FIRM-LEVEL INVESTMENT IN FRANCE AND THE UNITED STATES : AN EXPLORATION OF WHAT WE HAVE LEARNED IN TWENTY YEARS[1]

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Abstract

We review the changes in modelling strategy and econometric methodology when estimating a firm-level investment equation on panel data during the past twenty years, in order to assess which of these changes result from new estimation methods and changes in the practice of panel data econometrics, and which are "real" and due to the evolution of the economy. Thus our paper consists of a series of comparisons: a simple accelerator-profit specification versus one with error correction, traditional between- and within-firm estimation versus GMM estimation, the investment behavior of French firms versus that of U.S. firms, and investment behavior today versus ten to twenty years ago. Although the econometric advances have perhaps not been as successful as we had hoped, we do find some real change in firm behavior and some improvement in equation specification and interpretation during the past twenty years.

1 Introduction

Twenty years ago, at the first Econometrics of Panel Data Conference, whose anniversary we celebrated in June 1997, Robert Eisner and Gilles Oudiz presented some of the first panel data estimates of an investment equation at the firm level, Eisner for the United States and Oudiz for France.[1] Since then, a large number of advances have been made both in the econometric theory and in the econometric practice and technology for analyzing panel data. The anniversary of this conference seemed an appropriate moment to look back and ask what effects these advances have had on the estimation of this particular equation. Are we closer to having an investment equation that is a robust description of the investment behavior of industrial firms in developed economies? Has our new methodology really helped us to improve our understanding of the major determinants of firm investment?

In the past twenty years, important developments in the economic analysis and modeling of investment have also occurred, and many applied studies have been performed on investment at the firm level, based on increasingly available micro data sets. This vast literature, of which we will provide a very short overview by way of introduction, has given rise to a number of good surveys. Our intent here is not to propose one more of them, but to focus on the purely econometric issues. Our goal, which is ambitious enough, is to assess which changes in the firm investment relation over the past twenty years result from the new methods and practice of panel data econometrics, and which are "real" and due to the evolution of the economy. We will mainly consider two problems in the specification and estimation of the investment equation (which are closely related and affect many other panel data econometric analyses): the biases that may arise from the presence of (correlated) firm specific effects and the simultaneity biases that may arise from the joint determination of output and investment.

We first proceed with a careful comparison with the main results of the EISNER [1978] and OUDIZ [1978] papers, as well as with those in a related paper by MAIRESSE and DORMONT [1985], all of which are based on a standard accelerator-profit speci-

[1] For earlier work using firm panel data to estimate an investment equation, see GRUNFELD [1960] and KUH [1963] for the U.S. or ECHARD and HENIN [1970] for France.

fication of the investment equation and on traditional between- and within-firm type estimations. We conduct an analysis as much like theirs as possible, using four samples of large manufacturing firms, for two different time periods, a recent one and an earlier one, and for both France and the United States. We then present and discuss the results obtained for our four samples, but now using an improved error correction specification of the accelerator-profit equation and a more appropriate instrumental variables estimation method, the by now well-known GMM (Generalized Method of Moments). The main advantage of the error correction specification is in allowing us to identify the longer term and shorter run aspects of the investment relation separately; this aim was indeed one of the goals of Eisner, Oudiz and Mairesse-Dormont papers. GMM estimation should in principle be able to correct for the biases due to both the presence of correlated effects and simultaneity (and incidentally to those due to random errors of measurement). Our paper thus consists of a series of comparisons: a simple accelerator-profit specification versus one with error correction, traditional between- and within-firm estimations versus GMM estimations, the estimated investment behavior of French firms versus that of U.S. firms, and today versus ten to twenty years ago.

After a bird's eye view of the vast literature on firm investment, as it evolved in the past twenty years (section 2), we begin with a discussion of the familiar and fairly eclectic class of accelerator-profit investment equations (section 3), and with a brief presentation of the recent GMM estimation methods (section 4). This is followed by a description of our data sets, which cover about 400 to 500 manufacturing firms each, in France and the United States, for the 1968-1979 and 1979-1993 periods (section 5). Then we proceed to present the various sets of estimates based on these data-sets, assessing the differences with the estimates of Eisner for the United States 1961-68, Oudiz for France 1971-75, and Mairesse-Dormont for both countries 1970-79 (section 6), and the differences related to the choice of the error correction specification and the use of the GMM estimations (section 7). We conclude with a tentative discussion of what we have learned in twenty years from the advances in panel data econometrics.

2 The Changing Investment Equation

Over the past thirty years or so, a series of revolutions and evolutions have taken place in the modeling of investment, often driven by the lack of success of previous models in explaining very much of investment behavior and by a continuing or even increased interest in policies that affect the investment behavior of private firms. In our view, the major changes can be summarized in the following way: 1) a shift in attention away from macro modeling towards micro modeling, partly driven by data availability, but also by increasing awareness of the inappropriateness of time series data for the structural models that are of interest if one wishes to understand the fundamental determinants of investment; 2) the revolution caused by the influential paper of MODIGLIANI and MILLER [1958], who pointed out the irrelevance of financial considerations for investment decisions in some circumstances, followed by a counter-revolution due to the introduction of asymmetric information and agency costs into the theoretical models (JENSEN and MECKLING [1976], STIGLITZ and WEISS [1981], MYERS and MAJLUF [1984]); 3) with the move to the use of micro or panel data, increased understanding of the complexities introduced in econometric estimation by dynamics and by right hand side variables that may be endogenous, even if only with respect to the past history of the dependent variable. We discuss each of these changes briefly.

First, as has been well-documented by HASSETT and HUBBARD [1997] and HUBBARD [1997] among others, dissatisfaction with the empirical results obtained when macro-economic data are used to estimate investment relations derived from economic theory (that is, relations that focus on the cost of capital and expected returns as investment determinants) has led to a re-examination of the econometric assumptions underlying the macro investment relation. One does not have to examine these assumptions very carefully to reach the conclusion that the exercise is fundamentally flawed, owing to the obvious simultaneity between the dependent variable (investment or investment rates) and the independent variable (the relative price, cost of capital, or Tobin's q) in large semi-closed economies (which includes most developed economies during much of the post-World War II period). This simultaneity arises from the fact that the observed values of aggregate investment and its price trace out a sequence of equilibrium points that need not have any relationship to the investment demand relation supplied by economic theory. This is an old point, but it is frequently ignored in practice and in discussions of the "failure" of the aggregate investment literature (for a recent critique, see HALL's [1997] discussion of HASSETT and HUBBARD [1997]. If good instruments for the investment price existed that were uncorrelated with other aspects of the macro-economy, the solution to this particular problem would be easy, but these instruments have proved hard to find.

Thus as computing power and micro-level datasets became available, attention shifted to the estimation of investment equations using micro data. The papers presented at this conference twenty years ago are some of the earliest examples of this shift, but they themselves do not reflect the state of economic theory at the time, since they do not incorporate any explicit price information into the investment equation. In fact, in an economy with a fairly flexible capital market, such as those in most developed countries, variation in investment prices or the cost of capital is difficult to come by in the cross section dimension, so it is often ignored or subsumed in a series of time dummies in the regression equation. There are exceptions to this rule where such things as "exogenous" variation in tax exposure affect the cost of capital to individual firms, and these exceptions have been exploited in a series of papers reviewed in HUBBARD [1997].[2] In general, estimation of this type demonstrates sensitivity of investment to the price of investment when the equation estimated comes closer to satisfying the assumptions necessary for consistency.

