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Monetary Instability, the Predictability of Prices, and the Allocation of Investment: An Empirical Investigation Using U.K. Panel Data

By PAUL BEAUDRY, MUSTAFA CAGLAYAN, AND FABIO SCHIANTARELLI*

A major goal of macroeconomic policy throughout industrialized countries is to achieve low and stable inflation. There are many reasons advanced in defense of this goal. One reason often given is that lowering and stabilizing inflation improves the informational content of the price system and thereby favors a more efficient allocation of resources.¹ In this view, price stability allows investments to be more effectively channeled towards projects with the highest returns since the best opportunities are more easily identified. Although such a belief is widely shared, to our knowledge it has not been subject to close empirical scrutiny.

In this paper, we propose a framework aimed at giving empirical content to the idea that monetary instability, through its effect on the informational content of prices, adversely affects the allocation of investment. To this end, we present a simple macro model to illustrate the effect of monetary instability on the distribution of investment when agents are imperfectly informed about the fundamentals of the economy. In particular, we show that as monetary policy becomes more predictable and, as a consequence, individual relative prices become easier to forecast, the cross-sectional distribution of

investment should widen. The reason is that better quality information should lead to a more unequal distribution of investment across firms as the market takes advantage of more precise knowledge of different investment opportunities. In contrast, when prices are hard to predict, we should observe less cross-sectional variations in investment rates. Therefore, our framework predicts a negative association between the cross-sectional variance of investment and price uncertainty.

Our empirical work exploits a panel data set covering a large number of quoted U.K. companies over the period 1970–1990. We first document that there were substantial variations in the cross-sectional distribution of investment over this period. More specifically, the distribution of investment rates narrowed significantly during the 1970's and widened anew in the second part of the 1980's. The macroeconomic history of the United Kingdom, characterized by greater turbulence in the 1970's with frequent changes in monetary policy, followed by a greater stability in the 1980's, provides some *prima facie* evidence in favor of our story. The goal of our empirical work is therefore to supplement this suggestive evidence with more formal econometric evidence. In particular, we use our panel data set to construct two different tests of the hypothesis that monetary uncertainty, through its effect on the informational content of prices, affect the cross-sectional distribution of investment. The first test is based on examining whether aggregate inflation uncertainty can explain the variation in the cross-sectional distribution of investment, both at the aggregate and at the industry level. The second test examines the model's prediction that there should be a negative association between the cross-sectional variance of investment and the cross-sectional variance of the log of the profit rate. The reason for this predicted negative association is again quite intuitive. By reducing the

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¹ See, for instance, Milton Friedman (1977).

information content of prices, monetary instability reduces the capacity of investment flows to equate profit rates across firms. Therefore, a lower cross-sectional variance in investment should be associated with a higher cross-sectional variance of profit rates. In both cases, we take care to examine whether real forces, such as changes in the price of oil, could be driving our results. We find that the overall evidence is quite supportive of the framework we propose to explain the relationship between nominal uncertainty, the informational content of prices, and the time variations in the cross-sectional distribution of investment.²

The remainder of the paper is organized as follows. In Section I, we extend the seminal model developed by Robert E. Lucas, Jr. (1973) to show why variations in the predictability of monetary policy should lead to a negative comovement between the cross-sectional variance of investment and nominal uncertainty. Section II contains the empirical analysis. In particular, Section II, subsection A documents the nature of the time variation in the cross-sectional distribution of investment in our data. Section II, subsection B reports anecdotal evidence on our hypothesis, while Section II, subsection C presents econometric evidence on the relationship between the cross-sectional variance of the investment rate, on the one hand, and either the conditional variance of the inflation rate or of the log profit rate, on the other. Section III concludes the paper.

I. Cross-Sectional Distribution of Investment and Price Uncertainty: A Simple Model

In this section we develop a model to illustrate how macroeconomic uncertainty can affect the allocation of investment through its effect on the informational content of prices. We focus on variations in the predictability of monetary

policy as a source of changes in the informational content of market signals. In particular we show how the efficient use of information implies that as prices become more predictable, the cross-sectional variance of the investment rate increases.

The environment we consider modifies the island model used by Lucas (1973) in a manner that emphasizes the implications for investment as opposed to employment. As in Lucas (1973), consider an economy with a continuum of competitive markets indexed by z . The demand for goods in market z depends on relative prices, real balances, and an idiosyncratic disturbance $\varepsilon_t(z)$, as given by (1):

$$(1) \quad y_t^d(z) = m_t - p_t - \gamma(p_t(z) - p_t) + \varepsilon_t(z)$$

where $\gamma > 0$. In equation (1), as elsewhere in the paper, lowercase letters denote logs. Variables in levels are denoted by uppercase letters. The aggregate price level in the economy, p_t , is defined by $\int_0^1 p_t(z) dz$ and m_t denotes the money supply. The supply of output in each market is determined by a representative firm, where production depends only on capital $k_t(z)$ as given in (2):

$$(2) \quad y_t^s = \theta k_t(z).$$

The representative firm in market z faces a one-period delivery lag and, therefore, determines next period's capital stock by setting the expected marginal product of capital equal to the real user cost, which we assume to be constant and we denote by c .³ When making such an investment decision, the firm does not know the real price that will prevail. However, the firm can form an expectation of the real price of good z based on its observation of the current price of the good and all past information on the economy (which we denote by Ω_{t-1}). Therefore, profit maximization implies that the $t + 1$ capital stock in each market is given by equation (3):⁴

² The emphasis on the informational content of prices that characterizes our paper is also the distinguishing feature of the theoretical contributions by Laurence Ball and David Romer (1993) and Mariano Tommasi (1994). For related empirical contributions, see also Ball and Stephen G. Cecchetti (1990), who provide evidence that high levels of inflation are accompanied by greater uncertainty about future inflation. Moreover, Tommasi (1996) finds that higher inflation is associated with greater difficulties in predicting relative prices.

³ Since we do not endogenize the real rental cost of capital, the model is best interpreted as a partial-equilibrium model or as a model of a small open economy.

⁴ This step assumes that prices are lognormally distributed, which in equilibrium will be true. A detailed derivation of the results contained in this section is available from the authors upon request.

$$\begin{aligned}
 (3) \quad k_{t+1}(z) &= \frac{1}{1-\theta} \times E[(p_{t+1}(z) \\
 &\quad - p_{t+1} - c) | p_t(z), \Omega_{t-1}] \\
 &\quad + \text{Var}_{z,t}(p_{t+1}(z) - p_{t+1}).
 \end{aligned}$$

Equation (3) states that, because of the existence of a delivery lag, next-period capital stock depends positively upon expectations of relative prices for $t + 1$, based on the information available at t . Equation (3) also indicates that the capital stock increases with the conditional variance of the real price of good z . This last effect reflects Jensen's inequality, but is not important for our results since in equilibrium this variance does not vary with z .

