On the Interaction of Financial Frictions and Fixed Capital Adjustment Costs Evidence from a Panel of German Firms

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Abstract
This paper analyzes the interaction of financial frictions and non-convex adjustment costs. With non-convex adjustment costs firms infrequently carry out discrete investment projects. Therefore, financial variables may influence investment in two ways. Theoretically, they can alter the frequency at which investment projects are undertaken, or they can influence the size of the stock of capital a company wishes to hold in the long run. Empirically, finance has nearly no long-run influence on the stock of capital in the sample of German companies which this paper analyzes. By contrast, the influence of finance on investment decisions is substantial. Consequently, finance primarily affects investment frequencies and accordingly, financial factors and fundamental capital productivity strongly interact in the determination of investment.
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1 Introduction

Economists' knowledge of micro-level and aggregate investment is still far from being conclusive. The only thing seemingly well-established is the empirical rejection of the standard neoclassical investment model.[1] The question of which of the assumptions of the neoclassical model lead to its failure to what extent has yet to be answered. Beginning with Fazzari et al. (1988), the empirical literature has emphasized the role of financial factors in company-level investment. More recently, attention has been drawn to the role of non-convexities in investment technology.[2]

This paper empirically investigates the *interaction* of both of these deviations from the neoclassical model. Our analysis follows the so-called "gap-approach", which is essentially an error correction model for investment. This two-step approach first measures the difference ("gap") between the actual stock of capital and the capital stock a company would like to hold if there were no adjustment costs. In the second step, investment is regressed on this gap-measure. Especially in the context of financial frictions, this approach can generate new insights, since it sequentially estimates target capital levels and adjustment dynamics. Therefore, it allows us to differentiate between short and long run influences of financial variables. Additionally, the approach reveals whether abundant financial resources alter investment-rates mainly by directly shifting average investment-rates, or by changing the investment process in a more complex manner in interaction with fundamental investment incentives.

Despite the advantages of the "gap model" it has to be applied with some care. Cooper and Willis (2004) have recently shown that the model is somewhat sensitive to deviations from its basic assumptions. Hence without pre-testing the underlying assumption of the model, one may draw misleading conclusions from its estimation. The present paper takes this issue into account. The core of Cooper and Willis' argument is that a measurement-error problem may result if productivity has below unit-root serial correlation and is not directly observable. Therefore, we estimate productivity by exploiting all available firm-level data on employment, wages and sales, following the method developed in Cooper and Haltiwanger (2002) to minimize the measurement error

[1] See Caballero (2000).

[2] For evidence on non-convex adjustment costs, see Caballero et al. (1995), Doms and Dunne (1998), Cooper et al. (1999), Caballero and Engel (1999), Goolsbee and Gross (1997), Abel and Eberly (2002), or Cooper and Haltiwanger (2002).
The literature on financial frictions and their impact on investment has been surveyed by Hubbard (1998). Mairesse et al. (1999) also give a broad overview. More recent contributions are e.g. Kaplan and Zingales (1997), Gilchrist and Himmelberg (1998), Guarglia (1999), Cummins et al. (1999), or Erickson and Whited (2000).

in the first place. Secondly, we show that both *productivity and capital at the firm level exhibit a unit-root in our data, and are cointegrated,* as one would expect from theory.[3] Consequently, the cointegration error identifies the gap between the desired and the actual stock of capital, so that Cooper and Willis' criticism does not apply. Therefore, irrespective of the actual form of adjustment costs, investment can be estimated using the gap by a (non-linear) error correction model.[4]

For this error correction model of investment, it can be shown that under quadratic adjustment costs and a unit-root in shocks to productivity the error correction should be linear.[5] If adjustment costs are non-convex instead, higher order terms of the cointegration error become significant and the adjustment speed varies with the size of the gap between desired and actual stock of capital.[6] Empirically, not only the size of the gap determines the adjustment speed: Whited (2004) shows for US data that financially constrained firms invest much less frequently than unconstrained ones. Theoretically, the influence of finance on adjustment speed has been studied by Holt (2003) and in a companion paper to this one (Bayer, 2002). The latter paper also provides empirical evidence from an UK database and both papers show the potential importance of the interaction of a financial frictions and fixed adjustment costs. Therefore, we allow financial frictions to affect both, the adjustment process (investment) and the desired stock of capital. However, we only find a significant influence of finance on the adjustment process, so that finance only plays a role in the short run.[7]

To be able to differentiate between long-run and short-run influences of finance on investment and capital, the analysis has to rely on a stock measure of liquidity rather than on flow-measures. Blinder (1988) has pointed out that a stock measure is also preferable on theoretical grounds. As this stock measure, ideally one would use the line of credit for which we take the equity ratio as a proxy, whereby the equity ratio is the

[3] Our approach hence is similar to Caballero et al. (1995) but uses a different measure for productivity and hence for the desired stock of capital. Also, Caballero et al. estimate the cointegration relation using OLS, while we employ a panel dynamic LS method.

[4] Non-linear error-correction models can be understood as a generalization of threshold-cointegration models and linear error-correction models. Non-linear error-corection models have for example been recently applied to the analysis of financial data (Breitung and Wulff, 2001).

[5] See Rotemberg (1987) for a formal proof.

[6] Under the assumption of non-stationary productivity the gap approach hence appears preferable to a q-theoretic measures of investment incentives, since these measures are known to be problematic whenever stock-markets are not (perfectly) efficient (Cummins et al., 1999), whenever there are rents not related to the stock of capital (Merz and Yashiv, 2002), or when adjustment cost are not convex (Barnett and Sakellaris, 1999, p. 259).

[7] In the latest version of her paper, Whited (2004) reports a similar phenomenon for a simulated model of investment under non-convex adjustment costs and costly external finance.

Similarly, Guarglia (1999) finds that liquidity proxies and firm size is uncorrelated, wheras investment and liquidity is.

book value of equity over the book value of assets.

Like all measures of liquidity the equity ratio is partly endogenous and a result of past productivity shocks. Therefore, its endogeneity must be taken into account for the econometric analysis. In contrast to a flow measure, however, endogeneity will mostly play a role for the long-run analysis, but not that much for the short-run. The equity-ratio at the beginning of the investment period is pre-determined and should not strongly correlate with innovations to investment. This allows us to concentrate on endogeneity and (below unit-root) autoregressive behavior of equity for the long-run regression. We find an (insignificant) negative correlation of the equity ratio and the capital level when not controlling for endogeneity. Yet, when endogeneity is controlled for, the equity ratio still does not significantly correlate with the level of capital a company employs. Interestingly and maybe counterintuitively, however, the estimated long-run elasticity of capital with respect to the equity ratio increases when controlling for endogeneity. In the short-run investment regression, we control for endogeneity by exploiting just the variation of the equity ratio relative to its firm-specific long-run mean. This especially accounts for different baseline-access to capital markets across firms.

For investment, the following three results are found: First, the gap between desired and actual capital can explain a relatively large part of the variation in investment. Second, investment is a moderately convex function of the gap. Third, the financial condition has a significant short-run impact on investment decisions and this impact varies strongly with the size of the gap. A good financial status is complementary to a large gap. Figuratively speaking, finance is the grease in the investment process but not the fuel. It eases adjustment to the target level of capital, but from the estimation of the level equation we know it does not alter that target. Since finance has no predictive power for capital decisions in the long run, these results for investment cannot be attributed to a lack of measuring capital productivity correctly. If finance contained information about future investment prospects that was not contained in the productivity measure, this were reflected in a significant correlation of finance and the level of capital.

The remainder of this paper is organized as follows: Section 2 develops the theoretical grounds for the empirical analysis. It first reviews the recent debate between Cooper and Willis (2004) and Caballero and Engel (2004) on the gap approach and hence focuses our attention on the most critical steps and parts of the analysis. Secondly, the section sketches a possible extension of the gap approach to cover the influence of financial frictions. Thirdly, section 2 also introduces the method used by the present paper to measure productivity, and estimates the long-run optimal stock of capital and the investment equation. Section 3 gives a brief description of the data that has been used.

Section 4 presents the empirical results of both the regression for the optimal stock of capital and the investment regression. Section 5 compares these results to those of a companion paper for UK data (Bayer, 2002). Moreover possible extensions are discussed. Finally section 6 concludes and a data appendix follows.

2 The gap approach to capital adjustment

2.1 The gap model and non-convex adjustment costs–summarizing a current debate

Analyzing investment (and employment) data using the gap approach has been introduced by Caballero and Engel (1993) and Caballero et al. (1995, 1997). All three papers show that aggregate investment (employment) also depends significantly on the higher order moments of the distribution of fundamental investment incentives. Although the central focus of these papers is the aggregate consequences of non-convexities, their estimation procedure may also be interpreted as a test for these non-convexities.[8]

This view has recently come under criticism by Cooper and Willis (2004), and since the present paper also follows a gap approach a quick summary of the argument seems appropriate before laying out our own analysis. Cooper and Willis acknowledge that under the null hypothesis of non-convex adjustment costs the gap approach may be valid if firm productivity follows a random walk. However, they argue that the procedures used to measure the gap will result in a severe measurement error under the alternative hypothesis of convex adjustment costs and below unit-root serial correlation of productivity. This measurement error then causes the higher order moments of the micro-level gap-distribution to become significant when regressing investment on the first three moments of this gap distribution. Yet, the higher order moments should not be significant under the convex-cost alternative and Cooper and Willis find that higher-order moments are not significant when the gap is measured correctly. Consequently, Caballero and Engel's "test" based on the estimated parameters in the investment regression suffers from a lack of power.