The shift of attention to micro data means that two other considerations come to the forefront: first, in order to say something about aggregate investment using the micro estimates, it is necessary to understand the implications of micro behavior for aggregates. Given the selectivity of most micro samples, this task is not always trivial. Second, as we analyze smaller and smaller units (e.g., individual plants), it becomes

[2]See also HALL [1993], where variation in the tax price of R&D is used to estimate the responsiveness of R&D investment to changes in the cost of capital, yielding a price elasticity for R&D investment of one or greater.

more and more obvious that investment is a lumpy rather than smoothly continuous process and that we may need to take account of this in constructing our theoretical models. See CABALLERO [1997] for a thorough discussion of these two points.

The same time period that saw a shift of attention from macro-economic investment equations to micro has seen two major theoretical revolutions:[3] first, the influential theorem of MODIGLIANI and MILLER [1958] (hereafter M-M) demonstrated that in a world of perfect capital markets, investment decisions should not be affected by financing decisions or capital structure of individual firms, but only by the cost of capital faced in the market. This result implied that there should be no role for liquidity variables such as cash flow in the investment equation, except to the extent that they reflected future profit opportunities that were not otherwise accounted for by such things as sales growth. At the time M-M was published, there were already empirical results available that suggested a strong role for cash flow in the equation (MEYER and KUH [1957]), but the effect of the M-M proposition was to deflect attention for a time from the importance of cash flow or profits in the investment equation towards a more neo-classical view of the firm's investment decision, such as that in JORGENSON [1963]. Weaknesses in the empirical implementation of Jorgenson's model led among other things to the development of a literature that explicitly allowed for adjustment costs or delivery lags in investment. This literature culminated in the empirical Tobin's q literature (e.g. TOBIN [1969], SUMMERS [1981]), which attempts to provide a theoretically better measure of "price" or expected rate of return for investment than the current marginal product of capital used by Jorgenson (which was implemented in practice using ad hoc adjustment lags).

However, as is demonstrated by EISNER [1978], OUDIZ [1978], and MAIRESSE and DORMONT [1985] among others, interest in cash flow effects on investment never entirely waned during the period following the publication of Modigliani-Miller, and eventually theorists came to the rescue of those who continued to believe strongly

[3]We provide only a brief overview of well-trampled ground here. See CHIRINKO [1993], SCHIANTARELLI [1996], and HUBBARD [1997] for three excellent recent surveys of the theoretical and empirical developments in the estimation of the investment relation in the presence of asymmetric information or agency costs.

in the importance of firm liquidity for investment decisions. This rescue took the form of a series of papers beginning with that of JENSEN and MECKLING [1976] that demonstrated the breakdown of the M-M proposition in the presence of either asymmetric information between investors and the firm, or agency costs arising from the divergent goals of managers and shareholders. Holes in the theoretical barrier between investment and finance soon widened to permit a flood of empirical papers that explore various implications of the potential cost wedge between external and internal funds on the investment behavior of individual firms. Although it has become fairly clear from this work that cash flow plays an important role in the investment equation at the firm level, a role *consistent* with the presence of financial market imperfections in some (but not all) cases, definitive evidence that cash flow is not simply a proxy for news about expected future profits has sometimes been hard to come by. Our present work is no exception to this rule and we make no attempt to identify the source of the cash flow effect; we merely document its presence.

We now turn to issues that are specifically econometric. Investment equations typically display two phenomena that create difficulties when they are estimated using short panels: they are inherently dynamic because of adjustment costs and they usually involve some kind of expectations about the future profitability of investment. Both of these factors tend to introduce time series behavior in the left hand side variable or the disturbances that is related to the right hand side variables either contemporaneously or with a lag. Although the latter situation would normally create no difficulties for estimation, in the presence of firm specific but unobserved effects that are related to the right hand side variables, the addition of dynamics to the model renders the usual within firm estimator inconsistent (see NICKELL [1981] for computation of the asymptotic bias of the estimator in this case). Beginning with BALESTRA and NERLOVE [1966] and ANDERSON and HSIAO [1982], various instrumental variable solutions to the problem have been suggested in the literature, and the current state of this art is the use of fully efficient GMM estimators that allow for heteroskedasticity across firms, and serial correlation over time. We pursue this strategy in this paper.

During the past twenty or so years, the problem of incorporating expectations about the future profitability of investment in the equation has been confronted (or side-stepped) in several ways: 1) the afore-mentioned Tobin's q methodology, where the market value of the firm proxies for future investment opportunities; 2) the approach of ABEL and BLANCHARD [1986], where projections of future profits are used as a proxy; and finally, 3) the Euler equation approach,[4] which in essence removes the problem created by the need to construct expectations into the infinite future by first differencing the investment equation so that the current marginal product of capital (the capital-sales ratio if production has the Cobb-Douglas form) and expected one-period changes in adjustment costs are all that is needed to describe the change in expectations about the future profitability of capital at the level of an individual firm.[5]

Considerable experience with the fragility and implausibility of Euler equation estimates and a desire to stay close to the models estimated by Eisner and Oudiz leads us to focus on a different class of models from the three just discussed in the later part of this paper: error-corrected versions of the traditional accelerator. These models assume that sales and capital are proportional in the long run (as in the traditional accelerator), but that in the short run there may be complex dynamics relating the two as firms invest and disinvest in order to achieve their target capital. These dynamics are specified in an ad hoc distributed lag manner as an ADL(2,2) specification, and then the terms are rearranged to give the equation an error-correcting interpretation. The advantage of specifying the investment equation in this way is twofold: the error-correcting formulation explicitly separates the long run determinants of investment from short run adjustment lags, which can be informative if one believes that the simple neoclassical theory is likely to hold only in the long run. In addition, once a separation of this kind is made, it becomes straightforward to develop more realistic models for panels of firms where the long run relationship is the same across firms, but short run dynamics may vary (along the lines of recent theoretical work by PESARAN, SHIN, and SMITH [1997]). Recent empirical work that uses the error-correcting accel-

[4]See WHITED [1992], BOND and MEGHIR [1994], and BLUNDELL, BOND and MEGHIR [1996].

[5]As in all the panel data models described here, there will also be macro-economic changes in investment prices and interest rates that are subsumed in the time dummies. Only when there is variation in these quantities across firms for tax or other reasons will they enter the Euler equation, or any other investment equation for that matter.

erator specification includes BOND, ELSTON, MAIRESSE, and MULKAY [1997] and BOND, HARHOFF, and VAN REENEN [1997]; for estimation of such an equation using macro-economic data, see CABALLERO [1994].

To sum up, the current paper focuses on the changes in the micro-economic estimates of the investment equation from those of twenty years ago that are due to changes in econometric methodology. We do not consider aggregation issues, the consequences of recent theoretical advances in modeling lumpy and irreversible investment, nor do we search for "natural experiments" to identify the price or liquidity effects. This is not because all of these topics might not be important, but simply because of our desire to isolate the consequences of (supposedly) improved methodology holding other features of the research process constant.

3 Models of Investment

Our approach is in the spirit of BOND, ELSTON, MAIRESSE and MULKAY [1997], in that rather than focusing on finding the "correct" model of investment, we use a version of an error-correcting accelerator-profit model that encompasses the earlier literature in this area as exemplified in Eisner, Oudiz or Mairesse-Dormont, but that can also be related to the recent new wave of firm-level empirical work trying to ascertain the sensitivity of investment to financial constraints.

