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Investment decisions and the soft budget constraint

Evidence from a large panel of Hungarian firms

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Abstract

This paper investigates the investment behaviour of a large panel of Hungarian firms in the period 1989–99, in order to assess the impact of institutional and regulatory changes on the efficiency of credit allocation. We find that the role of financial factors for investment decisions has changed significantly after the introduction of major financial reforms, and that firms were affected differently depending on their ownership type. Reforms have hardened the budget constraint of private domestic firms, particularly small ones, and reduced informational problems for foreign-owned firms. State-owned firms remained subject to a soft budget constraint. In particular, small state firms became more sensitive to financial conditions, whereas large state firms were unaffected and kept operating under a soft budget constraint.

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1. Introduction

One of the key issues in the transition process of formerly centrally planned economies is the establishment of a functioning financial system that allows an efficient allocation of credit. This requires designing institutions and rules to impose financial discipline on firms that in the old system were subject to a soft budget constraint, due to the fact that loss-making firms could rely on external assistance by means of direct subsidies, favourable tax conditions, or bail-out credits.[1]

In the past two decades, the Hungarian financial system has undergone a number of major changes in order to increase its efficiency (see, for example, Halpern and Wyplosz, 1998, and Stephan, 1999). In particular, the banking sector reform was aimed at the separation of central banking and commercial banking functions, the restructuring of commercial banks and the definition of an appropriate regulatory framework. At the same time, the introduction of a new bankruptcy law was intended to enhance allocative efficiency and to provide agents with the appropriate incentives.

Institutional reforms *per se*, however, are not a sufficient condition for the achievement of an efficient credit allocation system. Once the new rules are created, agents must learn to play by the rules. In particular, in transition economies, lenders have to develop project appraisal and monitoring skills, while borrowers must learn to respond appropriately to the new system of incentives. Whether the reform process in transition economies has succeeded in establishing an efficient incentive-based financial system is an open and much-debated issue (for the Hungarian economy, see, for example, Bonin and Schaffer, 1995, 2002; Colombo and Driffill, 2003; Halpern and Körösi, 2001).

The objective of this paper is to examine whether reforms to the Hungarian financial system have been successful in increasing its efficiency. We investigate the investment behaviour of a large panel of Hungarian manufacturing and construction firms between 1989 and 1999, and examine whether the institutional and regulatory changes have succeeded in imposing a hard budget constraint, an issue that has been recently addressed by a number of studies for different transition economies (see, for example, Bratkowski *et al.*, 2000; Budina *et al.*, 2000; Lízal and Svejnar, 2002; Maurel, 2001; Schaffer, 1998; Sgard, 2001; Volchova, 2003).

Our study makes a number of contributions to the existing literature. First, it is the first analysis of the role of financial factors for investment decisions in the Hungarian economy based on a large and representative firm-level panel dataset (the sample covers about 80 percent of total employment and value added of the

[1] Under a soft budget constraint, '[. . .] the financial position of the state-owned firm is not without influence. Although there is a budget constraint that forces some financial discipline on the firm, it is not strictly binding, but can be "stretched" at the will of the higher authorities. In principle, the firm should cover expenditures from revenues made on the market. In practice, earnings from the market can be arbitrarily supplemented by external assistance.' Kornai (2000, p. 25). See also Kornai (1980, 1986).

Hungarian manufacturing sector). Second, the long time period covered by our dataset (1989 to 1999) allows us to compare firms' investment behaviour *before* and *after* the introduction of major financial reforms. We can therefore provide evidence not only on the extent to which firms face a soft budget constraint, but also on whether financial system reforms have affected the degree of rationing or softness of the budget constraint. Third, we provide evidence on investment behaviour by ownership type, analysing individually state-owned, private domestic and foreign-owned firms. Fourth, we use firm size and solvency to check the interpretation of the observed relationship between investment and liquidity indicators. Finally, we explicitly deal with firms' unobserved heterogeneity, estimating investment equations in first differences with a GMM estimator that exploits optimally the moment conditions available for each time period under the assumption of serially uncorrelated errors.

The main findings of the analysis can be summarized as follows. The role of financial factors for investment decisions has changed significantly after the introduction of major financial reforms, and firms were affected differently depending on their ownership type. In the post-reform period, small private firms came to face binding financial constraints, whereas state firms kept facing a soft budget constraint, although the investment decisions of small state firms became more sensitive to financial conditions. Foreign-owned firms were subject to a hard budget constraint in both periods, but became less sensitive to financial conditions after 1993, possibly indicating that reforms have been successful in lowering informational costs.

The remainder of the paper is structured as follows. Section 2 briefly describes the theoretical background, discussing the role of financial factors and informational asymmetries for investment decisions, and the different roles they play in market and transition economies. Section 3 reviews the main steps of the reform of the Hungarian financial system. Sections 4 and 5 describe the econometric methodology and the dataset, while Section 6 presents the results of the empirical analysis. Section 7 concludes with a discussion of the main findings of the analysis.

2. Credit rationing, soft budget constraints and investment decisions

Following Modigliani and Miller (1958), neoclassical theoretical analyses of the determinants of investment decisions generally abstracted from the role of firms' financial positions. More recently, however, the economics of asymmetric information has provided solid microeconomic foundations for the role of financial factors in determining investment levels. In the presence of informational asymmetries, the availability of internal funds allows firms to undertake investment projects without resorting to high-cost external finance. In addition, stronger balance sheet positions lower the cost of external finance. Firms' net worth positions therefore determine their capacity to obtain external funds and, as a consequence, their investment and production levels.[2]

At the empirical level, the evidence on the role of financial constraints for investment decisions can be traced back to the original work of Fazzari *et al.* (1988), who showed that the sensitivity of investment spending to financial positions is higher for firms *a priori* considered likely to be credit constrained. Subsequent studies have generally confirmed such findings, extending the analysis along several dimensions.[3] In this literature, a positive and significant relationship between investment and indicators of liquidity is taken as evidence that firms are credit constrained, whereas perfect capital markets would imply no such relation as internal and external financing would be perfect substitutes.

A similar approach has been followed to investigate the sensitivity of investment decisions to financial positions in transition economies. However, in contrast to market economies, in transition economies the absence of a positive and significant relationship between investment and financial indicators is not likely to indicate perfect capital markets: it rather suggests that firms are subject to a soft budget constraint, because they have access to external finance irrespective of their profitability. In other words, whereas in estimating cash flow-augmented investment equations for market economies the null hypothesis is perfect capital markets, in the case of formerly planned economies the null hypothesis is the presence of a soft budget constraint.

This approach has been followed in a number of recent papers that investigate investment decisions in transition economies. Lízal and Svejnar (2002) analyse firms' investment decisions in the Czech Republic between 1992 and 1998, finding that cooperatives and small firms are credit rationed, whereas large state-owned and private firms operate under a soft budget constraint. Budina *et al.* (2000), who examine the role of liquidity constraints for investment decisions of Bulgarian manufacturing firms in the 1993–95 period, find that size and financial structure help to determine the extent to which firms are credit constrained and that soft budget constraints continue to play a major role.[4]

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[2] See, for example, Bernanke and Gertler (1989), Greenwald and Stiglitz (1993) and Kiyotaki and Moore (1997). Other important theoretical works on the financial propagation mechanism include Calomiris and Hubbard (1990) and Gertler (1992).