In their reply to this criticism, Caballero and Engel (2004) provide two central arguments why Cooper and Willis' critique is somewhat misleading. Their first point is that only when the serial correlation of productivity shocks is dropped to unrealistically low levels, do the higher-order terms of the gap become significant regressors.[9] Their second

[8] Caballero and Engle (2004, p. 5) emphasize the central point of their analysis being a macro-data description conditional on the presumption that microeconomic behavior is driven by fixed costs. However, interpreting their aproach as a testing procedure appears to be more fruitful.

[9] This can be seen for example by inspecting Table 4 and 5a in Cooper and Willis (2003a, pp. 32). Cooper and Willis try two specifications for the adjustment costs, the first specification generates a half-

point is that the pure focus on statistical significance of the higher order moments is also misleading. They argue that the increase in the R^2-statistics when adding higher-order moments and most importantly the difference in adjustment speeds between large-gap and small-gap firms can still be used to test for non-convexities since they are not affected by Cooper and Willis' measurement-error argument.

Hence in summary, one may take three central points from the debate: First, when using the gap approach one should try carefully to minimize the potential sources of measurement errors when deriving productivity. Second, it is essential to check for a unit-root in productivity and capital and proceed with the gap approach only if there is a unit-root. And third, one should especially focus on adjustment-speed differences when interpreting the results of the investment equation using the gap.

2.2 A gap model with financial frictions

The gap model itself can be derived from the assumption of infrequent investment—e.g. the constant hazard model of Calvo (1983)—but also from quadratic adjustment costs and smooth adjustment over time (Sargent, 1978 and Rotemberg, 1987). Formally, the model may be described as follows. Let k^* denote the log of the stock of capital a company would hold if adjustment costs are set to zero *for one period*. In a world without financial frictions k^* would only depend on firm productivity ξ. When financial frictions distort firm decisions, financial means or "liquidity" e also influences k^*.

This desired level of capital $k^*(e,\xi)$ now is exactly the level of capital the firm adjusts too if the only costs of adjustment are fixed (see e.g. Caballero and Engel (1999) and Bayer (2002) for a microfoundation without or with financial frictions respectively). Thus upon investment, the gap is closed completely and the investment rate is simply the difference x of the current stock of capital k and the desired stock of capital k^* (in logs), $x := k^* - k$. Consequently, the expected investment rate is a compound of the gap x, which is mandated investment, and the probability of investment Λ. This probability can alternatively be interpreted as the adjustment speed in a convex adjustment cost model with continuous adjustment, see e.g. Sargent (1978). If we allow Λ to depend

life of the gap of one month, implying that at the end of the year the firm basically holds all the capital it would like to hold in the absence of adjustment costs. The second specification implies a half-life of one year, which still is a relatively fast adjustment speed.

First, the number of false rejections of the convex-cost model substantially decreases when adjustment costs increase and become more realistic. Second for the one-year half life model, out of 18 estimations with an autocorrelation of productivity being larger than 0.9, the estimation procedure only finds a higher-order term being falsely significant at the 10% level four times, while we would expect to find approximately two. Out of these four errors of first order, however, in three cases the sign of the higher-order term is negative.

on both liquidity e and the capital gap x, we obtain for the expected investment rate $i_t = k_t - k_{t-1}$:[10]

$$i\left(x,e\right) = \Lambda\left(x,e\right)x = \Lambda\left[x\left(e,\xi\right),e\right]x\left(e,\xi\right). \tag{1}$$

Now, taking first derivatives decomposes the effect of liquidity e on investment into a direct effect on the adjustment speed and an indirect effect via the optimal stock of capital. Formally, this is shown in the following equation

$$\frac{\partial i\left(x,e\right)}{\partial e} = \left(\frac{\partial \Lambda}{\partial x}\frac{\partial x}{\partial e} + \frac{\partial \Lambda}{\partial e}\right)x + \Lambda\frac{\partial x}{\partial e} = \underbrace{\frac{\partial x}{\partial e}\left[\frac{\partial \Lambda}{\partial x}x + \Lambda\right]}_{indirect} + \underbrace{\frac{\partial \Lambda}{\partial e}}_{direct}x, \tag{2}$$

In comparison, productivity affects investment only by altering the gap x :

$$\frac{\partial i\left(x,e\right)}{\partial \xi} = \frac{\partial i\left(x,e\right)}{\partial x}\frac{\partial x}{\partial \xi} = \frac{\partial x}{\partial \xi}\left[\frac{\partial \Lambda}{\partial x}x + \Lambda\right]. \tag{3}$$

Both equations together show that the indirect effect in (2) is an effect that is equivalent to a productivity change that alters the optimal capital level. Therefore, we can alternatively term the indirect effect a level effect, whereas the direct effect is a result of a change in the investment frequency

$$\frac{\partial i\left(x,e\right)}{\partial e} = \underbrace{\frac{\frac{\partial x}{\partial e}}{\frac{\partial x}{\partial \xi}}\frac{\partial i\left(x,e\right)}{\partial x}}_{\text{level effect}} + \underbrace{\frac{\partial \Lambda}{\partial e}}_{\text{frequency effect}}x \quad . \tag{4}$$

Moreover, the latter equation shows that the level effect itself decomposes into a term reflecting the sensitivity of investment to changes in the gap $\frac{\partial i(x,e)}{\partial x}$ and the term $\frac{\partial x}{\partial e}/\frac{\partial x}{\partial \xi}$ which can be interpreted as the marginal productivity of financial means (liquidity). Equivalently, this term can be interpreted as the marginal reduction in the user-cost-of-capital from a liquidity increase.[11]

The term $\frac{\partial \Lambda}{\partial e}x$ reflects the direct impact we assumed liquidity to have on the adjustment speed, but why should there be this direct influence? Empirically, Whited (2004) provides some evidence that the financial status influences adjustment hazards.[12] Theo-

[10] See Bayer (2002) for a detailed theoretical model that combines fixed adjustment costs and capital market imperfections.

[11] If one would think of liquidity altering mainly managerial decisions but not truly the cost structure, this term represents how liquidity changes the managerial discount rate.

[12] Whited (2004) analyzes the hazard rates for investment-spikes of financially constrained and unconstrained firms using Compustat data. Even though she concentrates on the *existence* of non-convexities and financial constraints rather than on the *interaction* of productivity and financial constraints, she

retically, a related effect has been shown by Holt (2003) and in a companion paper of this one (Bayer, 2002). That paper also assess this interaction empirically using a sample of UK-firms drawn from the Cambridge-Database and a different method in identifying the marginal productivity of capital.

In principle, the basic idea behind the frequency effect is relatively straightforward and may be illustrated in the following very stylized way: Suppose a firm completely leases its capital that does not depreciate. If liquidity now does not influence the user-cost of capital, then the stock of capital the firm adjusts to is only determined by productivity and does not depend on liquidity. Hence, there is no level effect, but liquidity may still influence the probability of adjustment: Suppose the firm pays some fixed costs upon investment, then a given amount of internal funds (liquidity) directly determines how often the firm can expect to adjust over a certain interval of time. For example, if liquidity e does not grow and the expected fixed cost of adjustment is C, this firm is endowed with e/C adjustment options. Since each option is more valuable if the number of options is small, a small number of options is equivalent to large adjustment costs. Consequently, for a given gap x the firm is more likely to adjust if liquidity is large, so that this establishes a frequency effect of liquidity.

Although, a given pattern of frequency effect and level effect might be related to more than one structural model of investment, this decomposition still carries information (as Cooper and Willis (2003) show for the standard gap model for employment). Directly related to the level-frequency decomposition, we can test two hypothesis about how the availability of (accumulated) internal funds influences investment activity. The first hypothesis reflects the long run neutrality of finance:

H_0^0 : Internal funds have no effect on the optimal stock of capital a company holds. This is equivalent to $\frac{\partial x}{\partial e} = 0$.

The second hypothesis accounts for the influence of equity on the investment process, this is:

H_0^1 : Investment reacts to changes in internal funds only because the optimal stock of capital is altered, i.e. $\frac{\partial i(x,e)}{\partial e} = \frac{\partial i(x,e)}{\partial x} \frac{\partial x}{\partial e}$.

If H_0^0 cannot be rejected, the Modigliani-Miller theorem holds in the long run. Moreover, an empirical non-rejection of H_0^0 also has an important econometric implication:

finds both evidence for increasing hazard-rates (and thus non-convexities) and a significant influence of financial constraints—which lower the hazard rates.

the financial variable cannot contain information on long-run capital productivity that is not included in the measure of productivity $\hat{\xi}$. If the financial variable e carries information on long-run capital productivity H_0^0 will be rejected irrespective of financial frictions being present or not. However, if H_0^0 is not rejected finance can still influence the transition path of the stock of capital if there is a frequency effect of liquidity. In this case hypothesis H_0^1 will be rejected. Note that (only) if H_0^0 holds true, H_0^1 simplifies to $\frac{\partial i(x,e)}{\partial e} = 0$.