Our base model implies that the long run capital stock of the firm is proportional to output:

$$k_{it} = \theta s_{it} + h_{it}, \tag{1}$$

where h_{it} is a function of the user cost of capital and the parameters of the production function and k_{it} and s_{it} denote the logs of capital and output.[6] This relationship is consistent with the simple neoclassical model of a profit-maximizing firm with a single type of capital, CES production function, and no adjustment costs, as shown in Appendix A.

[6]Note that throughout the paper K_{it} is the capital stock of firm i at the end of year t, while S_{it} and I_{it} are the sales and investment of firm i during year t.

We then specify a dynamic adjustment mechanism between k and s as an autoregressive-distributed lag of length two (an ADL(2,2) specification), which nests equation (1) as its long-run solution, and we also assume that variation in the user cost of capital/productivity term h_{it} can be controlled for in the estimation by including year-specific and firm-specific effects in estimation. These assumptions yield the following accelerator-type equation:

$$k_{it} = \alpha + \gamma_1 k_{i,t-1} + \gamma_2 k_{i,t-2} + \beta_0 s_{it} + \beta_1 s_{i,t-1} + \beta_2 s_{i,t-2} + \varepsilon_{it}, \tag{2}$$

where the disturbance ε_{it} contains firm and year-specific effects, as well as transitory shocks. Rewriting this equation in an error-correcting framework, we obtain

$$\begin{aligned} \Delta k_{it} &= \alpha + (\gamma_1 - 1)\Delta k_{i,t-1} + \beta_0 \Delta s_{it} + (\beta_0 + \beta_1)\Delta s_{i,t-1} \\ &\quad +(\gamma_2 + \gamma_1 - 1)(k_{i,t-2} - s_{i,t-2}) \\ &\quad +(\beta_0 + \beta_1 + \beta_2 + \gamma_2 + \gamma_1 - 1)s_{i,t-2} + \varepsilon_{it}. \end{aligned} \tag{3}$$

This expresses the growth rate of capital stock as a function of both growth rates and levels information: its own lagged growth rate, the growth in sales (current and lagged once), and an error correction term (the log of the capital-output ratio) and a scale factor (the log of sales). Writing the equation this way is convenient because the last two terms provide simple t-tests for error-correcting behavior and the hypothesis that θ is unity in the long run, while the first three variables capture the short-run dynamics. In estimation, we use the investment ratio $\frac{I_{it}}{K_{i,t-1}}$ as a proxy for the net growth in capital stock Δk_{it}.[7]

[7] We have:

$$\Delta k_{it} = \log\left[\frac{K_{it}}{K_{i,t-1}}\right] = \log\left[1 + \frac{\Delta K_{it}}{K_{i,t-1}}\right] \cong \frac{\Delta K_{it}}{K_{i,t-1)}} \cong \frac{I_{it}}{K_{i(t-1)}} - \delta$$

where δ is the (average) depreciation rate. The approximation of the growth in capital Δk_{it} by the net investment rate $\frac{I_{it}}{K_{i,t-1}} - \delta$ is likely to be fairly good, the median rate for our firms being quite small.

Note that the variation in δ now enters directly in the disturbance in addition to the cost of capital/productivity term h_{it} and that we are assuming that this variation can also be controlled in estimation by year and firm effects.

We finally augment equation (3) with the current and lagged ratio of cash flow to beginning of period capital stock in order to capture effects associated either with liquidity constraints or with changes in profitability that are not captured by the sales growth variables. Writing cash flow as CF, our final estimating equation is thus the following:

$$\begin{aligned}
\frac{I_{it}}{K_{i,t-1}} = {} & \alpha + (\gamma_1 - 1)\frac{I_{i,t-1}}{K_{i,t-2}} + \beta_0 \Delta s_{it} + (\beta_0 + \beta_1)\Delta s_{i,t-1} \\
& + (\gamma_2 + \gamma_1 - 1)(k_{i,t-2} - s_{i,t-2}) \\
& + (\beta_0 + \beta_1 + \beta_2 + \gamma_2 + \gamma_1 - 1) s_{i,t-2} \\
& + \delta_0 \frac{CF_{it}}{K_{i,t-1}} + \delta_1 \frac{CF_{i,t-1}}{K_{i,t-2}} + \delta_2 \frac{CF_{i,t-2}}{K_{i,t-3}} + \varepsilon_{it}.
\end{aligned} \tag{4}$$

The long-run properties of this specification depend only on the error correction coefficient $\rho = \gamma_1 + \gamma_2 - 1$ and the scale coefficient $\lambda = \beta + \rho$ with $\beta = \beta_0 + \beta_1 + \beta_2$. We expect that ρ will be negative, implying that if capital is less than its "desired level" future investment will be higher and conversely; and that λ will not be statistically different from zero, if there are constant returns to scale.[8] We would also expect that the sum of the coefficients on cash flow $\delta = \delta_0 + \delta_1 + \delta_2$, (or the corresponding long-run cash flow coefficient $-\delta/\rho$) will not be significant, if the cash flow variable captures only the transitory effects of liquidity constraints on firm investment. The dynamic properties of the equation depends on the values and profile of the individual coefficients. We can test for the presence of sales or cash flow by considering the joint significance of β_0, β_1 and β_2 or that of δ_0, δ_1 and δ_2. We can also test for the presence of lag two effects by looking at the joint significance of β_2, γ_2 and δ_2.

With the omission of the level terms in sales and capital stock and that of the lagged investment rate, our error correction model looks superficially like the traditional accelerator-profit model that was estimated in EISNER [1978], OUDIZ [1978] and MAIRESSE-DORMONT [1985]. However, the dynamic properties of the two specifications are different.[9] The traditional accelerator is derived by differentiating

[8] Or if Cobb-Douglas is a good enough approximation to the underlying production function: see Appendix A.

[9] The central reason for the difference is the presence or absence of the lagged dependent variable, of course. When it is present, as in the error-corrected version, we have a model that has a stochastic

equation (2), which destroys any equilibrium relationship between the level of variables k and s. On the other hand, the error correction specification (3) is just a reparametrization of the same equation (2) that retains information about the long-run equilibrium between k and s in addition to the short-run relationship between the rates of growth of the variables (see HENDRY, PAGAN, SARGAN [1984]). Moreover, when considering panel data estimation with specific firm effects, adding these effects in the accelerator model has very different consequences from adding them to the error correction model. In the former (accelerator) case, these effects correspond to *different trends* in the levels of capital and output, while in the latter (error correction) case these effects correspond to *different levels* of the capital-output ratios.