In order to solve for an equilibrium, we need to specify the properties of the two driving forces. First, the money-supply process is assumed to follow a possibly nonstationary autoregressive and heteroskedastic process, with innovations μ_t that are independent and normally distributed with mean zero and conditional variance τ_t^2 . The important element is that the money-supply process is characterized by time variation in its predictability since the conditional variance of μ_t is not assumed constant. The other driving forces in this economy are the relative demand disturbances, $\varepsilon_t(z)$. Each of these disturbances is assumed to be stationary and obey a first-order autoregressive process with autocorrelation parameter ρ , and innovation $v_t(z)$. These innovations are assumed to be independently and normally distributed with mean zero ($\int_0^1 v_t(z) dz = 0$) and constant variance, σ^2 . It is worth emphasizing that, because of delivery lags, the presence of serial correlation in $\varepsilon_t(z)$ is essential to the model, since it renders current information relevant for predicting next-period relative prices and, hence, for determining investment decisions.

An equilibrium for this economy is characterized by a pair of market-specific stochastic processes for capital and prices, such that: (a) given the allocation of capital, prices ensure the equality of supply and demand in each market; and (b)

given prices, the allocation of capital satisfies equation (3). Using the method of undetermined coefficients, and omitting the constant term, it is straightforward (but somewhat tedious) to show that in equilibrium the capital stock satisfy equation (4):⁵

$$\begin{aligned}
 (4) \quad k_{t+1}(z) &= \phi_1 \varepsilon_t(z) + (\phi_{2,t} - \phi_1) v_t(z) + \phi_{2,t} \gamma \mu_t,
 \end{aligned}$$

where $\phi_1 = [\rho/\theta + (1 - \theta)\gamma]$ and $\phi_{2,t} = [\rho\sigma^2/(\sigma^2 + \tau_t^2\gamma^2)\gamma(1 - \theta) + \theta\sigma^2]$.

Equation (4) states that the capital stock depends both on real factors, affecting relative demand, and on monetary innovations. The first element to note from equation (4) is that the sensitivity of capital expenditures to monetary innovations depends, through $\phi_{2,t}$, on the stability of monetary policy. In particular, an increase in monetary instability, as represented by an increase in τ_t^2 , leads to a fall in $\phi_{2,t}$, and hence in the size of the effect of monetary disturbance on the capital stock. This reflects the fact that when monetary policy is less predictable, firms interpret observed increases in demand as reflecting mainly monetary factors. The second and more important aspect to note from (4) is that an increase in monetary instability also leads firms to adjust less to innovations in relative demand. Hence, when monetary policy is more unstable, productive capacity will be less effectively targeted towards the sectors with high demand. It is this latter feature of the model that captures nicely the idea that monetary instability may reduce the efficient allocation of investment by reducing the informational content of prices.

In order to render the above observations empirically testable, we exploit the implication of equation (4) for the variance of the cross-sectional distribution of the investment rate. Using the fact that the investment rate can be approximated as the log difference of the capital stock, equation (4) implies that the cross-

⁵ Note that if $\rho = 0$, the capital stock equals a constant, which highlights the importance of persistence in $\varepsilon_t(z)$.

sectional variance of the investment rate can be written as:

$$\begin{aligned}
 (5) \quad & \text{Var}_z(I_{t+1}(z)/K_t(z)) \\
 &= \phi_1^2 \rho^2 \frac{(1-\rho)^2}{(1-\rho^2)} \sigma^2 + \phi_{2,t}^2 \sigma^2 \\
 &+ (\phi_{2,t-1} - \phi_1 \rho)^2 \sigma^2.
 \end{aligned}$$

From equation (5), it can be seen that cross-sectional distribution of the investment rate is related to both the variance of real shocks and the variance of monetary disturbances (through $\phi_{2,t}$). For our purpose, it is the effects of the time variation in the variance of monetary disturbances τ_t^2 which is of interest since it is this variance that reflects the effects of monetary instability. In fact, equation (5) implies that the cross-sectional variance of the investment rate depends upon the contemporaneous and once-lagged variance of monetary innovations. More specifically, an increase in τ_t^2 leads to a decrease in the variance of the investment rate, while the effect of τ_{t-1}^2 is ambiguous. As long as the persistence of the real shock, as captured by ρ , is not too large, the effect of τ_{t-1}^2 will also be negative. Therefore, the main insight we draw from (5) is that, when monetary policy becomes more stable and the predictability of prices improves, the cross-sectional distribution of the rates of investment should widen. This arises because investment is distributed more unevenly across firms as each of them responds more accurately to differences in demand conditions.

In order to implement an empirical test of this prediction, it is useful to consider the linear approximation of equation (5) as given below:

$$\begin{aligned}
 (6) \quad & \text{Var}_z\left(\frac{I_{t+1}(z)}{K_t(z)}\right) \\
 &= \beta_0 + \beta_1 \sigma^2 + \beta_2 \tau_t^2 + \beta_3 \tau_{t-1}^2
 \end{aligned}$$

where $\beta_2 < 0$. Equation (6) states that holding σ^2 and τ_{t-1}^2 constant, an increase in τ_t^2 is associated with a reduction in the cross-sectional variance of the investment rate.

In order to further highlight the implications of (6), it is useful to use the fact that in equilibrium τ_t^2 is equal to the conditional

variance of inflation.⁶ Therefore, we can replace τ_t^2 and τ_{t-1}^2 in (6) by the conditional variance of inflation, denoted by $\text{Var}_{t-1}(\pi_t)$. The resulting relationship is given by (7).

$$\begin{aligned}
 (7) \quad & \text{Var}_z\left(\frac{I_{t+1}(z)}{K_t(z)}\right) \\
 &= \beta_0 + \beta_1 \sigma^2 + \beta_2 \text{Var}_{t-1}(\pi_t) \\
 &+ \beta_3 \text{Var}_{t-2}(\pi_{t-1}).
 \end{aligned}$$

The advantage of equation (7) over equation (6) is that it directly relates investment behavior to a measure of the predictability of inflation. Moreover, equation (7) bypasses the need to select the appropriate monetary aggregate when examining the predictions of this model.⁷ For these reasons, equation (7) plays a central role in our empirical investigation.