2.3 The desired stock of capital

However, we can neither directly observe k^* nor the productivity ξ of capital. Typically, as a proxy for the marginal productivity of capital or directly for fundamental investment incentives Tobin's q, the ratio of firm-value over the replacement value of its assets has been employed. Yet, there are two major drawbacks in using Tobin's q, both related to a measurement error. The first issue concerns stock market bubbles, which lead to a measurement error with respect to the fundamental value of the firm.[13] The second source of measurement error in Tobin's q arises if there are more frictions than only the adjustment costs of capital. Then firm value includes all other rents the firm can exploit, but these rents might well be not related to the size of the stock of capital.[14]

Therefore, in this paper we directly derive the productivity of capital from sales, employment and wage data instead and employ the ideas developed in Cooper and Haltiwanger (2002) to measure capital productivity. As we can show that measured in this way, both the productivity and capital are non stationary variables which are cointegrated, there must exist a long run equilibrium relation between capital and productivity. Therefore, there also exists an equilibrium level of capital $k^{**}(\xi, e)$ for any level of productivity (and liquidity). This equilibrium level of capital equals the average stock of capital a company holds between two adjustments. If depreciation is constant over time, the target level of capital k^* and k^{**} also equal each other up to a constant that reflects the expected depreciation between two adjustments.

The derivation of k^* in Cooper and Haltiwanger's (2002) framework starts from the static optimization problem of a firm that employs capital K and labor L (denoted in straight levels not logs) and produces with a Cobb-Douglas production function. This firm generates revenues Y according to

$$Y = \Xi L^{\alpha} K^{\beta}. \tag{5}$$

[13] See for instance Cummins, Hassett and Oliner (1999) or Bond and Cummins (2000).

[14] See Gilchrist and Himmelberg (1998), Gomes (2001) or Merz and Yashiv (2002) for discussions of this topic.

For this revenue function, we assume that the firm has market power or that production is otherwise subject to decreasing returns of scale, so that $\alpha + \beta < 1$.[15] The variable Ξ represents total-factor productivity. If labor can be flexibly adjusted, the optimal employment decision is described by

$$wL = \alpha Y \tag{6}$$

with w being the wage per employee. Replacing L in the production function by optimal employment then yields

$$Y = \left[\Xi\left(\frac{\alpha}{w}\right)^{\alpha}\right]^{\frac{1}{1-a}} K^{\frac{\beta}{1-\alpha}}. \tag{7}$$

Therefore, the (log) productivity of capital is given by the first factor,

$$\xi := \ln\left(\left[\Xi\left(\frac{\alpha}{w}\right)^{\alpha}\right]^{\frac{1}{1-a}}\right). \tag{8}$$

For given parameters β and α, capital productivity ξ can be calculated directly after one has inferred Ξ from the production function, or indirectly from the production function and optimal employment according to (6). Taking logs from (7), we obtain

$$\xi = \ln(Y) - \frac{\beta}{1-\alpha}k. \tag{9}$$

Replacing Y according to (6) now yields the indirect measure of capital productivity, which is the measure Cooper and Haltiwanger (2002) use

$$\xi_{ind} = \ln\left(\frac{wL}{\alpha}\right) - \frac{\beta}{1-\alpha}k. \tag{10}$$

Theoretically, the direct and the indirect productivity measure should be the same, empirically they differ somewhat.[16] Hence, we take the average of both measures as the

[15] The assumption of decreasing returns to scale is well supported by the data, for every firm in the sample $\alpha + \beta < 1$ holds.

Furthermore, note that although firm indices like α_i are suppressed for notational convenience in the equations characterizing the empirical model and in its application the parameters α and β will be firm-specific.

[16] Reasons for the difference can be labour market imperfections or deviations from the assumption of a Cobb Douglas production function.

empirical measure of productivity later on:[17]

$$\hat{\xi} = \frac{\xi_{dir} + \xi_{ind}}{2}. \tag{11}$$

The static optimal stock of capital k^{***}, which is the stock of capital the firm would hold in the absence of any adjustment costs and financial frictions, can be determined by maximizing instantaneous profits, $Y - uc \cdot K$, which is earnings Y from (7) minus the cost of capital for which uc denotes the user cost. From this optimization, we can infer the (log) optimal static stock of capital, k^{***}, as

$$k^{***} = \frac{(1-\alpha)}{1-(\alpha+\beta)} \left[\gamma_1 \xi - \ln uc\right], \ \gamma_1 = 1. \tag{12}$$

Since there are adjustment costs and capital market imperfections k^{**} and k^{***} may differ. This is reflected by allowing the elasticity of capital to productivity to be different from $\frac{(1-\alpha)}{1-(\alpha+\beta)}$, i.e. the coefficient γ_1 of ξ is different from 1. Moreover, the financial friction can be reflected by assuming that the user cost uc depends on the aggregate risk-free rate of return r, a constant parameter μ that reflects firm-specific risk and other fixed differences among firms, and a term reflecting the impact of liquidity on capital cost, $\gamma_2 e$. Putting these items together, we obtain

$$\ln(uc) = r + \mu + \gamma_2 e \text{ and} \tag{13}$$

$$k^{**}(e, \xi) = \frac{(1-\alpha)}{1-(\alpha+\beta)} \left[\gamma_1 \xi - r + \mu + \gamma_2 e\right]. \tag{14}$$

2.4 Empirical specification and econometric methodology

2.4.1 Measuring the desired stock of capital

This now gives four parameters to be estimated for each company: α, β, and $\gamma_{1,2}$. The most straightforward approach would be to estimate the parameters α, β directly from the production function (again small letters denote logs)

$$y_{it} = \alpha_i l_{it} + \beta_i k_{it} + \ln(\Xi_{it}). \tag{15}$$

[17] The average is very close to the common factor obtained by factor analysis. Moreover, the results do not depend strongly on the use of either of the three alternative productivity measures.

Interestingly, using the direct productivity measure gives a slight increase in explanatory power for the investment regression at the cost of slightly less sensible estimates for the capital stock analysis compared to the indirect productivity estimate. The average of both estimates turns out to perform close to both the direct measure in the investment regression and close to the indirect measure in the levels regression.

However, both the dynamic structure of the data and technological heterogeneity among firms complicate the analysis. The expenditure shares for both labor and capital differ substantially between firms (see figure 1). Therefore, we must expect $\alpha_i \neq \alpha_j$, $\beta_i \neq \beta_j$ in general and cannot estimate (15) by panel estimation techniques. Moreover, in our sample y is non-stationary due to a unit-root in productivity, so that we cannot estimate the production function in levels. However, even in first differences one needs to account for endogeneity of capital and labor and has to use lagged first differences as instruments. With only between 5 and 35 observations per firm, this direct approach is rendered infeasible. Therefore, α_i and β_i are estimated as average expenditure shares.[18] Expenditures on labor are calculated as the product of the average wage per employee of a specific company times the number of employees. Expenditures on capital are calculated as the average depreciation rate of a specific company plus a 3% real interest rate times the actual stock of capital.

Thereafter, $\hat{\xi}$ is calculated as the mean of the indirect and the direct capital-productivity measure as described in section 2.3. Capital productivity $\hat{\xi}$ turns out to be I(1), just as capital is I(1). Therefore we can estimate $\gamma_{1,2}$ from the cointegration relationship

$$k_{it} = k_{it}^{**} + x_{it} = \frac{(1-\alpha_i)}{1-(\alpha_i+\beta_i)} \left[\gamma_1 \hat{\xi}_{it} - r_t + \mu_i + \gamma_2 e_{it}\right] + x_{it}. \tag{16}$$

The error term x_{it} turns out to be stationary, so that there is a cointegration relationship between productivity and capital; see table 2 in section 3 for unit-root tests. For the estimation we use the Panel-Full-Modified-OLS (PFM-OLS) estimator of Phillips and Moon (1999) and the Panel-Dynamic-OLS (PD-OLS) estimator of Kao and Chiang (2000), which both are asymptotically equivalent. For r_t and μ_i we control using fixed time and individual specific effects. The PFM / PD-OLS estimators unbiasedly estimate the parameters of I(1) variables in a cointegration relation. They also yield unbiased

[18] To see that this is a feasinble approach, note that L ,Y and K are endogenous unit-root processes, all driven in the long-run by Ξ. Thus, (6) gives—loosely speaking—a cointegration relation. So that OLS for each firm on this equation is super-consistent, but collapses to $\hat{\alpha}_i = \frac{1}{T_i} \sum \frac{w_{it} L_{it}}{Y_{it}}$ if we can write the estimation equation with a "heteroscedastic" error term

$$w_{it} L_{it} = \alpha Y_{it} + Y_{it} v_{it}.$$

If the error were to come in multiplicatively, the geometric mean were appropriate. Note that using the geometric mean instead, does not alter the results significantly.

Still, Caggese (2003, p. 10) has argued that this procedure will lead to biased results, if labor and capital employment decisions are constrained by some third variable, e.g. by financial constraints. In our case r might depend on the financial conditions. Nevertheless, as we find an only very minor influence of finance on the long-run stock of capital, the bias can be expected to matter only marginally. More formally, our estimates are consistent under H_0^0, the hypothesis we test.