The error correction specification also has the advantage of making our work comparable to much of the recent literature on firm-level investment that uses an Euler equation framework to look for evidence of "excess sensitivity to cash flow." The difference between our model and the typical Euler equation framework is that in the latter model, the adjustment costs are included in the firm's optimization problem. With a quadratic adjustment cost function, a squared term in I/K appears in the empirical equation, and the sales-capital ratio (Y/K) is used to proxy for the desired capital stock and expectation of future profitability, instead of the current and lagged growth rate of sales used in our model. Moreover, in the Euler equation, the lagged profit rate, proxied by operating income, is also included with a negative sign, because when it is transitorily high, firms will defer their investment so as to incur adjustment costs in a period of lower profitability. However, this variable generally has a positive coefficient in estimation, and some authors argue that this finding indicates the presence of liquidity or financial constraints on investment (see for example BOND and MEGHIR [1994]). In our error-correcting specification, in addition to measuring the sign of the coefficient, we can test whether the profit rate, computed using the gross profit or cash-flow, plays the role of a long-run determinant of investment, or whether it is only a short-run variable which can be interpreted as reflecting the transitory availability of

difference equation interpretation and can easily be derived from a partial adjustment mechanism, whereas when it is absent, as in the traditional accelerator, the model is of the type that HARVEY [1990] calls a transfer function model, and which is closely related to adaptive expectation models.

funds for investment purposes.

4 Methods of Estimation

Our econometric model is the usual linear regression model with fixed firm effects and year effects:

$$y_{it} = x_{it}\beta + \alpha_i + \delta_t + \varepsilon_{it} \qquad i = 1, ..., N\ ;\ t = 1, ..., T. \tag{5}$$

The disturbance in this investment equation will contain a variety of errors of specification: firms differ in their technology, the rate at which their capital depreciates, the rate of return required by financial markets, and in the construction of the accounting measures we use for estimation. In addition, the average depreciation rates and required rates of return for capital may change over time, which will imply changes in the average gross investment rate. For the most part, these effects will be captured by the presence of firm-specific and year-specific dummies.

Because T is small (in our case, between 6 to 9) and N is reasonably large (between 400 and 500), we estimate the δ_ts simply by including a full set of time dummies in all models, and focus instead on the treatment of the "permanent" differences across firms, the α_i. We consider and correct for estimation bias in β that arises from two different sources: the potential correlation of the α_i with the right hand side variables x_{it} and the potential endogeneity of the x_{it}s with respect to current or future shocks to the investment relation.

As in the earlier work of Oudiz, Eisner, and Mairesse and Dormont, the first problem is dealt with by the usual fixed effects transformation. We estimate the following variants of equation (5):[10]

[10] A bar over a variable denotes the time-average of the variables. $\bar{\alpha}$ is just the intercept of the model and should be interpreted as the average of α_i, as well as $\bar{\delta}$ is the average of the time dummies δ_t.

$$y_{it} = x_{it}\beta + \overline{\alpha} + \delta_t + (\alpha_i - \overline{\alpha} + \varepsilon_{it}) \qquad (TOTALS)$$

$$\bar{y}_{i\cdot} = \bar{x}_{i\cdot}\beta + (\overline{\alpha} + \bar{\delta}) + (\alpha_i - \overline{\alpha} + \bar{\varepsilon}_{i\cdot}) \qquad (BETWEEN)$$

$$y_{it} - \bar{y}_{i\cdot} = (x_{it} - \bar{x}_{i\cdot})\beta + (\delta_t - \bar{\delta}) + (\varepsilon_{it} - \bar{\varepsilon}_{i\cdot}) \qquad (WITHIN)$$

Thus the estimates of the slope coefficients β are average effects with no controls for the permanent differences in firm-level productivity or depreciations rates that may be correlated with output levels. The between and within estimates are based on an orthogonal decomposition of the variability of y and x into the variation across firm and the variation over time within a firm (note that the second and third equations add up to the first). Given the potential correlation of α_i and the x_{it}s, we do not expect the estimates of β based on these three equations to be the same (see e.g., MAIRESSE [1990]).

The differences between the total and between on the one hand (which contain α_i) and the within on the other (which does not) were used by Eisner and formalized by Mairesse-Dormont to characterize the differences between the long run (cross section) and short run (within firm) investment behavior. However, as we suggested in the previous section, under an error correction interpretation, one can obtain estimates of long run behavior even from the within firm regression, and these estimates differ from the traditional accelerator estimates obtained by Eisner and Oudiz.

To be more explicit, if the equilibrium relationship in levels includes a firm-specific effect, then taking the first difference to obtain the usual accelerator-profit specification should remove it. If one then computes the within estimator, as Eisner and Oudiz do, in effect one has allowed for a firm-specific *trend* in the original level equation. In contrast, when using the error-correction specification, the total estimator that pools all observations will be seriously biased because the error-term will contains the firm-specific effect that is correlated with the explanatory variables, while the within-firm transformation removes this firm-specific effect.

Although estimating the error correction model using the within firm transformation controls for these correlated effects, in a short panel such as ours these estimates can also be biased estimates of the "true" coefficients, for three reasons: (i) measurement error in the right hand side variables (which will probably impart a downward bias); (*ii*) simultaneity between contemporaneous cash flow, sales growth, and investment rates (which will probably impart an upward bias); and (*iii*) simultaneity between past values of the disturbance (which enters via the within transformation) and contemporaneous right hand side variables. And although the second source of bias is always there, the within transformation exacerbates bias caused by (*i*) and (*iii*). The solution proposed by many authors[11] is to use an instrumental variable estimator on the differenced version of the model, allowing for heteroskedasticity across firms and possible correlation of the disturbances over time, that is, to use the well-known GMM method of estimation.

We discuss each of these sources of bias in turn, beginning with *(iii)*. It is by now well known that in panels like ours, where the right hand side variables (output levels and lagged capital stock) are at best predetermined, the within transformation may introduce correlation between regressors and the disturbance (see CHAMBERLAIN [1982]). This occurs because the transformation puts all the past disturbances into the equation for the current year, and these are likely to be correlated with the x_{it}s through feedback effects, even if the contemporaneous or future disturbances are not correlated. The solution to this problem is to use transformations that avoid pulling in all the disturbances, such as differencing transformations of the form $y_{it} - y_{i,t-1}$, but these introduce their own problems in the form of magnifying the effects of serially uncorrelated measurement error (*(i)*, see GRILICHES and HAUSMAN [1986]). Because the bias due to simultaneity in the within dimension falls with T, whereas the bias due to measurement error remains the same, the within estimates can be preferred even in sample sizes such as ours. However, in situations where there is contemporaneous simultaneity *(ii)*, this argument fails and it will be appropriate to use the differenced transformation for estimation, and to instrument it by means of lagged values of the predetermined variables in order to control for both contemporaneous simultaneity and measurement error.[12]

[11] See for example BALESTRA and NERLOVE [1966], ANDERSON and HSIAO [1982], ARELLANO and BOND [1991], AHN and SCHMIDT [1995].

[12] When the disturbances ε are serially correlated to some order, the choice of instruments and

Because it is highly probable that investment rates, output, and cash flow are simultaneously determined, one of the major goals of this paper is to investigate the effects of using instrumental variable or Generalized Method of Moments estimation techniques on the estimated coefficients of the investment equation. We wish to assess whether the earlier estimates were subject to substantial biases from this source and whether our newer econometric techniques are able to help. Our GMM estimates are based on the following transformation of equation (5):

$$\Delta\varepsilon_{it} = (\varepsilon_{it} - \varepsilon_{i,t-1}) = (y_{it} - y_{i,t-1}) - (x_{it} - x_{i,t-1})\beta - (\delta_t - \delta_{t-1}).$$

Under the assumption that the disturbances are not serially correlated, we expect $\Delta\varepsilon$ to be orthogonal to the past history of the x and y variables (after the first lag), so that $y_{i,t-2}, y_{i,t-3}, \ldots, x_{i,t-2}, x_{i,t-3}, \ldots$ are available as instruments for $\Delta\varepsilon_{it}$. If the disturbances in level ε_{it} follow a moving average process of order one, the first valid instruments are found at lag 3 instead of lag 2 because the differenced disturbances follow an MA(2) process.

