC investment is extremely volatile. We pose the question of how well the traditional flexible-accelerator to investment-output model can explain this volatility, raising the issue of whether there is a role for profits in the determination of aggregate investment expenditure. We also make use of the international gold price as an indirect proxy for uncertainty[4]. Since gold is generally regarded as a low-risk hedge, movements in its price ought to reveal important information about market sentiment *vis à vis* other asset returns. Given that gold is highly inelastic in supply, we would expect price movements to reflect changes in demand and so, in contrast to other price series, we expect movements in the gold price to reflect more closely demand-driven substitution for other assets. Our results suggest important and significant effects for both real profits and the real gold price in both the short and the long-run. If we concentrate on a

[4] As a measure of uncertainty, this has the advantage that it has a global dimension and might therefore be construed as exogenous. There is also an issue over whether the correct terminology should be “risk” rather than “uncertainty”. Here we use them interchangeably. It is not possible to provide a compelling theory that links uncertainty and the price of gold. Company financial policy, and liquidity constraints create agency problems. Uncertainty adds to the difficulties for the parties interested in company performance, but as Chirinko points out makes it very difficult to derive structural econometric models. Often σ^2 is appended to theory models, and then a proxy is sought from the CAPM model or the measurement of conditional variance using ARCH techniques. We tried both of these methods, but with little success.

simpler bivariate model of investment and output, then we find no significant long-run relationship, and a short-run dynamic model that suffers from heteroscedasticity. Real profits and the real price of gold together can resolve this heteroscedasticity in our more general model, emphasising their importance for the specification of the empirical model, and consistent with the interpretation that heteroscedasticity is illustrative of increased uncertainty.

The remainder of the paper is structured as follows. Section 2 considers the irreversibility and uncertainty literature and its impact on the debate about modelling investment. Section 3 sets out a simple model which introduces a role for profits in the theoretical explanation of investment spending. Section 4 considers the long-run equilibrium of investment models for the industrial and commercial companies sector, after appraising the time series properties of the individual series. Section 5 develops a dynamic model of investment for the industrial and commercial companies sector. Section 6 concludes.

2. Irreversibility and Uncertainty

The recent literature on irreversible investment and uncertainty is concerned not so much with modelling the determinants of the steady-state level of fixed investment, but with the proposition that the timing of investment may be altered by uncertainty. There is, of course, nothing new in the suggestion that investor behaviour will be sensitive to uncertainty about future conditions. Indeed it is implicit in early adjustment costs models and in q-models. A key result, established by Hartmann (1972) and extended by Abel (1983), is that uncertainty over future input and output prices can increase the value of a marginal unit of capital and *increase* the level of investment. On the other hand other authors have subsequently demonstrated that the presence of irreversibility may lead to postponement of investment decisions (Bernanke 1983, MacDonald and Siegel 1986). With increasing uncertainty (or strictly speaking increasing variance in the distribution of the future rate of return from the project) the value of the "call option" to delay an investment project increases and the decision to invest is delayed. Increased uncertainty, other things equal, will *reduce* the current level of investment. This is illustrated in Figure 2 which shows two relationships between the rate of return and the final product price, P, which the firm takes as parametric but uncertain. The first shows the conventional net present value of an investment to produce goods for sale at P. The second includes the value of the option to delay the project. At low levels of P, the option to delay is more valuable than conventional net present value, and the two only become equal at or above a critical price level, P^*. We will therefore observe lower levels of actual investment than predicted by conventional theory at prices below P^*. It can also be shown (see Dixit and Pindyck, 1994) that increased uncertainty raises the option value curve and so shifts outwards the threshold price, P^*. Aggregate empirical evidence, to be discussed below, generally supports the existence of this second, negative relationship between uncertainty and investment (see Carruth, Dickerson and Henley, 1998).

Since the irreversible investment literature deals with the timing of investment and not necessarily the level of investment activity, a full model needs to incorporate an independent explanation of the "certainty-equivalent" determinants of investment. A further point concerns aggregation: if firm (or sectoral) fluctuations in uncertainty are not coincident then it is possible that these fluctuations will cancel each other out at the aggregate level, or that the uncertainty element will be submerged in the dynamic specification of any time series empirical model. However Bernanke (1983) makes two arguments for why the effects of uncertainty may not disappear in aggregate. Firstly macroeconomic factors, such as uncertainty about future interest, exchange and inflation rates or shocks in monetary, fiscal or regulatory policy regimes, may be important in determining firm-level decisions. Secondly if a firm is uncertain about whether a shock is transitory or permanent then it may delay investment decisions in order to learn more about its degree of permanence. Here there are the elements of a mechanism through which (genuinely) transitory shocks become converted into a more permanent fluctuation. As Bernanke points out, the propagating mechanism through which uncertainty affects long-run investment decisions is the irreversibility effect. However, the influence of irreversibility may be submerged in dynamic specification (timing of investment) in aggregate time series studies, so the quest to find good measures of aggregate uncertainty may be an elusive one. In other words, uncertainty may well be an important determinant of investment, but in aggregate empirical models attempts to find a role for explicit proxies may be fruitless, because the effects of uncertainty are already embedded in the modelling of investment dynamics, which are usually quite complex, and consistent with complex, possibly non-linear forms of dynamic adjustment.

There are a number of empirical issues here. One concerns the appropriate way to measure the effects of uncertainty. A number of approaches and proxy variables have been adopted in previous work. By and large all studies use a variance measure capturing volatility in output, inflation or exchange rates as a proxy for uncertainty. There is an apparent distinction in the literature between work which uses an unconditional variance, usually constructed using some form of moving moment method (Pindyck and Solimano 1993, Goldberg 1993, Favero, Pesaran and Sharma 1994) and work which uses an estimated conditional variance computed from preliminary estimation of some form of ARCH model (Huizinga 1993, Price 1995, 1996, Episcopos 1995).[5] On the other hand Ferderer (1993)

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[5] ARCH measures of uncertainty are dependent on the (mis)specification of the empirical model, Engle (1983), p. 287. It is quite common to use a univariate VAR on output or inflation, which display an ARCH error test failure, as the basis for estimating the conditional variance of these variables. This is not an adequate way of proxying uncertainty, because more

adopts a more direct approach by using the implied risk premium in the term structure of interest rates as a measure of uncertainty. We propose here the use of an indirect (rather than variance) proxy for uncertainty, namely gold prices.[6] As an indirect measure, gold is commonly regarded as a hedge commodity, the demand for which rises as returns on other less safe assets become more volatile. Thus, in aggregate, the price of gold will reflect investor sentiment. We would therefore expect a rising gold price to be coincident with rising uncertainty about investment returns. A time series plot of the growth in the real sterling-denominated gold price against real ICC fixed investment growth is illustrated in Figure 3. Particularly in the period from the late 1970s until 1990 there is some evidence of a negative association.

A further concern is the question of whether uncertainty has only a short-run or both a long and short-run effect on investment. Much will depend on how proxies for uncertainty are established and on the stationarity/non-stationarity of the relevant time series. This is an empirical issue, but it does suggest that any empirical methodology adopted should be able to distinguish long-run from short-run dynamic behaviour.