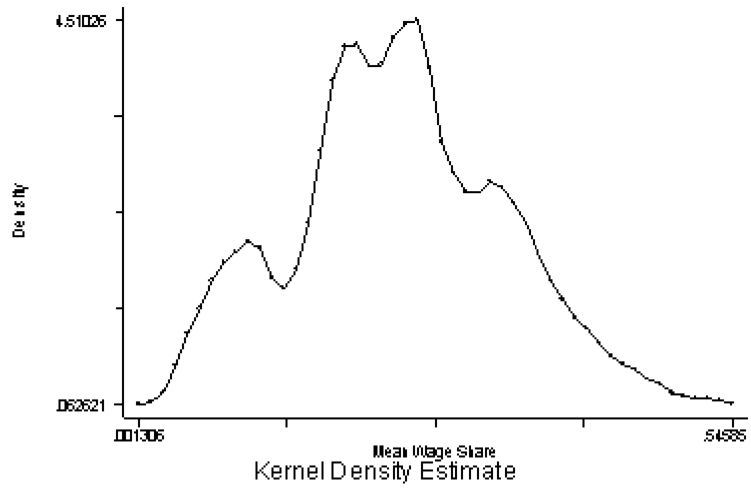

Figure 1: Firm–specific average wage shares

estimates for the parameters of stationary variables, if these are weakly exogenous. Yet, if there is lag dependency or if the contemporaneous shocks to equity and capital are correlated, the PFM-OLS estimator is likely to be biased.

Since productivity is clearly a unit-root process in our data, the parameter estimate $\hat{\gamma}_1^+$ is asymptotically unbiased. However, for our liquidity proxy, the equity-ratio, e_{it}, which is measured as the ratio of book value of equity over the book value of assets, the unit-root test rejects the unit root hypothesis and there might be an endogeneity or lag-dependency problem. To remove the contemporaneous correlation, one can replace e_{it} by e_{it-1} and argue that the beginning of period liquidity determines managerial discount factors. However, if there is lag dependency, the estimator remains biased.

As γ_1 can be estimated consistently in any case, there is another way to obtain estimates of γ_2. Since k and k^{**} are found to be cointegrated indeed, $k_{it} - \gamma_1 \hat{\xi}_{it} - r_t - \mu_i$ is stationary and can be expressed as

$$z_{it} := x_{it} + \gamma_2' e_{it} + \gamma_2'' e_{it-1} = \tag{17}$$

$$k_{it} - \hat{\gamma}_1^+ \hat{\xi}_{it} + \hat{r}_t + \hat{\mu}_i = \gamma_2' e_{it} + \gamma_2'' e_{it-1} + C^* (L) \psi_{it}, \tag{18}$$

with a moving-average error-term $C^* (L) \psi_{it}$ on the right hand side and a stationary cointegration error, z_{it}, on the left. In this equation, either γ_2' or γ_2'' is zero if the beginning-of-period equity ratio or the end-of-period equity ratio determines the managerial discount rate respectively. Similarly, the equity-ratio depends on its previous realization and on past and current capital imbalances, so that we have as a second equation

$$e_{it} = \rho e_{it-1} + \eta_1 z_{it} + \eta_2 z_{it-1} + \nu_{it}. \tag{19}$$

If now $\eta_1 = \gamma_2'' = 0$ and

$$\forall j \geq 0 : cov\left(\psi_{it-j}, \nu_{it}\right) = 0, \tag{COV}$$

or $\gamma_2' = 0$ and

$$\forall j \geq 0 : cov\left(\psi_{it-j}, \nu_{it-1}\right) = 0, \tag{COV*}$$

then the parameters in (19) can be consistently estimated as all regressors are predetermined.[19] In a second step (18) can be estimated, using the fixed-effects OLS-residual $\hat{\nu}_{it}$ as an instrument for e_{it}.

Since we would rather assume current capital imbalances influence the current equity ratio, but not the other way round, models with $\gamma_2' = 0$ are the preferred ones. Yet, this comes at the price that the assumption (COV*) is more restrictive: If x measures fundamental investment incentives imperfectly and current residual changes in equity reflect productivity, assumption (COV*) will be wrong and our estimates will be biased upwards. However, we can expect the endogeneity problem to be less pronounced for these instrumental variable estimators $\hat{\gamma}_2^{IV}$ than for the PFM-OLS estimate $\hat{\gamma}_2^{+}$.

2.4.2 Estimation of the investment function

Having estimated the long-run relation between capital, productivity and finance, we can turn towards estimating the investment equation. Similar to the standard error-correction framework, the error term is obtained as

$$x_{it} := k_{it}^{**} - k_{it-1} = \hat{\gamma}_1^{+}\hat{\xi}_{it} + \hat{\gamma}_2^{IV} e_{it-1} - (\hat{r}_t + \hat{\mu}_i) - k_{it-1}. \tag{20}$$

[19] Strictly speaking, this is only true if $T \to \infty$, as we use fixed effects OLS. Therefore, in small samples our estimates are biased. However, we are mainly interested in generating an instrument that is orthogonal to the within transformed variables, but contains information on e_t. Hence, one should not interpret the estimates of (19) structurally.

To avoid this problem at least for the estimation of (19) we additionally estimate a IV-regression of this equation in first differences, in which Δe_{it-1} is instrumented by e_{it-2} and Δe_{it-2}. Yet, we only obtain $\Delta\nu_{it}$ as error-term, which may be correlated with ξ_{it} under assumption (COV) as ξ_{it} and ν_{it-1} may be correlated. Hence, we use $\Delta\nu_{t+1}$ as instrument for e_t. The results are reported under IV-PDOLS, but are not significantly different from the ones obtained by fixed effects OLS. Nevertheless, notice that for (19) the small sample bias only vanishes if one assumes $\forall s, t : cov\left(\xi_{is}, \nu_{it}\right) = 0$.

The linear error-correction model results in the following investment equation

$$\Delta k_{it} = i_{it} = \alpha x_{it} + \sum_{j=1}^{L} \theta_j^T \Psi_{it-j} + \phi_{it}. \tag{21}$$

In this equation the Ψ-terms, $\Psi := \left(\Delta k \, \Delta\hat{\xi} \, \Delta e\right)^T$, pick up the short-run dynamics. Without short-run dynamics, this error-correction model corresponds to a micro-model with quadratic adjustment costs. More general forms of adjustment costs, including fixed costs, transform the adjustment speed parameter α into a function (here approximated by a polynomial) of (e, x) . Neglecting the short run dynamics, we obtain

$$i_{jt} = \left[\sum_{j=-1}^{p} \sum_{k=0}^{q} \alpha_{jk} \left(x_{it} - \bar{x}_{i.}\right)^j \left(e_{it} - \bar{e}_{i.}\right)^k\right] \left(x_{it} - \bar{x}_{i.}\right) + \phi_{it}.$$

If adjustment costs are non-convex, then the adjustment speed $\frac{\partial i}{\partial x}$ increases when the gap x becomes larger than its average value $\bar{x}_{i.}$.

Besides this semi-parametric estimation for investment, we also carry out a nonparametric analysis. This allows us, to obtain direct inference on the *average* derivatives of expected investment with respect to finance and fundamentals. To analyze the data non-parametrically, the data are pooled after individual fixed effects have been removed. The derivatives are then calculated by applying a local linear kernel-estimator to the data. For this estimation, two estimators are most prominent candidates: One is Li et al.'s (1998) (analytic) estimator from the local linear regression in which the average derivative is computed by taking the sample average over the pointwise estimates of $\hat{\beta}_{x,e}(q)$, $q := (x, e)$. These pointwise estimates are generated by weighted least squares on

$$y_i = m(q) + \beta_x(q)(x_i - x) + \beta_e(q)(e_i - e) + u_i, \; q := (x, e). \tag{22}$$

The weights themselves are computed using a kernel-function.[20] Alternatively, numerical derivatives can be used, which are obtained as

$$\widetilde{\beta}_k(q) = \frac{\hat{m}\left(q + \frac{1}{2}h_{q,k}\mathsf{e}_k\right) - \hat{m}\left(q - \frac{1}{2}h_{q,k}\mathsf{e}_k\right)}{h_{q,k}}, \; q := (x, e), \tag{23}$$

where $h_{q,k}$ is the (variable) window-width used to generate kernels at evaluation point q, e_k is the k-th unit-vector, and $\hat{m}$ again is the weighted least squares estimate.[21]

[20] This estimator is asymptotically equivalent to Rilstone's (1991) estimator.

[21] In most cases the numerical estimator has better small sample (and asymptotic bias) properties

Average derivatives are computed as sample-means of $\widetilde{\beta}_{x,e}(q)$. Both the analytic and the numerical average derivative estimator converge with parametric rates. In the non-parametric analysis, a Gaussian-product kernel has been employed. To generate window width $h_{q,k}$, an adaptive two-stage estimator for the window width was used, starting with a fixed window width of $s_k n^{-1/4}$, in the first stage, in which s_k stands for the standard deviation of argument k and n is the number of observations.[22]

3 Brief description of the data

The data that we analyze come from the "Bonner Stichprobe" which is a sample of annual company accounts of German companies. Most of which are large listed stock companies. The data covers the time-period 1960 to 1997. The panel is unbalanced and contains 694 companies (observational units) and 18943 observations in total. Thus, the average time of a company in the sample is 28.7 years.

The database includes complete profit- and loss-statements as well as annual accounting data. Moreover, for the allmost all company years data on average wages and salaries as well as on the number of employees are reported.