In the estimation, we pay some attention to the validity of the instruments. There are two problems with GMM or instrumental variables methods: first, the instruments should be uncorrelated with the error terms, or the orthogonality conditions should be satisfied by the data, the exogeneity requirement; second, the instruments should have a strong correlation with the regressors of the model, the relevance requirement. The former is usually verified by a classical Sargan test which tests the over-identifying restrictions (see HANSEN [1982]) while the latter, called the problem of "weak" instruments, is too often neglected in empirical papers. NELSON and STARTZ [1990] propose the use of R-squared statistics from the regression of the regressors on the instruments, and BOUND, JAEGER and BAKER [1995] advocate a related F-statistic. A. R. HALL, RUDEBUSH and WILCOX [1996] suggest the use of canonical correlations to assess this problem. They propose testing the smallest canonical correlations between the set of regressors and the set of instrumental variables in order to detect the relevance of the instruments. We report the result of using all three of these diagnostics

estimation technique becomes more complex but the basic idea is the same, and the validity of the instruments can easily be tested in this setting.

for some of our results in Section 7 of the paper. But first we describe our data and present some panel data estimates for the accelerator-profit model.

5 Data

The construction of our data samples is described in Appendix B, which also gives some detail on the characteristics of the four balanced samples (two each for France and the United States, 1968-1979 and 1979-1993) that are used for the analysis.

In Table 1, we present descriptive statistics for the variables that are used in our regressions. Note first that the median firm size in terms of employment is about ten times as large in the United States sample as in the French sample, in spite of the fact that our samples consist of the largest manufacturing firms in each country for the most part. These tables also confirm that the U.S. firms have higher profit and investment rates on average, and that their capital intensity has increased between the two periods, while that for the French firms has remained constant. Some of these differences across countries reflect the relative sizes of the economies (roughly 3 to 1) and differences in the sampling frame: the U.S. sample consists of publicly traded manufacturing firms that report their accounting data to the Securities and Exchange Commission, whereas the French sample, which is drawn from a very large database, is closer to a Census of manufacturing firms, and includes a number of non-quoted firms.

Figure 1 compares the evolution over time of the medians of our key variables for France and the United States: the investment rate I/K, the capital-output ratio K/S, the growth of sales, and the operating income-capital ratio OPINC/K for the first period and the operating income-capital (OPINC/K) and cash flow-capital (CF/K) ratios for the second period. The two countries display similar behavior overall between 1968 and 1993. The break in the series in 1979 is due to the composition of the samples which do not include the same firms. Especially in the United States, the earlier sample includes somewhat larger firms.

The growth rates of sales and the investment rates are low on average. After a rapid fall in the investment rate for France during the earlier period, it seems to be somewhat more stable in the second period, while in the United States, the investment rate for the sample firms declines throughout the whole period. These patterns are consistent with those for the profit rates: in the U.S., the profit rates decline over the whole period, while in France they are roughly constant and at a lower level than in the U.S. Although it is possible that this result reflects differences in the way cash flow is computed in the two countries, it does appear that higher investment rates are associated with higher profit rates and lower investment rates with lower profit rates, even at this aggregate level.

The capital intensity for firms in the United States grows far more rapidly than for firms in France during the period. One explanation for this contrast is that in France the price of equipment goods is not adjusted for the quality (or the productivity) of new investment (especially investment in computers and related equipment).[13] Therefore higher prices of equipment goods relative to older equipment lower the "real" investment and the stock of capital, even though in principle, these higher prices are due to quality change. This could explain while the capital-output ratio for France increases less than in the U.S. A second reason could be the difference in the benchmark value for the capital stock used the two countries. For both countries, although they differ in accounting schemes and practices, we have used the net stock of assets in the first year of data for each firm as a benchmark. However, adjusting this benchmark value did not lead to significant changes in the estimates that we present below.

6 The Traditional Accelerator : Now and Then

In this section we compare the estimates obtained on our four data samples to those of Eisner, Oudiz, and Mairesse-Dormont, using the same accelerator-profit specification of the investment equation and computing the same estimates. There are, however, many variants in the precise way in which the analysis can be carried out, and hence

[13]This is not the case for the U.S. where the price of intermediate investment goods is lowered by the falling price of computer equipment, for which the National Income Accounts use a quality-adjusted hedonic price index.

some differences in its actual implementation by these authors themselves and some other differences in its replication by us, that we could not fully avoid. These variants mainly concern the definitions and measures of the sales and profit variables, and in what exact form and with how many lags they enter the estimated equation. In fact, several of them have already been documented in the Eisner, Oudiz and Mairesse-Dormont papers, and we have also investigated a good number (as many as we could and thought might matter!). Most of them turn out to have negligible or little effect on the estimates, and only a few appear to be of some possible real significance. We have been able to control for these cases by precisely reproducing the earlier analyses, with the notable exception of the profit measure, for which we do not have some of the relevant data for the earlier period.

The estimates reported by both Eisner and Oudiz are based on a cash-flow measure of profit, while those reported by Mairesse-Dormont are based on an operating income measure. We have only this last measure for our earlier period samples but we have both measures for our recent period samples. Experimenting with cash flow versus operating income on the recent samples shows that the magnitude of the operating income coefficients tends to be smaller than that of the cash flow coefficients by a factor of about one third to one half (which is what would be expected on the basis of the difference of their sample means), but that their statistical significance and the overall fit of the investment equation is very much the same.[14]

There are also differences in the way the investment, profit and sales variables are "normalized" to enter the estimated equation, which turn out not to really matter. Eisner measures the firm current investment and profit rates relative to the firm fixed assets in a given year (1957) rather than in the beginning of the current year t (or end of the previous year t-1), as Oudiz, Mairesse-Dormont and us in this paper. He

[14]Note also that in Eisner the cash-flow measure is net of depreciation (i.e., net profit), while the one in Oudiz is gross of depreciation (i.e., gross profit). Like Oudiz (see his section 4.1), we prefer the gross cash flow measure, which corresponds to the internal funds available for investment. Moreover, the depreciation figures reported in the firm accounts reflect in part their economic situation and their dividend policy; in the case of France, they are computed on the basis of fiscal service lives which are much shorter than the economic service lives. Eisner and Oudiz also report estimates of the accelerator-profit model in which they enter the accounting depreciation rate as a separate variable in addition to the net profit rate. Although this may have some significant impact on the individual coefficient estimates (depending on the type of estimates), the basic picture remains about the same.

also normalizes the current change in firm sales by an average of firm sales around the same year, while Oudiz normalizes it by a moving average of the current and previous two years firm sales, and Mairesse-Dormont and we simply take the log differences in firm sales. Eisner reports that taking the measures used by Oudiz instead of his preferred ones leaves his "*major results essentially undisturbed,*" and we have confirmed this assertion using the Mairesse-Dormont measures rather than those of Eisner and Oudiz.[15]

We have also experimented with lag lengths, finding as did our predecessors that one or two lags (and the current value) was enough for profits, and that six lags for sales like Eisner rather than three like Oudiz and Mairesse-Dormont could to some extent matter.