We now want to highlight a second implication of this model, which in the empirical section will allow us to further exploit information contained in the firm-level panel data set. In particular, recall that in response to increased nominal uncertainty, investments will be less efficiently channeled towards the most profitable opportunities. Correspondingly, the model predicts that increased nominal uncertainty should lead to an increased dispersion of the *ex post* profit rates since investment flows are *de facto* less likely to equalize profit rates. Hence, these two observations imply that an increase in the cross-sectional variance of the *ex post* profit rate

⁶ This can be easily seen once it is noticed that the equilibrium price process is given by:

$$\begin{aligned}
 p_t(z) = & m_t + \frac{1}{\gamma} \varepsilon_t(z) - \frac{\theta}{\gamma} \phi_1 \varepsilon_{t-1}(z) \\
 & - \frac{\theta}{\gamma} (\phi_{2,t-1} - \phi_1) v_{t-1}(z) - \theta \gamma \phi_{2,t-1} \mu_{t-1}
 \end{aligned}$$

and that the mean of $\varepsilon_t(z)$ across markets equals zero.

⁷ Given the changes in methods of monetary control and in the financial structure in the United Kingdom over the 20-year period covered in our investigation, it would be next to impossible to select a money aggregate which has been used continuously as a policy target and that bears a constant relationship with demand.

should be associated with a decrease in the variance of the investment rate. To be more precise, the model implies that in equilibrium the cross-sectional variance of the logarithm of the profit rate, is given by equation (8):⁸

$$(8) \quad \text{Var}_z \left(\log \frac{R_{t+1}(z)}{K_{t+1}(z)} \right) = \frac{\sigma^2}{\gamma^2} + \left(\frac{\rho}{\gamma} - \phi_{2,t} \left(\frac{\theta}{\gamma} + (1 - \theta) \right) \right)^2 \sigma^2.$$

In (8), $R_{t+1}(z)$ denotes operating profits defined as real revenues before capital costs. The main element to notice from equation (8) is that, again through $\phi_{2,t}$, the cross-sectional variance of the profit rate will be positively related to monetary instability. In other words, the partial derivative of $\text{Var}_z(\log R_{t+1}(z)/K_{t+1}(z))$ with respect to τ_t is positive. Therefore, using equation (8), equation (6) can be rewritten to express (as a linear approximation) the relationship between the cross-sectional variance of the investment rate and the cross-sectional variance of the log profit rate. This relationship is given by equation (9):

$$(9) \quad \text{Var}_z \left(\frac{I_{t+1}(z)}{K_t(z)} \right) = \beta'_0 + \beta'_1 \sigma^2 + \beta'_2 \text{Var}_z \left(\log \frac{R_{t+1}(z)}{K_{t+1}(z)} \right) + \beta'_3 \text{Var}_z \left(\log \frac{R_t(z)}{K_t(z)} \right),$$

with the prediction that $\beta'_2 < 0$. Equation (9) indicates that an increase in the cross-sectional variance of the profit rate should be associated with a contemporaneous decrease in the cross-sectional variance of the investment rate, since, as the informational content of prices goes down, investment flows should be less effective

⁸ This follows from the equilibrium price equation (see footnote 5) and from

$$\begin{aligned} \log R_{t+1}(z)/K_{t+1}(z) \\ = p_{t+1}(z) - p_{t+1} - (1 - \theta)k_{t+1}(z). \end{aligned}$$

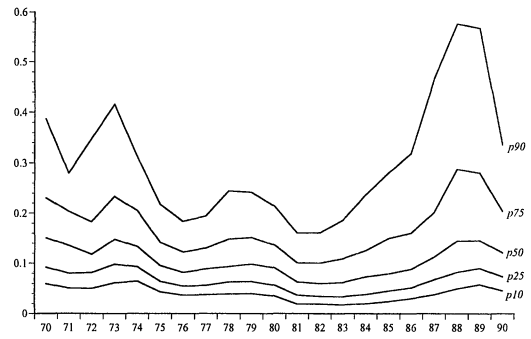


FIGURE 1. INVESTMENT RATE FOR THE FULL SAMPLE, ALLOWING FOR ENTRY AND EXIT

Notes: The lines (from top to bottom) represent the 90th, 75th, 50th, 25th, and 10th percentiles, denoted respectively by p90, p75, p50, p25, p10.

in equating profit rates across firms. This prediction is interesting also from a methodological point of view, since it provides an example where a model’s prediction relates to comovements between cross-sectional variances.

II. Empirical Results

The model presented in Section I suggests the existence of a negative relationship between nominal uncertainty and the cross-sectional variance of the investment rate. This section contains the empirical evidence on this hypothesis. The analysis is based on an unbalanced sample of 988 quoted U.K. companies in the manufacturing sector for which complete and consistent balance-sheet data are available for at least three consecutive years. The source of the data is the Datastream file that contains accounting information for quoted U.K. companies.

A. Trend and Cyclical Distributional Changes of the Investment Rate

We begin by discussing the empirical evidence on the evolution of the cross-sectional distribution of U.K. firms’ investment rate over the period 1970–1990. The objective of this exercise is to document both the trend and cyclical movements that occurred during the 1970’s and the 1980’s. In order to provide a visual summary of the distributional changes that have occurred, we display in Figure 1 the

first decile, the first quartile, the median, the third quartile and the ninth decile of the investment rate.⁹

Three main features are revealed. First, concentrating on trend movements, it is evident that investment rates decreased during the 1970's. Second, the decrease was, however, greater for the ninth decile and for the third quartile, so that we detect a substantial narrowing from the top of the distribution of investment rate over the 1970's. Both the ninth decile and the third quartile dropped by approximately 45 percent comparing the investment rate in 1970 with that of 1980. During the 1980's we observe the opposite phenomenon with the dispersion of the distribution of the degree of leverage gradually increasing starting from 1983 and, more dramatically, between 1986 and 1988. This is because of a faster increase in the investment rate for the ninth decile, and also for the third quartile than that of the median up to 1988. Comparing 1981 with 1988, we observe that the ninth decile and the third quartile increased, respectively, by approximately 260 percent and 185 percent. Finally, at a cyclical frequency, we observe a decrease in investment rates for firms in periods corresponding to most of the recessionary episodes, leading to a decrease in dispersion. This is true for the recession of 1971, 1974–1975, 1980–1982 and the recession starting in 1990.

