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Financial factors and investment in Belgium, France, Germany, and the United Kingdom: A comparison using company panel data

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FINANCIAL FACTORS AND INVESTMENT IN BELGIUM, FRANCE, GERMANY, AND THE UNITED KINGDOM: A COMPARISON USING COMPANY PANEL DATA

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Abstract—We construct company panel data sets for manufacturing firms in Belgium, France, Germany, and the United Kingdom, covering the period 1978–1989. These data sets are used to estimate empirical investment equations, and to investigate the role played by financial factors in each country. A robust finding is that cash flow and profits terms appear to be both statistically and quantitatively more significant in the United Kingdom than in the three continental European countries. This is consistent with the suggestion that financial constraints on investment may be relatively severe in the more market-oriented U.K. financial system.

I. Introduction

THERE is now a large microeconometric literature that investigates the role of financial factors in company investment decisions. Most studies find that financial variables such as cash flow help to explain investment spending. For some econometric models of investment, this relationship should not occur under the null hypothesis that company investment spending is not affected by financial constraints, and the evidence of *excess sensitivity* to cash flow is interpreted as suggesting the importance of such constraints. It is sometimes suggested that these financial constraints on investment may be the outcome of asymmetric information between firms and suppliers of finance. The excess sensitivity of investment to financial variables has been found to be less important for certain types of firms, such as those with close relationships to banks in Japan and Germany, and those which pay out high dividends in the United Kingdom and the United States.[1]

Once we move away from a model of perfect capital markets in which financial decisions and real investment decisions are separable, we raise the possibility that different financial systems may have different effects on company investment. Heterogeneity across countries has been well documented, for example in patterns of investment finance, corporate ownership patterns, corporate governance rules, the market for corporate control, and the relative importance of different financial markets and institutions.[2] Differences between Anglo-American *market-based* and German or Japanese *bank-based* systems have received particular attention. It is sometimes suggested that the arm's-length relation between firms and suppliers of finance that tends to characterize the market-oriented systems may be less effective in dealing with problems of asymmetric information. If so, one consequence may be a higher cost premium for the use of external sources of investment finance, and a more severe effect of financing constraints. Perhaps surprisingly, there has been little investigation of whether these differences between financial systems may be related to differences in the effect of financial constraints on investment.[3]

The aim of this paper is to begin an econometric investigation of this question. We construct company panel data sets for manufacturing firms in Belgium, France, Germany, and the United Kingdom, covering the period 1978–1989. These data sets are used to estimate empirical investment equations and to investigate the role played by financial factors in each country. The main focus of the investigation is to compare results for the same investment model across different countries, rather than to compare competing econometric specifications within each country. We therefore emphasize results that appear to be robust to the choice of model specification. The models are estimated using GMM methods which control for biases due to unobserved firm-specific effects and endogenous explanatory variables. Some OLS and within-groups results are reported for comparison, and suggest the importance of controlling for these biases.

We estimate two types of investment equations, an error-correction model and an Euler-equation specification. The Euler equation is a structural model, explicitly derived from a dynamic optimization problem under the assumption of symmetric, quadratic costs of adjustment. This has the advantage that, under the maintained structure, the model captures the influence of current expectations of future profitability on current investment decisions; and it can therefore be argued that current or lagged financial variables should not enter this specification merely as proxies for expected future profitability. Unfortunately, this maintained structure of adjustment costs is very restrictive, as has been emphasized in the recent literature on investment under irreversibility.[4] We therefore also estimate a reduced-form

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[1] See *inter alia* Fazzari, Hubbard, and Petersen (1988), Hoshi, Kashyap, and Scharfstein (1991), Bond and Meghir (1994), and Elston (1998). Schiantarelli (1996), Hubbard (1998), and Mairesse, Hall, and Mulkay (1999) provide surveys of this literature.

[2] See *inter alia* Mayer (1988, 1990) and Edwards and Fischer (1994).

[3] Hall et al. (1998) report interesting comparative results for samples of U.S., Japanese, and French firms. Bond, Harhoff, and Van Reenen (1999) report results for U.K. and German companies which complement those in this paper. See Mairesse and Dormont (1985) for an earlier comparative study.

[4] See, for example, Dixit and Pindyck (1994) and Caballero (1999).

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error-correction model, in which the long-run formulation for the level of the capital stock is specified to be consistent with a simple model of the firm's demand for capital, but in which the short-run investment dynamics are found from an empirical specification search, rather than being imposed a priori.

It is important to consider what can be learned about the effects of financing constraints from the comparison of results across countries. For any one sample, it is well known that a significant coefficient on (say) a cash flow variable can be accounted for by cash flow acting as a proxy for omitted expected future profitability variables, rather than signaling that fluctuations in the availability of internal finance affect investment as a result of financing constraints.[5] However, to the extent that the relationship between current cash flow and expected future profitability is similar across countries, it may be that large and significant differences across countries in the estimated coefficients on cash flow variables are more likely to reflect differences in the effects of financing constraints. This is essentially the type of *sample-splitting* test that has been widely used in the literature on financing constraints and investment since the work of Fazzari et al. (1988), where the sample-splitting criterion we focus on is the country where the firm is based. We also confirm that there are not big differences across our four samples in the ability of current or lagged cash flow variables to forecast future cash flow or sales growth.

Recently Kaplan and Zingales (1997) have criticized this methodology and presented a counterexample in which a firm that is more financially constrained, in the sense of facing a greater cost premium for the use of external finance, displays lower sensitivity of investment to fluctuations in cash flow.[6] Two observations on this critique are worth noting here. First, Kaplan and Zingales's analysis is conducted under the alternative hypothesis that all firms face financial constraints, and the question is whether differences in cash flow sensitivities are informative about differences in the degree of these constraints. It remains the case in their model that a firm facing no financial constraint (no cost premium for external finance) would display no excess sensitivity to cash flow. To the extent that we find insignificant cash flow effects in country A, and large and significant cash flow effects in country B, this remains consistent with the interpretation that financing constraints are unimportant in country A but may have a significant effect on investment in country B. Secondly, Kaplan and Zingales's example is derived in a model with no adjustment costs of any type; to the best of our knowledge, the robustness of their result has not been demonstrated in more realistic settings with adjustment costs or other impediments to capital-stock adjustment.

There are important differences between firms in the four countries we consider, and in the nature of the accounts data that were available for this study. The U.K. data refer to the consolidated accounts of company groups that are traded on the London Stock Exchange. The accounts data available for corporations in the other countries are generally not consolidated. The German data also refer only to stock-market-quoted firms. The data for France and Belgium cover a wider range of firms that report accounts in those countries, and include some unquoted companies as well as some subsidiaries of larger firms. As the proportion of corporate activity accounted for by firms quoted on national stock markets varies considerably across these countries, it would not necessarily be desirable to restrict attention to quoted firms. It would certainly be desirable to have more comparable accounting data, preferably on a consolidated basis, but such data were not available for the French and Belgian samples. However, we were able to consider a subsample of independent French firms, for which the issue of consolidation is irrelevant, and a subsample of German companies for which consolidated accounts were available, to investigate the effect of this accounting difference on our results. The effect of other differences between accounting rules in the four countries is minimized by using recorded cash flows wherever possible, and by estimating values for the capital stock from the investment flows on a standard basis for each sample.

Partly as a result of these differences in the data available, the U.K. and German companies we study tend to be much larger than the French and Belgian companies.[7] Nevertheless we find that our investment, capital-stock, and sales series have rather similar time series properties across the four data sets. However, the results of the econometric investment models reveal some interesting differences between the four countries. Financial variables are found to be insignificant in Belgium, and to have small and only weakly significant effects in Germany and France, but to have large and highly significant effects in the United Kingdom. This finding is consistent with the suggestion that financial constraints on investment may be more severe in the more market-oriented U.K. financial system than in the continental European countries.

The remainder of the paper is organized as follows. Section II briefly describes the investment models that we estimate in this study; section III describes the data sets we use; section IV presents our empirical results; and section V concludes with a discussion of these findings.

[5] See, for example, Fazzari, Hubbard, and Petersen (1988). Although this possibility is particularly transparent for reduced-form investment equations, which make no explicit attempt to control for expected future profitability, it will also apply to structural investment equations that are not correctly specified.

[6] See also Kaplan and Zingales (2000) and Fazzari et al. (2000).

[7] We use a subsample of the largest firms in France to investigate the sensitivity of our results to firm size.