3. Empirical Approach

Broadly speaking, the pre-irreversibility, theoretical literature on investment divides itself into two main strands: the flexible accelerator approach and the neoclassical approach. In addition to these main strands is a subsidiary literature concerned with the importance of

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appropriate (different) models of output or inflation processes may not exhibit ARCH error failures. However the discovery of a better specification, which passes the ARCH test, does not mean that the impact of uncertainty in economic behaviour has disappeared.

[6] In preliminary work we also investigated using the Confederation of British Industry's quarterly business optimism series as a more direct proxy for uncertainty. We could find no

corporate liquidity in determining investment. Accelerator models posit a desired capital-output ratio, and construct some form of dynamic adjustment mechanism to relate actual gross investment to output. The neoclassical approach argues that such a simple bivariate specification ignores substitution possibilities among factor inputs within the firm, and hence suggests that factor prices, notably the user cost of capital, should be taken into account. The data-based methodology of Cuthbertson and Gasparro (1995) as well as Bean (1981a) integrates neoclassical theory with a general-to-specific approach to the parameterisation of the dynamic relationships between the variables, and we adopt a similar approach here.

Suppose we start from a constant elasticity of substitution, constant returns to scale technology, and suppose that demand for capital is subject to a mean zero disturbance, γ[7]. The factor demand for capital will be given by:

$$K = A\, Y\, [(1 + 1/\varepsilon)/C\,]^{\sigma} \exp(\gamma), \qquad (1)$$

where A is the scale parameter, K is capital, Y is output, C is the user cost of capital, ε is the output price elasticity of demand, and σ is the elasticity of substitution. We can think of γ as capturing the size of the wedge between the user cost of an extra unit of capital and its present value that arises from the option value of refraining from investing, as illustrated in Figure 2. The familiar identity describing the dynamic evolution of capital stock is:

$$I = \delta K + dK, \qquad (2)$$

significant relationship with ICC investment. It is, of course, a moot point as to whether an increase in optimism necessarily implies a reduction in uncertainty.

[7] We adopt a static framework to illustrate the basic arguments that lead to a reduced form empirical specification. Dynamic optimisation of the representative firm's intertemporal value function is a more popular theory model, but is more appropriate to the estimation of structural Euler equations.

where I is gross investment, d represents the change and δ is the depreciation rate. Assume the existence of a steady state in which capital stock and output are growing at a constant rate g, then from (2):

$$I = (g + \delta)K. \tag{3}$$

Substituting (1) into (3) eliminates the capital stock term and gives:

$$I = A\,(g + \delta)Y\,[(1 + 1/\varepsilon)/C\,]^{\sigma} \exp(\gamma). \tag{4}$$

If we assume that the elasticity of demand, ε, is constant, and allow a general polynomial lag structure to model dynamic adjustment costs, then the discrete time analogue of equation (4) is:

$$a_1(L)\Delta i_t = a_0 + a_2(L)\Delta y_t + a_3(L)\Delta c_t + a_4(L)\Delta\gamma_t + b_1 i_{t-j} + b_2 y_{t-j} + b_3 c_{t-j} + b_4 \gamma_{t-j} + u_t, \tag{5}$$

where a_1 to a_4 are polynomials in the lag operator, L, u_t is the error term, and lower case letters denote logs. The parameterisation of the log levels terms allows the imposition of a long-run steady-state solution in a conventional equilibrium correction form (ECM).

We can relax the assumption of a constant elasticity of demand in an empirical specification by using the Lerner condition identity between $1 + 1/\varepsilon$ and the price-cost margin. Movements in profit margins may therefore reflect changes in market power and so in the firm’s derived demand for capital. Parameterisation of this within the time series investment function is attractive since profit margins may exhibit considerable cyclical variability (Rotemberg and Saloner 1986, Bils 1987); and, if aggregate industrial structural conditions display secular movement, then we may expect to observe long-run movements in ε. However, any aggregate empirical profit margin construct such as a profit-sales ratio presents a problem for empirical representation here in that it is an I(0) series. Therefore, we use a measure of real profits (π) since this displays the same integration properties as real investment and output. This leads to an empirical formulation of the following form:

$$a_1(L)\Delta i_t = a_0 + a_2(L)\Delta y_t + a_3(L)\Delta c_t + a_4(L)\Delta\gamma_t + a_5(L)\Delta\pi_t + b_1 i_{t-j} + b_2 y_{t-j} + b_3 c_{t-j} + b_4 \gamma_{t-j} + b_5 \pi_{t-j} + u_t. \tag{6}$$

Implicit in modelling investment in terms of the cost of capital is the assumption that capital markets are perfect; so, for example, the firm is indifferent between rented capital goods and capital goods bought with borrowed funds. Furthermore, and perhaps more importantly, the firm is assumed to be indifferent between internal and external financing of funds, since external finance is not rationed and is available at the same opportunity cost as internal finance. By contrast, as mentioned in the introduction, recent microeconometric evidence has provided a number of alternative justifications for the general hypothesis that profits may, in part, determine investment. The stock of finance available to the firm will be a determinant of investment expenditure under less than perfect capital markets. In addition to the suggestion that current profits might capture information about investment sentiment about the future, real profits might play an important long-run role in such a model by proxying retained earnings. While these different justifications suggest different measures of profitability, given the high collinearity that exists between pre-tax and post-tax profit, net and gross profit, profit including dividends and retained profits, and stock market valuation (Tobin’s q), it seems unlikely that time-series modelling of aggregate investment will allow us to discriminate econometrically between the different explanations for the role of real profits.

One further problem is the possible collinearity between profitability and output. When comparing long-run equilibrium outcomes for a profit-maximising firm under imperfect competition, in a stable industrial structure, one should observe a monotonic relationship between output and profits (once output is determined profit must be determined). Hence any profit-based theory becomes empirically indistinguishable from the simple bivariate investment-output model. However, it is certainly not clear that at all points in time all firms will be observed to be in long-run equilibrium. For example Costrell (1983) shows that changes in profitability affect investment but that the direction of this relationship depends on whether risk aversion is an increasing or decreasing function of income, and so on whether agents prefer risk-free holdings of money or bonds to risky investment. Moreover it is not the case that profits and output are collinear either in a cyclical or secular sense as Figure 4 illustrates[8].

As already discussed, in order to capture the impact of uncertainty we use a measure of uncertainty based on the real gold price. In the results presented below, this is derived as the dollar price of gold, converted to sterling and deflated by the GDP deflator. The alternative of simply deflating the dollar gold price by the UK GDP deflator made no difference to our findings, so we are confident that our results are not being explained as the impact of exchange rate movements on investment.