In order to make certain that these are genuine changes and do not only reflect the variation in the composition of the sample, we have partitioned the sample into two subperiods: 1970–1980 and 1980–1990. For each subperiod, we have analyzed the change in the percentiles for different subsamples that differ with respect to the conditions imposed on firms' entry and exit in the panel. In the first sample, we allow for both entry and exit. In the second, we allow entry but not exit by choosing only the firms

that remained in the panel until the final year of each subperiod. In the third sample, we do not allow entry but we allow exit choosing only the firms that were in the panel in the initial year of each subperiod. Finally in the fourth sample, we allow neither entry nor exit by considering firms with observations for all the years in each subsample.¹⁰ Regardless of the restrictions we imposed on entry and exit, we found a marked reduction in dispersion in the 1970's for all subsamples. Perhaps it would not be surprising to see such reduction in dispersion in a sample that allows firms to exit, since as the time passes surviving firms become more alike. However, we obtain the same results when we do not allow for exit. Similarly, the widening of the distribution between 1987 and 1989 was found to be robust to the treatment of entry and exit of firms. However, the increase in dispersion from 1984 to 1988 is more marked for the samples that allow for new entrants, compared to the samples in which entry is not allowed. The decrease in dispersion starting in 1989 is also more pronounced. This says that new entrants effect the dispersion of investment rates in an important way. For this reason, in our econometric work we restrict our samples to firms that are observed for at least seven consecutive years, thereby assuring that no entry occurs after 1983.

B. Anecdotal Evidence

Our model predicts that a more unstable and uncertain macroeconomic environment, because of its adverse effect on the informational content of prices, should lead to a cross-sectional distribution of the investment rate characterized by less dispersion.¹¹ We will start

¹⁰ Note that, given the nature of Datastream, entry and exit in and from the panel cannot be equated with the birth and death of a firm.

¹¹ There are other explanations of the time-series evolution of the cross-sectional distribution of investment, besides the one we have provided. Potential candidates include explanations based on asymmetric information and agency costs, or on fixed adjustment costs. For instance, it may be argued that in bad times there is a flight to quality away from riskier firms (as in Ben Bernanke et al., 1996) and this may imply a decrease both in the mean value of the investment rate and a compression from the top of the distribution. Alternatively, times of higher aggregate nominal uncer-

⁹ The investment rate is calculated by dividing investment expenditure (in constant prices) by the replacement value of the (end of previous period) capital stock (also in constant prices). Investment expenditure on fixed assets in current prices is available in Datastream. The replacement value of the capital stock in current prices is obtained using the perpetual inventory method (see Richard Blundell et al. [1992] for details). To go from current to constant value (real) figures for investment and capital we have used industry-specific price indices for investment.

from an anecdotal approach and argue that the main macroeconomic developments in the United Kingdom during the periods covered by our panel are indeed consistent with the changes in the cross-sectional distribution of investment. The main theme we want to emphasize is that the 1970's were characterized by a high and increasing degree of macro volatility while the 1980's saw a return to a more stable environment, particularly with regard to the inflation rate. Clearly the two oil price shocks were very important in giving rise to the initial upward pressure on prices in the 1970's. However, the increase in inflation is intimately related to the macroeconomic policy response adopted by the government in its attempt to keep unemployment from rising in the context of wage resistance by unions. We focus on this policy response and, in particular, on the consequences of an increase in the uncertainty that characterizes such response in the wake of the adverse supply shocks.

The United Kingdom was one of the countries for which the adjustment to the adverse supply shocks of the 1970's was particularly difficult. The annual rate of inflation in the United Kingdom was in the double digits for most of the 1970's and exceeded 20 percent in 1975. Moreover, the inflation rate in the 1970's was not only higher on average than in the 1980's but also more variable. Neither the policies of the Conservative governments, between 1970–1974, nor the Labour governments, between 1974–1979, were successful in controlling inflation, although different strategies were attempted, including statutory and voluntary income policies during some of the periods. Over this period there were frequent changes of economic policy strategies by both sets of administrations, which may have contributed to uncertainty about the evolution of prices and other macro aggregates. The overall lack of permanent success in the struggle to control inflation was due in part to the real wage resistance exhibited by the unions, to which income

policy brought only temporary relief. It also reflected the postwar political consensus in the attempt by the governments to maintain high levels of employment. This commitment, shared, for the most part, by all political parties, gave both fiscal and monetary policy a stop-and-go quality as the governments tried to navigate among slowdowns in economic activity with the attending unemployment problem, the occurrence of balance-of-payment crises, and the resurgence of inflation. We have already commented about inflation. The two main contractionary episodes followed the oil price shocks at the end of 1973 and in 1979 and the contemporaneous slowdown of world trade. There was also a significant contraction in 1971.

The victory by the Conservative Party in the 1979 elections changed the rules of the game in terms of economic policy. Following the collapse of the previous Labour government's income policy during the winter of discontent of 1978–1979, the new Conservative government led by Mrs. Thatcher pursued a restrictive monetary and fiscal policy, that in addition to the effect of the oil price increase and the slowdown in world demand, created a very severe contraction that lasted until 1982. It has been debated how restrictive monetary policy actually was during that period, since the growth rate of M3 often overshot the target ranges set by the government. However, the behavior of narrower money aggregates, like M0, and the appreciation of the pound sterling suggests that monetary policy was indeed tight. Moreover, the fiscal policy stance was also contractionary, and indeed the control of the Public Sector Borrowing Requirement was the main instrument through which the government attempted to reduce the rate of growth of money.¹² Whatever the final judgment on how the initial contraction was achieved, it is clear that the government eventually succeeded in establishing its credibility in the anti-inflation strategy.

The willingness to live with high levels of unemployment was a departure from the postwar consensus and, together with legislation

tainty could also be times of higher idiosyncratic uncertainty. In the presence of fixed cost, this may lead to hesitation on the part of firms to invest and to a low variance of the investment rate. See Ricardo J. Caballero (1992) and Caballero and Eduardo M.R.A. Engel (1991) for models with such nonconvexities.

¹² See Willem H. Buiter and Marcus Miller (1981, 1983), Kent G. P. Matthews and Patrick Minford (1987), and Charles Bean and James Symons (1989) for an evaluation of Margaret Thatcher's policies.

aimed at decreasing the power of the unions, succeeded in achieving low and steady inflation rates. By 1982 the rate of inflation was down to 5.4 percent and reached 3 percent in 1986. The years between 1982 and 1989 were a period of steady expansion in output. The evolution of the money growth rate suggests that monetary policy was not particularly expansionary for most of the 1980's. Finally, the period of expansion came to a close at the end of the 1980's and was followed by the beginning of another recession, starting in 1990, that was reflected in a negative growth rate of GDP for manufacturing in that year. The change in the volatility and predictability of the economic environment that occurred over the 1970's and 1980's, as suggested by our brief summary of the macroeconomic developments in the United Kingdom, fits together nicely with the observed variations in the dispersion of the investment rate if one is willing to accept that the overall turbulence is associated, at least in part, with monetary instability.