II. The Empirical Investment Equations

We estimate two different econometric models of company investment, which allows us to consider the sensitivity of our empirical findings to the choice of model specification. The models we use are an error-correction model and an Euler equation. These are described in the following sections.

A. An Error-Correction Specification

The error-correction model was introduced into the investment literature by Bean (1981). The basic idea is to nest a long-run specification for the firm's demand for capital within a regression model that allows a flexible specification for short-run investment dynamics to be estimated from the data.

We start from the assumption that, in the absence of adjustment costs, the desired capital stock can be written as a log linear function of output and the cost of capital. Letting k_{it} denote the (natural) log of the desired capital stock for firm i in period t, y_{it} denote the log of output, and j_{it} denote the log of the real user cost of capital, we write the desired capital stock as

$$k_{it} = a_i + y_{it} - \sigma j_{it}. \tag{1}$$

This is consistent with profit maximization subject to constant returns to scale and a CES production function, and nests the possibility of a fixed capital-output ratio ($\sigma = 0$); the log linear formulation with $\sigma = 1$ is also consistent with a Cobb-Douglas production function, with or without constant returns to scale. The firm-specific intercept (a_i) may reflect a firm-specific markup parameter in a monopolistic competition framework, or a firm-specific distribution parameter in the production function.

To account for slow adjustment of the actual capital stock to the desired capital stock, we nest this within a general dynamic regression model. Implicitly this assumes that the firm's desired capital stock in the presence of adjustment costs is proportional to its desired capital stock in the absence of adjustment costs,[8] and that the short-run investment dynamics are stable enough over the sample period to be well approximated by the distributed lags in the regression model. We also assume that variation in the user cost of capital can be controlled for by including both time-specific and firm-specific effects. For example, if we consider an autoregressive-distributed lag specification with up to second-order dynamics [an ADL(2,2) model], we have

$$\begin{aligned} k_{it} = {} & \alpha_1 k_{i,t-1} + \alpha_2 k_{i,t-2} + \beta_0 y_{it} + \beta_1 y_{i,t-1} \\ & + \beta_2 y_{i,t-2} + d_t + \eta_i + v_{it}, \end{aligned} \tag{2}$$

where d_t is a time dummy, η_i is an unobserved firm-specific effect and v_{it} is an error term. Here we require the restriction $(\beta_0 + \beta_1 + \beta_2)/(1 - \alpha_1 - \alpha_2) = 1$ to be consistent with the long-run unit elasticity of capital with respect to output implied by equation (1).

It is convenient to reparameterize this ADL model in error-correction form. Imposing the long-run unit-elasticity restriction, this gives

$$\begin{aligned} \Delta k_{it} = {} & (\alpha_1 - 1)\, \Delta k_{i,t-1} + \beta_0\, \Delta y_{it} + (\beta_0 + \beta_1)\, \Delta y_{i,t-1} \\ & - (1 - \alpha_1 - \alpha_2)(k_{i,t-2} - y_{i,t-2}) \\ & + d_t + \eta_i + v_{it}. \end{aligned} \tag{3}$$

The validity of this restriction on the long-run properties of the model will be investigated in our empirical analysis. Error-correcting behavior then requires that the coefficient on the error-correction term $(k_{i,t-2} - y_{i,t-2})$ be negative, so that a capital stock above its desired level is associated with lower future investment, and vice versa.

Finally, letting I_{it} denote gross investment, K_{it} the capital stock, and δ_i the (possibly firm-specific) rate of depreciation, we use the approximation $\Delta k_{it} \approx I_{it}/K_{i,t-1} - \delta_i$ to obtain a specification for the investment rate. To investigate the role of financial variables, we include additional current and lagged cash flow (C_{it}) terms. The error-correction model we estimate then has the form

$$\begin{aligned} \frac{I_{it}}{K_{i,t-1}} = {} & \rho \frac{I_{i,t-1}}{K_{i,t-2}} + \gamma_0\, \Delta y_{it} + \gamma_1\, \Delta y_{i,t-1} \\ & + \phi(k_{i,t-2} - y_{i,t-2}) + \pi_0 \frac{C_{it}}{K_{i,t-1}} \\ & + \pi_1 \frac{C_{i,t-1}}{K_{i,t-2}} + d_t + \eta_i + v_{it}. \end{aligned} \tag{4}$$

As we emphasized in the introduction, the interpretation of these additional financial variables in this type of investment equation is ambiguous. Although a significant cash flow effect could reflect the presence of financial constraints on investment, it is also possible that such terms would be significant in the absence of financial constraints. By the presence of financial constraints on investment, we mean a situation where a *windfall* increase in cash flow—one that conveyed no new information about future profitability or investment opportunities—would nevertheless be associated with a rise in investment spending. However, if the firm faces strictly convex adjustment costs, for example, one can show that current investment depends on expected future changes in the desired stock of capital.[9] Then if information on cash flow helps to forecast output, for example, such information on cash flow will help to explain investment spending in a reduced-form model, even in the absence of financial constraints. For this reason we focus on differences

[8] A similar proportionality assumption is used by Caballero, Engel, and Haltiwanger (1995).

[9] See Nickell (1978, chapter 11), for example.

in the effects of cash flow variables between samples of firms located in different countries, subject to the qualifications noted in the introduction. We also investigate directly whether current or lagged cash flow variables forecast future sales growth or profitability differently across our four data sets.

B. An Euler-Equation Specification

Whereas the error-correction model can be regarded as an empirical generalization of the first-order condition for the optimal capital stock in a static factor demand model [such as equation (1)], the Euler equation is based on an explicit theoretical generalization of this first-order condition to the case of strictly convex costs of adjustment. Euler-equation models were introduced into the investment literature by Abel (1980). The version of the Euler-equation model we estimate is based on Bond and Meghir (1994). This is a relation between investment rates in successive periods, derived from dynamic optimization in the presence of symmetric, quadratic adjustment costs. Under these assumptions, and as long as we assume that expectations are formed accordingly, the Euler-equation model has the advantage of controlling for all expectational influences on the investment decision. Evidence of misspecification associated with the role of financial variables in this model is less easily explained away as merely capturing an expectational influence.[10]

The firm is assumed to maximize the present discounted value of current and future net cash flows. Letting L_{it} denote variable factor inputs, w_{it} the price of variable factors, p^I_{it} the price of investment goods, p_{it} the price of output, β^t_{t+j} the nominal discount factor between period t and period $t + j$, δ the rate of depreciation, $F(K_{it}, L_{it})$ the production function gross of adjustment costs, $G(I_{it}, K_{it})$ the adjustment cost function, and $E_t(\cdot)$ the expectation operator conditional on information available in period t, the firm solves

$$\max E_t\left(\sum_{j=0}^{\infty} \beta^t_{t+j} R(K_{i,t+j}, L_{i,t+j}, I_{i,t+j})\right) \qquad \text{s.t.} \quad K_{it} = (1-\delta)K_{i,t-1} + I_{it}, \tag{5}$$

where $R_{it} = p_{it}F(K_{it}, L_{it}) - p_{it}G(I_{it}, K_{it}) - w_{it}L_{it} - p^I_{it}I_{it}$.

The Euler equation characterizing the optimal investment path relates marginal adjustment costs in adjacent periods. This can be written as

$$-\left(\frac{\partial R}{\partial I}\right)_{it} = -(1-\delta)\beta^t_{t+1}E_t\left(\frac{\partial R}{\partial I}\right)_{i,t+1} + \left(\frac{\partial R}{\partial K}\right)_{it}. \tag{6}$$

Assuming competitive markets and that $F(K_{it}, L_{it})$ is constant returns to scale, and specifying $G(I_{it}, K_{it}) = \frac{b}{2}[(I/K)_{it} - c]^2K_{it}$, this can be expressed as

$$\left(\frac{I}{K}\right)_{it} - \alpha_1\left(\frac{I}{K}\right)^2_{it} = \alpha_2E_t\left(\frac{I}{K}\right)_{i,t+1} + \alpha_3\left[\left(\frac{\Pi}{K}\right)_{it} - J_{it}\right] + \alpha_0, \tag{7}$$

where

$$\Pi_{it} = p_{it}F(K_{it}, L_{it}) - p_{it}G(I_{it}, K_{it}) - w_{it}L_{it}$$

is the gross operating profit and J_{it} is the real user cost of capital. The coefficients α_1, α_2, and α_3 can be shown to be positive, and the left-hand side of equation (7) is increasing in $(I/K)_{it}$. Current investment is positively related to expected future investment and to the current-average-profits term (reflecting the marginal profitability of capital under constant returns), and negatively related to the user cost of capital. An attractive feature of the Euler-equation model is that all relevant expectational influences are captured by the one-step-ahead investment forecast.