Firms which are holding companies ("Holdinggesellschaft"), or groups ("Konzerngesellschaft") have to be removed from the sample. Their company accounts basically duplicate company accounts of operating companies that are also recorded in the data or summarize the accounting information of the companies within the group only partially.[23] Additionally, we have to drop a few firm years for which data seems inconsistent with usual accounting standards (e.g. negative depreciation, very high appreciation). This leaves us with a sample of about 10000 observations.

If a firm series is split into two parts by removing a single observation (due to data inconsistency) and if both parts are long enough to be sensibly analyzed, the second part of the series is identified as a new firm. If the missing observation separated the series into a very short and a longer one, the short one was completely removed, i.e. only firms with five or more consecutive observations remain in the sample. Additionally, single observations were removed, if the investment rate differed from its mean by 5 times the standard deviation (removing 11 observations), if the current equity ratio (in logs) differed from the firm mean log-equity ratio by 4 standard deviations (39 observations),

(Ullah and Roy, 1998). However, its asymptotic variance is not yet known (Pagan and Ullah, 1999).

[22] See Pagan and Ullah (1999) for details on the non-parametric estimation techniques.

Moreover, note that in comparison to pointwise derivative estimations, this choice of window width leads to substantial undersmoothing.

[23] For example there is "RWE Holding AG" which has no other economic activity but holding 100% of the stock of "RWE AG". The former is traded on the stock market whereas the latter is not.

Table 1: Descriptive Statistics "Bonner-Stichprobe"

Variable	Obs.	Mean	Std. Dev.	Min	Max
investment-rate	9770	0.210	0.120	-0.140	0.8254
capital	9770	164.9	491.5	0.036	6893.2
equity-ratio	9770	0.403	0.139	0.016	0.9371
real wage	9770	14.80	4.969	1.756	36.908
total value added (turnover)	9770	464.7	1415	0	20584
No. Employees	9770	6009	17813	4	215800

or if the turnover-change differed from the mean by 6-times the standard deviation (16 observations). Moreover, firms were excluded, if their average wage-share or proxied average cost-of-capital share exceeded 70% (removing 122 observations). This leaves us with 449 firms and a total of 9770 observations, making an average of 21.75 accounting years per firm.

The stock of capital series has been generated using the perpetual inventory method, investment, wages and profit were deflated using the producer-price index for investment goods. The descriptive statistics are presented in Table 1. Although Goolsbee and Gross (1997) report that assuming a homogeneous capital good biases the estimated investment function towards a linear specification, it is necessary to do so in this paper, since many firms do not report stock and depreciation of land and buildings and machinery separately.

However, although we measure investment rates on the basis of a homogenous capital good, investment rates still exhibit moderate excess skewness and kurtosis. That kurtosis and skewness is only moderate reflects the fact that most firms in the sample are aggregates of many plants, but only at the plant level is investment highly lumpy. Still, we find that 17.4% of all firm years exhibit an investment-spike and these spikes account for 36.1 % of all investment, when using the widely employed cut-off value of 30% for the definition of such a spike.

Since Cooper and Willis (2004) have emphasized the importance of the assumption of a unit-root in productivity for the gap model, we test this assumption. The time-dimension of the sample that we use is only moderate compared to the number of observational units and the sample is unbalanced. Therefore, the Breitung-Meyer (1994) unit-root test has been chosen. Table 2 reports the results. The hypothesis of a unit root cannot be rejected for capital, revenues (turnover), the number of employees, and for the measure of capital-productivity $\hat{\xi}$ that has been described in the previous sections. We can however reject the unit root hypothesis for the equity ratio. Also for the

Table 2: Breitung-Meyer Unit-Root Tests

Variable	estim. root	sign. of $\alpha \geq 1$
log No. Employees	1.034	1
log turnover	1.010	1
log capital	1.008	1
log equity-ratio	0.965	0
$\hat{\xi}$	1.010	1
x	0.965	0

cointegration error x, we can reject the null of a unit-root. Consequently, productivity $\hat{\xi}$ and capital must be cointegrated.

4 Empirical results

4.1 Long-run optimal stock of capital

The cointegration relation (16) can be estimated using the Panel-Dynamic-OLS-Estimator (PD-OLS) of Kao and Chiang (2000), the Panel-Full-Modified-OLS-Estimator (PFM-OLS) of Phillips and Moon (1999) , or OLS controlling for fixed effects. The PD-OLS estimator usually puts no cross-sectional restriction on the short-run dynamics. However, allowing for heterogeneous short-run dynamics in the PD-OLS regression means including more than 2300 parameters (for 2 lags and 2 leads of first differences). Therefore, the PD-OLS regressions assume a homogeneous short-run dynamics in all but two specifications (PD-OLS-Ind). In these two specifications industry specific short-run-dynamics are assumed.[24] All regressions control for fixed time- and firm-effects. Table 3 presents the main results of the four PD-OLS, two PFM-OLS and the OLS regressions.

Model PD-OLS(-Ind)-1 includes 3 lags and leads of $\Delta\hat{\xi}$, while model PD-OLS(-Ind)-2 only includes 2 lags and leads, but of both Δe and $\Delta\hat{\xi}$. Standard errors are calculated on the basis of the PFM-OLS estimate, using the average number of observations per firm for the respective calculation of the standard error. Standard errors are generated for both, the case where the regressors are I(1) and the case where the regressors are I(0). Significance is indicated on the I(1) basis.

Although γ_1 is significant in all regressions and the estimates are reasonably large, the parameter of productivity γ_1 is clearly smaller than 1. This means that the static target level of capital and the dynamic optimal level of capital differ somewhat. One reason for this could be that a fixed percentage of revenues has to be attributed to a not modelled

[24] The industry variable provided in the Bonn Database has been used for classification. This variable splits up the database in 52 different industries. Note however, that this variable does not coincide with SIC.

Table 3: Single-Stage Cointegration regressions

Estimator and Model	PD-OLS-1	PD-OLS-2	PD-OLS-Ind-2	PD-OLS-Ind-1
γ_1 $\left(\hat{\xi}\right)$	0.7626***	0.7404***	0.7442***	0.7717***
γ_2 (log equity-ratio)	-0.0386	-0.0275	-0.0144	-0.0175
No. of Parameters	39	43	378	535
No. of Observations	6612	7447	7447	6612[a]
No. of Firms	383	416	383	412

Estimator and Model	PFM-OLS	PFM-OLS (e_{t-1})	OLS	FM-std. err. I(1)
γ_1 $\left(\hat{\xi}\right)$	0.6938***	0.6009***	0.6442***	0.02768
γ_2 (log equity-ratio)	-0.0345	-0.0484*	0.03813	0.02544
No. of Parameters	39	39	39	std. err. I(0)
No. of Observations	9289	8823	9767	γ_1 .02802
No. of Firms	442	442	442	γ_2 .02763

***/**/* indicate significance at the 1/5/10% level

(quasi-)rent. This would deterministically drive up the productivity measure. Another reason for $\gamma < 1$ could be that wages endogenously react to productivity growth in the long-run. In both cases, the gap between desired and actual capital is still recovered by the regression. Moreover, the estimates for γ_1 are in line with the estimates Caballero et al. (1995) obtained for their cost-of capital proxy.[25]

For the estimate of γ_2 evidence is mixed assuming that the log equity-ratio is I(1). However, the unit-root test clearly rejects this hypothesis. Therefore, we use a formula analogous to the general one provided in Phillips (1995, p. 1038, eq. 14) to determine the standard error of the I(0) regressor e. Now, the standard error increases slightly, since the parameter estimates of I(0) variables are not super-consistent. In consequence, hypothesis H_0^0 cannot be rejected, which means the equity-ratio has no influence on the optimal-stock of capital. Moreover, all estimates except the OLS one have a negative sign. This means that higher equity ratios lead to lower optimal stocks of capital, which is contradictory to most of the earlier empirical financing-constraints literature. Also this seems inconsistent with the "wealth effect on the cost-of-capital" explanation that now is common in a number of theoretical (macro-)models since the seminal contribution of Bernanke et al. (1999).

Yet, our regressions have not controlled for the endogeneity of the equity-ratio. The fixed effects only remove a different baseline access to capital markets. Additionally, the

[25] However, if adjustment costs strongly dampen the variation of capital, it is well known, that our estimator will underestimate γ_1 by construction (Caballero, 1997, p. 8).

PD-OLS and the PFM-OLS account for short-run correlations. (See Phillips (1995) for details).

However, if there is lag dependency in the equity ratio, or if the contemporaneous shocks to equity and capital are correlated, the PFM-OLS and the PD-OLS estimates $\hat{\gamma}_2^+$ are likely to be biased. The parameter estimate for productivity $\hat{\gamma}_1^+$ remains asymptotically unbiased in any case because of the unit root in productivity $\hat{\xi}$. The contemporaneous correlation problem can be reduced by replacing e_t by e_{t-1}. This is even the better specification theoretically, if liquidity at the beginning of the period determines managerial discount factors. The resulting estimates are reported in column PFM-OLS (e_{t-1}). The estimate for γ_2 decreases further, becomes smaller and is now weakly significant. However, if there is lag dependency, the estimator still remains biased.