In addition to the variants in the measurement of variables, the earlier papers also differ in the variety of estimates on which they have chosen to report. Eisner thus presents five types of estimates: total overall, total with year dummies, total with industry-year dummies, between firm and within firm (which in his terminology, also adopted by Oudiz, he respectively calls: firm overall, firm cross sections across industries, firm cross sections within industries, cross sections of firm means across industries, and firm time series). He also discusses a few others that were done using data aggregated to the industry level rather than at the firm level. Oudiz considers four out of the five firm level estimates, omitting the totals with industry-year dummies. By contrast, Mairesse-Dormont limit themselves to two: the total overall and between firm estimates, which they also call first differences or year growth rates estimates, and long differences or average growth rates estimates, where their terminology refers to the original capital-sales relationship and not to the derived investment equation.[16] For the same reasons they did (see their section 2.3 and footnote 11), we choose to

[15] There is one last difference in the normalization of the investment and profit variables between our predecessors and us: while they use a gross fixed assets measure (based on the gross book values of the firm), we prefer a net value, as computed by the permanent inventory method (see Appendix B), which is more in line with the underlying economic model (see Appendix A). Mairesse-Dormont report estimates showing that this choice can make a sizeable difference: see below in the text.

[16] Note that Mairesse-Dormont include industry dummies in their total and between regressions, and that they run the between regressions without the lagged variables, which are highly collinear with the current values, being computed as the firm average growth rates over overlapping periods. See, however, footnote 19 below.

focus on these two estimates, although we also document the within firm estimates for the sake of completeness.[17]

As a result of these various considerations, our most comparable estimates to Eisner, Oudiz and Mairesse-Dormont are shown in Tables 2, 3, and 4.[18] Let us begin by comparing the estimates for the earlier period and then go on to look at our estimates for the earlier and latter periods. We first note that the sales accelerator effects are about the same for Eisner and our early U.S. sample, .70 in between and .55 in total, and also quite close for Oudiz and our early French sample, roughly .50 in between and .35 in total. We see to the contrary that these effects differ significantly between our early samples and the Mairesse-Dormont estimates, which are much lower. However, one finds a nearly complete explanation in Appendix A of their paper, where they report much higher estimates when they normalize investment and (profit) by the same net measure of capital as we do (i.e., computed by the permanent inventory method), rather than by the gross book value of fixed assets.[19] Thus, in this paper we obtain total and between estimates of the accelerator which all basically agree among the earlier studies and ours, and are significantly higher in the U.S. samples than in the French ones.

The comparison of the profit effects at first sight is less satisfactory. Comparing our results to Mairesse-Dormont, the main differences in the cross-sectional estimates are that the cash flow coefficients are lower in the new data in the U.S. and especially in France.[20] The conclusion is that even when we use the same methodology as the

[17]These estimates are based on the deviations of the year growth rates from their firm averages, and they thus involve a "double differentiation" of the basic capital-sales relation, implicitly assuming (correlated) firm trends in this relation in addition to the (correlated) firm effects. See the discussion in Section 3 of the paper.

[18]In order to be as comparable as possible, our early samples in Tables 2 and 3 do not cover the eight year period 1971-79, as in the rest of the paper: the U.S. sample only covers 1974-79, allowing us to have 6 lags for the sales growth rates as in Eisner, and the French sample covers 1971-75, which is the same period as Oudiz. The corresponding estimates computed for the 1971-79 samples do not differ much in fact.

[19]The estimates in Appendix A of Mairesse-Dormont are the following: .50 in between and .35 in total for France, and .60 in between and .35 in total for the U.S. They are thus practically the same as ours for France, and somewhat smaller, but much less so, for the U.S. This last discrepancy, however, is largely due to the fact that the Eisner estimates and ours given in Table 2 are obtained with six lags for sales, while the Mairesse-Dormont estimates and ours in Table 4 are obtained using only three lags for sales.

[20]Careful comparison of Tables 2 and 3 shows that the between estimates for the sales coefficient

earlier work, we obtain slightly different estimates for these coefficients.

7 The Error Correction Specification: Usual and GMM Estimates

In this section of the paper we present estimates for equation (4), which augments the basic accelerator model for investment with error correction and scale terms. In our view, the advantage of this specification over the traditional accelerator is that it enables the results to be interpreted in terms of a separation into long run and short run behavior. That is, the long run relationship between capital stock and output (θ) is captured by the error correction terms ($k_{i,t-2} - s_{i,t-2}$ and $s_{i,t-2}$) and the ratio of their two coefficients measures the deviation of θ from unity (which corresponds either to a Cobb-Douglas production function or to constant returns in a CES production function):

$$\frac{\beta_0 + \beta_1 + \beta_2 + \gamma_1 + \gamma_2 - 1}{\gamma_1 + \gamma_2 - 1} = 1 - \frac{\beta_0 + \beta_1 + \beta_2}{1 - \gamma_1 - \gamma_2} = 1 - \theta.$$

It is easy to see that under the null of $\theta = 1$, the level of output will enter the equation only via the error correction term and not separately. The short run adjustment dynamics will be captured by the individual behavior of the β's and γ's.

We present estimates of this model for our two sub-periods and two countries using several econometric estimation methods: total and within regression for comparison to the earlier work, and then GMM estimates using first-differenced data with several sets of instruments.

The total and within estimates are shown in Table 5. In all cases, the standard errors

tend to be quite a bit higher in Table 2 and the estimates for cash flow somewhat lower in the last two periods (for which the dates in the two tables are exactly comparable, unlike the dates for the first period). This is because the model estimated in Table 3 corresponds to the precise Dormont-Mairesse specification:

$$\overline{\frac{I_{i.}}{K_{i,.-1}}} = \beta\overline{\Delta y_{i.}} + \delta\overline{\frac{CF_{i.}}{K_{i,.-1}}} + \alpha + \overline{\varepsilon_{i.}}$$

That is, lagged values of average sales growth and average profit rates are omitted from the specification. Although they are highly collinear with the current values (see the standard errors in Table 2), they do have some additional explanatory power for investment rates.

of estimate are lower and the R-squares considerably higher than in the estimates of the traditional accelerator model using the same method of estimation and the same data in Tables 2 and 3. In the case of the total estimates, this finding is because of the presence of the lagged investment rate in the equation (which to some extent proxies for a fixed firm effect). In the case of the within estimates (where the improvement in R-squared is the most dramatic), it seems to be due to the error correction term, which is quite significant and negative. Thus the error-corrected version of the traditional accelerator fits the data better and more parsimoniously (fewer lags of sales are required) than a model without error correction.

Looking at the individual coefficient estimates, we see that the long run coefficient of sales in levels is approximately one, that is, s_{t-2} does not enter the equation in any of the total models, and the error correction term is small (on the order of $-.01$) and of the right sign, except for the older period in the U.S. In contrast, the within estimates have a long run coefficient of sales equal to about 0.6 and a good-sized error correction term (on the order of -0.3 in the first period in both countries and -0.2 in the second period). Accompanying these changes in long-run behavior is a change in the coefficient of the lagged investment rate, which is sizable in the totals and zero in the within estimates. Note also that adding the lagged investment rate to the totals equation reduces the impact of sales considerably compared to the traditional accelerator. The implication of these results is that both French and U.S. firms have permanent differences in their gross investment rates (perhaps due to differences in capital depreciation in different sectors), but that in the cross section capital is still roughly proportional to sales. A second finding, which we will explore later in the paper using GMM estimation, is that constant returns does not hold once we correct for these permanent differences in investment rates.

