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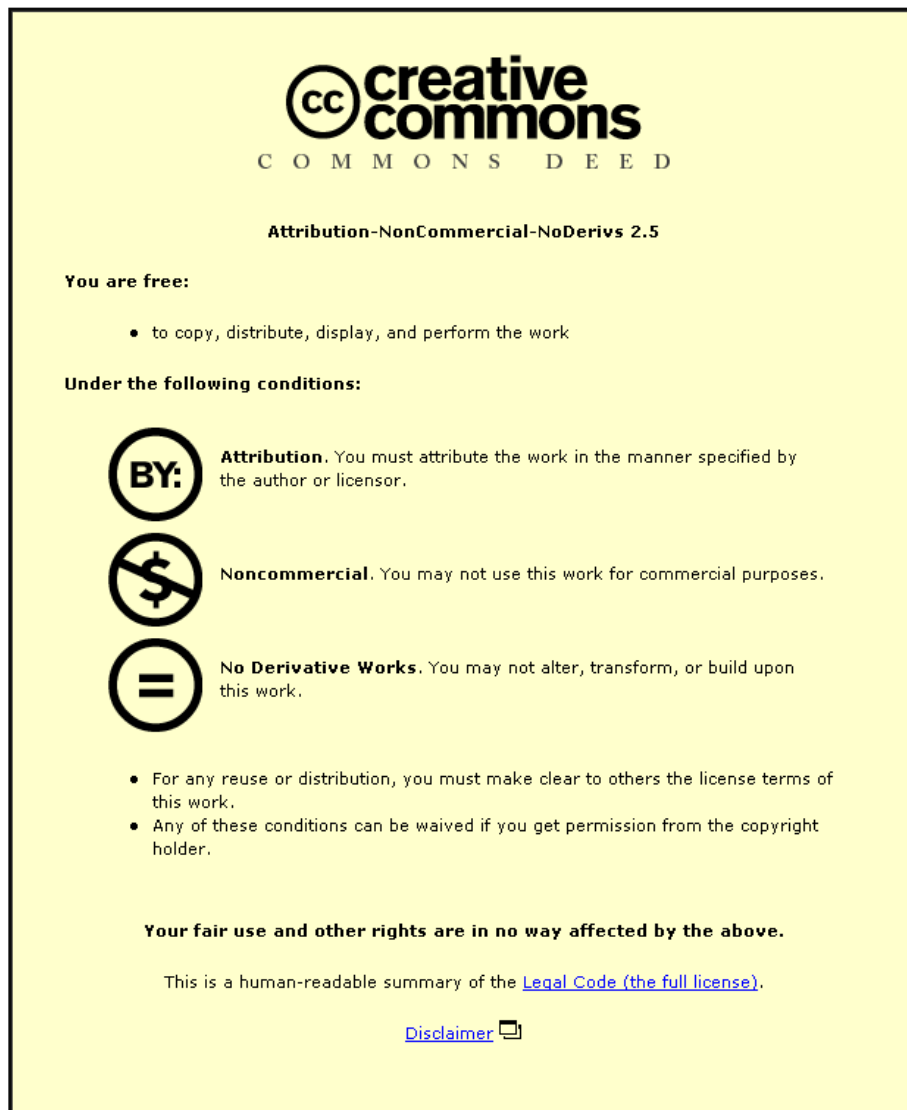
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# Financial Restraints and Private Investment: Evidence from a Nonstationary Panel

Mauro Costantini<sup>a</sup>, Panicos Demetriades<sup>b,\*</sup>, Gregory James<sup>c</sup> and Kevin Lee<sup>d</sup>

<sup>a</sup>Department of Economics and Finance, Brunel University, UK

<sup>b</sup>Department of Economics, University of Leicester, UK

<sup>c</sup>School of Business and Economics, Loughborough University, UK

<sup>d</sup>School of Economics, University of Nottingham, UK

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## Abstract

We employ recently developed panel data methods to estimate a model of private investment under financial restraints for 20 developing countries using annual data for 1972-2000. We show that the qualitative nature of the results varies depending on whether we take into account cross-country effects. When we allow for cross-sectional dependence, investment displays more sensitivity to world capital market conditions and exchange rate uncertainty. A perhaps even more surprising result is the finding that countries that managed to suppress domestic real interest rates without generating high inflation enjoyed higher levels of private investment than those that would have been obtained under liberalised conditions.

*Keywords:* financial restraints, private investment, nonstationary panels, cross-sectional dependence

*JEL Classification:* O16, G18, G28

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\*Corresponding author: Panicos Demetriades. Address: Department of Economics, University of Leicester, LE1 7RH, UK. Tel.: +44 (0) 116 252 2835. Fax: +44 (0) 116 252 2908. Email: pd28@le.ac.uk.