To implement this model, we replace the unobserved $E_t(I/K)_{i,t+1}$ by the realized $(I/K)_{i,t+1}$ plus a forecast error, and take this $(I/K)_{i,t+1}$ term to the left-hand side to obtain an econometric model that is linear in variables.[11] We also replace the cost-of-capital term by time effects and firm-specific effects, and include a term in the output/capital ratio that may be introduced either by nonconstant returns to scale or by monopolistic competition in the product market. The resulting empirical specification is

$$\left(\frac{I}{K}\right)_{i,t+1} = \beta_1\left(\frac{I}{K}\right)_{it} - \beta_2\left(\frac{I}{K}\right)^2_{it} - \beta_3\left(\frac{\Pi}{K}\right)_{it} + \beta_4\left(\frac{Y}{K}\right)_{it} + d_{t+1} + \eta_i + v_{i,t+1}. \tag{8}$$

Unlike the error-correction model, the structure of the Euler-equation model should control for the influence of financial variables on expectations of future profitability, at least in the case of symmetric, quadratic adjustment costs. Under the null of no financial constraints, it can be shown that $\beta_1 \geq 1$, $\beta_2 \geq 1$, $\beta_3 > 0$, and (under constant returns to scale) $\beta_4 \geq 0$.[12] Under the alternative, investment spending

[10] The same comment applies to the Q-model. We do not consider a Q-model here, because we wish to include unquoted companies in our samples, and because stock-market data were not available in all four countries. Another possibility would be to construct a measure of marginal q based on econometric forecasts of future cash flow rates, as in Abel and Blanchard (1986) and Gilchrist and Himmelberg (1995). However, this approach does not have a structural interpretation in the usual adjustment-costs model, as discussed in footnote 5 of Abel and Blanchard (1986). Some problems with investment Euler equations are discussed in Oliner, Rudebusch, and Sichel (1995).

[11] Notice that this normalization switches the signs on $(\Pi/K)_{it}$ and J_{it}. Other normalizations are considered by Johansen (1994) and Whited (1992).

[12] It is possible to show that $\beta_1 = (1 + c\nu)/\psi$, $\beta_2 = (1 + \nu)/2\psi$, $\beta_3 = \mu/(b\psi)$, and $\beta_4 = (\mu - \nu)/(b\psi)$, where $\psi = \beta^t_{t+1}(1 - \delta)(p_{t+1}/p_t)$ is treated as constant, μ is the markup coefficient in a monopolistic compe-

is positively related to cash flow or profits through the effect of financial constraints. The basic Euler equation (8) is then misspecified. Since the gross operating profits term $(\Pi/K)_{it}$ in equation (8) will be highly correlated with cash flow, the prediction of a negative sign on this term may be expected to fail in the presence of financial constraints.

The main aim of our study is to investigate whether robust results are obtained across countries from these alternative investment models, not to evaluate them as rival specifications. It would not make much sense to compare, in terms of goodness of fit, the largely empirically derived error-correction model with the more structural Euler equation. Moreover, the validity of these equations is not mutually exclusive. The Euler equation is not inconsistent with the CES assumption used to obtain the error-correction model; and the error correction model is not incompatible with the assumption of symmetric, quadratic adjustment costs.[13] If we were completely confident that adjustment costs take this particular form, there would be little gained by considering the error-correction model in addition to the Euler equation. Otherwise the error-correction model can be regarded as an empirical approximation to some more general adjustment process, albeit at the cost of compounding influences from the expectations-formation mechanism and the adjustment dynamics into the same reduced-form regression coefficients.

III. Data

We use panel data on company accounts covering the period 1978–1989.[14] All firms have their main activity in manufacturing, and firms with fewer than 100 employees in their first year in our sample were excluded. Firms that had engaged in major merger or acquisition activity were also excluded wherever possible, as the standard models of investment may not characterize these discrete adjustments very well.

The U.K. sample comprises 571 firms quoted on the London Stock Exchange for which consolidated accounts data were available from Datastream. Some of these companies have branches and subsidiaries overseas, whose activities will be included in these data. The French and Belgian samples comprise 1,365 firms and 361 firms respectively, for which unconsolidated accounts data have been collected by INSEE in France and the central bank in Belgium. These need not be stock-market-quoted companies, and may include subsidiaries of foreign companies.[15] The German sample comprises some 228 quoted *Aktiengesellschaft* (AG) corporations, for which unconsolidated accounts data were available from the Bonn Data Bank. This sample contains most of the quoted manufacturing AG firms in Germany for which sufficient years of data were available.

The main variables we use are flows of investment, sales, and gross operating profits. Investment spending is obtained from the account of sources and uses of funds, and not inferred from changes in the balance sheet. For Germany and the United Kingdom, we construct a measure of cash flow by adding back reported depreciation to reported profits net of interest and taxes. For Belgium and France, we obtain an equivalent measure of cash flow by subtracting costs of materials, the wage bill, interest, and taxes from sales. For comparability across countries, we use real sales as a proxy for output, even though a measure of value added was available from the company accounts in the Belgian and French data. Experiments showed that very similar results were obtained for these countries when this measure was used instead of sales.

A measure of the stock of capital at current replacement cost was estimated from the flow data on investment using a standard perpetual inventory method for each sample. The starting value was based on the net book value of tangible fixed capital assets, adjusted for previous years' inflation. Subsequent values were obtained using accounts data on investment and disposals, national price indices for investment goods prices, and a depreciation rate of 8% assumed to be common to all countries. Further details of this calculation can be found in the data appendix. We have also experimented with other measures of the capital stock based on the reported gross and net book values of tangible assets; our results remained very similar when using these alternative measures.

Table 1 presents some basic features of these data sets. The size distribution of all the samples is highly skewed, with mean employment 2–7 times higher than median employment. The U.K. and German firms are clearly much larger on average than those in our French and Belgian samples, and the former samples are also more skewed. The French and Belgian firms had similar employment levels on average in 1985, but the French sample contains some much larger firms than the Belgian sample.

Table 2 reports the mean values of the variables used in our econometric analysis between 1981 and 1989. The investment rates (I/K) appear very similar on average in these data sets. However, the average growth of real sales (Δy) is larger on average for Belgium and the United Kingdom than for our French and German samples. The cash flow rates (C/K) and the gross operating profit rates

tition framework, and ν is the returns to scale of the gross production function.

[13] See Nickell (1985) for further discussion of the links between adjustment costs and error-correction models.

[14] We do not have access to French and Belgian data for earlier years. Data are not readily available on a consistent basis for German firms after 1989, following a major reform of German accounting law (*Bilanzrichtliniengesetz*). We thus maintain this common sample period to facilitate comparison of results across the four countries.

[15] In principle, the investment of French or Belgian subsidiaries of U.K. companies could appear in both samples, although this is unlikely to be very common.

TABLE 1.—SOME BASIC FEATURES OF THE DATASETS

Statistic	Belgium	France	Germany	U.K.
Size of the Samples (1981–1989)				
Firms	361	1,365	228	571
Observations	2,571	9,485	1,797	4,036
Obs./firm	7.12	6.95	7.88	7.07
Employment in 1985				
Mean	777	939	6,944	6,342
Std. dev.	1,300	3,838	22,215	21,117
Q25	231	180	425	394
Median	399	332	980	983
Q75	732	659	2,991	3,261
Maximum	11,670	91,049	202,200	312,000

(Π/K) are quite similar across the four samples. The sales/capital ratio (Y/K) is smaller on average in our U.K. sample, which may reflect in part the average size of the firms in the samples and in part the netting out of intragroup sales in these consolidated accounts.

IV. Empirical Results

A. Time Series Properties

We begin our empirical investigation by considering the time series properties of our main variables. This is important for two reasons. First, the long-run unit elasticity of capital with respect to output, imposed in our error-correction specification, implies that if the logs of both capital stock and real sales measures are $I(1)$ variables and the log of the real user cost of capital is $I(0)$, then the log of the capital/sales ratio should also be an $I(0)$ variable [see equation (1)], implying that the logs of capital and real sales are cointegrated. We can test this restriction by considering the unit root properties of the capital, sales, and capital/sales ratio. Secondly, random-walk properties for any of the variables used in the investment equations could lead to an identification problem for our preferred GMM estimates, which rely on lagged levels of the series providing informative instruments for subsequent first differences.