C. Econometric Evidence

This anecdotal approach provides some prima facie evidence that is consistent with the implications of our simple model. In order to provide more formal evidence, we will first estimate equation (7), which relates the cross-sectional variance of investment to the conditional variance of aggregate inflation.¹³ In our empirical application, we have used the log difference of the monthly CPI as a measure of aggregate inflation. Formal testing and estimation by Maximum Likelihood of models that allow for a time-varying conditional heteroskedasticity, over the period 1961–1990, suggest that the latter indeed changes over time and that an ARCH(1) model captures such changes in a simple and adequate way. A summary of the

¹³ John Huizinga (1993) presents empirical evidence for the United States on the relationship between inflation and real price uncertainty, and on their effect on the level of aggregate investment. Gary Ramey and Valerie A. Ramey (1995) analyze, using aggregate country-level panel data, the impact of output volatility and uncertainty on the GDP growth rate and on the investment share of GDP. The emphasis in our paper is, instead, on the effect of uncertainty on the firm-level cross-sectional variance of investment.

TABLE 1—CONDITIONAL VARIANCE OF THE INFLATION RATE: ARCH(1) MODEL

α_0	α_1	Lagrange Multiplier test	Number of observations
0.139 (0.016)	0.557 (0.104)	4.81	348

Notes: The model is $\Delta \log CPI_t = \beta_0 + \sum_{i=1}^{12} \beta_i \Delta \log CPI_{t-i} + u_t$, and $h_t = \alpha_0 + \alpha_1 u_{t-1}^2$. It is estimated using Maximum Likelihood. Standard errors are in parentheses. The estimation period is January 1962–December 1990. Monthly dummies are also included in the equation. The Lagrange Multiplier test for ARCH(1) effects is distributed as $\chi^2(1)$.

econometric results are reported in Table 1. Our estimated model for inflation is: $\Delta \log CPI_t = \beta_0 + \sum_{i=1}^{12} \beta_i \Delta \log CPI_{t-i} + u_t$, where u_t equals $\varepsilon_t \sqrt{h_t}$, ε_t is a zero mean, unit variance independently and identically distributed (i.i.d.) process, and $h_t = \alpha_0 + \alpha_1 u_{t-1}^2$ denotes the conditional variance. The standard Lagrange Multiplier (LM) test cannot reject the presence of ARCH(1) effects. Additional lags of u_t^2 were not found to contribute significantly to the variance, when entered unrestrictedly. Similar conclusions are reached from the subsidiary regressions used to calculate the LM test for ARCH(p) effects.

The estimated conditional variances for inflation suggest that the 1970's were turbulent years compared to the 1980's. These results complement those obtained for U.K. inflation by Robert F. Engle (1982) who finds significant ARCH effects during the 1960's and 1970's. Figure 2 plots both the cross-sectional variance of investment and our estimate of the lagged conditional variance of inflation, \hat{h}_{t-1} , against time as to illustrate the negative comovement between these two variables (their correlation coefficient equals -0.41 with a standard error of 0.2).

Another more complex model for the $\Delta \log CPI$ that appears to be reasonably consistent with the data is an LGARC(1,1) model. Note that we also found similar ARCH effects when the quarterly implicit GDP deflator is used as the measure of aggregate price. Since the results from estimating equation (7) are very similar when using the estimated time-varying conditional variance from any of the three possible models mentioned above, we present only those

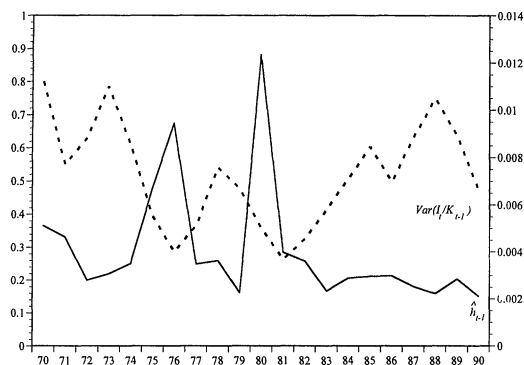


FIGURE 2. CONDITIONAL VARIANCE OF INFLATION AND CROSS-SECTIONAL VARIANCE OF THE INVESTMENT RATE

Notes: The continuous line denotes the lagged conditional variance of inflation, \hat{h}_{t-1} (left axis). The broken line denotes the cross-sectional variance of the investment rate, denoted by $\text{Var}(I_t/K_{t-1})$ (right axis).

based on the simpler ARCH(1) model for $\Delta \log CPI_t$. Moreover, as the model for inflation is estimated using monthly data, we have taken 12-month averages of the estimated conditional variances.

In Table 2, column (1) we present the results of the OLS regression of the cross-sectional variance of the investment rate on the contemporaneous and once-lagged estimated average conditional variance of inflation, denoted by \hat{h}_t and \hat{h}_{t-1} respectively.¹⁴ In calculating the variance of the investment rates we have used only the firms with at least seven years of observations, which precludes entry in the sample after 1983.¹⁵ In this and the following regressions, we have allowed for an intercept shift after 1980. This is meant to capture any decade long shifts in the variance σ^2 .

The important result to note for our purposes is that the coefficients of \hat{h}_t and \hat{h}_{t-1} are both negative and are jointly significantly different

¹⁴ The error term in the estimated equation is meant to capture changes in σ which are uncorrelated with monetary uncertainty and which are not captured by the additional regressors we include in our robustness checks.

¹⁵ This leaves us with 729 firms. We have deleted firm-year observations that were outliers in terms of the investment rate or the profit rate (more specifically observations for the investment rate above the 97.5 percentile and observations for the profit rate below the 2.5 percentile and above the 97.5 percentile). The results are robust to changes in the cutoff points.

TABLE 2—TIME-SERIES RELATIONSHIP BETWEEN THE CONDITIONAL VARIANCE OF INFLATION AND THE CROSS-SECTIONAL VARIANCE OF THE INVESTMENT RATE

Regressors	(1)	(2)	(3)	(4)
\hat{h}_t	-0.005 (0.002)	-0.005 (0.002)	-0.005 (0.002)	-0.005 (0.002)
\hat{h}_{t-1}	-0.007 (0.002)	-0.007 (0.002)	-0.007 (0.002)	-0.004 (0.002)
1970's dummy	-0.001 (0.001)	-0.001 (0.001)	-0.001 (0.001)	-0.001 (0.006)
Percentage change in oil price	—	0.005 (0.005)	—	—
Conditional variance of percentage change in oil price	—	—	6.8e-6 (1.6e-4)	4.0e-5 (1.4e-4)
Lagged dependent variable	—	—	—	0.352 (0.174)
$F_{2,n}$	9.00	8.31	7.99	6.60
[p-value]	[0.002]	[0.003]	[0.004]	[0.009]
\bar{R}^2	0.44	0.44	0.41	0.56
Standard error of the regression	0.0016	0.0016	0.0012	0.0014
Durbin-Watson test	1.25	1.14	1.22	1.77
Lagrange Multiplier test	—	—	—	0.16
Number of observations	21	21	21	20