Therefore, we employ the two-step approach that has been developed in section 2. This instrumental-variable approach basically builds on the assumption that there is a triangular structure in the gap-equity relation, see equations (17) - (19). Table 4 presents the two-step estimates. Again, we cannot reject hypothesis H_0^0 of no long run influence of the equity ratio on the desired levels of capital. The estimated coefficient for equity is even closer to zero as the coefficient increases (from negative towards zero) if we control for endogeneity. This increase may be explained as follows: when high productivity (high gap) firms increase their stock of capital, they finance this increase of the capital stock with internal funds. This drives down the equity ratio and yields the found negative correlation of the equity ratio and the capital stock if endogeneity is not controlled for.

Now when endogeneity is controlled for, the estimated parameter γ_2 has a cost of capital interpretation. Since the estimated coefficients $\gamma_{1,2}$ equal $\frac{\partial x}{\partial \xi}$ and $\frac{\partial x}{\partial e}$ up to a common constant, we can obtain the elasticity of the cost of capital with respect to finance as $\frac{\frac{\partial x}{\partial e}}{\frac{\partial x}{\partial \xi}} = \frac{\gamma_2}{\gamma_1}$ (see section 2 equation (4)). Based on the estimated parameters this elasticity lies approximately between -0.02 and 0.065. Hence, finance is negligible for the long-run firm decisions. Consequently, if there was no additional frequency effect for finance on investment, i.e. H_0^1 holds true and $\frac{\partial i}{\partial e} = \frac{\partial x}{\partial e}\frac{\partial i}{\partial x}$, finance also would not matter for investment.[26]

This hypothesis is tested next. To be as conservative as possible in testing for H_0^1, we use the PD-OLS-Ind-1 model with $\gamma_2' = 0$ for estimating the investment-function, since the corresponding estimates for the long-run influence of equity are among the largest ones for a '$\gamma_2' = 0$'-model.

[26] There is also an important technical implication of this result. Since the stock of capital and the equity-ratio are uncorrelated in the long run, the equity ratio e–our financial variable–cannot have much predictive value for future productivity, once current productivity is controlled for.

Table 4: Two-Stage Cointegration regressions

Model	$\gamma_2'' = \eta_1 = 0$	$\gamma_2' = 0$
Estimator	PFM-OLS	PFM-OLS
γ_1 [a]	0.6938***	0.6009***
γ_2	-0.0118	0.0187
η_1	0 (assumed)	-0.0177
η_2	0.0119	0.0285**
ρ	0.8057***	0.8058***
No. of Observ.	9364	8897
Estimator	PD-OLS-Ind-1	PD-OLS-Ind-1
γ_1 [a]	0.7717***	0.7714***
γ_2	-0.0074	0.0310
η_1	0 (assumed)	-0.0067
η_2	0.0107	0.0164
ρ	0.8056***	0.8056***
N	8897	6720
Estimator	PD-OLS-2	PD-OLS-2
γ_1 [a]	0.7404***	0.7404***
γ_2	-0.0091	0.0287
η_1	0	-0.0086
η_2	0.1121	0.0184
ρ	0.8056***	0.8056***
N	6720	6720
Estimator	IV-PD-OLS-Ind-2	IV-PD-OLS-Ind-2
γ_1 [a]	0.7442***	0.7442***
γ_2	0.0475	-0.0311
η_1	0 (assumed)	-0.0176
η_2	0.0058	0.0068
ρ	0.5436***	0.5458***

***/**/* indicate significance at the 1/5/10% level,

[a] Std. Err. from first stage PFM-OLS is 0.02833

4.2 Investment behavior

4.2.1 Parametric Analysis

Consequently, for the analysis of investment the gap between desired and actual stock of capital is measured as

$$x_{it} := k_{it}^{*} - k_{it-1} = \hat{\gamma}_1^{+}\hat{\Pi}_{it} + \hat{\gamma}_2^{IV} e_{it-1} - (\hat{r}_t + \hat{\mu}_i) - k_{it-1}, \tag{24}$$

where the estimates of $\hat{\gamma}_1^{+}$, $\hat{\gamma}_2^{IV}$ from PD-OLS-Ind-1 are used. This gap can be interpreted as mandated investment. Fixed effects are removed by subtracting firm-specific means from all variables. This especially controls for inter-firm differences in the optimal capital imbalance which result from different target levels of capital. In contrast, it is not obvious how to treat aggregate shocks and estimate $\hat{r}_t$. The coefficients of time-dummies used in the cointegration regression would also pick up the state of aggregate mandated investment (which is to the most extent driven by productivity). However, aggregate mandated investment should not be subtracted from the individual mandated investment as both together determine the actual investment of a firm non-linearly. Hence, we project the series of the time-specific effects obtained from the cointegration-regression on a series of real-interest rates and take these projections as estimate of $\hat{r}_t$.[27] Last, we need to determine which equity ratio to use. Since the equity ratio at the beginning of the investment period is predetermined and also more likely to shape managerial decisions, the equity-ratio used in the investment regressions is the equity ratio in the opening balance of a firm. Table 5 presents the regression results for the investment equation [28]

$$i_{jt} = \left(\sum_{j=-1}^{p} \sum_{k=0}^{q} \alpha_{jk} \left(x_{it} - \bar{x}_{i.}\right)^{j} \left(e_{it} - \bar{e}_{i.}\right)^{k} \right) \left(x_{it} - \bar{x}_{i.}\right) + \phi_{it}. \tag{25}$$

The direct influence of e $(j = -1)$ reflects that the firm-average mandated investment (gap) $\bar{x}_{i.}$ and the target level are only equal up to a constant since depreciation deterministically opens a gap between actual and desired capital between two adjustments.

The parametric estimates for the investment function show a moderate degree of

[27] The correlation between the real-interst-rates and the time-specific effects is quite low. This reflects the fact, that the aggregate (average) capital-imbalance is mainly driven by productivity and / or demand-shocks, which vary more than the real interest-rate.

[28] To preclude the possibility that our results are driven by extreme observations of mandated investment, we remove all observations from the sample which deviate by more than 4 standard deviations from the firm-specific average in the capital-imbalance measure.

Table 5: Short-Run Parametric Estimates

	Model 1		Model 2	
Variable	Coefficient	(Std. Err.)	Coefficient	(Std. Err.)
x	0.1495***	0.0058	0.1504***	0.0042
x^2	0.0650***	0.0112	0.0662***	0.0111
x^3	-0.0416*	0.0231	-0.0429***	0.0073
x^4	-0.0415***	0.0113	-0.0414**	0.0110
x^5	-0.0204	0.0162	–	–
e	0.0246*	0.0091	0.0162***	0.0045
e^2	0.0666**	0.0258	0.0329**	0.0132
e^3	-0.1430	0.0796	–	–
e^4	-0.0792*	0.0592	–	–
e^5	0.2718	0.1274	–	–
xe	0.0358***	0.0130	0.0347***	0.0122
$(xe)^2$	-0.0630	0.0439	-0.0840**	0.0398
xe^2	0.0152	0.0288	–	–
x^2e	0.0240	0.0216	–	–
$const$	-0.0060***	0.0014	-0.0052***	0.0013
Adj. R^2	0.1977	–	0.1973	–
No. Obs.	8973	–	8973	–

***/**/* significant at the 1/5/10% level

convexity with respect to mandated investment, x. The average second order derivative $\frac{\partial^2 i}{\partial x^2}$ equals 0.126.[29] Moreover, the investment function becomes concave when x is about as large as one standard deviation. Although the investment function should be convex with non-convex adjustment cost, this convexity of the investment function should not be taken at face value as evidence against the fixed adjustment cost model. It more likely reflects the fact, that most companies in the sample are multi-establishment/multi-plant firms, so their individual investment function rather equals an average over many investment functions of different plants with mean capital imbalance x. Due to this fact—and as for example Whited (2004) or Goolsbee and Gross (1997) argue—the observed investment function becomes less curved. Figure 2 plots the shape of the estimated investment function for 1.5 standard deviations around the means of x and e.

The relatively strong influence of the equity ratio reveals that finance effects investment mainly through altering the adjustment frequency. This is also reflected by an important interaction between fundamental capital imbalance x and the financial variable e. This frequency effect is much harder to interpret in a model with convex costs

[29] Average parametric derivatives are calculated by differencing the estimated function (Model 2) and then averaging over the observation-wise calculated derivatives.

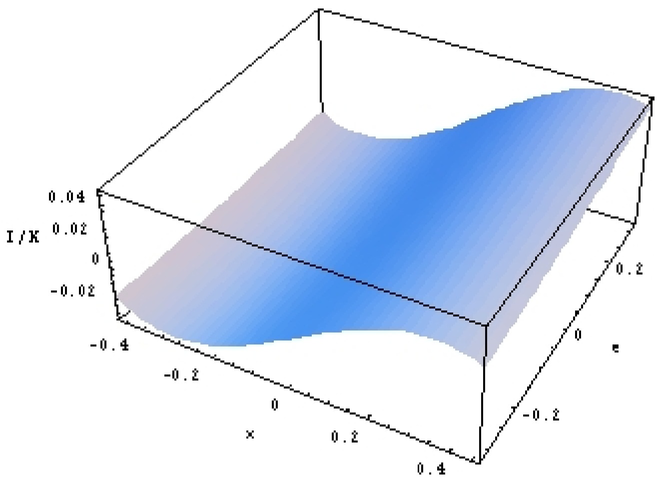

Figure 2: Investment function, e and x between 1.5 std. errors

than in a model with non-convex adjustment costs. Intuitively in a convex cost model with liquidity dependent cost of capital, the influence of the equity-ratio should be completely captured by the previously estimated long-run effect. In line with this intuition and with our results, Whited (2004) reports that for a simulated model with non-convex adjustment cost and financial frictions the financial frictions only influence investment hazards but not optimal capital levels.