Given that our panels contain large numbers of firms observed for a small number of time periods, and that asymptotic distributional approximations rely on the number of firms becoming large with the number of periods assumed fixed, we can test the null hypothesis of unit roots by estimating simple AR(1) specifications using ordinary least squares (OLS) and considering conventional t-tests.[16] Table 3 reports these OLS estimates and heteroskedasticity-consistent asymptotic standard errors for the log of the capital stock, real sales, and capital/sales series. Within-groups and first-differenced GMM estimates are also reported for comparison, and will be discussed further below.[17]

For our capital-stock series the unit root hypothesis is not rejected for three of the four samples, and for our real-sales series the unit root hypothesis is not rejected for two of the four samples. As the point estimates exceed 0.99 in the remain-

TABLE 2.—MEANS (STANDARD DEVIATIONS) OF VARIABLES USED IN ESTIMATION (PERIOD: 1981–1989)

Variable	Belgium	France	Germany	U.K.
I_t/K_{t-1}	0.125	0.110	0.122	0.117
	(0.107)	(0.089)	(0.079)	(0.110)
Δy_t	0.027	0.010	0.005	0.033
	(0.145)	(0.125)	(0.143)	(0.177)
$(k - y)_{t-2}$	−0.993	−0.926	−0.827	−0.658
	(0.612)	(0.546)	(0.564)	(0.499)
C_t/K_{t-1}	0.178	0.119	0.160	0.134
	(0.144)	(0.136)	(0.093)	(0.106)
I_{t-1}/K_{t-1}	0.111	0.103	0.117	0.102
	(0.076)	(0.067)	(0.061)	(0.073)
$I_{t-1}/K_{t-1})^2$	0.018	0.015	0.017	0.016
	(0.027)	(0.022)	(0.020)	(0.027)
Π_{t-1}/K_{t-1}	0.239	0.222	0.218	0.198
	(0.163)	(0.176)	(0.124)	(0.134)
Y_{t-1}/K_{t-1}	3.247	2.868	2.578	2.186
	(2.195)	(1.665)	(1.629)	(1.260)

TABLE 3.—AR(1) MODELS FOR k, y, AND $k - y$

	ln (Capital Stock), k			
	Belgium	France	Germany	U.K.
OLS	0.990	0.999	1.003	1.000
	(0.002)	(0.001)	(0.002)	(0.001)
Within	0.845	0.861	0.849	0.924
	(0.019)	(0.010)	(0.019)	(0.015)
GMM	0.802	0.716	0.556	0.742
	(0.052)	(0.027)	(0.064)	(0.045)
Sargan	0.000	0.000	0.233	0.269
	ln (Real Sales), y			
	Belgium	France	Germany	U.K.
OLS	0.996	1.003	1.001	0.993
	(0.003)	(0.001)	(0.002)	(0.002)
Within	0.688	0.763	0.664	0.801
	(0.029)	(0.013)	(0.033)	(0.018)
GMM	0.375	0.667	0.090	0.887
	(0.195)	(0.150)	(0.270)	(0.077)
Sargan	0.081	0.011	0.567	0.000
	Error Correction Term, $k - y$			
	Belgium	France	Germany	U.K.
OLS	0.967	0.964	0.961	0.958
	(0.006)	(0.003)	(0.007)	(0.006)
Within	0.607	0.668	0.557	0.685
	(0.028)	(0.013)	(0.039)	(0.023)
GMM	0.785	0.935	0.664	0.976
	(0.078)	(0.039)	(0.152)	(0.085)
Sargan	0.556	0.031	0.081	0.703

[16] See Bond, Nauges, and Windmeijer (2001) and Hall and Mairesse (2001). OLS estimates of autoregressive coefficients will be biased upwards under the alternative of a stationary AR(1) process with unobserved firm-specific effects, so finding OLS estimates significantly below 1 provides evidence against the unit root hypothesis. The test procedure could be misleading if higher-order dynamics are present, or if the series are measured with error. However, we obtained very similar results considering OLS and IV estimates of AR(2) specifications.

[17] All results were computed using DPD98 for Gauss (Arellano and Bond, 1998), and year dummies are included in all the specifications we report. The GMM results reported are one-step estimates. All standard errors reported are asymptotically robust to heteroskedasticity.

ing cases, it seems reasonable to regard both series as $I(1)$ or borderline $I(1)$ variables. However, for the log of the capital/sales ratio we find that the unit root hypothesis is rejected for all four countries. This suggests that the capital-stock and real-sales series are cointegrated, consistent with the long-run proportionality imposed in our error-correction model.

Stationarity of this error-correction term is also important for the estimation of our error-correction model of investment. Consistent estimation of dynamic models like equation (4) or (8) should allow for the presence of unobserved firm-specific effects and the fact that our panels cover relatively few time periods. Lagged dependent variables are necessarily correlated with firm-specific effects, so OLS levels estimates are subject to an omitted variable bias. The within-groups estimator, which is OLS after transforming the data to deviations from firm means, eliminates the firm-specific effects, but the transformation itself induces a simultaneity bias in autoregressive models estimated from short panels. In the baseline AR(1) specification

$$x_{it} = \alpha x_{i,t-1} + (1 - \alpha)\eta_i + v_{it}, \tag{9}$$

which approaches a random walk as $\alpha \to 1$, the OLS estimate of the autoregressive coefficient α can be shown to be biased upwards for $\alpha < 1$, whereas the within-groups estimate can be shown to be biased downwards.[18]

To avoid these biases, our main results are estimated using a first-differenced GMM estimator which eliminates the firm-specific effects by differencing the equations, and then uses lagged values of endogenous variables as instruments. If the error term (v_{it}) in levels is serially uncorrelated, then the error term in first differences is MA(1), and instruments dated $t - 2$ and earlier should be valid in the differenced equations. Under this assumption, consistent parameter estimates can be obtained. If the error term in levels is itself MA(1), then only instruments dated $t - 3$ and earlier will be valid; and so on. We test the validity of the instruments used by reporting both a Sargan test of the overidentifying restrictions, and direct tests of serial correlation in the residuals.[19]

One potential concern with this estimator relates to identification and weak instruments. In the AR(1) model (9), first differences Δx_{it} are uncorrelated with past information if $\alpha = 1$, so the first-differenced GMM estimator does not identify α in this case. More generally, Blundell and Bond (1998) show that the correlation between Δx_{it} and lagged levels of the series becomes weak as $\alpha \to 1$, and the first-differenced GMM estimator can be subject to large finite-sample biases. Simulations reported in Blundell and Bond (1998) indicate that the first-differenced GMM estimates of α are biased downwards in this case.[20]

[18] See respectively Hsiao (1986) and Nickell (1981).

[19] See Arellano and Bond (1991) for further details of these procedures.

[20] Nelson and Startz (1990) and Staiger and Stock (1997) provide more general analyses of the effects of weak instruments on instrumental variables estimators.

TABLE 4.—AR(1) MODELS FOR I/K, Δy, AND C/K

	Belgium	France	Germany	U.K.
Investment Rate, I_t/K_{t-1}				
OLS	0.332	0.383	0.386	0.418
	(0.028)	(0.018)	(0.040)	(0.030)
Within	0.077	0.057	0.133	0.068
	(0.028)	(0.019)	(0.043)	(0.027)
GMM	0.261	0.191	0.302	0.173
	(0.048)	(0.023)	(0.056)	(0.044)
Sargan	0.002	0.862	0.436	0.659
Real Sales Growth, Δy				
OLS	0.031	0.080	−0.027	0.213
	(0.050)	(0.024)	(0.059)	(0.022)
Within	−0.149	−0.111	−0.173	0.020
	(0.043)	(0.020)	(0.052)	(0.022)
GMM	−0.025	0.014	−0.094	0.117
	(0.055)	(0.027)	(0.070)	(0.029)
Sargan	0.173	0.635	0.314	0.573
Cash Flow Rate, C_t/K_{t-1}				
OLS	0.715	0.695	0.629	0.893
	(0.024)	(0.015)	(0.034)	(0.016)
Within	0.266	0.307	0.174	0.553
	(0.034)	(0.020)	(0.035)	(0.028)
GMM	0.371	0.347	0.291	0.706
	(0.067)	(0.033)	(0.057)	(0.093)
Sargan	0.110	0.062	0.049	0.691