4. Data Description and Long-Run Analysis

Definitions and sources of the various data series are explained in the Appendix. Table 1 reports summary statistics and augmented Dickey-Fuller unit root tests for the real ICC investment, real GDP, real ICC profits, the real sterling gold price, a real long-term interest rate as the cost of capital (yield on 20 year gilts). All of the series, with the exception of the real interest rate, are expressed in natural logs. The descriptive statistics reveal that

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[8] Strictly speaking, the comparison should be between profits and ICC output, rather than GDP, but ICC output is difficult to measure (see Anderson, 1981), so GDP may be a reasonable instrument.

average annual investment growth over the sample period was 2.6%, slightly higher than the annual growth in profits and output which were 2.1% and 2.5% respectively. Growth in the real gold price was of a very similar order of magnitude at 2.1% per annum. The table also shows that for all the levels series we do not reject the null hypothesis of a unit root. Stationarity is present in all the I(1) levels cases, when the series are expressed as quarterly or annual differences. As the raw series are not seasonally adjusted, we use annual differences in the short-run dynamic estimation. This can also be justified on the grounds that annual percentage changes are more likely to be contained in agents' information sets; however, similar results can be obtained with the variables defined as quarterly changes.

Granger (1981, 1986) established an important result between cointegrated regressions and error or, nowadays, equilibrium correction mechanisms, ECM. This is the idea that for any group of cointegrated variables an ECM representation holds, and the converse is also true. This link has ensured the continuing popularity of ECM parameterisations of time series econometric relationships, and is the empirical methodology adopted by Cuthbertson and Gasparro despite their setting out of a dynamic optimisation, theory model. With multivariate long-run relationships there are a number of ways to proceed. One possibility is to adopt the Engle-Granger (1987) method, and to estimate either a static long-run relationship or solve for the static solution to an ADL dynamic long-run relationship. This method has been criticised because it ignores the possibility of multiple cointegrating vectors but, in practice, it is seen to work quite well, and can give eminently sensible results (albeit of a reduced form nature). An alternative is to use Johansen's multivariate technique (Johansen and Juselius, 1990), even though the underlying structural model is not developed analytically. However, the possibility of multiple cointegration vectors can lead to severe identification problems, requiring the researcher to provide an economic interpretation of the relationships that are identified.

Moreover, the number of significant cointegrating vectors found is often dependent on the length of the lags chosen for the VAR, so a judicious choice of lag length can serve to disguise these kinds of difficulties. Given the relative advantages and disadvantages of the different methods, we compare the results from all of them in our empirical analysis.

Tables 2 and 3 report the results of the long-run analysis of the data. Table 2 restricts the empirical model to examine investment and output alone which parallels the earlier work of Bean (1981a) who estimated an investment-output ECM model[9], while Table 3 also includes profits, the interest rate and the price of gold. Panel (a) uses Johansen's analysis to provide evidence of the rank of the long-run matrix of a vector autoregression representation. On the basis of the model reduction methodology employed by PC-Fiml V8.1 (Doornik and Hendry, 1995) it was established that lags above order five were not jointly significant and so the results reported are derived from a VAR which uses lags up to and including order 5. For Table 2, the eigenvalue statistic does not reject the null of non-cointegration (rank = 0) at the 5 per cent level (using Osterwald-Lenum critical values). For Table 3, it rejects the null of non-cointegration, but indicates that the rank of the long-run matrix is not above one. The trace statistics confirm these results, although they only reject non-cointegration at the 10% level in Table 3.

If we impose a reduced rank of one for both models, then we obtain the vectors reported in column 3 of panel (b) in Tables 2 and 3. In both cases, these are very close to the unrestricted rank cointegrating vector. Columns 1 and 2 report alternative OLS estimates of a

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[9] This model could be expanded to include the real interest rate, but this would not change the essence of the results. In fact when we do this the results look worse because the real interest rate enters with the wrong sign. This is also true of the dynamic model results of Table 4.

long-run relationship between the appropriate variables, the first based on a static regression and the second as the long-run solution to a 5th order ADL specification. For the simple investment-output model, the results of both the static and the dynamic OLS estimation are very close to the Johansen result, but the results do not support a valid cointegrating relationship according to the ADF(4) test and MacKinnon's response surface critical values in either case. For Table 3, a similar result holds for columns 1 and 2.

Thus the simple investment-output model does not appear to cointegrate using Johansen's methodology, although it looks to be close for the dynamic OLS results. Whereas the more general model is cointegrated for Johansen estimation, but not for the other two OLS cases. Furthermore the inclusion of the real price of gold is important to the Johansen results of Table 3. If it is removed as a potential regressor, then cointegration is rejected at the 5% level amongst the remaining four variables.[10] Casual diagrammatic investigation of the residuals for all 3 columns of Table 2 and Table 3 shows that they are quite similar with considerable mean reversion, indicative of stationarity. It is well known that the cointegration tests have low power, so the foregoing results suggest that it is not unreasonable to investigate the short-run dynamic properties of the data for the investment-output and more general models.

For Table 3 there are some differences in the size of the coefficients in the long-run vector from each method. In the OLS methods, the long-run investment-output elasticity is close to or above unity and the long-run investment-profits elasticity is rather smaller in size. In the Johansen analysis this ranking is reversed and the long-run proportional impact of

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[10] These results are not reported and are available from the authors on request.

profits on investment is larger than that of output. Each long-run solution finds a small negative impact for the real interest rate.[11] We also find a small negative long-run impact on investment from increases in the real gold price. However it should be noted that neither the interest rate nor the gold price are statistically significant in the dynamic ADL equation.[12] Given these differences we explore the use of the residuals from all three methods as a measure of equilibrium correction in a dynamic model, following the methodology of Engle and Granger (1987). However, for Table 2, we use only the dynamic OLS residuals since the three sets of results are so close in their estimation of the long-run investment-output elasticity.

5. Dynamic Modelling

The Engle-Granger approach is to parameterise a short-run dynamic representation of the data generating process through specifying a relationship between lag polynomials of the stationary, differenced variables and lagged values of the ECM term derived from the long-run representation. A general to specific modelling strategy may then allow the discovery of a parsimonious model structure. Table 4 reports the results of such an exercise using equilibrium correction from the dynamic OLS results of Table 2. While this simple model performs well enough, it is noticeable that it fails the heteroscedasticity diagnostic. This is consistent with increased unmodelled volatility in investment arising from changes in investor uncertainty, though this is not the only possible explanation. As such we will concentrate on

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[11] Initial work attempted to measure a cost of capital series which incorporated estimates of the impact of investment subsidies and corporate taxation on the real interest rate, but like ICC output it is hard to combine the appropriate information across different sectors of the economy over long periods of time. As such the real interest rate outperformed (encompassed) our cost of capital measure.

[12] This conclusion is supported by F-tests of the joint significance of the terms in each variable: for R, $F(6,97) = 1.74$; for g, $F(6, 97) = 0.98$, CV 5% = 2.17.

the results of Table 5, which uses the equilibrium corrections from each of the three long-run vectors in Table 3 (note that real gold price information is used in determining these equilibria and ECMs). Results in each case are reported for the full period and for a sub-period up to 1985 quarter 4 to investigate forecast performance in the final ten years of the sample.