[3] A number of articles have considered different datasets for the United States (Calomiris and Hubbard 1995), countries other than the United States (Blundell *et al.*, 1992; Chirinko and Schaller 1995; Devereux and Schiantarelli 1990; Hoshi *et al.*, 1991), alternative sample split criteria to identify credit constrained firms (Oliner and Rudebusch, 1992; Whited, 1992), and alternative model specifications (Bond and Meghir, 1994; Hubbard *et al.*, 1995). See also Hayashi and Inoue (1991) and Gertler and Gilchrist (1994).

[4] Among related studies, Volchova (2003) estimates an accelerator model for Russian industrial firms in 1996 and 1997, finding that firms in unregistered groups invest a larger proportion of their retained earnings relative to the rest of the economy, whereas Bratkowski *et al.* (2000), examine survey data for Czech, Hungarian and Polish newly established private firms to assess the presence of credit constraints, and conclude that imperfections in capital markets do not seem to restrain the growth of new private firms.

For the Hungarian economy, to our knowledge, there is no firm-level study of investment decisions based on a large panel dataset, with the only exception of Maurel (2001), who analyses company accounts between 1992 and 1998, and finds that credit rationing applies to all categories of firms (foreign-, private- and domestically-owned). The focus of that study, however, is on the role of investment in improving the technical efficiency of firms, as measured by total factor productivity. Another work based on a large panel of Hungarian firms is the one by Sgard (2001), who adds to the large body of literature on foreign direct investment, finding that between 1992 and 1999 foreign equity is associated with higher productivity levels and substantial positive spillover effects on aggregate TFP growth. More recently, Perotti and Vesnaver (2004) investigate the financing of investment in a sample of 56 listed Hungarian firms between 1992 and 1998, finding evidence of significant financial constraints with the exception of foreign-owned firms.[5]

Against this background, our study contributes to the existing literature on investment decisions in transition economies not only by filling the gap for the Hungarian economy, but also by exploiting the richness of our firm-level dataset to explicitly address the efficiency effect of financial sector reforms on state, private, and foreign enterprises. To this purpose, before presenting the results of the empirical analysis, the next section briefly reviews the recent transformation of the Hungarian financial sector, examining in particular the banking sector reform and the bankruptcy law.

3. Financial sector reform in Hungary

One of the key elements of the reform of the Hungarian financial system was the restructuring of the banking sector. A two-tier banking system had already been established in Hungary in 1987, when three state-owned banks had taken over commercial functions from the National Bank of Hungary, which retained central banking functions. Under the planned system, the monobank did not operate on the basis of profit considerations, and its portfolio included a high share of non-performing loans, which were inherited by the newly established commercial banks. After 1990, new commercial banks entered the market, even though foreign participation remained relatively low until 1994. The average quality of the loans portfolio remained low, due to both the absence of an appropriate regulatory

[5] Other important related studies for the Hungarian economy include Bonin and Schaffer (1995), who provide an assessment of the banking and bankruptcy reforms on the basis of a survey of 200 manufacturing firms, and Halpern and Körösi (2001), who estimate frontier production functions to investigate the impact of competition on the efficiency of the corporate sector. More recently, Colombo (2001) and Csermely and Vincze (2003) examine the determinants of the capital structure of Hungarian firms, finding evidence of imperfections that constrain firms in the achievement of their optimal capital structure.

framework to enforce prudential lending practices, and the lack of expertise by bank managers, which often resulted in bad lending decisions. The interplay of these factors resulted in a series of banking crises in the early 1990s. The reaction of the Hungarian authorities was twofold: a program of bank consolidation was started in 1992, and a strict bankruptcy law was enacted to enforce hard budget constraints on firms.

The consolidation programme foresaw first recapitalization and then privatization of existing banks. As a result of recapitalization, the fraction of bad loans started to decrease in 1994, and it has steadily declined thereafter, reaching levels comparable to those of Western economies. The privatization of banks took off between 1994 and 1995 with the government selling strategic shares to foreign banks and other foreign investors. Over the period 1994–2000, direct state ownership fell from 65 percent to less than 20 percent, whereas the share of foreign-owned banks rose from 20 to 80 percent (see Abel and Bonin, 2000).

It is commonly agreed that the banking sector reform in Hungary has been to a large extent successful in establishing an efficient system of independent and financially strong commercial banks (see, for example, Halpern and Wyplosz, 1998, and Stephan, 1999). A key factor of success was the outward orientation of the reform, with foreign banks being allowed not only to become shareholders of domestic institutions, but also to establish their own subsidiaries. The presence of foreign ownership had positive spillover effects, increasing competition in the sector and introducing innovative technologies and higher quality banking services. More advanced banking skills enabled foreign firms to screen and monitor loans more efficiently, contributing significantly to the reduction of bad loans.

The new bankruptcy law was the second pillar of the legislative shock therapy implemented in 1992. It established two possible tracks, liquidation and reorganization, both of which allowed for the continuation of the firm after restructuring.[6] In addition, it imposed an automatic trigger that required a firm to file for reorganization if it was unable to repay any debt to any creditor within 90 days of the debt becoming due. The motivation for the strictness of the bankruptcy law and for the automatic trigger was the Hungarian authorities' concern about two main problems: creditor passivity and soft budget constraints.

The bankruptcy code was therefore engineered with the primary objective of improving the state of payments discipline and hardening budget constraints, particularly through the limitation of inter-enterprise debt arrears. Nevertheless, as observed by several authors, (see, for example, Bonin and Schaffer, 2002), the severity and strictness of the bankruptcy law was probably excessive, as it represented one of the major causes of the credit crunch that hit the Hungarian economy

[6] Despite the difference in the denomination of the two tracks, in both cases the firm was given the opportunity to reorganize and restructure. The actual difference was that, under the reorganization track, control remained with the incumbent management while reorganization took place, whereas under the liquidation track, control was transferred to liquidators.

in 1992–93.[7] As a consequence, the bankruptcy law was amended in late 1993, one of the most important changes being the abolition of the automatic trigger. Despite these changes, the reforms implemented in 1992–93 provided Hungary with one of the strictest banking and bankruptcy regimes in Eastern Europe.

4. Methodology

The relevance of financial factors for corporate investment decisions is commonly investigated by adding financial indicators, such as cash flow, to empirical specifications derived from a real investment model. The estimated coefficients for the financial indicators are interpreted as a measure of the sensitivity of investment to financial constraints. In this study we estimate an accelerator model of investment demand (see Appendix 2 for details):

$$\left(\frac{I_{i,t}}{K_{i,t-1}}\right) = \beta_0 + \beta_1\left(\frac{I_{i,t-1}}{K_{i,t-2}}\right) + \beta_2\left(\frac{Y_{i,t}}{K_{i,t-1}}\right) + \beta_3\left(\frac{Y_{i,t-1}}{K_{i,t-2}}\right) + \varepsilon_{i,t} \tag{1}$$

where I denotes gross investment, measured as the change in the level of net fixed assets plus depreciation, K is capital stock (net fixed assets), Y indicates net sales, as a proxy for output, and $\varepsilon_{i,t} = \alpha_i + \gamma_i + \eta_{i,t}$, with α_i representing firm-specific effects, γ_i time-specific effects, and $\eta_{i,t}$ the idiosyncratic component of the error term.