One indication of whether these biases are likely to be serious in practice can be obtained by comparing the GMM estimates of α in these AR(1) models with the OLS levels and within-groups estimates, which are likely to be biased upwards and downwards respectively. For the log of the capital/sales-ratio series in table 3, we find that the GMM estimates lie above the within estimates in all four countries.[21] In contrast, for the capital-stock series the GMM estimates lie below the within estimates in all four samples, and this occurs for the sales series in three of the four samples. We tentatively conclude that this finite-sample bias due to weak instruments is unlikely to be a major problem provided we impose the long-run unit-elasticity restriction and work with the stationary capital/sales-ratio series in our error-correction model.[22]

In table 4 we report alternative estimates of AR(1) specifications for the remaining variables used in our error-correction model. For all three variables we clearly reject the hypothesis of a unit root, and in all cases the first-differenced GMM estimates lie above the corresponding within-groups estimates. There do not appear to be serious

[21] The instruments used to compute these GMM estimates were lagged levels of the series dated $t - 2$ and $t - 3$. We report p-values for the Sargan test (that is, the probability of generating the calculated Sargan statistic under the null of valid instruments), which indicate a possible bias in our results for the log of the capital/sales ratio in France.

[22] Blundell and Bond (2000) also find that imposing a constant-returns-to-scale restriction can improve the properties of the first-differenced GMM estimator in the context of estimating a production function.

TABLE 5.—ERROR-CORRECTION MODELS: GMM FIRST DIFFERENCES, $t - 2$ INSTRUMENTS

	Belgium	France	Germany	U.K.
I_{t-1}/K_{t-2}	0.003	0.027	0.096	−0.015
	(0.053)	(0.032)	(0.075)	(0.049)
Δy_t	0.189	0.151	0.017	0.179
	(0.073)	(0.043)	(0.036)	(0.065)
Δy_{t-1}	0.234	0.119	0.118	0.055
	(0.048)	(0.023)	(0.034)	(0.033)
$(k - y)_{t-2}$	−0.216	−0.115	−0.134	−0.071
	(0.047)	(0.023)	(0.039)	(0.038)
C_t/K_{t-1}	−0.055	−0.033	0.180	0.520
	(0.087)	(0.066)	(0.071)	(0.168)
C_{t-1}/K_{t-2}	0.080	0.086	−0.013	0.025
	(0.050)	(0.028)	(0.040)	(0.120)
m1	−7.04	−12.25	−6.46	−7.70
m2	−0.04	−1.01	−1.78	−1.55
Sargan	0.518	0.046	0.318	0.223

biases associated with the use of lagged levels as instruments for subsequent first differences of these series in our samples. The investment-rate series are found to be positively autocorrelated in all four countries, with our preferred estimates of α being in the range 0.2–0.3. We find no persistence in real sales growth, except in the United Kingdom, but more persistence in cash flow rates, particularly in the United Kingdom.

B. Investment Equations

Table 5 reports our GMM results for an error-correction model of the form outlined in equation (4). The instruments used were the lagged values of all right-side variables dated $t - 2, t - 3, \ldots, t - 6$, which allows for contemporaneous correlation between these variables and shocks to the investment equation, as well as correlation with unobserved firm-specific effects.[23]

We find that the error-correction terms are correctly signed and significantly different from zero in all four countries, and that sales growth has a positive short-run effect on investment rates that is statistically significant in all four samples. Perhaps more interesting are the differences across countries in the coefficients on the cash-flow variables. In our Belgian sample, we find that neither current nor lagged cash flow has a significant effect on investment, even in this reduced-form specification.[24] At the other extreme, in our U.K. sample we find that current cash flow has a large positive and highly significant coefficient. Results for the German and French samples are intermediate. Current cash flow has a significant coefficient in Germany, but the long-run effect of an increase in cash flow on the firm's capital stock is much smaller than in the United Kingdom. Lagged cash flow has a statistically significant coefficient in France, but the long-run effect of cash flow on the capital stock is even smaller in this case. Based on our results for the error-correction model, the sensitivity of investment spending to fluctuations in cash flow appears to be much greater in the United Kingdom than it is in Belgium, France, or Germany.[25]

For comparison, table 6 reports the within-groups estimates for this model. Noting from equation (3) that the coefficient on the error-correction term, $\alpha_1 + \alpha_2 - 1$, depends on the sum of the autoregressive parameters in the underlying ADL specification (2), we can first observe that the GMM estimates of $\alpha_1 + \alpha_2$ lie above the within-groups estimates in three of the four samples. The within-groups estimates suggest significant positive coefficients on current sales growth in Germany and on current cash flow in Belgium. These differences from the GMM estimates are both consistent with the possibility of simultaneity biases affecting the within-groups results. However, the result that cash flow has a larger effect in the U.K. sample is again found in these within-groups estimates. We have experimented with a variety of alternative instrument sets and more general dynamic specifications, and found that this pattern is a robust feature of the results for our error-correction models of investment.

One possibility noted earlier is that cash flow plays a more prominent role in this kind of reduced-form investment equation in one sample simply because current or lagged cash flow variables are more useful for forecasting future sales growth or future profitability in that sample. In part this is the motivation for considering structural Euler-equation models, which control for the influence of expected future profitability on current investment decisions, albeit under a restrictive assumption about the form of adjustment costs. We can also consider this issue more

[23] That is, both current sales and current cash flow are treated as potentially endogenous variables in the investment equation. In addition to the Sargan test, we also report direct tests for first-order (m1) and second-order (m2) serial correlation in the differenced residuals. These are asymptotically standard normal under the null of no serial correlation, and indicate the expected first-order moving-average serial correlation in our first-differenced residuals.

[24] The joint test that the coefficients on the current and lagged cash-flow terms are both zero has a p-value of 0.273.

[25] It is perhaps worth noting that we found significant cash-flow effects for all four countries in more restricted reduced-form specifications. There is some indication that the cash-flow variables proxy for omitted dynamics in these simpler dynamic specifications.

TABLE 6.—ERROR-CORRECTION MODELS WITHIN GROUPS

	Belgium	France	Germany	U.K.
I_{t-1}/K_{t-2}	−0.034	−0.068	−0.015	−0.087
	(0.029)	(0.020)	(0.053)	(0.027)
Δy_t	0.137	0.148	0.068	0.177
	(0.021)	(0.010)	(0.017)	(0.019)
Δy_{t-1}	0.148	0.124	0.113	0.057
	(0.023)	(0.010)	(0.016)	(0.015)
$(k - y)_{t-2}$	−0.175	−0.157	−0.164	−0.092
	(0.020)	(0.010)	(0.022)	(0.017)
C_t/K_{t-1}	0.130	0.008	0.209	0.250
	(0.038)	(0.013)	(0.035)	(0.052)
C_{t-1}/K_{t-2}	0.030	0.068	0.016	0.185
	(0.032)	(0.011)	(0.026)	(0.053)
m1	−7.50	−13.16	−6.50	−8.37
m2	−0.74	−1.88	−1.95	−2.02

directly by estimating simple forecasting models for sales growth and cash flow rates, as a function of lagged cash flow and other variables. Table A1 in the results appendix reports OLS estimates of a simple forecasting model for real sales growth, whilst table A2 reports corresponding results for a simple model of the cash flow rate (C_t/K_{t-1}).

Lagged investment rates predict future sales growth in all four countries. Consistent with the AR(1) specifications in table 4, lagged sales growth is informative only in the United Kingdom. Lagged cash flow does help to forecast sales growth in three of the four samples, but is not noticeably more informative in the U.K. sample than in Germany or Belgium. Lagged sales growth predicts future cash flow rates in all four countries, and cash flow rates are negatively related to past investment. Lagged cash flow does help to forecast future cash flow rates in all four countries. Consistent with the AR(1) models in table 4, there is some evidence that the cash flow rate is more persistent in the U.K. sample than in the other three countries. However, the sum of the coefficients on the two cash-flow terms is not much higher for the United Kingdom than for Belgium, where we find no significant effect of cash flow in our error-correction investment models. These results do not suggest that cash flow would play a very different role in conditioning expectations of future sales growth or future profitability in the United Kingdom from that in the three continental European countries, although this may account for part of the difference in the cash-flow effects found in our reduced-form investment equations.