Notes: Dependent variable: cross-sectional variance of investment rate, $\text{Var}_{z_t}(I_{t+1}(z)/K_t(z))$. \hat{h}_t : Yearly average of the conditional variance of inflation from a monthly ARCH(1) model of $\Delta \log CPI_t$. All equations contain a constant. Estimation period: 1970–1990. Standard errors of coefficients are in parentheses. The F -statistic corresponds to the test of joint significance of \hat{h}_t and \hat{h}_{t-1} (the degrees of freedom of the denominator differ across columns). The Lagrange Multiplier test is for first-order serial correlation, and is distributed as $\chi^2(1)$.

from zero, with a marginal significance level smaller than 1 percent. Even though the conditional variances are generated regressors, the coefficient estimates are consistent and the joint test of significance is perfectly valid (while the individual t -statistics are not).¹⁶ Although the theory presented in Section I has strong predictions mainly for the coefficient associated with \hat{h}_t , the estimates presented in Table 2 suggest that lagged nominal uncertainty also has a

¹⁶ See Adrian R. Pagan (1984). The appropriateness of the F -test for $H_0: \beta_2 = 0, \beta_3 = 0$ follows from Theorem 1, (ii). See also Pagan [1986 Theorem 4, (i)].

negative effect. To gain some intuition about the magnitude of the effects, note that these estimates imply that a 50-percent increase in the (conditional) standard deviation of the inflation rate generates approximately a cumulative 25-percent increase in the standard deviation of the investment rate after three periods.

It is immediately relevant to ask whether the results reported in column (1) of Table 2 reflect nominal uncertainty or, alternatively, could more reasonably be interpreted as reflecting time variation in real forces (that is, changes in σ^2 in the context of our model) which are not controlled for in the regression.¹⁷ For example, during this time period, the price of oil changed dramatically and varied in predictability. However, such changes are likely to have caused the efficient distribution of investment to widen, since it is efficient to distribute investment more unevenly when there is more divergence in profit opportunities. In any case, in order to see whether such forces may be driving the observed negative relationship between the variance of investment and the conditional variance of inflation, in columns (2) and (3) of Table 2 we add respectively as regressors the percentage change in the real price of oil or the conditional variance of the price of oil [estimated from an ARCH(1) model, using quarterly data].¹⁸ As can be seen from the table, these additional regressors do not reduce the strength of the observed relationship between the variance of investment and nominal uncertainty. It is also of interest to examine whether this observed negative relationship is robust to allowing for alternative dynamic specifications of equation (7). Although the theory developed in Section I suggests the need to include only two lags of the conditional variance of inflation, extensions of the model which introduce frictions in the adjustment of the capital stock

would most likely imply a richer dynamic specification.¹⁹ Moreover, the Durbin-Watson test for the models in columns (1)–(3) are somewhat low and lie in the uncertainty region. Hence, we have experimented with a richer dynamic specification. As an example we report in column (4) of Table 2, the results obtained when we include the lagged dependent variable in the model with the conditional variance of oil prices. The results suggest that the lagged dependent variable has additional explanatory power. However, the significant negative relationship between the cross-section variance of investment and our measure of nominal uncertainty appears robust to allowing a more flexible dynamic specification. This conclusion also holds when we add a further lag of the conditional variance of inflation as an additional regressor.

Obviously, the results presented in Table 2 can be criticized on the grounds of their exclusive time-series nature and the limited number of observations. In particular, over the 1970's and 1980's, there has been substantial changes in the real forces governing investment both within and across industries and these changes are difficult to control for in a pure time-series framework. In order to address this issue, in Table 3 we present panel regression results where the unit of observation is the industry. This approach allows us to better control for heterogeneity in the effect of additional regressors (such as oil prices), and to test the theory at a different level of aggregation, by focusing on the intrasectoral distribution of investment. Note that in terms of the theory set out in Section I, the relationship given by equation (7) should hold at the industry level as well as the aggregate level, with the only difference being that σ^2 should now be indexed by industry and interpreted as the within-industry variance in real demand.

¹⁷ The model derived in Section II can be extended to allow for changes in the real variance, σ^2 . In such case it can be shown that the conditional variance of the investment rate will depend on an infinite sum of past values of the real variance.

¹⁸ We have experimented with different timing for the conditional variance of the rate of growth in the real oil price (or the rate of growth itself). The conclusions are identical, whether its contemporaneous or the lagged value (or both) are included. We report only the former results.

¹⁹ For example, we explored the case where capital is assumed to adjust only gradually to its optimal level according to the stock-adjustment principle, and where all individuals take this into account when making inferences. In this case, we found that the cross-section variance of the investment rate depends on an infinite moving average of the conditional variance of inflation; with the maintained prediction that the once-lagged condition variance of inflation negatively effects the cross-sectional distribution of investment.

TABLE 3—TIME-SERIES RELATIONSHIP BETWEEN THE CONDITIONAL VARIANCE OF INFLATION AND THE CROSS-SECTIONAL VARIANCE OF THE INVESTMENT RATE AT THE INDUSTRY LEVEL

Regressors	(1)	(2)	(3)	(4)	(5)	(6)
\hat{h}_t	-0.005 (0.001)	-0.005 (0.001)	-0.006 (0.001)	-0.005 (0.001)	-0.003 (0.001)	-0.003 (0.002)
\hat{h}_{t-1}	-0.004 (0.001)	-0.004 (0.001)	-0.004 (0.001)	-0.004 (0.001)	-0.003 (0.001)	-0.003 (0.001)
Lagged dependent variable	0.271 (0.069)	0.240 (0.074)	0.275 (0.078)	0.271 (0.072)	0.296 (0.070)	2.58 (0.071)
Industry dummies interacted with:						
Percentage change in oil price	No	Yes	No	No	No	No
Conditional variance of percentage change in oil price	No	No	Yes	No	No	No
Percentage change in user cost of capital	No	No	No	Yes	No	No
Growth rate of GDP	No	No	No	No	Yes	No
Inflation rate	No	No	No	No	No	Yes
$F_{2,n}$	14.12	13.59	13.85	13.69	4.01	3.44
[p -value]	[0.000]	[0.000]	[0.000]	[0.000]	[0.002]	[0.034]
\bar{R}^2	0.26	0.31	0.29	0.28	0.34	0.31
Standard error of the regression	0.0030	0.0031	0.0031	0.0031	0.0030	0.0031
Number of observations	180	180	180	180	180	180

Notes: Dependent variable: cross-sectional variance of investment rate at the industry level (nine industrial sectors). \hat{h}_t : Yearly average of the conditional variance of inflation from a monthly ARCH(1) model of $\Delta \log CPI_t$. Estimation period: 1970–1990. Standard errors of coefficients are in parentheses. The F -statistic corresponds to the test of joint significance of \hat{h}_t and \hat{h}_{t-1} (the degrees of freedom of the denominator differ across columns). Industry dummies and a 1970's dummy are included in all regressions.