This specification, known as the modified accelerator model, is based on the level of sales, rather than the change in sales as in the traditional accelerator model. Theoretically, the modified accelerator can be derived from a standard neoclassical model with 'putty clay' capital (Oliner *et al.*, 1993), or from an accelerator model with delivery and installation lags (Abel and Blanchard, 1988). Empirically, it has been used, among others, by Fazzari *et al.* (1988), Vermeulen (2002) and, for transition economies, Lízal and Svejnar (2002). Oliner *et al.* (1995), using quarterly data for the aggregate business sector in the United States, find that the modified accelerator outperforms the traditional accelerator model both in sample and out of sample.

Equation (1) reflects firms' investment demand and implicitly assumes perfectly elastic credit supply or, in the case of a transition economy, a soft budget constraint. In order to account for the possibility that firms face constraints in obtaining external financing, we augment the basic equation with lagged values of cash flow (see, for example, Fazzari *et al.*, 1988 and Bond *et al.*, 2003):

[7] Moreover, the emphasis on payment discipline created several distortions: as pointed out by Mitchell (1998) and Bonin and Schaffer (2002), the automatic trigger was not based on a measure of insolvency, but rather on a measure of illiquidity. As a consequence, even profitable and viable firms would be forced to enter reorganization if they had overdue payables, independently of their amount.

$$\left(\frac{I_{i,t}}{K_{i,t-1}}\right) = \beta_0 + \beta_1\left(\frac{I_{i,t-1}}{K_{i,t-2}}\right) + \beta_2\left(\frac{Y_{i,t}}{K_{i,t-1}}\right) + \beta_3\left(\frac{Y_{i,t-1}}{K_{i,t-2}}\right) + \beta_4\left(\frac{CF_{i,t-1}}{K_{i,t-2}}\right) + \varepsilon_{i,t} \tag{2}$$

where cash flow is measured by adding depreciation to profits after interest, tax and preference dividends.

The use of investment/cash flow sensitivities to assess the role of financing constraints has been widely debated (see, for example, Fazzari *et al.*, 2000; Kaplan and Zingales, 1997, 2000). In particular, Kaplan and Zingales (1997) argue that investment/cash flow sensitivities do not increase monotonically with the degree of financing constraints, and that most of the *a priori* sorting criteria used in the literature to identify financially constrained firms are theoretically ambiguous. Our approach is not subject to this criticism, given that we focus on the comparison of investment/cash flow sensitivities in the pre-reform and post-reform periods. The sorting criterion is temporal, so that we do not split the cross-sectional distribution of firms, but compare the role of cash flow in different periods for all firms in the sample.

It is also important to observe that, as discussed in Section 2, when testing the significance of cash flow for the investment levels of domestic firms in a transition economy, the null hypothesis is the presence of a soft budget constraint: the absence of a significant relationship between investment and cash flow is likely to indicate that firms are subject to a soft budget constraint, since they have access to external finance irrespective of their profitability.

In estimating Equation (2), the presence of the lagged dependent variable, which is correlated with the firm-specific component of the error term, implies that the OLS estimator is inconsistent even if the idiosyncratic component of the error term is serially uncorrelated. The *within* transformation, although eliminating the fixed effects, does not solve the problem, as it introduces correlation between the lagged dependent variable and the time averaged idiosyncratic error term (the same problem would apply to the random effect-GLS estimator).

An alternative solution for the correlation with the fixed effects is to first difference the data. The effect of differencing, however, is not only to eliminate the individual effects, but also to produce a first-order moving average error term. This, in turn, introduces correlation between the lagged dependent variable and the differenced error term, thus posing the problem of the selection of the appropriate instruments. To solve this problem, we follow the approach suggested by Arellano and Bond (1991) who developed a generalized method of moments (GMM) estimator that uses lagged levels of variables as instruments.

The GMM estimator optimally exploits all the moment conditions specified by the model. More lagged instruments become available for the differenced equations as we consider later cross-sections of the panel.[8] As the number of valid

[8] Very remote lags are unlikely to be informative instruments, and hence we did not use all available moment restrictions, but instruments dated $t-2$ to $t-6$. All the results reported are qualitatively robust to the choice of the instrument set.

instruments depends on the serial correlation of the idiosyncratic component of the error term, it is essential to verify the assumption of serially uncorrelated errors. To this purpose, we report the m_1 and m_2 statistics, which test for first- and second-order serial correlation in the residuals.[9] We also report p-values for the Sargan test of over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity, where k is the number of over-identifying restrictions.[10] We report one-step coefficient estimates, and test statistics based on heteroskedasticity consistent standard errors (see Arellano and Bond, 1991).

5. Data

The analysis presented in this paper is based on a large dataset of about 18,000 Hungarian firms from 1989 to 1999 (see Appendix 1 for details). The dataset contains information on balance sheet and income statement items, employment, export, ownership, regional location and industry classification at the four-digit level. From the original dataset we selected companies whose main activity was in the manufacturing and construction sectors. The resulting sample represents about 80 percent and 35 percent of total employment in the Hungarian manufacturing and construction sectors, respectively.[11]

Information about the distribution of equity ownership allows us to identify separately state-owned, private domestic and foreign firms. State and private firms were classified by majority ownership or, in the absence of an absolute majority stake, by relative majority. Foreign firms were defined as firms with more than 25 percent of capital owned by foreign firms. Firms with changing ownership over time were assigned to the ownership type they belonged to for the largest span of the sample.[12]

We applied a number of checks to account for possible data inconsistencies. First, we eliminated companies with illogical figures, such as negative sales, capital

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[9] Both statistics are asymptotically distributed as standard normal under the null of no serial correlation. If the assumption of no autocorrelation for the errors in levels is correct, so that second order lags of variables are valid instruments, the null hypothesis should be rejected for m_1 (because of the negative autocorrelation induced by first-differencing) but not for m_2.

[10] We also report the z_1 statistic, a Wald test of joint significance of the reported coefficients (asymptotically distributed as a χ^2_k under the null of no relationship, where k is the number of coefficients tested), and the z_2 and z_3 statistics, performing Wald tests of the joint significance of the coefficients of the time and industry dummies, respectively. Estimation was carried out using the DPD program (Arellano and Bond, 1988) with GAUSS version 5.5.

[11] In terms of value added, the firms in the sample represent about 81 and 34 percent of the total Hungarian manufacturing and construction sectors, respectively.

[12] It should be noted that this assignment procedure did not have a significant impact on the definition of ownership types, given that 75 percent of the firms did not change ownership, and 20 percent of the firms changed ownership type in only one or two years over the 11-year sample period.