Still, in our data the fundamental investment incentives x explain most of the (explained) variation in investment and the adjusted R^2 is notably large for an investment regression. This also shows that the quality of $\hat{x}$ is relatively good as a measure of investment incentives.[30] Compared to Cooper and Haltiwanger's (2002) results for the reduced form investment equation, the R^2-statistics is substantially larger. Since we have borrowed the general technique to measure productivity from their paper but allowed for technological heterogeneity, the increase in R^2 may potentially be interpreted as evidence for technological heterogeneity being non-negligible.

In table 6 average derivatives $\frac{\partial i}{\partial e}$ for the parametric model are reported. The results are in line with the frequency-effect interpretation of short term influences of equity on investment introduced in section 2: Equity has a much larger effect if there are strong fundamental investment incentives anyway. Interestingly, the Kaplan and Zin-

[30] Typical R^2 statistics in most (homogeneous) investment regressions (using q or some other estimator for productivity) range in between 5 and 10%. See for instance Cooper and Haltiwanger (2002) or Barnett and Sakellaris (1999).

Table 6: Parametric estimates of average derivative $\frac{\partial i}{\partial e}$

	$x \leq 0$	$x > 0$
$e \leq 0$	-0.0001	0.0110
$e > 0$	0.0194	0.0328

gales (1997) result, that financially constrained firms are less sensitive to changes in liquidity is replicated. Firms with below "normal" equity exhibit a far lower average derivative with respect to equity.

4.2.2 Non-parametric Analysis

As the parametric analysis naturally depends on the choice of the functional form a nonparametric analysis has also been employed. Additionally, this allows us to obtain direct inference on the derivatives of *expected* investment with respect to finance and fundamentals. To analyze the data non-parametrically, the data are pooled after individual fixed effects have been removed. The estimation techniques have been outlined in section 2.3.

Table 7 reports average derivative estimates for both the direct and the numerical estimator. Additionally, a robust estimator $\overset{\simeq}{\bar{\beta}}{}^{cens}$ is calculated as the mean of all point-wise derivative estimates within ± 5 standard deviations around the mean $\overset{\simeq}{\bar{\beta}}$. This estimator is not affected by outliers generated by undersmoothing and low density in the tails of the distribution of x, e, and i.

Table 7: Average nonparametric derivative estimates

(a): Full sample

	$\bar{\beta}_x$	$\bar{\beta}_e$	$\overset{\simeq}{\bar{\beta}}_x$	$\overset{\simeq}{\bar{\beta}}_e$	$\overset{\simeq}{\bar{\beta}}{}^{cens}_x$	$\overset{\simeq}{\bar{\beta}}{}^{cens}_e$
overall	0.137	0.018	0.134	0.020	0.138	0.018
std. error[31]	0.0009	0.0013	–	–	–	–

(b): Stratified

Case		$\bar{\beta}_x$	$\bar{\beta}_e$	$\overset{\simeq}{\bar{\beta}}_x$	$\overset{\simeq}{\bar{\beta}}_e$	$\overset{\simeq}{\bar{\beta}}{}^{cens}_x$	$\overset{\simeq}{\bar{\beta}}{}^{cens}_e$
$x > 0$	$e > 0$	0.163	0.043	0.152	0.050	0.159	0.044
	$e \leq 0$	0.137	-0.021	0.126	-0.018	0.129	-0.016
$x \leq 0$	$e > 0$	0.124	0.043	0.121	0.043	0.127	0.039
	$e \leq 0$	0.126	-0.003	0.141	-0.002	0.141	-0.002

The overall speed of adjustment, measured by the derivative of the investment rate with respect to the capital imbalance x, is at 0.137 again rather low. This speed of adjustment is equivalent to an overall half-life of 4.49 years for a gap between desired and actual stock of capital. In comparison, the average derivative with respect to equity is, at 0.0142, quite substantial. An effect of that size would result in a long-run elasticity of capital k with respect to liquidity e of approximately 0.1 in the linear error correction model model (21). This is substantially larger than the previously obtained long-run estimate. Moreover the estimated influence of finance on investment is an effect additionally to the one the equity-ratio has on x. Therefore, the hypothesis H_0^1 of no frequency effect is equivalent to a test of $\beta_e = 0$. Consequently, we clearly have to reject H_0^1 in favor of the alternative hypothesis of a frequency effect of liquidity on investment. Therefore, one can expect finance, e, and fundamental investment incentives, x, to interact in a complex manner.

This is validated by looking at the sample stratified by values of e and x: Firms with abundant financial ($e > 0$) resources that wish to increase capital ($x > 0$) adjust 27% faster than they do when financial resources are scarce ($e \leq 0$ and $x > 0$). If firms wish to decrease the stock of capital ($x \leq 0$) there is no such strong effect of finance, at best larger internal funds make those firms more reluctant to decrease their stock of capital. Additionally the stratified non-parametric estimates reveal a non-linear relationship of investment and the gap only for firms with above normal equity ratios. The picture remains the same irrespective of the actual choice of the estimator.

5 Discussion

The aim of this paper—as formulated in section 2—was to test hypothesis H_0^0 and H_0^1, the "long-run" and "short-run versions" of the Modigliani and Miller theorem, if one likes. With respect to these hypotheses we can state the following:

1. H_0^0 can at best be rejected on weak grounds, so in the long-run finance does not seem to matter.

2. Since the estimated short-run influence of equity, measured by the average derivative, is both substantial and significant, H_0^1 has to be rejected. This holds true, as we only tested for *additional* short-run influences of liquidity on investment.

Hence, the question arises, why there is the additional short-run effect that has been found. Inspecting the short-run parametric and nonparametric estimates, we find a substantial interaction of finance and fundamentals in determining investment.

To further condense the results and to give them a more intuitive appeal, table 8 presents (geometric) means of pointwisely calculated half-lifes of capital imbalances. These are calculated as $\frac{\ln 0.5}{\ln\left(1-\hat{\beta}(x)\right)}$. As the pointwise derivatives exhibit large variation, and sometimes obtain negative values or are larger than 1, the derivatives are

Table 8: Average half-lifes of capital-imbalances (in years)

	$x \leq 0$	$x > 0$
$e \leq 0$	5.06	4.41
$e > 0$	5.19	4.07

re-estimated with a three times larger window width. Again the larger x, the faster is investment, and if firms wish to invest, more equity speeds up investment.

A potential shortcoming of our analysis might be that we omitted short run dynamics to keep the empirical model simple and close to the theoretical gap model. However, one may argue that the non-linearities found are a mere result of the omitted dynamic links between changes in productivity, capital, and the equity-ratio.[32] Table 9 presents the regression results from a model similar to the one in table 5 but augmented by some short-run dynamics.

Though the point estimates change, the overall structure of the estimated error-correction, i.e. investment function, remains the same. Moreover, the serial correlation as measured by the parameter on Δk_{it-1} is small although significant. Hence, our results seem—at least to a certain extent—robust to the inclusion of short-run dynamics. However, the levels of significance of the terms involving equity drop. Yet, they jointly remain highly significant.

We may also briefly compare the results obtained for the German data in the preceding section, with the results for the UK data analyzed in the companion paper (Bayer, 2002). As the Cambridge DTI-Database, on which is the UK data used, does not contain wage data for a large number of firms, but yearly data on firm-specific subsidies, the "within variation" (i.e. after controlling for fixed effects) in subsidies has been used to estimate the coefficient on the user cost of capital θ instead of the one on productivity γ_1. Table 10 cites the regression results for this cointegrating relation. The coefficient for the user-cost of capital is insignificantly different from its neoclassical benchmark value 1 and the coefficient κ of the equity-ratio is small but statistically significant and still larger than the one we obtained for the German sample.

For the investment function, again the estimated derivative with respect to the equity ratio is much larger than the coefficient in the cointegration relation; Table 11 reports the estimated average derivatives of the investment function. $\bar{b}^{FE}$ is Ullah and Roy's

[32] While a structural interpretation for including the lagged change in the stock of capital could be a delivery lag, an interpretation for other short-run dynamics is far from obvious; and even if we find a significant short-run dynamics, this could well be due to an imperfect approximation of the true functional form which is picked up by the first-differences of the equity-ratio and productivity.