Table 3 reports results for the fixed-effect regression of the cross-sectional variance in industry-level investment on the conditional variance of inflation. Our data set allows us to group observations into nine industries, which makes for a total of 189 observations. For space sake, we only present results for the more demanding specifications which include the lagged dependent variable. In all cases we impose the restriction, which is not rejected by the data, that the coefficients of \hat{h}_t and \hat{h}_{t-1} are the same across industries. This should allow us to estimate the effect of the conditional variance of inflation more precisely. As can be seen in column (1), the negative relationship between the cross-sectional variance of investment and the conditional variance of inflation appears to be a phenomenon that also holds at the industry level. In the following columns we add, in order, the percentage change in oil prices or their conditional variance, the percentage change in the user cost of capital, the GDP growth rate,

and the aggregate inflation rate. Note that we allow changes in the variables to have different impacts across industries in the hope of better controlling for industry-specific responses to real forces. However, imposing equal responses does not alter our basic conclusions.

We find once again that controlling for oil price shocks [see columns (2) and (3)] has no substantial impact on this relation. This last result strongly confirms that variations in oil prices are unlikely to be the reason why we observe a negative relationship between the conditional variance of inflation and the variance of investment. In column (4) we control for changes in the user cost of capital. Our measure of the user cost of capital allows for changes in tax parameters and is constructed with reference to an institutional investor that uses retention as a source of finance. Since the real user cost of capital varied considerably over this period, it could potentially be an important factor in explaining movement in the cross-

sectional distribution of investment. However, controlling for industry-specific effects of the user cost of capital has very little effect on the observed negative relationship between the variance of investment and the conditional variance of inflation. In essence, the same holds true when allowing for industry-specific effects of the growth rate of real GDP. This latter case is of particular interest since one can easily conjecture scenarios where a faster-growing economy is associated with both more stable prices and a more unequal distribution of investment. Moreover, changes in GDP may be correlated with changes in the real variance. Nevertheless, the estimates in column (5) suggest that such a story is not enough in and of itself to explain our results, that is, inflation uncertainty appear to affect the distribution of investment above and beyond the effect accounted for by GDP growth.²⁰ Finally in column (6) we control for the inflation rate (measured as $\Delta \log CPI_t$). Not surprisingly, the inflation rate and the conditional variance of inflation tend to move together and consequently our estimated effect of the conditional variance of inflation on the cross-sectional distribution of investment is somewhat weaker in this case. However, the test of the joint significance of \hat{h}_t and \hat{h}_{t-1} suggests that the hypothesis that their coefficients are jointly zero can be rejected at a 5-percent significance level.

Overall, the results so far provide considerable support for the hypothesis that aggregate price uncertainty, through its effect of reducing the informational content of relative prices, may adversely effects the process of investment allocation. In order to further examine the plausibility of this view, we now analyze the relationship between the cross-sectional variance of the investment rate and the cross-sectional variance of the log profit rate, as suggested by equation (9).

²⁰ If we include as a regressor the industry-specific mean investment rate or the real return on equity, measured as the difference in the change in the log of the stock market index and $\Delta \log CPI_t$, our results still hold. The same holds true if we change the dependant variable from being the cross-sectional variance of investment rates to being the interdecile range of investment rates. This modification allows us to highlight that our results are not being governed by outliers but are likely reflecting pervasive changes in the distribution of investment.

The results of regressing the variance of the investment rate on the contemporaneous and lagged variance of the logarithm of operating profits relative to capital are presented in Table 4 and a plot of the two series against time in Figure 3 (their correlation coefficient equals -0.54 with a standard error of 0.19).²¹ Column (1) presents estimates for our baseline regressions in which only a 1970's dummy is included to control for possible changes in σ . The coefficients on the variance of the profit rate are both negative and jointly significantly different from zero at the 1-percent level. Note that the coefficient of the contemporaneous variance of the logarithm of the profit rate is much larger (and more significant) than the coefficient of its lagged value. In this case, a 50-percent increase in the standard deviation of the log profit rate is approximately associated to a 55-percent cumulative increase in the standard deviation of the investment rate. These results are again supportive of the predictions of our theoretical model. In column (2), we add the lagged value of the dependant variable as a regressor and find that it does not have a significant effect. The same result is found (although not reported in the table) if a further lag of the variance of the log profit rate is added to the equation. Since, in addition, the Durbin-Watson statistic in column (1) does not suggest any dynamic misspecification, we omit reporting the results with more general dynamics for the remainder of this section.

Columns (3), (4), (5), and (6) examine the robustness of this result with respect to the inclusion of the change in the price of oil, of the rate of change in the user cost of capital, of growth rate in GDP, and the inflation rate. As can be seen, the percentage change in the user cost of capital have only negligible effects on the relationship between the variance of the investment rate and the variance of the profit rate. The coefficients of the growth rate in the price of oil, the growth in GDP, and the inflation rate play a larger positive role. However, the

²¹ Profits (in current prices) are calculated by adding depreciation and interest payments to net income. They are then deflated by an industry-specific, value-added deflator to obtain real profits. The latter are then divided by the end-of-period real value of the capital stock to obtain the profit rate.

TABLE 4—TIME-SERIES RELATIONSHIP BETWEEN THE CROSS-SECTIONAL VARIANCE OF THE LOG PROFIT RATE AND THE CROSS-SECTIONAL VARIANCE OF THE INVESTMENT RATE

Regressors	(1)	(2)	(3)	(4)	(5)	(6)
$\text{Var}(\log(R/K)_{t+1})$	-0.015 (0.004)	-0.012 (0.005)	-0.015 (0.004)	-0.015 (0.005)	-0.010 (0.004)	-0.010 (0.004)
$\text{Var}(\log(R/K)_t)$	-0.005 (0.005)	-0.002 (0.006)	-0.004 (0.004)	-0.005 (0.005)	-0.009 (0.004)	-0.008 (0.004)
1970's dummy	0.004 (0.001)	0.003 (0.002)	0.004 (0.001)	0.004 (0.001)	0.004 (0.001)	0.002 (0.001)
Lagged dependent variable	—	0.216 (0.268)	—	—	—	—
Percentage change in oil price	—	—	0.008 (0.004)	—	—	—
Percentage change in user cost of capital	—	—	—	-4.9e-06 (0.003)	—	—
Growth rate of GDP	—	—	—	—	0.045 (0.013)	—
Inflation rate	—	—	—	—	—	-0.025 (0.007)
$F_{2,n}$	8.22	2.44	9.84	3.89	10.76	9.51
[p value]	[0.003]	[0.121]	[0.002]	[0.005]	[0.001]	[0.002]
\bar{R}^2	0.42	0.40	0.51	0.38	0.66	0.65
Standard error of the regression	0.0016	0.0016	0.0014	0.0016	0.0012	0.0012
Durbin-Watson test	1.71	1.68	1.78	1.71	1.80	1.77
Lagrange Multiplier test	—	0.32	—	—	—	—
Number of observations	20	20	20	20	20	20