Turning to the cash flow coefficients, we find that they are significantly different from zero in all the estimates except the within estimates for France in the second period. In the totals, the profit or cash flow rate has a very high but imprecisely determined positive effect on the level of the capital stock in the long run, except in the earlier period for the U.S. In general, the within estimates have much larger cash flow coefficients, again except for France in the second period, and implied long run effects that are of the same order of magnitude as sales. As in the earlier work, we find support for the hypothesis that there are significant cash flow effects for investment, either due to liquidity constraints or to the role of cash flow as a signal of future profitability; these effects are somewhat lower in the second period for both countries.

Having established that the error-corrected accelerator is a somewhat better model of investment than the traditional accelerator using conventional estimation methods, we now turn to estimates based on the GMM methodology applied to panel data. In principle, GMM estimates are preferred here for several reasons: first, the evidence in Tables 2 to 5 suggests strongly that firm effects are present and that they are correlated with the right hand side variables sales, capital stock, and cash flow rates, so that we would like to control for these effects by differencing. Second, as discussed previously, the right hand side variables in our model are likely to be endogenous with respect to the differenced disturbance, so we need to use an instrumental variable estimator of some kind. Finally, the heterogeneity of our data with respect to size and other characteristics suggests that estimators that allow for heteroskedastic disturbances such as GMM should be preferred. However, as we shall see below, using such a robust estimator has a cost also, in terms of reduced precision.

Table 6 presents the results of estimating the first differenced version of the error correction model in (4), using two different instruments sets for both countries and both periods, while Table 7 shows statistics about the relevance of these two sets of instruments. The first set of instruments includes lagged values of the right hand side variables (I/K, Δs, and CF/K lagged 3 through 6 times) and the time dummies as instruments, and the second set includes the first set plus some predetermined variables ($\Delta(k-s)$, Δs and $\Delta(CF/K)$) at time $t-2$. The choice of the instruments was made after a careful look at different possibilities, including the use of levels and/or differenced information and the choice of the predetermined variables.

Our first set of instruments is valid even if there is a first-order moving-average process for the error term in the model, while the second requires a white-noise process for the error terms if the instruments are to be valid. As suggested by ARELLANO and

BOND [1991], we test for validity using a Lagrange Multiplier test for serial correlation in the residuals. Because there are 12 years of data in the first period and 15 years of data in the second, but we use 9 estimating equations in each period, the number of orthogonality conditions using the first set of instruments is 99 and 117 for the first and second period respectively, and 102 and 120 for the second set of instruments.

We find that the 1971-79 French data reject both the validity of the orthogonality conditions and validity of using lag 2 levels as instruments (because the disturbances in the equations are serially correlated at least to order 2), whereas for the U.S., the exclusion of the lag 2 predetermined variables leads to an acceptable estimate for the first period. Therefore it is probably necessary to view the earlier estimates with somewhat more suspicion than those for the 1985-93 period, where both the Sargan Test and the Lagrange Multiplier Test for second order serial correlation are accepted using both sets of instruments and for both countries.

In Table 7, we assess the relevance of the instruments by looking at the multiple correlation coefficient between each regressor and the corresponding set of instruments. The R-squares are generally low but the F-tests are still quite significant in most cases. However, NELSON and STARTZ [1990] advocate the use of a value of about 2.0 for the F-test (corresponding to a very small significance level when the number of instruments is 100). If we use a significnce level this low, there are no acceptable instruments for sales growth for France in the second period, unless we include sales growth at time $t-2$ as an instrument. This is confirmed by inspection of the canonical correlations (see A. R. HALL, RUDEBUSCH and WILCOX [1996]). Even though they are often weak for all samples, the are insignificant only for France in the second period (1985-93). In this sample, two of the regressors (probably sales growth) have no relevant instruments in the sets of 120 or 117 instruments.

Comparing the GMM estimates that purport to correct for simultaneity bias and measurement error with the within firm estimates in Table 5 reveals that there is little evidence of this kind of bias, at least in the estimates for 1985-93, where the Sargan Test is easily accepted. The main difference in the estimates is that the cash flow or profit rate coefficients have disappeared completely in the second period in both countries. Although cash flow is not a significant determinant for either country and for both periods because of the very large standard errors, it seems to have disappeared completely as a long run or a short run determinant of investment during the eighties and the early nineties, as we saw earlier in the within estimates, but only for France. The current effect of the cash-flow is negative in the second time period and this is not entirely wiped out by the lagged positive coefficients.[21] A test of significance of the three cash-flow parameters is rejected in the second period, while they are jointly significant in the first period for both countries. As we saw earlier, some of the effect of the profit rate in the accelerator-profit model was eliminated by using an error-correcting model and the remainder seems to have been removed by instrumenting for simultaneity between investment and cash flow, at least in the second period.

The sales coefficients are roughly the same for the within and GMM estimates but the former are much more precise. This is especially the case for France in the second period where the long-run coefficient is not significant, perhaps because there are no relevant instruments for sales growth in the instrument set. We cannot reject constant returns to scale for the U.S. firms, although we can for the French firms (t-statistics on the s_{t-2} coefficient of -2.4 and -2.3).

Thus these estimates reinforce the conclusion that whatever cash flow or profit effects might have existed twenty years earlier are greatly reduced in the late 1980s and early 1990s, especially for our French sample of large and fairly long-lived manufacturing firms, and also for our U.S. sample, but only when we instrument cash flows.

8 Conclusions

Our motivation for this study was to assess the effect of changes in modeling and estimating investment equations during the past twenty years, focusing on the implications of improvements in panel data econometrics for the estimation of investment equations using firm level data. Although we gave an overview of the evolution of the

[21] Note that this precisely what one would expect from the Euler equation interpretation of the investment equation: when output (sales) is in the model, the contemporaneous coefficient of the profit rate should be negative.

theoretical modelling of investment at the firm level from the traditional Jorgensonian approach through q-theory to the modern Euler equation specification, we have chosen to concentrate on a fairly robust error-corrected accelerator-profit model and explore the effects of changes in the empirical specification and econometric methodology.

We began with the traditional accelerator-profit model estimated with classical least-squares methods in the Total, Within or Between dimensions, using methods that were relatively new at the time of the first conference on Panel Data in 1977. We compared this econometric specification to one that adds an error-correction mechanism and showed that it was both more parsimonious and was able to disentangle the long-run from the short-run behavior of investment in a theoretically consistent way (unlike the somewhat imprecise intuition that cross sectional (between) estimates represent the long run and time series (within) estimates the short run.

The change in the specification from an accelerator model to an error corrected accelerator model because we cannot compare the estimates of the two specifications directly. That is, the accelerator-profit model is derived by differentiating the equilibrium relationship, whereas the error-corrected version retains the levels relationship as a part of the model. The implication is that firm-specific effects have a different meaning in the two models: for the accelerator, they imply heterogeneous growth rates for the capital stock, whereas for the error-corrected version, they imply heterogeneous capital-output ratios.

We have also investigated the use of modern GMM methods for the estimation of the model. GMM should correct for simultaneity biases coming from the endogeneity of variables or the presence of correlated firm-specific effects. But we find that the estimation results are not dramatically different when we use GMM instead of within estimation. The potential gain obtained by instrumenting the regressors is offset by a large imprecision in the estimated parameters due to the use of internal (to the model) and therefore "weak" instruments.