In general the Table 5 results are very robust to the alternative methods of deriving the ECM term. The Engle-Granger ECM model (columns 1 and 2 using static OLS residuals from Table 3, column 1) yields the smallest regression standard error. On the basis of the reported encompassing test statistics model 1 is preferred to models 3 and 5 and model 3 is in turn preferred to model 5. In all cases the models satisfy the reported diagnostic tests for autocorrelation (AR), autoregressive-conditional heteroscedasticity (ARCH), and for heteroscedasticity due to omitted squared regressors (Hetero x_i^2) and due to omitted cross-products (Hetero $x_i.x_j$). The normality tests reveal slight evidence of significant non-normality on the sub-sample estimates for the Engle Granger ECM, column 2.

The significant presence of the lagged dependent variable ($\Delta_4 i_{t-1}$) points to a degree of autoregression in the determination of short-run movements in investment, consistent with the earlier discussion on how the irreversibility idea may manifest itself in the timing of investment and on to the dynamic structure of an aggregate time series model of investment spending. The largest short-run impact on investment comes through changes in output ($\Delta_4 y_t$); the impact elasticity varies between 0.76 and 0.94 depending on specification and sample period. Changes in profits have a significant though quantitatively smaller impact; the elasticity varies between 0.08 and 0.17 and is largest in the Johansen ECM model. The long-run rate of interest has a quantitatively small but significant short-run impact. The gold price also has a negative and significant short-run impact on investment; the elasticity varies between -0.06 and -0.09. The two equilibrium correction variables are allowing lagged information in the disequilibrium levels of the data to control for short-run dynamics, over and above the lagged dependent variable dynamics of the investment growth process. This follows a suggestion of Phillips and Loretan (1991). The sum of the two coefficients on the equilibrium correction terms reveals the speed of adjustment towards equilibrium. This is generally quicker in the Engle-Granger ECM models than in the Johansen ECM model.

The even numbered columns in Table 5 reveal the forecast performance of the models. In each case the model coefficients are pleasingly robust to sample period and this is confirmed by the forecast Chow test statistics. The overall forecast performance, given by the $\chi^2(40)$ statistic for the significance of the difference between actual and forecast predictions, is good. Figure 5 provides a visual presentation of the forecast performance of the model in column 2 of Table 5. It is evident that the predictions track the data very well, and remain within the error bounds throughout the 10 year forecast period, 1986q1 to 1995q4. Finally, the encompassing tests of Table 5 favour the disequilibrium errors generated from the OLS estimated long-run relationships against the system based Johansen errors.

6. Conclusions

Recent work has attempted to address the appropriate way to capture the impact of uncertainty on investment, usually relying on some initial parameterisation of price or output variance, however these conditional variance approaches adopt an inappropriate methodology, relying on misspecification of an empirical model to generate a conditional variance measure (see footnote 5). Uncertainty may well be an important determinant of investment but, in aggregate empirical models, finding significant, explicit proxies may be difficult because the effects of uncertainty are already implicit in the modelling of the dynamics of investment. In this paper we have employed a more indirect, and arguably, more interesting proxy for uncertainty - namely the gold price. The size of the relationship between investment and gold price movements is small but significant - a long-run elasticity of around -0.07 (Tables 3 and 5). Disaggregated analysis may find a stronger relationship at the company level for the UK, but this requires further research.

We have also addressed the question of whether there is an aggregate role for corporate financial performance, as found in recent micro-econometric studies. Our results show that corporate profits significantly explain long and short-run movements in aggregate investment. A 10 per cent increase in corporate profits is associated with up to a 1.6 per cent short-run increase in fixed investment by the UK industrial and corporate sector. In the long-run the impact could be as high as 8.5 per cent. Aggregate analysis cannot reveal the source of this relationship, but it is consistent with a world of less than perfect capital markets in which industrial and commercial companies may be liquidity constrained, and heavily reliant on internally generated funds to fund long-term fixed investment activity. It is also consistent with current profits capturing investor expectational effects. While we find a significant, negative impact for long-term interest rates, it is a quantitatively small one. One conclusion to draw from this is that changes in corporate taxation policy may have a much larger impact on investment through the availability of internal corporate funds than through any impact on the cost of capital. Similarly we found on investigation of a simple investment-output model (Tables 2 and 4), that investment and output may not cointegrate but more importantly that there was some evidence of heteroscedasticity in the dynamic model. This was suggestive of a misspecification, perhaps due to the omitted effects of uncertainty, or missing variables like real profits or the real price of gold. The inclusion of real profits and the real gold price appear to satisfactorily eliminate this misspecification problem.

Table 1: Summary Statistics and Unit Root Tests

	Mean	
Levels: 1964q3 to 1995q4		ADF(4) H_0: I(1) CV: 5%=-2.884, 1%=-3.483
I	8.961	-0.842
Y	11.44	-1.113
R	2.770	-2.760
π	9.426	-0.685
g	5.368	-1.777
First differences: 1964q4 to 1995q4		ADF(4) H_0: I(2) CV: 5%=-2.885, 1%=-3.483
ΔI	0.0069	-4.341
Δy	0.0053	-4.645
ΔR	1.6070	-6.400
$\Delta\pi$	0.0071	-4.589
Δg	0.0050	-4.780
Fourth differences: 1965q3 to 1995q4		ADF(4) H_0: I(2) CV: 5%=-2.885, 1%=-3.485
Δ_4I	0.0263	-3.912
Δ_4y	0.0207	-3.127
Δ_4R	0.1390	-4.256
$\Delta_4\pi$	0.0249	-3.843
Δ_4g	0.0206	-3.456

Notes: i = log(real investment), y = log(real GDP), R = real interest rate, π = log(real profits), g = log(real gold price). Estimation is by PC-Give 8.1

Table 2: Long-run Analysis of Investment-Output Model

(a) Johansen Analysis for reduced rank of VAR system (1964q2 - 1995q4)

	Eigenvalue statistic	CV		Trace statistic	CV	
		95%	90%		95%	90%
H_0: rank = 0	9.79	14.1	12.1	11.22	15.4	13.3
H_0: rank ≤ 1	1.43	3.8	2.7	1.43	3.8	2.7

Note: Tests derived from 5^{th} order VAR, after suitable reduction using PC-Fiml 8.1

(b) Estimated cointegrating vectors

	(1) OLS static regression (1963q1-1995q4)	(2) Solved static long-run solution to ADL(5) dynamic model (1964q2 - 1995q4)	(3) Reduced Rank Johansen VAR(5) (1964q2 - 1995q4)
i	-1	-1	-1
intercept	-9.838	-10.95	
	(0.515)**	(1.92)**	
y	1.643	1.737	1.713
	(0.045)**	(0.168)**	
ADF(4) H_0: I(1)	-3.00	-3.06	
	(5% CV = -3.34)	(5% CV = -3.34)	
AR F(5,110)		0.662	1.94
ARCH F(4,107)		2.942*	5.225**
Normality χ^2 (2)		1.382	5.648
Hetero F(22,92)		2.001*	2.656**
Reset F(1,96)		2.346	N.A.