To investigate this possibility further, table 7 reports GMM estimates for the Euler-equation specification set out in equation (8). The instrument set used here includes variables dated $t - 2$, which were found to be invalid instruments for this specification in an earlier study using U.K. data by Bond and Meghir (1994).[26] The coefficients on the lagged investment terms are correctly signed in all four samples, but much smaller in absolute value than suggested by the derivation of this model in the absence of financial constraints on investment.[27] The coefficient on the gross-operating-profits term is positive in all four cases, and significantly different from zero for France and the United Kingdom, which is contrary to the theoretical prediction under the null of no financial constraints. This effect is again much stronger in the U.K. sample, and the overidentifying restrictions are rejected, as would be expected if the model is seriously misspecified. Our results for the U.K. sample are very similar to those obtained using $t - 2$ instruments by Bond and Meghir (1994).

TABLE 7.—EULER-EQUATION MODELS: GMM FIRST DIFFERENCES, $t - 2$ INSTRUMENTS

	Belgium	France	Germany	U.K.
$(I/K)_{t-1}$	0.426	0.366	0.388	0.434
	(0.092)	(0.039)	(0.111)	(0.079)
$(I/K)^2_{t-1}$	−0.466	−0.465	−0.328	−0.715
	(0.248)	(0.106)	(0.305)	(0.197)
$(\Pi/K)_{t-1}$	0.008	0.059	0.024	0.214
	(0.032)	(0.012)	(0.031)	(0.040)
$(Y/K)_{t-1}$	0.024	0.009	0.013	0.005
	(0.005)	(0.005)	(0.006)	(0.007)
m1	−9.77	−17.91	−8.16	−10.36
m2	0.22	−0.08	−1.71	−1.96
Sargan	0.398	0.349	0.387	0.031

TABLE 8.—EULER-EQUATION MODELS: GMM FIRST DIFFERENCES, $t - 3$ INSTRUMENTS

	Belgium	France	Germany	U.K.
$(I/K)_{t-1}$	0.507	0.600	0.129	0.767
	(0.311)	(0.139)	(0.259)	(0.311)
$(I/K)^2_{t-1}$	−1.073	−1.304	−0.371	−1.854
	(0.916)	(0.414)	(0.774)	(0.670)
$(\Pi/K)_{t-1}$	0.057	0.040	0.102	0.185
	(0.056)	(0.024)	(0.058)	(0.056)
$(Y/K)_{t-1}$	0.024	0.016	0.006	0.004
	(0.007)	(0.005)	(0.006)	(0.010)
m1	−4.78	−8.02	−3.24	−5.11
m2	−0.89	−0.47	−2.57	−1.85
Sargan	0.362	0.246	0.509	0.056

In table 8 we report GMM estimates of the Euler-equation model using only instruments dated $t - 3$ and earlier. The exclusion of instruments dated $t - 2$ substantially reduces the precision of the parameter estimates. In our smallest sample, for Germany, we then fail to identify any significant investment dynamics. However in the other three samples, the coefficients on the lagged investment terms increase in absolute value towards the values that should characterize the adjustment of capital in the quadratic adjustment costs model. For Belgium and France, the coefficient on gross operating profits is not significantly positive in these results. For the United Kingdom, however, the positive coefficient on the profits term remains large and highly significant, and the validity of the overidentifying restrictions remains doubtful.

For each of the investment models we have considered, the cash-flow and profits variables appear to play a much more important role in the sample of U.K. firms than in the remaining countries. Although the U.K. sample contains much larger firms than the French and Belgian samples, and some previous studies have found stronger evidence of financial effects on investment among larger firms,[28] we can be reasonably confident that this finding is not driven by differences in firm size. First, the size distribution of firms in our German sample is quite similar to that in our U.K. sample (see table 1), but we find much weaker effects from financial variables in our German results. Secondly, we used a subsample of large French companies to investigate the effect of firm size directly.

[26] The instruments used were the lagged values of all right-side variables dated $t - 2$, $t - 3$, . . . , $t - 6$. We also checked the time series properties of the variables included in the Euler-equation specification, and confirmed that all these series appear to be stationary in our samples.

[27] See the discussion of these coefficients following equation (8).

[28] See Devereux and Schiantarelli (1990).

This subsample consists of 234 firms which had at least 1,000 employees in their first year. It is more comparable to our U.K. and German samples, with mean and median employment in 1985 of 3,819 and 1,794 respectively. The results of estimating each of the investment models for this subsample are reported in tables A3 and A4 of the results appendix. They show almost no significant differences from the results for the full French data set, and there is no indication that the large French firms are more (or less) affected by financial constraints than the smaller French firms.

Finally we investigated the effect of using consolidated or unconsolidated accounts data on our results. Recall that our data for the U.K.—where we have found the strongest effects from cash flow—are consolidated accounts for company groups, whereas our data for the remaining countries are unconsolidated accounts for individual corporations. Clearly there is a possibility that the investment spending by a subsidiary is constrained by the cash flow of the company group as a whole, rather than by the cash flow of the subsidiary itself, and that this will not be detected by regressing the subsidiary's investment on its own cash flow using unconsolidated data. Notice that in this case we would be underestimating the effect of financial constraints in France, Germany, and Belgium, rather than overestimating the importance of financial constraints in the United Kingdom.

To investigate this possibility, we used a subsample of 437 independent French firms that are not subsidiaries of larger companies,[29] and a subsample of 87 German companies for which consolidated company accounts were also available in the Bonn Data Bank.[30] The firms in the independent French subsample tend to be small, with mean and median employment in 1985 of 214 and 184 respectively. The firms in the consolidated German sample are much larger on average than in any of the other samples we have used, with mean and median employment in 1985 of 37,317 and 7,669 respectively.

The results of estimating our investment equations on these subsamples are also reported in tables A3 and A4 of the results appendix. The results for the independent French subsample are mixed, with no significant cash-flow effects found in the error-correction model, but with less satisfactory results for the Euler equation than found for the large French firms. For the subsample of consolidated German accounts, the coefficient on cash flow in the error-correction model is higher than for our main sample of unconsolidated German accounts, but this coefficient is estimated imprecisely. Overall these comparisons do not suggest that the differences between our results for the United Kingdom and for Belgium, France, and Germany are primarily driven by this difference in the level of aggregation at which the company data are available, although it would be interesting to investigate this issue further.

V. Conclusion

A consistent pattern emerges from our results. For Belgian companies, we find no significant cash flow effects in the error-correction model, and we do not reject the Euler-equation specification derived under the null of no financial constraints. For U.K. companies, we find a large and significant cash flow effect in the error-correction model, and we clearly reject the basic Euler-equation specification; in particular, the large and significant positive coefficient on the gross-operating-profits term is inconsistent with the basic Euler equation. The results for French and German companies are intermediate: in the error-correction model, we find statistically significant cash flow effects, but much smaller than those found in the United Kingdom. A robust result across both models is that the sensitivity of investment to financial variables is both statistically and quantitatively more significant in the United Kingdom than in France, Germany, or Belgium. This difference does not appear to be accounted for by differences in the size distribution of firms, or in the nature of the company-accounts data available for the United Kingdom.

The significance of cash-flow terms in the error-correction model could in principle reflect expectation formation rather than financial constraints.[31] However, we do not find big differences across countries in the power of cash-flow variables to forecast future sales growth or future profitability in the context of simple econometric forecasting equations. This interpretation is also less appealing in the context of the Euler-equation model, where the model derived under the null of no financial constraints is strongly rejected in the U.K. sample.[32]

The availability of internal finance appears to have been a more important constraint on company investment in our sample of U.K. firms than in our samples of continental European firms over the period 1978–1989. This finding is consistent with the suggestion that the market-oriented financial system in the United Kingdom performs less well in channeling investment funds to firms with profitable investment opportunities than do the continental European financial systems. However, we caution that we have not tested this hypothesis directly, and our results are doubtless consistent with other interpretations. The accounts data avail-

[29] These independent firms are defined as not being subsidiaries of either a French or a foreign company during the sample period, and as not having subsidiaries themselves.

[30] Because the sample size is much smaller, our results for the consolidated German subsample do not use instruments dated $t - 5$ or $t - 6$.

[31] As pointed out by Kaplan and Zingales (1997), the difference between a large cash flow coefficient in the United Kingdom and smaller (but significant) cash flow effects in France and Germany does not necessarily indicate that financing constraints are more severe in the United Kingdom. However, the contrast between significant cash flow effects in the United Kingdom and insignificant cash flow effects in Belgium is more compelling.