Table 9: short-run parametric estimates, dynamics-augmented

Variable	Model 1 Coefficient	(Std. Err.)	Model 2 Coefficient	(Std. Err.)
Δk_{it-1}	0.1627***	0.0064	0.1630***	0.0064
Δe_{it}	-0.0204**	0.0064	-0.0210***	0.0064
Δe_{it-1}	0.0276***	0.0073	0.0281***	0.0073
$\Delta \Pi_{it}$	0.0340***	0.0073	0.0340***	0.0073
$\Delta \Pi_{it-1}$	0.0458***	0.0074	0.0459***	0.0074
$(k^*_{it} - k_{it-1})$	0.1586***	0.0060	0.1580***	0.0044
$(k^*_{it} - k_{it-1})^2$	0.0394***	0.0116	0.0405***	0.0115
$(k^*_{it} - k_{it-1})^3$	-0.0579**	0.0251	-0.0458***	0.0081
$(k^*_{it} - k_{it-1})^4$	-0.0255*	0.0132	-0.0237*	0.0128
$(k^*_{it} - k_{it-1})^5$	0.0067	0.0191	–	–
e_{it-1}	0.0194**	0.0091	0.0105**	0.0049
e^2_{it-1}	0.0604**	0.0260	0.0306**	0.0134
e^3_{it-1}	-0.1302	0.0805	–	–
e^4_{it-1}	-0.0809	0.0595	–	–
e^5_{it-1}	0.2252	0.1310	–	–
$(k^*_{it} - k_{it-1})\, e_{it-1}$	0.0239*	0.0135	0.0205	0.0128
$[(k^*_{it} - k_{it-1})\, e_{it-1}]^2$	-0.0550	0.0501	-0.0972**	0.0432
$(k^*_{it} - k_{it-1})\, e^2_{it-1}$	0.0373	0.0315	–	–
$(k^*_{it} - k_{it-1})^2\, e_{it-1}$	0.0084	0.0240	–	–
const	-0.0094***	0.0013	-0.0088***	0.0013
Adj. R^2	0.2637		0.2635	
No. Obs.	8153		8153	

***/**/* significant at the 1/5/10% level

Note that $k^*_{it} := \hat{\gamma}^+_1 \hat{\Pi}_{it} + \hat{r}_t + \hat{\mu}_i + \gamma''_2 e_{it-1}$, $x := (k^*_{it} - k_{it-1})$

Table 10: Estimates from the cointegration regression (UK-sample, PFM-OLS, Within)

	κ	θ	Observations
PFM-OLS[a]	0.079***	0.98***	5944
prelim. OLS	0.070	1.08	
std. err. I(1)	0.026	0.07	
std. err. I(0)[b]	0.035	0.14	
OLS	0.082	0.71	7147

[a] Only observational units have been used for which (outliers removed) 5 or more observations are available.

[b] The standard errors are obtain as panel analogues to Phillips (1995, p. 1033ff).

Table 11: UK-sample–Average first-order derivatives of the investment rate $i(e, z)$

Number of observations: $N = 6950$				std. deviation
derivative	$\widetilde{\bar{b}}$	$\bar{b}$	$\bar{b}^{FE}$	for $\bar{b}$
$\frac{\partial \hat{i}}{\partial \hat{x}}$	0.5057	0.5146	0.5341	0.0068
$\frac{\partial \hat{i}}{\partial e}$	0.1555	0.1588	0.1782	0.0097

(1998) fixed effects estimator.

Interestingly, the response of the investment-rate to changes in equity is even larger for the UK sample. The sensitivity of investment to liquidity is still larger for the UK sample when the smaller half-lives of capital imbalances in the UK are taken into account for the comparison to the German sample. This result is similar to what Bond et al. (2003) report. In their analysis they focus on cross-country differences in the influence of liquidity on investment and find that financial factors appear to be much more influential for a sample of UK firms than for a sample of German firms. However, in our case the estimation procedures for the German and the UK sample slightly differs, so that any differences in estimates have to be interpreted with more care than in the case of Bond et al. (2003).

6 Conclusion

In this paper the interaction of fixed capital adjustment costs and financial frictions was studied empirically. To do so, a proxy for the productivity of investment was obtained. This proxy explicitly accounted for the technological heterogeneity of the observed firms. Since capital productivity follows an I(1) process in the analyzed sample and is cointegrated with capital, we have performed a two step non-linear cointegration analysis for capital and investment. From the estimation of the long-run relation, we found that liquidity is hardly correlated with the choice of capital. Accordingly, in the long-run the hypothesis that finance does not matter or put differently the Modigliani-Miller theorem holds, respectively cannot be rejected.

However, the picture substantially changes, if the effect of finance on investment decisions is analyzed. Larger equity ratios starkly increase the speed of adjustment of capital to its equilibrium level. This means that financial considerations primarily have intertemporal substitution effects for investment. Firms endowed with more financial means do not invest more, but they invest more often (and in smaller amounts) than firms which have lesser financial means. Figuratively, finance is the grease but not the fuel of investment.

This finding, just like the others of our analysis, obviously hinges on the quality of the proxy used for capital productivity and later for mandated investment. The derived measure of mandated investment (gap) can explain a large fraction of the variation in investment. This suggests that the proxy can be considered as reasonable. Moreover and since the proxy explicitly allows for heterogeneity, the relatively good quality of the proxy suggests that there is indeed a substantial degree of technological heterogeneity across firms. On the more formal side, the econometric issues raised by Cooper and Willis (2004) do not apply to our data since productivity is non-stationary. Therefore, the differences in adjustment speeds at high and low mandated investment can be interpreted structurally as evidence for non-convex adjustment cost reconfirming our a priori assumption of non-convex costs. Moreover, the differences in the adjustment speed are not only econometrically but also economically significant.

These estimation results raise the question of what to conclude for economic primitives. Although the estimation technique of the present paper does not recover the parameters of economic primitives, such as the adjustment cost function, themselves and hence does not allow us to draw strong structural conclusions directly, some economic structures are more compatible with our results than others. That finance primarily influences the adjustment speed but not the level of the stock of capital, for example, intuitively seems to be not compatible with a model in which the managerial discount factor depends strongly on the financial situation of a company.

Similarly our results have some policy implications although they come from a reduced form model: Suppose there are shocks to the balance sheet positions of firms (e.g. through exchange rates as in Céspedes et al. (2000), Aghion et al. (2001) or Devereux and Lane (2001)), then this paper's results predict a strong short run real impact but a weak long-run impact, which is somewhat different to the financial accelerator model of Bernanke et al. (1999). Moreover, this impact will depend on the position of the economy along the business cycle. Thus policies that influence the balance sheet (shocks) will be rated differently along the business cycle. Therefore, policy makers, central banks for example, need to take into account both, the fundamental economic investment incentives for companies and their financial situation to forecast the effectiveness of a given policy. For tax policy the results also provide some interesting detail. A tax system that encourages higher equity ratios may be welfare enhancing. Firms in such systems would adjust their stock of capital to its desired level more frequently, so that the aggregate allocation of real capital becomes more efficient, at least in the partial model studied.

7 Data Appendix

The dataset used is the "Bonner Stichprobe", a sample of annual company accounts of German companies. To the very most, these companies are large listed stock companies. As explained in the main text, some companies have to be removed since their accounts are only consolidated accounts of other companies in the sample (holding companies or goups) and do not contain actual information on individual economic activities.

Additionally we remove observations for which the data seemed inconsistent with usual accounting standards (e.g. negative depreciation, very high appreciation) or otherwise seemed to be mis-reportings (like changes in the stock of capital by more than factor 10) are removed from the sample, sample size drops substantially to 9969 observations. If removing a single observation (due to data inconsistency) splits a firm-series into two parts which are long enough to be sensibly analyzed, the second part of the series is identified as a different firm. If the missing observation separated the series into a very short and a longer one, the short one was completely removed, i.e. only firms with five or more consecutive observations remain in the sample. Additionally, single observations were removed, if the investment rate differed from the mean by 5 times the standard deviation (removing 11 observations), differed from the firm specific log-equity ratio by 4 standard deviations (39 observations), or if the turnover-change differed from the mean by 6-times the standard deviation (16 observations). Moreover, firms were excluded, if their average wage-share or proxied average cost-of-capital share exceeded 70% (removing 122 observations). This leaves us with 449 firms and a total of 9770 observations, making an average of 21.75 accounting years per firm. In many cases series for "land and buildings" and "machinery" were not reported separately over the full sample period. Therefore "capital" is identified as "total tangible fixed assets" ("Sachanlagevermögen"). The equity ratio is defined as the sum of all assets minus total liabilities (as reported in the balance "Verbindlichkeiten") devided by the sum of all assets ("Bilanzsumme"). However, the sum of all assets is corrected for the different valuation of tangible assets following the perpetual inventory method instead of taking their book value.

Depreciation rates were generated as reported depreciation relative to the reported stock of capital before depreciation. For a number of firm-years the data contains capital sales as well as gross investment. For some firm years only investment net of capital sales are reported. The stock of capital used for the analysis was generated by the perpetual inventory method. Investment was deflated by the producers-price index for investment goods. To account for sales of capital, it was assumed that in case capital is sold, the capital stock of each vintage is reduced by the same fraction. Thus we obtain for the capital series (in real terms):

$$K_{it} = K_{it-1}\left(1-\delta_{it}\right)\left(1-\frac{CS_{it}}{\widehat{K}_{it}+CS_{it}}\right)+\frac{I_{it}}{P_t}, t > T_i$$

$$K_{iT_i} = \frac{\widehat{K}_{iT_i}}{P_{T_i}}.$$

Here $\widehat{K}_{it}$ is the reported stock of capital of firm i at time t. CS_{it} are reported capital-sales and I_{it} is reported investment, P_t is the price-index, T_i is the year when firm i joins the sample. Wages, profits etc. were also deflated using the producer-price index for investment goods as well.

By using the perpetual inventory method, problems induced by a change in accounting standards in 1987 are partly avoided. However, the perpetual inventory method leads to different (mostly larger) stocks of capital than reported. Thus the book-value of equity was adjusted as well.

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