Finally what has changed in twenty years in the economic determinants of the firms' investment behavior? Our primary finding is that the profit or cash flow rate no longer enters the firm-level investment equation in either the United States or France, once we control for permanent differences across firms in investment behavior and the bias from measurement error and simultaneity. That is, although we find that the long run impact of sales on the desired level of capital is nearly one in the United States and positive but very imprecisely determined in France in our 1985-93 period, we find no role for the profit rate or cash flow in the long run relationship and almost none in the short run. This result contrasts with our result for the 1970s in both countries, even when we use the newer model and estimation methods.

The most disappointing feature of our investigation here is the low precision we obtain when using the newer GMM methods of estimation that are intended to correct for simultaneity bias, measurement error bias, and firm effects. In spite of this, we do not find the results reported here discouraging. On the contrary, our view is that they demonstrate the considerable progress made in the last twenty years in our understanding of the investment relation, leading to improvements in modeling and interpretation, and to a better comprehension of what can and cannot be measured. The investment equation is one of the most difficult relations to estimate empirically and it should not surprise us that progress is slow.

We see several ways in which one might make future progress, even within the framework we have outlined here (that is, a traditional smooth adjustment mechanism for the firm-level capital stock), and we intend to pursue them in future work. First, recent work by ARELLANO and BOVER [1995] and BLUNDELL and BOND [1995] has suggested that if equation (5) is the true model, it should also be possible to instrument the totals equation (which contains the fixed firm effect α_i) by lagged differences of the x's and y's, since these presumably do not contain any firm effects that are constant across time. Our initial attempts at this kind of specification typically found that these estimates were somewhat different from estimates obtained using levels of x and y to instrument differences of u, which implies that the assumptions required for consistency do not hold in our data. In a future version of this paper, we plan to test for the validity of these moment conditions using our differenced model as the maintained hypothesis.

Second, our samples of manufacturing firms are typically heterogeneous and thus far we have forced them all into a "one size fits all" investment model; it seems implausible to expect an automobile firm and a supplier of small-scale computer equipment to adjust their capital stock at the same rate and in response to the same shocks. Recent work by PESARAN, SHIN, and SMITH [1997] suggests a reasonable generalization of our model to allow for some heterogeneity across firms. They propose estimating a model of long-run relationships in heterogeneous dynamic panels by specifying a cointegrating relationship that is the same for all observations in the panel, but allowing the short-run adjustment dynamics to vary across the units. If we have a long enough time series for each firm, this is a very reasonable way to enrich our model of investment behavior by requiring the capital stock-sales ratio to be constant in the long-run, but to allow the speed of adjustment to this target vary across firms. Preliminary exploration using this model in our data yielded quite plausible results with a long-run coefficient of unity (when the data are in logarithms) and a range of short-run coefficients which were quite reasonable.

Finally, BOND, HARHOFF, and VAN REENEN [1997] present evidence that the investment behavior of R&D-performing and non-R&D-performing firms differs in the United Kingdom and Germany. In future work, we plan to investigate both whether this fact is also true in France and the United States, and whether R&D investment itself displays behavior similar to that described for investment in this paper.

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APPENDIX

A Derivation of the model.

The firm solves the problem:

$$Max \sum \beta^t \left[p_t f(K_t) - p_t^I I_t \right] \quad s.t. \quad K_{t+1} = (1-\delta)K_t + I_t.$$

When the prices of output and capital are constant over time, this yields the steady state solution

$$f'(K_t) = \frac{p_t^I}{p_t}(r+\delta),$$

where δ is the depreciation rate and r is the interest rate implicit in $\beta = (1+r)^{-1}$. When prices are allowed to vary over time, the solution is

$$f'(K_t) = \frac{p_t^I}{p_t}\left(r\frac{p_{t-1}^I}{p_t^I} + \delta - \frac{\Delta p_t^I}{p_t^I} \right) = C_t.$$

Thus there is an additional term in the relative price of investment that comes from the capital gain or loss on the existing capital. In panel data estimation, this will imply that year effects belong in the equation, regardless of whether real or nominal values of capital and output are used.

If the production function is Cobb-Douglas $f(L_t, K_t) = A_t L_t^{\beta} K_t^{\alpha}$, we obtain:

$$K_t = \alpha \frac{S_t}{C_t}$$

or in logarithmic form:

$$k_t = s_t + h_t \quad \text{where} \quad h_t = \log(\alpha) - c_t$$

More generally for a CES production function where σ and ν are respectively the elasticities of substitution and scale:

$$f(L_t, K_t) = A_t \left[\beta L_t^{\frac{\sigma-1}{\sigma}} + \alpha K_t^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}\nu},$$

we obtain:

$$k_t = \left(\sigma + \frac{1-\sigma}{\nu} \right) s_t + h_t \quad \text{where} \quad h_t = cst. - \sigma c_t.$$

Note that $k_t = s_t + h_t$ if there is constant returns to scale ($\nu = 1$) **or** if the elasticity of substitution is unity ($\sigma = 1$) corresponding to the Cobb-Douglas production function.

B Data Appendix

B.1 New Samples

The data for the U.S. firms are drawn from the 1995 Standard & Poor edition of Compustat; the active and research files for Annual Industrial and Full Coverage firms were merged (using both the current 1976-1995 files and the historical 1957-1976 files). Firms that are incorporated in a foreign country were deleted, as were wholly-owned subsidiaries and non-publicly-traded firms.

The data for French firms are drawn from the balance sheets and income accounts in the BIC-SUSE database of INSEE, covering firms with more than 20 employees from all industries during the period 1978-1993. This data reporting is compulsory, and based on the fiscal statements of the firms concerned. The companies may be publicly-traded on stock markets, but many of them are not.

The variables we used were the following:

- S: Total sales or turnover in millions of dollars or in thousands of French francs (FRF).
- E: Number of employees during year in thousands for US and at the date of accounts for France.
- I: Investment in fixed capital in millions of dollars or in thousands of FRF.
- CF: After-tax cash-flow = Net profit plus depreciation allowances, in millions of dollars or in thousands of FRF.
- K: Net capital stock in volume (at the end of the period) in millions of dollars or in thousands of FRF, computed by a perpetual inventory method with a constant rate of depreciation ($\delta = 8\%$):

  $$K_t = (1-\delta)K_{t-1} + \frac{I_t}{P_t^I}.$$

  The benchmark at the first year is the net fixed assets revaluated.

S and CF have been deflated by a sectorial production price index, while I has been deflated by the sectorial price of equipment goods (P^I).

In general, ratios and growth rates of these variables were trimmed so that one percent of the observations in the tails of each variable were removed.

In Table B1, we give the industrial composition of the firms in the second period: although the proportion of firms in the high-tech sectors such as computers, pharmaceuticals, chemicals, electrical machinery, and aircraft is about the same in the two countries, firms in the United States are more likely to have an R&D program (true for 74 percent as opposed to 56 percent in France) and those in high-tech are more likely to be in computers and instruments than in chemicals and pharmaceuticals. There are a variety of other minor differences documented in the table.

B.2 Old Samples

The old samples for the period 1968-1979 for France and the U.S. were described in MAIRESSE [1990]. The capital stock has been computed in the same manner as above with a 8% depreciation rate.