Notes: Standard errors in brackets. ** denotes significance at 1%, * at 5%, but not that the static OLS standard errors are not reliable. Estimation by PC-Give and PC-Fiml 8.1. The diagnostic tests in column 3 refer to the investment equation in the system. The reduction tests accept a 5^{th} order system, but the single equation diagnostics and vector diagnostics all worsen with the reduction from 8^{th} to 5^{th} order, as we would expect. The ADF critical values for columns 1 and 2 are based on Mackinnon's response surface estimates.

Table 3: Long-Run Analysis of the Full Model

(a) Johansen Analysis for reduced rank of VAR system (1964q2 - 1995q4)

	Eigenvalue statistic	CV 95%	CV 90%	Trace statistic	CV 95%	CV 90%
H_0: rank = 0	37.9	33.5	30.9	67.4	68.5	64.8
H_0: rank ≤ 1	14.8	27.1	24.7	29.5	47.2	44.0

Note: Tests derived from 5th order VAR, after suitable reduction using PC-Fiml 8.1

(b) Estimated cointegrating vectors

	(1) OLS static regression (1963q1 – 1995q4)	(2) Solved static long-run solution to ADL(5) dynamic model (1964q2 - 1995q4)	(3) Reduced Rank Johansen VAR(5) (1964q2 - 1995q4)
i	-1	-1	-1
intercept	-8.571	-6.767	
	(0.643)**	(1.732)**	
y	1.296	0.934	0.617
	(0.099)**	(0.293)**	
R	-0.006	-0.012	-0.034
	(0.003)*	(0.009)	
π	0.332	0.549	0.850
	(0.059)**	(0.185)**	
g	-0.076	-0.018	-0.038
	(0.021)**	(0.052)	
			$\chi^2(4) = 34.94$
ADF(4) H_0: I(1)	-3.151	-3.13	
	(5% CV = -4.41)	(5% CV = -4.41)	
AR F(5,92)		0.390	1.57
ARCH F(4,89)		0.804	3.471*
Normality $\chi^2(2)$		6.256*	2.842
Hetero F(58,38)		0.504	0.971
Reset F(1,96)		1.016	N.A.

Notes: Standard errors in brackets. ** denotes significance at 1%, * at 5%. Estimation by PC-Give and PC-Fiml 8.1. The $\chi^2(4)$ statistic in column 3 is a likelihood ratio test for the joint significance of y, R, π and g. The diagnostic tests in column 3 refer to the investment equation in the system. The reduction tests accept a 5th order system, but the single equation diagnostics and vector diagnostics all worsen with the reduction from 8th to 5th order, as we would expect. The ADF critical value for columns 1 and 2 is based on Mackinnon's response surface estimates.

Table 4: Dynamic Equilibrium Correction Models of the Investment-Output Model

Dependent variable: $\Delta_4 i_t$

	Dynamic Engle-Granger ECM	
	(1)	(2)
	65q2 - 95q4	65q2 - 85q4
Intercept	0.002	-0.005
	(0.007)	(0.009)
$\Delta_4 i_{t-1}$	0.491	0.461
	(0.06)**	(0.075)**
$\Delta_4 y_t$	0.997	1.193
	(0.22)**	(0.284)**
ecm_{t-3}	0.312	0.266
	(0.083)**	(0.103)*
ecm_{t-4}	-0.553	-0.555
	(0.085)**	(0.106)**
σ_u	0.0570	0.0606
R^2	0.645	0.633
5^{th} order AR	1.402	0.689
4^{th} order ARCH	1.341	1.060
Normality $\chi^2(2)$	2.486	3.941
Hetero x_i^2	2.363*	1.713
Hetero $x_i.x_j$	1.432	1.192
Reset	1.936	2.515
Forecast $\chi^2(40)$		29.27
Chow F(40,75)		0.668

Notes: Standard errors in brackets. ** denotes significance at 1%, * at 5%.

Estimation by PC-Give 8.1.

Table 5: Dynamic Equilibrium Correction Models - Real Gold Price

Dependent variable: $\Delta_4 i_t$

	Engle-Granger ECM		Dynamic Engle-Granger ECM		Johansen ECM	
	(1)	(2)	(3)	(4)	(5)	(6)
	65q2 - 95q4	65q2 - 85q4	65q2 - 95q4	65q2 - 85q4	65q2 - 95q4	65q2 - 85q4
intercept	-0.003	-0.009	-0.005	-0.012	-1.219	-1.249
	(0.007)	(0.008)	(0.007)	(0.009)	(0.373)**	(0.471)**
$\Delta_4 i_{t-1}$	0.474	0.432	0.466	0.440	0.517	0.526
	(0.061)**	(0.073)**	(0.063)**	(0.077)**	(0.067)**	(0.082)**
$\Delta_4 y_t$	0.842	0.936	0.763	0.930	0.759	0.831
	(0.240)**	(0.305)**	(0.245)**	(0.319)**	(0.263)**	(0.353)**
$\Delta_4 R_t$	-0.003	-0.003	-0.004	-0.004	-0.006	-0.005
	(0.001)*	(0.001)*	(0.001)**	(0.002)*	(0.002)**	(0.002)*
$\Delta_4 \pi_t$	0.103	0.083	0.149	0.132	0.166	0.142
	(0.043)*	(0.048)	(0.045)**	(0.051)*	(0.050)**	(0.059)*
$\Delta_4 g_t$	-0.085	-0.092	-0.073	-0.076	-0.063	-0.061
	(0.026)*	(0.030)*	(0.026)**	(0.031)*	(0.028)*	(0.034)
ecm_{t-3}	0.214	0.109	0.148	0.079	0.122	0.105
	(0.081)**	(0.098)	(0.071)*	(0.086)	(0.055)*	(0.066)
ecm_{t-4}	-0.584	-0.675	-0.501	-0.547	-0.330	-0.319
	(0.083)**	(0.106)**	(0.076)**	(0.097)**	(0.067)**	(0.084)**
σ_u	0.0560	0.0576	0.0572	0.0604	0.0617	0.0670
R^2	0.667	0.680	0.652	0.647	0.596	0.567
5^{th} order AR	0.725	0.364	0.538	0.282	1.182	0.558
4^{th} order ARCH	1.629	1.269	1.811	1.654	1.883	1.911
Normality $\chi^2(2)$	3.979	7.262*	5.149	5.855	4.449	4.903
Hetero x_i^2	1.319	1.118	1.104	1.005	1.384	1.441
Hetero $x_i.x_j$	0.822	0.570	0.698	0.494	0.891	0.944
Reset	2.830	4.498*	1.152	1.297	0.007	0.105
Forecast $\chi^2(40)$		43.199		32.408		22.918
Chow F(40,75)		0.839		0.702		0.558
Encompassing Tests:						
	(1) v (5)		(3) v (1)		(5) v (1)	
Cox N(0,1)	0.850		-2.54*		-7.274**	
Ericsson N(0,1)	0.839		2.37*		6.198**	
Sargan $\chi^2(2)$	0.835		5.64		20.911**	
	(1) v (3)		(3) v (5)		(5) v (3)	
Cox N(0,1)	0.226		1.803		-6.415**	
Ericsson N(0,1)	-0.220		1.809		5.670**	
Sargan $\chi^2(2)$	0.672		3.966		19.329**	

Notes: Standard errors in brackets. ** denotes significance at 1%, * at 5%.