Table 1. Median values by subsample

Sample	*I/K*	*Y/K*	*CF/K*	*D/A*	*E*	No. of obs.
Overall	0.15	4.32	0.20	0.43	79.64	25,202
1989–93	0.10	4.34	0.16	0.40	65.50	9,568
1994–99	0.18	4.30	0.23	0.44	86.50	15,634
Small firms	0.19	5.79	0.27	0.47	44.13	12,607
Large firms	0.13	3.50	0.16	0.39	184.57	12,595
Low leverage	0.11	3.40	0.20	0.27	91.43	12,606
High leverage	0.22	5.76	0.21	0.62	69.59	12,596

***Note*:** See Section 5 for details on data sources and definitions of variables. I = investment; K = capital; Y = sales; CF = cash flow; D = total debt; A = total assets; E = employment.

or employment. After computing the main variables, we also eliminated companies for which any of the variables of interest (investment, sales, cash flow) fell below the 2.5th percentile or above the 97.5th percentile of its distribution.[13] We then excluded companies with incomplete (discontinuous) time series and required that at least four consecutive annual observations on each of the main variables were available for the firms included in the final sample. These criteria left us with an unbalanced panel of 4,333 firms for a total of about 25,000 observations between 1989 and 1999.

Table 1 reports median values for the main variables used in the investment equations (investment, sales and cash flow relative to capital, and leverage and employment) for the overall sample, and grouping firms according to sample period (pre- and post-1993), size, and leverage.[14] For comparative purposes, Tables 2 to 4 present the same statistics by ownership type: state-owned, private domestic and foreign-owned companies, representing about 41, 39, and 20 percent of the whole sample, respectively.

Comparing the overall median values by ownership type, investment and cash flow are lowest in state-owned firms (0.10 and 0.15) and highest in foreign-owned firms (0.24 and 0.25). A similar pattern also applies to leverage (0.41 and 0.46 for state- and foreign-owned firms, respectively). The sales–capital ratio is highest for private domestic firms (5.23) and lowest for foreign firms (3.20). Finally, private domestic firms are generally smaller (the median employment is 71 against 80 for the whole sample).

[13] This check was necessary to control for the presence of outliers and the occurrence of major mergers or acquisitions. All the results were qualitatively robust to the choice of the trimming threshold.

[14] Small firms are defined as those whose employment is below the median value of the distribution for the whole sample. Low/high leverage firms are defined as those whose debt–asset ratio is below/above the median value for the whole sample.

Table 2. Median values by subsample: State firms

Sample	I/K	Y/K	CF/K	D/A	E	No. of obs.
Overall	0.10	4.16	0.15	0.41	80.00	10,331
1989–93	0.08	4.83	0.16	0.42	44.50	6,195
1994–99	0.13	3.36	0.15	0.40	132.42	4,136
Small firms	0.14	6.56	0.25	0.44	30.50	5,177
Large firms	0.08	3.13	0.11	0.39	251.00	5,154
Low leverage	0.08	3.20	0.16	0.26	92.20	5,166
High leverage	0.14	5.68	0.15	0.62	67.40	5,165

***Note*:** See Section 5 for details on data sources and definitions of variables. I = investment; K = capital; Y = sales; CF = cash flow; D = total debt; A = total assets; E = employment.

Table 3. Median values by subsample: Private firms

Sample	I/K	Y/K	CF/K	D/A	E	No. of obs.
Overall	0.16	5.23	0.23	0.43	71.09	9,949
1989–93	0.10	4.22	0.17	0.33	83.75	2,577
1994–99	0.18	5.63	0.26	0.46	66.06	7,372
Small firms	0.22	6.39	0.29	0.50	46.63	4,985
Large firms	0.11	4.47	0.19	0.36	123.88	4,964
Low leverage	0.09	4.03	0.20	0.26	84.73	4,980
High leverage	0.27	7.07	0.27	0.64	60.67	4,969

***Note*:** See Section 5 for details on data sources and definitions of variables. I = investment; K = capital; Y = sales; CF = cash flow; D = total debt; A = total assets; E = employment.

Focusing on subsamples within groups, investment, cash flow and leverage rise substantially in the 1994–99 subperiod, and this rise is particularly evident for private domestic firms. Small firms are characterized by higher median values for all the indicators, both in the overall sample and by ownership type. The disaggregation by leverage indicates that high-debt firms are characterized by higher investment and sales, and lower employment levels.

6. Results

Our empirical strategy consists of estimating the cash flow augmented accelerator specification in Equation (2), both in the whole sample of firms and in subsamples

Table 4. Median values by subsample: Foreign firms

Sample	*I/K*	*Y/K*	*CF/K*	*D/A*	*E*	No. of obs.
Overall	0.24	3.20	0.25	0.46	109.14	4,922
1989–93	0.25	2.67	0.17	0.43	112.20	796
1994–99	0.24	3.29	0.27	0.46	108.14	4,126
Small firms	0.25	3.33	0.27	0.47	58.50	2,464
Large firms	0.24	3.09	0.24	0.45	261.50	2,458
Low leverage	0.21	2.76	0.29	0.33	111.25	2,465
High leverage	0.27	3.91	0.21	0.61	102.38	2,457

***Note*:** See Section 5 for details on data sources and definitions of variables. I = investment; K = capital; Y = sales; CF = cash flow; D = total debt; A = total assets; E = employment.

defined by ownership type. We start by examining the relationship between cash flow and investment in the entire 1989–99 sample period. Next, we focus on how this relationship changes across the pre-reform and post-reform subsamples, and perform a number of robustness checks. Finally, in order to obtain a sharper interpretation of the results, we explore within subperiods differences across subsamples of firms defined according to size and leverage.

Table 5 presents estimates of the basic accelerator investment equations for the overall sample (column 2) and by ownership type (columns 3–5). Concerning the model specification, the diagnostic statistics are generally supportive of the validity of the instruments. In all equations, the m_2 statistic does not reject the hypothesis of no second-order serial correlation, whereas the m_1 statistic indicates significant (negative) first-order serial correlation. Both results are to be expected if the errors in levels are serially uncorrelated, which is a necessary condition for $t - 2$ lags to be valid instruments. In addition, with the only exception of the equation for the overall sample, the Sargan test does not reject the validity of the instruments used.

In the overall sample, lagged investment is positive and significant, and the coefficients for sales are significant and consistent with accelerator effects. A similar pattern applies to the estimates for the ownership subsamples, with the exception of the equation for state firms. It is interesting to observe that foreign firms display the highest investment persistence, with a point estimate for the lagged dependent variable (0.16) that is close to the ones observed for Western economies in similar specifications (see, for example, Bond *et al.*, 2003). Both sets of dummies (industry- and year-specific) are jointly significant in all the equations.