Notes: Dependent variable: cross-sectional variance of investment rate, $\text{Var}_z(I_{t+1}(z)/K_t(z))$. The cross-sectional variance of the log profit rate is denoted by $\text{Var}(\log(R/K)_t)$ in the tables. All equations contain a constant. Estimation period: 1970–1990. Standard errors of coefficients are in parentheses. The F -statistic corresponds to the test of joint significance of $\text{Var}(\log(R/K)_{t+1})$ and $\text{Var}(\log(R/K)_t)$. The Lagrange Multiplier test is for first-order serial correlation, and is distributed as $\chi^2(1)$.

sum of the coefficients on the variance of the profit rate and its lagged value change very little and they remain jointly significant at the 1-percent level.

It is worth noting that the results presented in Table 4 are particularly suggestive of the imperfect-information story we are advocating as opposed to a financial-constraints story. For example, if the alternative hypothesis is that the time variation in the cross-sectional distribution of investment is due mostly to changes in the importance of credit constraints, then we would expect the cross-sectional variance of the (log) profit rate (a proxy for the availability of internal funds for investment) to be positively related to the cross-sectional variance of the investment rate. As seen in Table 4, this is not the case. In order to further insure that our imperfect-information explanation is kept separate from one based on financial constraints, we

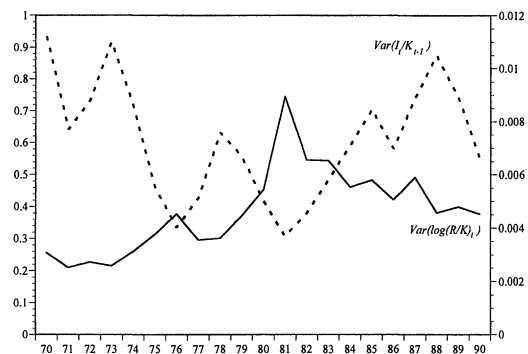


FIGURE 3. CROSS-SECTIONAL VARIANCE OF THE LOG OF THE PROFIT RATE AND CROSS-SECTIONAL VARIANCE OF THE INVESTMENT RATE

Notes: The continuous line denotes the cross-sectional variance of the log of the profit rate, denoted by $\text{Var}(\log(R/K)_t)$ (left axis). The broken line denotes the cross-sectional variance of the investment rate, denoted by $\text{Var}(I_t/K_{t-1})$ (right axis).

TABLE 5—ROBUSTNESS OF THE TIME-SERIES RELATIONSHIP BETWEEN THE VARIANCE OF THE LOG OF THE PROFIT RATE AND THE CROSS-SECTIONAL VARIANCE OF THE INVESTMENT RATE, EXCLUDING POSSIBLY FINANCIALLY CONSTRAINED FIRMS

Regressors	(1)	(2)	(3)
Var(log(R/K) _{t+1})	-0.018 (-3.329)	-0.024 (-3.057)	-0.019 (-2.186)
Var(log(R/K) _t)	-0.005 (-0.908)	-0.003 (-0.361)	-0.008 (-0.858)
1970's dummy	0.003 (2.778)	0.002 (2.208)	0.003 (2.209)
F _{2,n}	8.19	7.13	5.26
[p-value]	[0.004]	[0.006]	[0.018]
R ²	0.33	0.37	0.28
Standard error of the regression	0.0016	0.0016	0.0018
Durbin-Watson test	1.67	1.56	1.50
Number of observations	20	20	20

Notes: Dependent variable: cross-sectional variance of investment rate. Var(log(R/K)_t) denotes the cross-sectional variance of the log profit rate. All equations contain a constant. Estimation period: 1970–1990. Standard errors of coefficients are in parentheses. The *F*-statistic corresponds to the test of joint significance of Var(log(R/K)_{t+1}) and Var(log(R/K)_t). In column (1), only firm-year observations with real sales above the first quartile are included. In column (2), only firm-year observations with positive dividends are included. In column (3), only firm-year observations with dividend-pay-out ratio above the first quartile are included.

report in Table 5 results for our baseline regression in the case where we restrict attention to firms that are less likely to be financially constrained. In particular, in column (1) of Table 5 the sample is restricted to firm-year observations with real sales above the first quartile of the distribution. In column (2) we restrict attention to observations with positive dividends and in column (3) we restrict attention to observations characterized by a dividend-payout ratio (dividends divided by cash flow) above the first quartile of the distribution. In all three cases the coefficients on the variance of the log profit rate change very little and our conclusions still hold.

III. Conclusions

This paper has focused on the effect of the informational content of prices on the distribution of investment across firms. In order to explain the distributional movements in the

investment rate, we have presented a simple model whose objective is to formalize how nominal uncertainty may affect the allocation of investment. The model implies that as the general price level becomes more predictable, the firm's own relative price, and hence its profit opportunities, become easier to forecast. This should cause the cross-sectional distribution of investment to widen, since improved information allows firms to channel investment towards the most profitable opportunities. Correspondingly, our empirical work has concentrated on explaining the changes in the dispersion of the investment rate in the United Kingdom over the period 1970–1990, where we have documented that dispersion decreased in the 1970's and widened in the 1980's. We have also documented that the variance of the investment rate moved procyclically. The overall changes are consistent with anecdotal evidence about greater macroeconomic policy uncertainty in the 1970's compared to the 1980's.

Formal support for the model comes from two sources. First, econometric estimation shows that the conditional variance of inflation is inversely related to the cross-sectional variance of the investment rate, calculated for the entire manufacturing sector or within more narrowly defined industries. The implications of the model also receive empirical support from the significant negative correlation between the variance of the investment rate and the variance of the log of the profit rate calculated using panel data. In both cases we have taken care to examine whether real forces, such as oil price changes or changes in real GDP growth, may be driving our results. This was not found to be the case. Our conclusions also hold after controlling for changes in the user cost of capital. Hence, we believe that this paper provides some clear and intuitive evidence in support of the view that monetary instability, through its effect on the information content of prices, may hinder the efficient allocation of investment.

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