Estimation by PC-Give 8.1.

Figure 1: Real UK Corporate Fixed Investment Growth

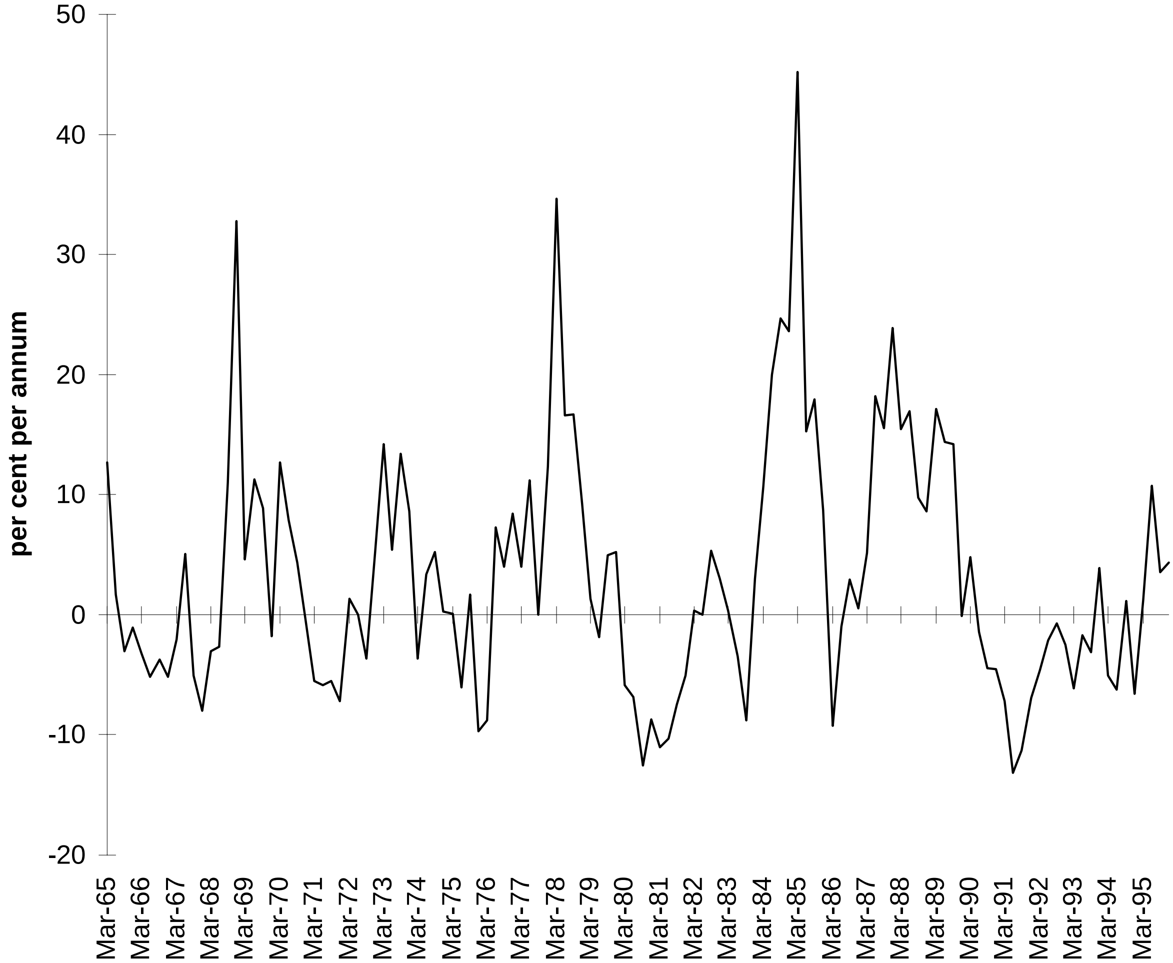


Figure 2: Option Values and the Decision to Invest

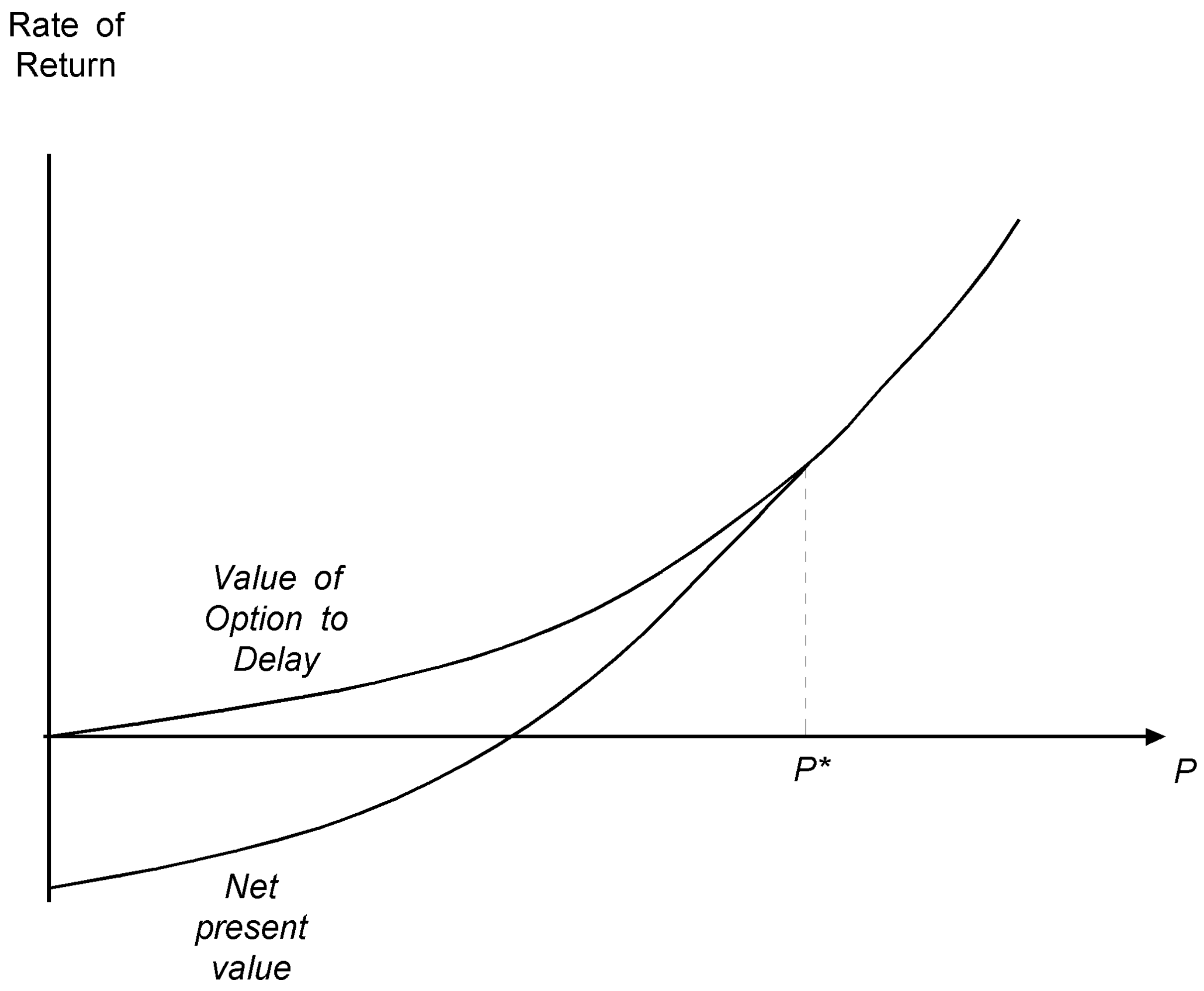


Figure 3: Investment and Gold Price Growth

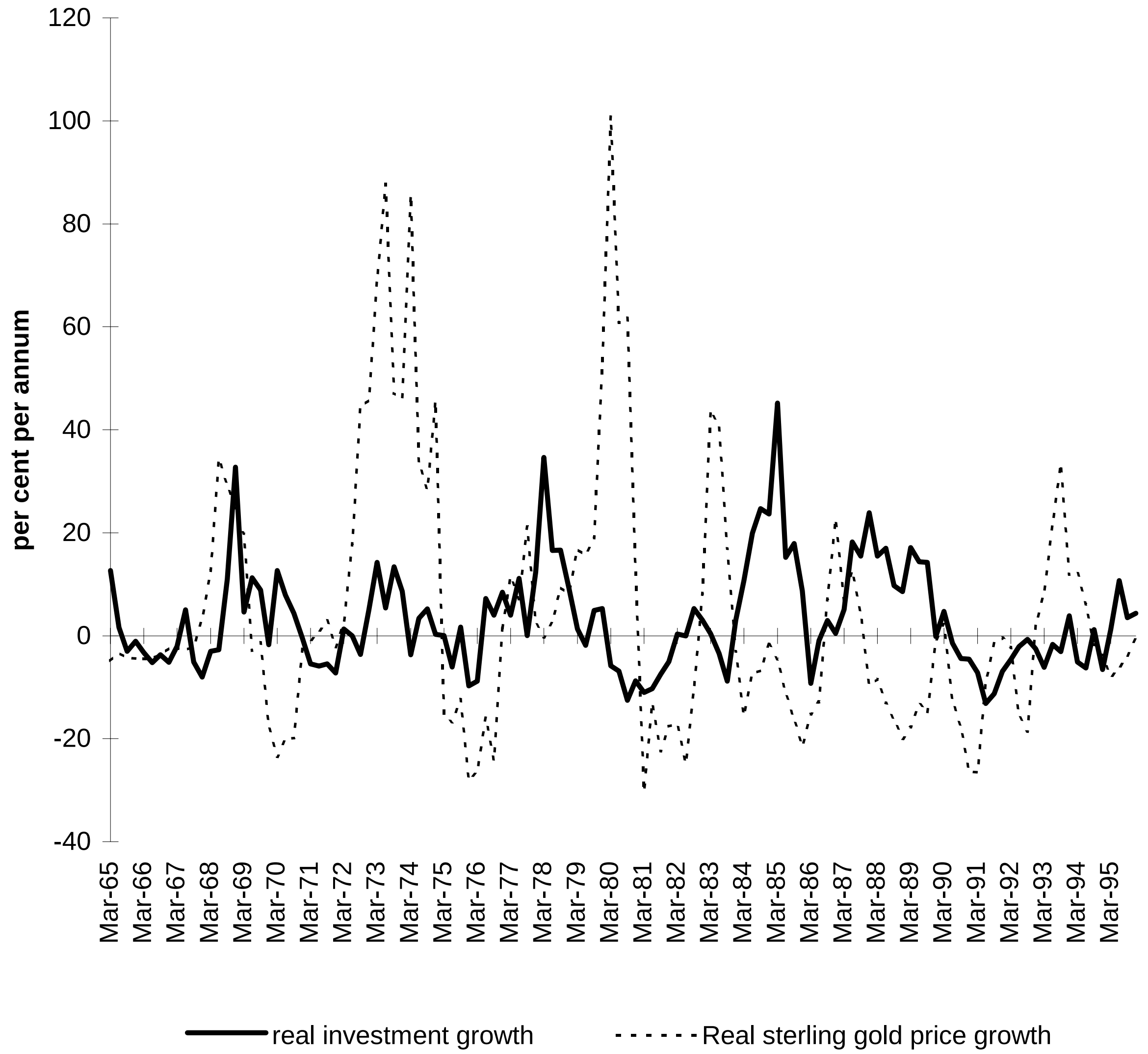


Figure 4: Real GDP and Profits Growth

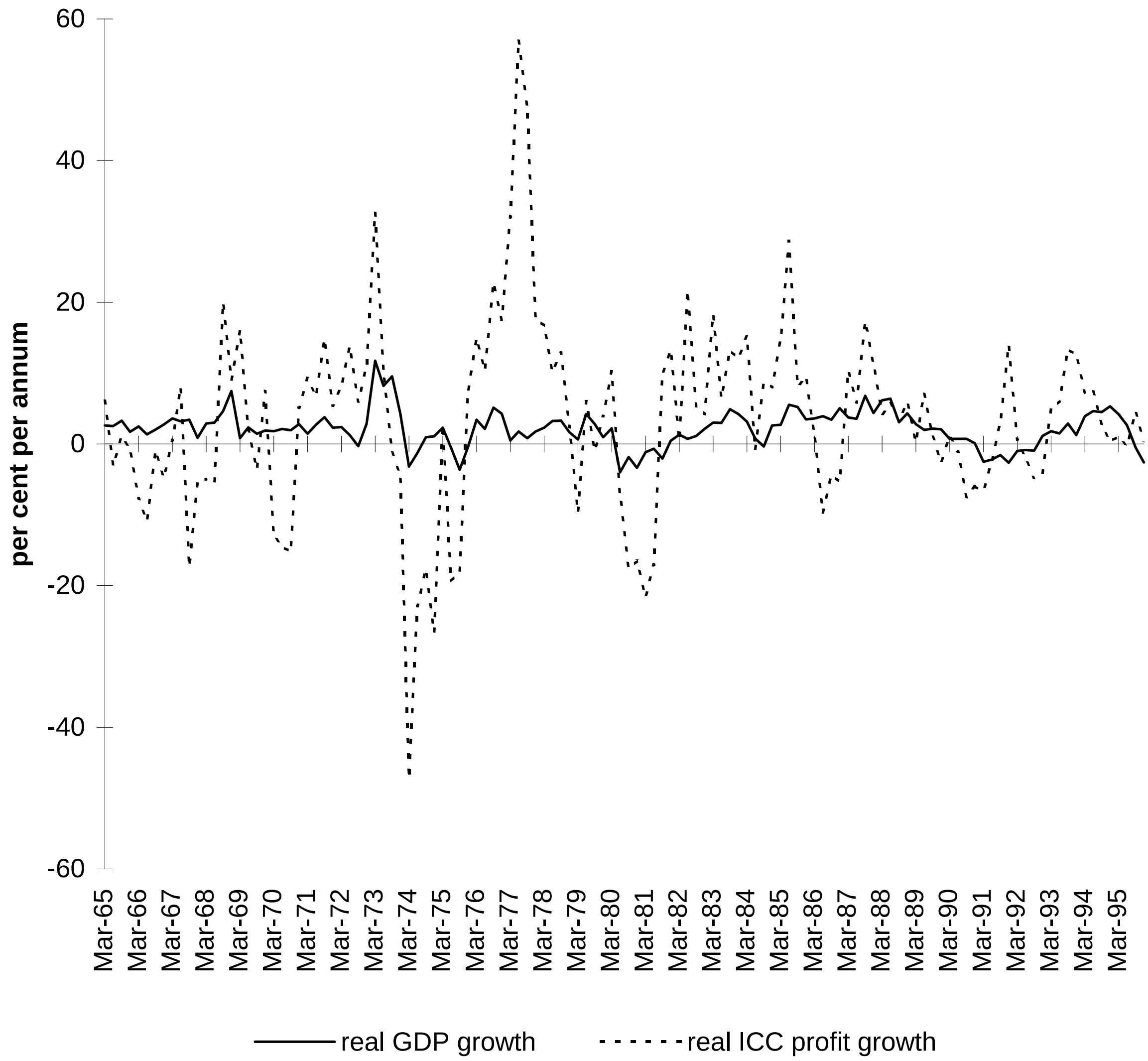
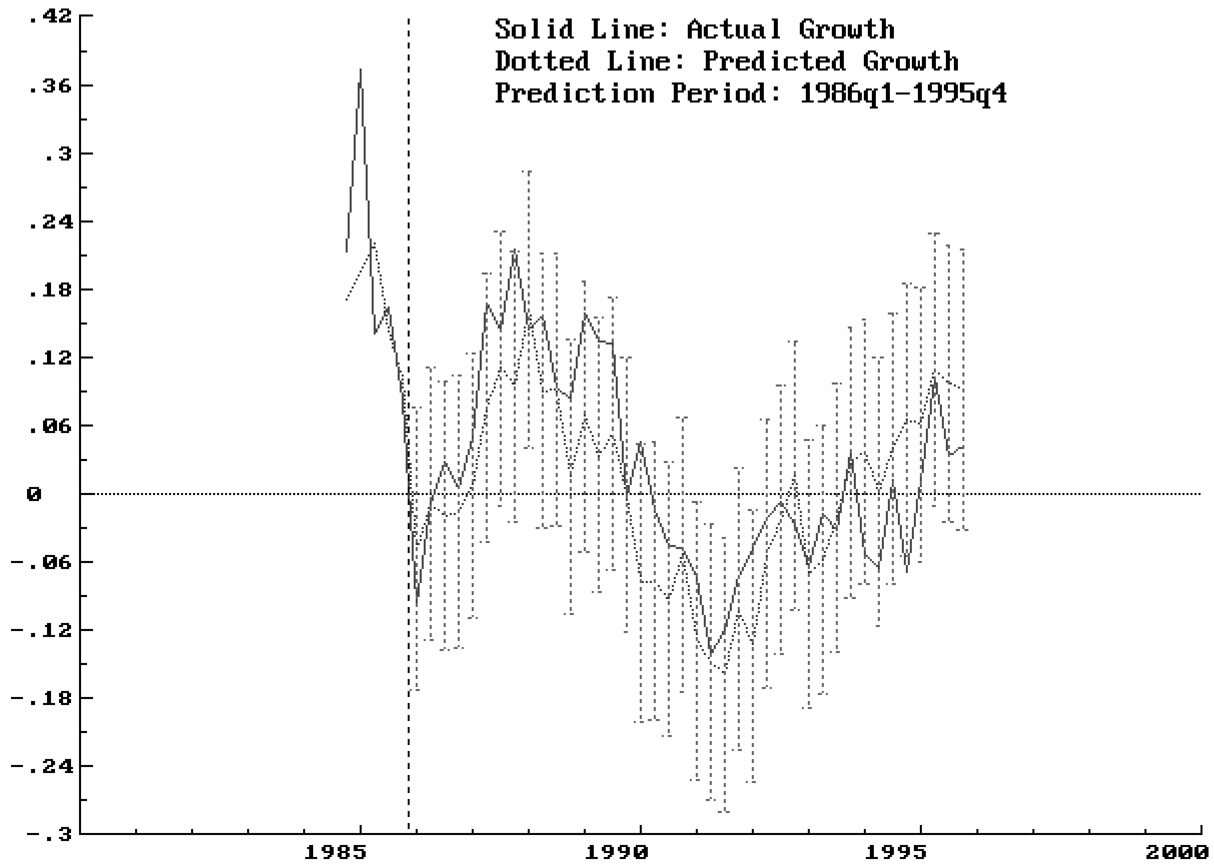


Figure 5 - Forecast of Investment Growth (Table 3 Column 2)



Appendix - Data Sources and Definitions

All series are for the period 1963:1 to 1995:4 and seasonally unadjusted.

Investment: Gross fixed capital formation in Industrial and Commercial companies (source ONS databank). The published data is only available in current price form and so it was deflated using the implicit deflator for private sector gross fixed capital formation, which is available in current prices and at constant 1990 prices.

Real GDP Real GDP at factor cost, 1990 prices (source ONS databank). Initially we calculated an output series for the ICC sector based on the methodology described in Anderson (1981), but we found that it was encompassed considerably by using real GDP as the measure of output.

Real profits: This is defined as post-tax gross ICC sector trading profits plus rent and non-trading income net of stock appreciation at 1990 prices. The starting point was gross ICC sector trading profits net of stock appreciation at current prices (source ONS Databank). Added to this was ICC sector rent and non-trading income (source ONS Databank), in order to obtain a measure of total income. A post-tax series was obtained by subtracting the published ICC taxes on income (payments) series. This latter series displays a very pronounced seasonal pattern with a pronounced seasonal peak in the first quarter of each year, indicating the high degree of coincidence of the timing of corporate tax payments. It was therefore converted to a centred four quarter moving average.

Interest Rate: 20 year long dated government bond yield, deflated by GDP deflator (source ONS databank). We initially calculated a cost of capital series, following King (1974), but found that it was outperformed by using a simple long-run real interest rate.

Gold Price: International gold price (quarterly average), converted to sterling using the sterling US/ $ exchange rate, and deflated by the GDP deflator (source Datastream)

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