The cash flow coefficient is positive and significant in the overall sample, thus leading to rejection of the soft budget constraint null hypothesis. In the equations by ownership, it is interesting to observe that the coefficient for cash flow is lowest and not significant for state firms, whereas it is positive and significant for both

Table 5. Investment equations: Overall

Regressors	Overall	State	Private	Foreign
$(I_{i,t-1}/K_{i,t-2})$	0.100*	−0.013	0.131*	0.159*
	(7.74)	(−0.52)	(9.00)	(7.37)
$(Y_{i,t}/K_{i,t-1})$	−0.009*	−0.009*	−0.006*	−0.011
	(−4.21)	(−3.19)	(−3.18)	(−1.80)
$(Y_{i,t-1}/K_{i,t-2})$	0.007*	0.004*	0.008*	0.015*
	(6.03)	(2.86)	(8.85)	(3.41)
$(CF_{i,t-1}/K_{i,t-2})$	0.025*	0.003	0.041*	0.039*
	(3.41)	(0.24)	(4.16)	(2.78)
m_1 (1st order aut.)	0.000	0.000	0.000	0.000
m_2 (2nd order aut.)	0.676	0.868	0.815	0.819
Sargan test	0.000	0.383	0.119	0.998
z_1 (overall)	0.000	0.005	0.000	0.000
z_2 (time dummies)	0.000	0.000	0.000	0.000
z_3 (industry dum.)	0.000	0.000	0.030	0.000
No. of obs.	16,536	5,893	7,167	3,476

Notes: Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t-2)$ to $(t-6)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. t-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 4,333.

private and foreign firms. These preliminary findings are therefore consistent with the hypothesis that, in the whole 1989–99 period, Hungarian state firms faced a soft budget constraint, whereas private and foreign firms were subject to binding financial constraints. This hypothesis, however, deserves further investigation.

We therefore move to the analysis of how the investment behaviour of firms has been affected by financial reforms. We interact the cash flow variable with a dummy variable (and its complement to 1) that equals 0 up to (and including) 1993 and 1 thereafter, in order to compare the sensitivity of firms' investment behaviour before and after financial markets reforms. The choice of 1993 as the cut-off year for the sample split is based on a number of reasons. First, even though the bankruptcy and banking laws were introduced during 1992, a number of amendments were made during 1993, such as the elimination of the automatic trigger in the bankruptcy law. Second, it is reasonable to assume that the new regime displayed its effects only after some time from the introduction of the new regulations. Third, the loan consolidation programmes aimed at dealing with the bad debt problem

Table 6. Investment equations: Pre-post reforms (1993)

Regressors	Overall	State	Private	Foreign
$(I_{i,t-1}/K_{i,t-2})$	0.093*	−0.016	0.125*	0.158*
	(7.11)	(−0.61)	(8.24)	(7.32)
$(Y_{i,t}/K_{i,t-1})$	−0.009*	−0.009*	−0.005*	−0.011
	(−4.15)	(−3.14)	(−2.78)	(−1.85)
$(Y_{i,t-1}/K_{i,t-2})$	0.007*	0.005*	0.008*	0.015*
	(6.09)	(2.86)	(8.26)	(3.38)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}$	0.002	−0.004	0.006	0.064*
	(0.21)	(−0.31)	(0.32)	(2.17)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}93}$	0.056*	0.035	0.056*	0.037*
	(6.61)	(1.94)	(4.94)	(2.53)
$(CF/K)^{post} - (CF/K)^{pre}$	(4.17)	(1.77)	(2.27)	(−0.87)
m_1 (1st order aut.)	0.000	0.000	0.000	0.000
m_2 (2nd order aut.)	0.594	0.840	0.691	0.813
Sargan test	0.000	0.422	0.237	0.998
z_1 (overall)	0.000	0.007	0.000	0.000
z_2 (time dummies)	0.000	0.000	0.000	0.000
z_3 (industry dummies)	0.000	0.000	0.016	0.000
No. of obs.	16,536	5,893	7,167	3,476

***Notes*:** Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t-2)$ to $(t-6)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. t-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 4,333.

were implemented throughout 1992 and 1993 (see Bonin and Schaffer, 1995). Finally, at the empirical level, 1993 is preferable to 1992 as it produces subsamples of similar size (5 and 6 years, respectively).[15]

The results for the pre- and post-reform subperiods, presented in Table 6, are revealing. Looking at the overall sample of firms, in the pre-reform period the cash flow coefficient is close to zero and not significant, whereas in the post-reform period it is highly significant and about twice as large (0.056) as the corresponding estimate for the whole period (0.025). This finding seems to suggest that financial market reforms have indeed hardened budget constraints. The disaggregation by

[15] It should be observed that, as shown below, the results reported are robust to the choice of the cut-off year.

ownership is also particularly informative. For both state and private firms, the cash flow coefficient is close to zero and not significant before 1993, but it rises substantially after 1993. However, for private firms only the sensitivity of investment to financial conditions becomes strongly significant after 1993, whereas it is smaller and marginally significant for state firms. Note also that the change in the coefficients is significant only for private firms. For foreign firms the picture is quite different: investment is significantly affected by cash flow both before and after 1993, but the coefficient actually falls in the second period.

On the whole, these results indicate that financial reforms significantly affected the investment behaviour of all firms, but in different ways depending on the ownership type. Private firms come to face binding financial constraints in the post-reform period. State firms appear to keep facing a soft budget constraint, although their investment decisions become more sensitive to financial conditions. Foreign firms are subject to a hard budget constraint in both periods, but become less sensitive to financial conditions, possibly indicating that reforms might have been successful in lowering informational costs.

In order to verify the validity of these results, we perform a number of robustness checks. First, we consider the possibility that the changes in cash flow coefficients across subperiods might be due to differences in sample size: the number of observations available in the two subperiods is indeed quite different, if we take into account the fact that two cross-sections are lost at the beginning of the sample (1989–90) because of differencing and taking lags. We therefore consider an alternative definition of pre- and post-reform periods, selecting 1995 as the threshold year. This implies that the effective subperiods contain 5 and 4 cross-sections, respectively. The results, presented in Table 7, confirm and qualify those presented in Table 6 for the 1993 sample split: the cash flow coefficient is not significant throughout the sample period for state firms, it rises significantly after 1995 for private firms and, for foreign firms, it falls over time and is actually significant before 1995 but not significant thereafter.[16]

Second, we investigate the robustness of the results with respect to the choice of the instrument set. This is a particularly important issue in the present context, given that the sample split for the cash flow variable implies that less moment conditions are available for the first period than for the second period.[17] We therefore also present estimates obtained using instruments dated $t-2$ and $t-3$, so that the number of moment conditions is (almost) the same for each cross-section

[16] The results for private and foreign firms are consistent with those presented in Table 3, given the different subsample definition: as 1994 and 1995 are now part of the pre-reform period, cash flow becomes significant in the first subsample for private firms and not significant for foreign firms in the second subsample.

[17] In particular, for $t = 89, \ldots, 99$, if the instruments are dated $t-2$, to $t-6$, the cross-sections available as instruments are CF_{89} for ΔCF_{91}, CF_{89} to CF_{90} for ΔCF_{92}, CF_{89} to CF_{91} for ΔCF_{93}, and so on until CF_{93} to CF_{97} for ΔCF_{99}.