[32] Previous research using U.K. data also indicates that this excess sensitivity of investment to cash flow is concentrated among observations on low-dividend-paying companies, which is consistent with the presence of relatively severe financial constraints on investment spending for these firms. See Bond and Meghir (1994).

able for this study were not as consistent across countries as we would have liked. Moreover, models of financial constraints predict that investment is only constrained when desired investment exceeds the supply of internal finance; it may simply be that our results reflect transient differences in the frequency of this event within our samples, rather than deeper differences in the effects of different financial systems. Discriminating between these alternative interpretations will require more detailed comparative analyses of the investment behavior of different types of companies across countries; we believe this to be an important challenge for future research.

REFERENCES

Abel, A. B., "Empirical Investment Equations: An Integrative Framework," in K. Brunner and A. Meltzer (Eds.), *On the State of Macroeconomics*, *Carnegie-Rochester Conference Series* 12 (1980), 39–93.

Abel, A. B., and O. J. Blanchard, "The Present Value of Profits and Cyclical Movements in Investment," *Econometrica* 54 (1986), 249–273.

Arellano, M., and S. R. Bond, "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations," *Review of Economic Studies* 58 (1991), 277–297.

——— "Dynamic Panel Data Estimation Using DPD98 for Gauss: A Guide for Users," Institute for Fiscal Studies mimeograph, London, http://www.ifs.org.uk/staff/steve_b.shtml (1998).

Bean, C. R., "An Econometric Model of Manufacturing Investment in the UK," *Economic Journal* 91 (1981), 106–121.

Blundell, R. W., and S. R. Bond, "Initial Conditions and Moment Restrictions in Dynamic Panel Data Models," *Journal of Econometrics* 87 (1998), 115–143.

——— "GMM Estimation with Persistent Panel Data: An Application to Production Functions," *Econometric Reviews* 19 (2000), 321–340.

Bond, S. R., D. Harhoff, and J. Van Reenen, "Investment, R&D and Financial Constraints in Britain and Germany," IFS working paper W99/5 (1999).

Bond, S. R., and C. Meghir, "Dynamic Investment Models and the Firm's Financial Policy," *Review of Economic Studies* 61 (1994), 197–222.

Bond, S. R., C. Nauges, and F. Windmeijer, "Unit Roots and Identification in Autoregressive Panel Data Models: A Comparison of Alternative Tests," Institute for Fiscal Studies mimeograph, London (2001).

Caballero, R. J., "Aggregate Investment," in J. Taylor and M. Woodford (Eds.), *Handbook of Macroeconomics* Vol. 1B, (North-Holland, 1999).

Caballero, R. J., E. Engel, and J. Haltiwanger, "Plant-Level Adjustment and Aggregate Investment Dynamics," *Brookings Papers on Economic Activity* 1995:2 (1995), 1–54.

Devereux, M. P., and F. Schiantarelli, "Investment, Financial Factors and Cash Flow: Evidence from UK Panel Data," in R. G. Hubbard (Ed.), *Asymmetric Information, Corporate Finance and Investment* (Chicago: NBER and University of Chicago Press, 1990).

Dixit, A. K., and R. S. Pindyck, *Investment under Uncertainty* (Princeton University Press, 1994).

Edwards, J. S. S., and K. Fischer, *Banks, Finance and Investment in Germany* (Cambridge: Cambridge University Press, 1994).

Elston, J. A., "Investment, Liquidity Constraints and Bank Relationships: Evidence from German Manufacturing Firms," in S. Black and M. Moersch (Eds.), *Competition and Convergence in Financial Markets: The German and Anglo-American Models, Advances in Finance, Investment and Banking, Volume 5* (Elsevier Science North-Holland, 1998).

Fazzari, S. M., R. G. Hubbard, and B. C. Petersen, "Financing Constraints and Corporate Investment," *Brookings Papers on Economic Policy* 1 (1988), 141–195.

——— "Investment–Cash Flow Sensitivities Are Useful: A Comment on Kaplan and Zingales," *Quarterly Journal of Economics* CXV:2 (2000), 695–705.

Gilchrist, S., and C. P. Himmelberg, "Evidence on the Role of Cash Flow for Investment," *Journal of Monetary Economics* 36 (1995), 541–572.

Hall, B. H., and J. Mairesse, "Testing for Unit Roots in Panel Data: An Exploration Using Real and Simulated Data," University of California mimeograph, Berkeley (2001).

Hall, B. H., J. Mairesse, L. Branstetter, and B. Crepon, "Does Cash Flow Cause Investment and R&D: An Exploration Using Panel Data for French, Japanese and United States Scientific Firms," in D. B. Audretsch and R. Thurik (Eds.), Innovation, Industry Evolution, and Employment, (Cambridge, UK: Cambridge University Press, 1999).

Hoshi, T., A. Kashyap, and D. Scharfstein, "Corporate Structure, Liquidity and Investment: Evidence from Japanese Industrial Groups," *Quarterly Journal of Economics* 106 (1991), 33–60.

Hsiao, C., *Analysis of Panel Data* (Cambridge: Cambridge University Press, 1986).

Hubbard, R. G., "Capital-Market Imperfections and Investment," *Journal of Economic Literature* XXXVI (1998), 193–225.

Johansen, F., "Investment and Financial Constraints: An Empirical Analysis of Norwegian Firms," Statistics Norway discussion paper 109 (1994).

Kaplan, S. N., and L. Zingales, "Do Investment–Cash Flow Sensitivities Provide Useful Measures of Financing Constraints?" *Quarterly Journal of Economics* CXII:1 (1997), 169–216.

——— "Investment–Cash Flow Sensitivities Are Not Valid Measures of Financing Constraints," *Quarterly Journal of Economics* CXV:2 (2000), 707–712.

Mairesse, J., and B. Dormont, "Labour and Investment Demand at the Firm Level: A Comparison of French, German and U.S. Manufacturing, 1970–79," *European Economic Review* 28:2 (1985), 201–231.

Mairesse, J., B. H. Hall, and B. Mulkay, "Firm-Level Investment in France and the United States: An Exploration of What We Have Learned in Twenty Yeras," *Annales d'Economie et de Statistique* 55/56 (1999), 27–67.

Mayer, C. P., "New Issues in Corporate Finance," *European Economic Review* 32:5 (1988), 1167–1189.

——— "Financial Systems, Corporate Finance and Economic Development," in R. G. Hubbard (Ed.), *Asymmetric Information, Corporate Finance and Investment* (Chicago: NBER and University of Chicago Press, 1990).

Nelson, C. R., and R. Startz, "The Distribution of the Instrumental Variable Estimator and Its *t*-Ratio When the Instrument is a Poor One," *Journal of Business and Economic Statistics* 63 (1990), 5125–5140.

Nickell, S. J., *The Investment Decisions of Firms* (Cambridge: Cambridge University Press, 1978).

——— "Biases in Dynamic Models with Fixed Effects," *Econometrica* 49 (1981), 1399–1416.

——— "Error Correction, Partial Adjustment and All That: An Expository Note," *Oxford Bulletin of Economics and Statistics* 47 (1985), 119–130.

Oliner, S., G. Rudebusch, and D. Sichel, "New and Old Models of Business Investment: A Comparison of Forecasting Performance," *Journal of Money, Credit and Banking* 27 (1995), 806–826.

Schiantarelli, F., "Financial Constraints and Investment: A Critical Survey of the International Evidence," *Oxford Review of Economic Policy* 12:2 (1996), 70–89.

Staiger, D., and J. H. Stock, "Instrumental Variables Regression with Weak Instruments," *Econometrica* 65 (1997), 557–586.

Whited, T. M., "Debt, Liquidity Constraints and Corporate Investment: Evidence from Panel Data," *Journal of Finance* 47:4 (1992), 1425–1460.

DATA APPENDIX

The company datasets for the four countries are obtained from different sources. For France, we obtain data on large and medium-size firms from INSEE, which collects them from fiscal sources of the Ministry of Finance. Therefore these French data are the unconsolidated accounts of these corporations. For Belgium, we have access to the unconsolidated balance sheets and income accounts of a selected sample of large and medium-size Belgian corporations. This sample has been built by the

Banque National of Belgium, which collects the data from the Commerce Court. In fact, by law, all Belgian firms must register their annual accounts, which are sent to the National Bank of Belgium. These French and Belgian companies need not be stock-market-quoted, and may include subsidiaries of foreign companies.