Table 7. Investment equations: Pre-post reforms (1995)

Regressors	Overall	State	Private	Foreign
$(I_{i,t-1}/K_{i,t-2})$	0.098*	−0.013	0.130*	0.156*
	(7.58)	(−0.53)	(8.88)	(7.27)
$(Y_{i,t}/K_{i,t-1})$	−0.009*	−0.009*	−0.005*	−0.012
	(−4.18)	(−3.15)	(−2.94)	(−1.91)
$(Y_{i,t-1}/K_{i,t-2})$	0.007*	0.005*	0.008*	0.015*
	(6.05)	(2.82)	(8.67)	(3.38)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}95}$	0.015	−0.001	0.028*	0.071*
	(1.53)	(−0.07)	(2.14)	(3.67)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}95}$	0.050*	0.026	0.056*	0.028
	(5.04)	(1.29)	(4.22)	(1.71)
$(CF/K)^{post} - (CF/K)^{pre}$	(2.69)	(1.19)	(1.55)	(1.82)
m_1 (1st order aut.)	0.000	0.000	0.000	0.000
m_2 (2nd order aut.)	0.730	0.853	0.840	0.710
Sargan test	0.000	0.387	0.151	0.999
z_1 (overall)	0.000	0.012	0.000	0.000
z_2 (time dummies)	0.000	0.000	0.000	0.000
z_3 (industry dummies)	0.000	0.000	0.027	0.000
No. of obs.	16,536	5,893	7,167	3,476

Note: Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t-2)$ to $(t-6)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. *t*-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 4,333.

before and after the threshold year.[18] Table 8 reports the coefficients for cash flow, considering either 1993 or 1995 as threshold dates for the sample split. The estimates are very close to the ones presented in Tables 6 and 7 for all specifications. The parameters are estimated less precisely, as expected, using the smaller instrument set, resulting in lower *t*-statistics. However, all the results for the role of cash flow in the two periods are robust to use of the smaller instrument set. This indicates that the results presented in Tables 6 and 7 are not affected by the smaller number of moment conditions available for the first period.

[18] Indeed, the first period still has one less instrument than the second since only one instrument (CF_{89}) is available for 1991.

Table 8. Investment equations: Instruments dated $t - 2$, $t - 3$

Regressors	Overall	State	Private	Foreign
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}$	0.002	−0.001	0.004	0.059
	(0.16)	(−0.07)	(0.21)	(1.47)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}$	0.055*	0.029	0.055*	0.029
	(6.39)	(1.61)	(4.70)	(1.85)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}95}$	0.012	0.001	0.023	0.060*
	(1.17)	(0.07)	(1.73)	(2.78)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}95}$	0.053*	0.022	0.060*	0.021
	(5.33)	(1.10)	(4.34)	(1.18)
$(CF/K)^{post\text{-}93} - (CF/K)^{pre\text{-}93}$	(4.03)	(1.38)	(2.29)	(−0.72)
$(CF/K)^{post\text{-}95} - (CF/K)^{pre\text{-}95}$	(3.10)	(0.90)	(1.99)	(−1.47)
Number of observations	16,536	5,893	7,167	3,476

***Note*:** Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t - 2)$ to $(t - 3)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. t-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 4,333.

A further robustness check is necessary in order to consider the possibility that, due to the high turnover of firms in the overall sample, the differences in the estimated cash flow coefficients before and after 1993 might actually reflect the different composition of the subsamples. We therefore estimate the investment equations on a balanced sample containing only firms that are present throughout the 1989–99 period. The results, presented in Table 9, indicate that the effect of financial liberalization on investment behaviour is not spurious: cash flow coefficients rise and become statistically significant in the post-reform period, both in the overall sample and in the ownership disaggregation, with the only exception of foreign-owned firms.[19]

One potential problem with testing the role of financial constraints using liquidity indicators such as cash flow is that these variables may be capturing the effect of other determinants, such as expectations about the profitability of investment projects, to the extent that they are not already captured by sales. The solution generally adopted in the literature relies on firms' cross-sectional heterogeneity,

[19] It should be observed, however, that in the balanced sample the number of observations for state and foreign firms is quite low (117 and 459, respectively), so that the results of the corresponding equations should be interpreted with care.

Table 9. Investment equations: Pre-post 1993 (balanced sample)

Regressors	Overall	State	Private	Foreign
$(I_{i,t-1}/K_{i,t-2})$	0.031	−0.232	−0.004	0.158*
	(1.24)	(−1.39)	(−0.01)	(2.62)
$(Y_{i,t}/K_{i,t-1})$	−0.002	−0.003*	−0.008*	−0.007
	(−0.04)	(−3.76)	(−4.65)	(−1.37)
$(Y_{i,t-1}/K_{i,t-2})$	0.002	−0.004	0.009*	0.022*
	(0.40)	(−0.25)	(2.76)	(2.79)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}$	0.053	0.029	0.062	0.012
	(1.63)	(0.90)	(1.86)	(0.10)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}93}$	0.112*	0.114*	0.086*	0.048
	(5.24)	(3.39)	(3.87)	(1.47)
$(CF/K)^{post\text{-}93} - (CF/K)^{pre\text{-}93}$	(2.15)	(3.87)	(1.23)	(1.07)
m_1 (1st order aut.)	0.000	0.098	0.002	0.001
m_2 (2nd order aut.)	0.162	0.188	0.423	0.757
Sargan test	0.643	0.942	0.451	1.000
z_1 (overall)	0.000	0.000	0.000	0.000
z_2 (time dummies)	0.001	0.032	0.001	0.110
z_3 (industry dummies)	0.000	0.000	0.000	0.000
No. of obs.	2,403	117	1,827	459

***Notes*:** Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t-2)$ to $(t-6)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. t-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 267.

exploiting the fact that the sensitivity of investment spending to changes in financial positions should be higher for firms believed to face significant agency costs.[20] Empirical studies of investment behaviour thus usually split the sample into groups according to a number of criteria considered *a priori* to identify financially constrained firms, including dividend policy, age, size, industrial group, bond rating, stock listing, and ownership structure.

[20] An alternative solution is to assume that investment opportunities are captured by the Q ratio (see, e.g., Blundell *et al*., 1992; Hayashi and Inoue, 1991; Schaller, 1990). However, apart from the practical consideration that the construction of Tobin's Q ratio is substantially more data demanding, it is difficult to determine the extent to which an average estimate of Q actually reflects expected profitability.

We follow a similar approach in order to control for the possibility that the different sensitivity of investment to cash flow before and after the reforms might be reflecting a change in the unobserved determinants of investment demand, such as expected profitability. We therefore present estimates obtained with a further disaggregation, within the pre- and post-reform subsamples, according to firm size and solvency, respectively. Similarly to the case of informational problems in market economies, size can be expected to matter for budget constraints in a transition economy. Large firms are more likely to face a soft budget constraint, as there are stronger social and political pressures to grant them external support in case of financial difficulties. Because of their influence, large firms are more likely to have access to external assistance by means of direct subsidies or bail-out credits.[21] As a consequence, an insignificant cash flow coefficient for large firms can be interpreted as an indication of soft budget constraints, rather than efficient capital markets (see, for example, Lízal and Svejnar, 2002, and Budina *et al.*, 2000). On the other hand, leverage, measured by the debt–asset ratio (non-equity assets over total assets), is not expected to be related to the tightness of the budget constraint.[22]

The results for the disaggregation by size, presented in Table 10, are quite interesting: in the post-reform period cash flow is positive and significant only for small firms. Looking at the results by ownership type, cash flow is significant for small private firms and (marginally) for small state and foreign firms. This indicates that the hardening of the budget constraint following financial market reforms only affected small private firms and, to a lesser extent, small state firms. Large firms, on the contrary, were largely unaffected irrespective of their ownership type.

The results for the disaggregation by solvency (based on the debt–asset ratio), presented in Table 11, indicate that in the post-reform period the sensitivity of investment to financial conditions is higher and significant for low-leverage firms. However, if we consider the equations by ownership, the significance of the estimated relationship is not related to leverage: in the post-reform period cash flow does not affect investment levels of state firms, whereas it does affect those of private and (marginally) foreign firms, irrespective of their debt levels. This result is consistent with the hypothesis that the budget constraints of small private and state-owned firms became more binding after the introduction of financial reforms, given that leverage, differently from size, is not expected *a priori* to be related to the tightness of the budget constraint.

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[21] 'There is a peculiar disparity in the treatment of large and small state-owned firms. [. . .] Large firms are much more successful in lobbying for favours, particularly for investment resources. Some of them are in great financial trouble; nevertheless large credits or subsidies are granted to them' (Kornai, 2000, p. 29).

[22] The sample split is obtained by allowing the cash flow coefficient to take different values in the two subsamples, by interacting cash flow with the appropriate dummy variable.

Table 10. Investment equations: Disaggregation by size

Regressors	Overall	State	Private	Foreign
$(I_{i,t-1}/K_{i,t-2})$	0.037*	−0.029	0.094*	0.157*
	(2.35)	(−1.07)	(5.32)	(7.03)
$(Y_{i,t}/K_{i,t-1})$	−0.010*	−0.008*	−0.005*	−0.012
	(−4.58)	(−3.05)	(−2.88)	(−1.92)
$(Y_{i,t-1}/K_{i,t-2})$	0.005*	0.004*	0.007*	0.014*
	(4.65)	(2.73)	(7.08)	(3.48)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}_{large}$	−0.105	−0.281	−0.063	0.125
	(−0.62)	(−1.64)	(−1.10)	(1.84)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}_{small}$	0.007	0.007	0.020	0.033
	(0.42)	(0.45)	(0.58)	(0.82)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}93}_{large}$	0.031	−0.016	0.029	−0.010
	(0.70)	(−0.38)	(0.75)	(−0.12)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}93}_{small}$	0.045*	0.069	0.059*	0.049
	(2.09)	(1.82)	(3.24)	(1.77)
$CF^{post}_{large} - CF^{pre}_{large}$	(0.81)	(1.52)	(1.47)	(−1.21)
$CF^{post}_{small} - CF^{pre}_{small}$	(1.56)	(1.55)	(1.06)	(0.39)
m_1 (1st order aut.)	0.000	0.000	0.000	0.000
m_2 (2nd order aut.)	0.010	0.249	0.681	0.798
Sargan test	0.016	0.847	0.431	0.997
z_1 (overall)	0.000	0.000	0.000	0.000
z_2 (time dummies)	0.000	0.000	0.000	0.006
z_3 (industry dummies)	0.002	0.009	0.021	0.007
No. of obs.	16,536	5,893	7,167	3,476

Notes: Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t-2)$ to $(t-6)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. t-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 4,333.

7. Concluding remarks

This paper presented an empirical investigation of the investment behaviour of a large panel of Hungarian manufacturing and construction firms in the period from 1989 to 1999. We examined the role of financial factors for corporate investment

Table 11. Investment equations: Disaggregation by leverage

Regressors	Overall	State	Private	Foreign
$(I_{i,t-1}/K_{i,t-2})$	0.040*	−0.028	0.089*	0.139*
	(2.52)	(−0.98)	(4.81)	(6.28)
$(Y_{i,t}/K_{i,t-1})$	−0.010*	−0.010*	−0.006*	−0.010
	(−4.95)	(−3.52)	(−3.81)	(−1.79)
$(Y_{i,t-1}/K_{i,t-2})$	0.005*	0.004*	0.007*	0.012*
	(4.41)	(2.13)	(7.35)	(3.49)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}_{low\text{-}debt}$	0.053	−0.106	0.033	0.033
	(0.60)	(−0.77)	(0.39)	(0.48)
$(CF_{i,t-1}/K_{i,t-2})^{pre\text{-}93}_{high\text{-}debt}$	−0.016	0.051	−0.001	0.093
	(−0.38)	(0.87)	(−0.02)	(1.12)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}93}_{low\text{-}debt}$	0.212*	0.150	0.113*	0.073*
	(2.99)	(1.36)	(2.35)	(2.58)
$(CF_{i,t-1}/K_{i,t-2})^{post\text{-}93}_{high\text{-}debt}$	0.030	0.031	0.047*	0.044
	(1.34)	(0.99)	(2.48)	(1.87)
$CF^{post}_{low\text{-}debt} - CF^{pre}_{low\text{-}debt}$	(1.36)	(1.47)	(0.87)	(0.59)
$CF^{post}_{high\text{-}debt} - CF^{pre}_{high\text{-}debt}$	(0.96)	(−0.29)	(1.11)	(−0.55)
m_1 (1st order aut.)	0.000	0.000	0.000	0.000
m_2 (2nd order aut.)	0.468	0.579	0.339	0.825
Sargan test	0.000	0.787	0.300	0.999
z_1 (overall)	0.000	0.000	0.000	0.000
z_2 (time dummies)	0.000	0.000	0.000	0.006
z_3 (industry dummies)	0.000	0.119	0.050	0.001
No. of obs.	16,536	5,893	7,167	3,476

***Notes*:** Dependent variable: $I_{i,t}/K_{i,t-1}$. GMM one-step estimates in first differences.
Instruments: $(t-2)$ to $(t-6)$ lags of I/K, Y/K, CF/K.
Year and industry dummies included in all equations. t-statistics in round brackets (heteroskedasticity consistent standard errors). Asterisks indicate significance at the 5 percent level.
The bottom part of the table reports P-values for the corresponding test statistics.
Sample period: 1991 to 1999. Overall number of firms: 4,333.

decisions by ownership type before and after the introduction of major financial reforms, and explored differences across subsamples of firms defined according to size and leverage.

Our results indicate that financial reforms have significantly affected the investment behaviour of Hungarian firms. The effects, however, were different depending on firms' ownership. Both state-owned and domestic private firms faced a soft budget constraint before 1993. In the post-reform period, however, while private firms came to face binding financial constraints, state-owned firms remained subject to a soft budget constraint, although their investment decisions became more sensitive to financial conditions. The response of foreign-owned firms was quite different: they were subject to a hard budget constraint in both periods, but became less sensitive to financial conditions, possibly indicating that reforms have been successful in lowering informational costs. These results were found to be robust to a number of consistency checks.

Splitting the sample further by size and leverage, we found that in the post-1993 period budget constraints have become binding for small private firms and, to a lesser extent, small state firms. Large firms, on the contrary, continued to face a soft budget constraint irrespective of their ownership type. The fact that the post-1993 relationship between financial conditions and investment for Hungarian domestic firms depends on size but not on leverage can be taken as a further indication that financial reforms displayed their effects through the hardening of the budget constraint for small private and state firms.

Overall, our results extend and qualify those obtained by previous studies for Hungary and other transition economies. On the one hand, the persistent absence of liquidity constraints suggests that, despite the introduction of major financial reforms, large state-owned firms operated under a soft budget constraint throughout the 1990s. This result is consistent with the findings of Lízal and Svejnar (2002) for the Czech Republic and Budina *et al.* (2000) for Bulgaria.