The U.K. sample comprises the consolidated accounts of companies quoted on the London Stock Exchange, which were obtained from Datastream. Finally, the German data have been collected at the University of Bonn. They include basically all the quoted manufacturing *Aktiengesellschaft* (AG) corporations for which sufficient years of data were available. These are all quoted firms, but the data available are unconsolidated accounts for these corporations.[33]

We use panel data on company accounts covering the period 1978–1989, even though for the United Kingdom and Germany longer time series were available. All firms have their main activity in the manufacturing sector, and firms with fewer than 100 employees in the first year of observation were excluded. The initial samples of firms which satisfied these requirements were 1,473 firms for France, 410 firms for Belgium, 600 firms for the United Kingdom, and 287 firms for Germany.

We attempted to use variables that are reasonably comparable across countries, even though the national accounting definitions are not precisely the same. The French and Belgian accounts have very similar definitions of the main variables. The U.K. and German accounts provide more limited information on costs, and the U.K. accounts report commercial depreciation rather than fiscal depreciation.

The main variables we use are flows of investment, sales, gross operating profits, and cash flow. Investment spending is obtained from the account of sources and uses of funds, and not inferred from changes in the balance sheet. We use sales as a proxy for output. For the French and Belgian data, a measure of value added was also available from the accounts. All the flow variables were deflated using output price indices at the sectoral level.

A measure of the stock of capital at current replacement cost $p_t^I K_t$ was estimated from the flow data on investment $p_t^I I_t$ using a standard perpetual inventory method, in a similar way for each sample:

$$p_t^I K_t = (1 - \delta)\, p_{t-1}^I K_{t-1} \frac{p_t^I}{p_{t-1}^I} + p_t^I I_t,$$

where

K_t = capital stock,
p_t^I = price of investment goods,
I_t = real investment,
δ = depreciation rate (8%).

The starting value was based on the net book value of tangible fixed capital assets in the first observation within our sample period, adjusted for previous years' inflation. For France, Belgium, and Germany, where the reported net book value of assets subtracts the fiscal depreciation allowed for tax purposes rather than commercial depreciation, we have corrected this measure by taking into account accelerated fiscal depreciation. This correction lowers the value of accumulated depreciation, and thus increases the net book value of assets. Subsequent values were obtained using accounts data on investment and disposals, national price indices for investment-goods prices, and a depreciation rate of 8% assumed to be common to all countries.

For France and Belgium, we construct a measure of gross operating profits Π by subtracting the total wage bill from value added. The measure of cash flow C is then computed from gross operating profits by subtracting payments of interest and taxes. This method was not possible for the United Kingdom and Germany, because we do not have data on value added. In these cases we computed cash flow by adding back reported depreciation to reported profits (net of interest and taxes). Gross operating profits were then obtained by adding back interest payments and taxes to this measure of cash flow. These top-down and bottom-up methods should yield equivalent measures of gross operating profits and cash flow.

After computing the main variables used in the investment models, we excluded observations where the change in sales suggested that a major merger or acquisition (or disposal) had occurred, since it is not clear that such large adjustments would be well characterized by the usual investment models. We also excluded observations which appeared to contain substantial outliers. Specifically, observations were discarded if the investment rate exceeded 1, if real sales increased or decreased by more than a factor of 3, or if the observed ratio of either sales, gross operating profits, or cash flow to the capital stock fell in the first or the last centile of the empirical distribution for each country. In these cases we retained the longest available time series of consecutive annual observations for the firms affected. We also required that at least six consecutive annual observations be available for the firms included in our final samples. These criteria resulted in a loss of respectively 7.3%, 11.9%, 5.0%, and 21.3% of our initial observations for France, Belgium, the United Kingdom, and Germany.

[33] For example, Audi, which is an almost wholly owned subsidiary of Volkswagen, nevertheless has a separate listing.

RESULTS APPENDIX

TABLE A1.—FORECASTING MODELS FOR SALES-GROWTH
DEPENDENT VARIABLE Δy_t; OLS

	Belgium	France	Germany	U.K.
I_{t-1}/K_{t-2}	0.114	0.113	0.162	0.360
	(0.033)	(0.021)	(0.071)	(0.049)
I_{t-2}/K_{t-3}	0.067	0.050	−0.014	−0.051
	(0.030)	(0.020)	(0.064)	(0.044)
Δy_{t-1}	−0.021	0.037	−0.064	0.093
	(0.053)	(0.027)	(0.057)	(0.028)
Δy_{t-2}	−0.027	−0.008	0.000	0.014
	(0.036)	(0.020)	(0.048)	(0.021)
C_{t-1}/K_{t-2}	0.111	0.018	0.150	0.183
	(0.039)	(0.021)	(0.066)	(0.068)
C_{t-2}/K_{t-3}	−0.034	0.023	−0.046	−0.169
	(0.035)	(0.020)	(0.057)	(0.072)

TABLE A2.—FORECASTING MODELS FOR CASH-FLOW
DEPENDENT VARIABLE C_t/K_{t-1}; OLS

	Belgium	France	Germany	U.K.
I_{t-1}/K_{t-2}	−0.091	−0.050	−0.039	−0.113
	(0.028)	(0.017)	(0.026)	(0.018)
I_{t-2}/K_{t-3}	−0.033	0.007	−0.039	0.023
	(0.019)	(0.015)	(0.025)	(0.018)
Δy_{t-1}	0.057	0.090	0.053	0.022
	(0.018)	(0.014)	(0.015)	(0.009)
Δy_{t-2}	−0.008	0.005	0.016	0.011
	(0.016)	(0.012)	(0.023)	(0.007)
C_{t-1}/K_{t-2}	0.607	0.558	0.442	0.930
	(0.032)	(0.020)	(0.038)	(0.034)
C_{t-2}/K_{t-3}	0.179	0.174	0.241	−0.014
	(0.036)	(0.019)	(0.036)	(0.035)

TABLE A3.—ERROR-CORRECTION MODELS: GMM FIRST DIFFERENCES,
$t - 2$ INSTRUMENTS

	France, Large Firms	France, Independent	Germany, Consolidated
Firms	234	437	87
Observations	1,440	2,544	520
I_{t-1}/K_{t-2}	0.027	−0.034	0.003
	(0.064)	(0.067)	(0.057)
Δy_t	0.204	0.155	−0.064
	(0.050)	(0.046)	(0.060)
Δy_{t-1}	0.058	0.201	0.072
	(0.038)	(0.047)	(0.057)
$(k - y)_{t-2}$	−0.106	−0.195	−0.068
	(0.030)	(0.049)	(0.053)
C_t/K_{t-1}	−0.036	−0.048	0.393
	(0.066)	(0.069)	(0.157)
C_{t-1}/K_{t-2}	0.088	0.065	−0.112
	(0.048)	(0.041)	(0.086)
m1	−5.11	−7.24	−3.94
m2	1.01	−0.21	0.83
Sargan	0.157	0.564	0.894

TABLE A4.—EULER-EQUATION MODELS: GMM FIRST DIFFERENCES

	France, Large Firms		France, Independent		Germany, Consolidated	
	$t-2$ Inst.	$t-3$ Inst.	$t-2$ Inst.	$t-3$ Inst.	$t-2$ Inst.	$t-3$ Inst.
$(I/K)_{t-1}$	0.567	0.844	0.311	0.212	0.315	0.202
	(0.092)	(0.249)	(0.064)	(0.207)	(0.114)	(0.223)
$(I/K)^2_{t-1}$	−0.853	−1.680	−0.405	−0.248	−0.603	−0.538
	(0.247)	(0.802)	(0.190)	(0.658)	(0.235)	(0.591)
$(\Pi/K)_{t-1}$	0.094	0.055	0.054	0.070	−0.106	−0.080
	(0.024)	(0.037)	(0.019)	(0.042)	(0.060)	(0.076)
$(Y/K)_{t-1}$	−0.005	−0.001	0.013	0.016	0.031	0.039
	(0.004)	(0.008)	(0.005)	(0.008)	(0.011)	(0.012)
m1	−6.91	−4.63	−10.62	−5.51	−4.92	−3.65
m2	1.53	0.97	−0.36	−0.57	0.17	−0.28
Sargan	0.399	0.698	0.192	0.093	0.833	0.481