On the other hand, financial reforms seem to have significantly improved the efficiency of credit allocation to the private sector in at least two respects: budget constraints became binding for private domestic firms, particularly small ones, and informational costs became less relevant for foreign-owned firms. This finding, based on a sample of firms representing about 80 percent of total Hungarian manufacturing employment and value added, confirms and extends the results obtained by Perotti and Vesnaver (2004) for the period from 1992 to 1998.[23] Future research will determine whether the further transformation of the financial sector will allow the imposition of hard budget constraints on Hungarian firms irrespective of their ownership type.

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Appendix 1

The dataset

The dataset used in this paper contains company account data for a large cross-section of Hungarian firms observed from 1989 to 1999. The dataset is based on two sources: a dataset collected by the Hungarian Ministry of Finance that contains information on all firms that paid corporate or profit taxes from 1989 to 1996, covering the majority of Hungarian firms; and a second dataset from the Hungarian Central Statistical Office that contains end of year financial statements of medium and large firms, from 1992 to 1999. Merging the information from the two sources we obtained firm-level annual time series between 1989 and 1999 for balance sheet and income statement variables, plus information on ownership, employment, export, regional location and industry identification at the four-digit level.

Firms are identified by their identification numbers. It should be observed that when a firm is split, due to restructuring or privatization, a branch or a part of it normally keeps the same identification number of the original firm, while a different identification number is assigned to the other parts or branches. Although the original firm and the branch that keeps the same identification number are *de facto* different firms, in the sample they are recorded as the same firm.

The dataset can be considered highly representative of the overall Hungarian economy. Table A1 provides some information on the sample coverage by reporting total employment and value added in the sample as a percentage of the whole economy. The firms contained in the sample account for 78 percent and over 80 percent of total employment and value added respectively in the manufacturing sector. In other sectors such as agriculture and services the degree of representativeness is lower, reflecting the higher number of small firms.

Table A1. Sample coverage

Sector	Employment	Value added
Agriculture	37.2	27.3
Mining	41.0	82.7
Manufacturing	78.0	81.2
Electricity, Gas, Water	93.8	97.0
Construction	36.1	34.3
Trade, Tourism	35.8	40.3
Transport	72.1	58.9
Finance	29.8	9.4
Public Administration	4.3	4.7

Note: Total employment and value added in the sample as a percentage of the Hungarian economy (1995).

Table A2. Sample representativeness by size

Size class	Dataset	Whole economy
0–10	13.6	75.5
11–20	5.0	12.0
21–50	21.2	7.0
51–300	49.4	4.5
300–	10.8	1.0

Note: Distribution of employment by employment size class (1995).

Table A3. Sample representativeness by sector

Sector	Dataset	Whole economy
Agriculture	7.6	8.0
Mining	1.0	0.9
Manufacturing	43.0	23.1
Electricity, Gas, Water	6.3	2.6
Construction	5.4	5.9
Trade, Tourism	14.2	15.6
Transport	15.9	8.7
Finance	4.4	6.0
Public Administration	2.2	29.2

Note: Distribution of employment by sector (1995).

If we consider size representativeness, medium and large firms are over-represented in the dataset compared to the overall economy. As shown in Table A2, firms in the smallest size class (0–10 employees) account for over two thirds of the total number of firms in the Hungarian economy, whereas in our dataset they account for only 13.6 percent of the total. Looking at the distribution by sector, Table A3 shows that the manufacturing sector is over-represented (43 percent, against 23.1 percent in the whole economy).

Appendix 2

The accelerator model

In the accelerator model, investment is determined by setting the marginal product of capital equal to marginal cost. For a given technology, the optimal level of the capital stock can be obtained, and investment fills the gap between the optimal and current capital stock. Under a number of simplifying assumptions, the demand for capital can be expressed as a function of the level of output and the user cost of capital. Formally, the model can be derived from firms' maximization of profit (see Cho, 1996):

$$V_t = \sum_{i=0}^{\infty} \beta_{t+i} \left[p_{t+i} F(K_{t+i}, L_{t+i}) - w_{t+i} L_{t+i} - p^I_{t+i} I_{t+i} \right] \quad \text{(A1)}$$

subject to

$$K_{t+i} = (1-\delta) K_{t+i-1} + I_{t+i} \quad \text{(A2)}$$

where β is the discount rate, p the price of output, K the capital stock, L the labour input, w the wage rate, I is gross investment, p^I the price of investment goods, and δ the rate of depreciation. The first-order condition for the capital stock implies that

$$\left(\frac{\partial F}{\partial K} \right)_t \cong \frac{p^I_t}{p_t} (r_{t+1} + \delta - \pi^I_{t+1}) = J_t \quad \text{(A3)}$$

where J is the *user cost* of capital, r the nominal rate of return and π^I the inflation rate for investment goods. Intuitively, at the optimum capital stock, the marginal product of capital equals the cost of using an additional unit of capital.

If the production function has constant elasticity of substitution (*CES*):

$$Y_t = [\alpha K^\rho + (1-\alpha) L^\rho]^{\frac{1}{\rho}} \quad \text{(A4)}$$

then

$$F_K = \alpha Y^{(1-\rho)} K^{(\rho-1)} \tag{A5}$$

so that the first-order condition implies

$$K_t^* = \alpha^{\frac{1}{1-\rho}} Y_t J_t^{-\sigma} = A Y_t J_t^{-\sigma} \tag{A6}$$

where σ is the elasticity of substitution between capital and labour. This gives the desired *stock* of capital as a function of sales and the user cost of capital. Investment decisions should be aimed at achieving this optimal level of capital.

If the production function is Cobb-Douglas ($\sigma = 1$), we obtain

$$K_t^* = \alpha \left(\frac{Y_t}{J_t} \right). \tag{A7}$$

Under the further assumption that there is no substitution between capital and labour ($\sigma = 0$), or that J_t is constant,[24] then

$$K_t^* = \alpha Y_t \tag{A8}$$

Investment is then given by

$$I_t = \alpha Y_t - (1 - \delta) K_{t-1} \tag{A9}$$

and dividing by K_{t-1} we obtain an accelerator investment model, where investment is not affected by the user cost of capital:

$$\left(\frac{I_t}{K_{t-1}} \right) = \beta_0 + \beta_1 \left(\frac{Y_t}{K_{t-1}} \right). \tag{A10}$$

This equilibrium relationship can be modified to account for gradual adjustment of the actual capital stock to the desired capital stock (changes in output):

$$\left(\frac{I_t}{K_{t-1}} \right) = \beta_0 + \beta_1 \left(\frac{I_{t-1}}{K_{t-2}} \right) + \beta_2 \left(\frac{Y_t}{K_{t-1}} \right) + \beta_3 \left(\frac{Y_{t-1}}{K_{t-2}} \right) \tag{A11}$$

(see, for example, Fazzari *et al.*, 1988). To test for the presence of financial constraints this basic specification can be augmented with lagged cash flow (as a ratio of the capital stock).

[24] Alternatively, it can be assumed that the variation in the user cost of capital is captured by time-specific or firm-specific effects in the error term.