# 1 Introduction

In the early 1970s, McKinnon (1973) and Shaw (1973) put forward the idea that financial repression – i.e. government imposed controls on lending and deposit rates, capital controls, and directed credit - had a negative impact on investment and growth by suppressing domestic saving and distorting the allocation of credit. While their views were vigorously challenged by a range of critics,<sup>1</sup> their main policy recommendation for financial liberalisation gained momentum among policy makers in both developing and developed countries. As a result, the last forty years have witnessed a gradual removal of financial restraints worldwide with increased movement of capital around the globe.<sup>2</sup>

Both these developments are likely to influence the behaviour of private investment. Increased international capital flows are likely to result in a relaxation of borrowing constraints for many firms, leading to credit expansion.<sup>3</sup> Under fully liberalised conditions the price of credit for many, if not all, firms will rise, making their investment plans more sensitive to the price of credit and no longer sensitive to the availability of credit. Under partial liberalisation or continued financial repression, however, some firms may continue to have access to subsidised credit while others may have access to more expensive international loans. Does the retention of financial restraints under these circumstances deter or promote investment? In other words, once a country moves away from complete financial repression - where the only source of credit for private investment is the domestic banking system - can the provision of cheaper, albeit rationed, domestic credit help stimulate private investment? This is the question we address in this paper. In order to do so, we employ a theoretical model of investment which assumes that firms have access to quantity-constrained domestic loans that are cheaper than those they can obtain from international capital markets.<sup>4</sup> This accommodates the idea that increased international capital flows might

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<sup>1</sup>See for example, Arestis and Demetriades (1999), Diaz-Alejandro (1985), Hellman *et al.* (2000), Singh (1997), Stiglitz (1994), Taylor (1983), Van Wijnbergen (1983).

<sup>2</sup>Abiad and Mody (2005) document the gradual reduction of financial restraints around the world while Lane and Milesi-Ferretti (2005) document the increase in financial openness.

<sup>3</sup>In some circumstances, such credit expansion can also feed consumption and lead to asset price bubbles (see, for example, Gylfason *et al.* 2010).

<sup>4</sup>The model is based on Demetriades and Devereux (2000).

have relaxed borrowing constraints for many firms while, at the same time, some firms may have continued to benefit from access to cheaper policy loans. We operationalise the model in a multi-country setting and derive five variants of a private investment equation including a baseline neoclassical model without financial restraints. To estimate the investment equations, we employ recently developed nonstationary panel methodologies that allow for cross-sectional dependence across countries. The presence of dependence across countries is a plausible hypothesis in a world characterised by growing real and financial inter-linkages, which we test by appropriate econometric procedures.

Our sample includes 20 developing countries over the period 1972-2000. The econometric analysis consists of three steps. First, unit root tests for cross-sectionally dependent panels are applied. Second, the existence of a cointegrating relationship among the variables is investigated, fully allowing for cross-section dependence. Third, the Fully Modified Ordinary Least Square (FMOLS) estimator developed by Bai and Kao (2006) is used to estimate the investment equations. We contrast our results with those obtained using the pooled FMOLS estimator of Pedroni (2000) which assumes cross-sectional independence.

Our findings confirm the importance of taking into account cross-country dependence. We find that when we allow for cross-sectional dependence, investment displays more sensitivity to world capital market conditions and exchange rate uncertainty. Perhaps more surprisingly, we find that ‘repressing’ domestic real interest rates resulted in higher levels of private investment than those that would have been obtained under more ‘liberalised’ conditions. This finding, which contrasts sharply with the McKinnon-Shaw prediction, complements a growing literature on the possible negative effects of financial liberalisation on the channels of economic growth. Stiglitz (1994) provides a unifying theoretical rationale for such effects, drawing on information asymmetries in financial markets which provide scope for meaningful government interventions. Singh (1997), drawing on Keynes (1936) and a large body of empirical evidence, emphasises the negative effects that emanate from stock market volatility. Demetriades and Luintel (2001) provide evidence of positive effects of financial re-

straints on South Korea's financial development, reflecting lack of competition in the banking system. More recently, Andrianova *et al.* (2008 and 2010) provide evidence suggesting that bank privatisation - one of the main pillars of financial liberalisation – has been negatively associated with both financial development and growth, reflecting poor regulation. Last but not least, recent work by Ang (2010) suggests that financial liberalisation had a negative effect on technological deepening by distorting the allocation of human capital.

The paper is organised as follows. Section 2 describes the modelling framework. Section 3 discusses econometric methodology and empirical results. Section 4 summarises and concludes.

## 2 The modelling framework

### 2.1 Theoretical underpinnings

The dynamic investment equations estimated in this paper are based on the theoretical model put forward by Demetriades and Devereux (2000), henceforth D&D. D&D use a microeconomic model of a representative firm's investment decision under financial restraints as their starting point. The model suggests a structural relationship between the optimal capital stock and the 'modified' cost of capital which is then used to derive a long-run theory-consistent aggregate investment equation that takes into account the presence of financial restraints. The rest of this section provides a brief outline of the D&D approach.

The main assumption of D&D is that the official banking system is unable to satisfy the entire demand for investible funds because of the presence of an interest rate ceiling which restricts the supply of funds à la McKinnon-Shaw (see also Fry, 1994). The model departs from the McKinnon-Shaw tradition, however, in that it assumes the existence of an 'alternative' financial market in which firms can borrow freely, albeit at an interest rate that is higher than the official lending rate. Their interpretation of the alternative market is that it is the world capital market although it could also be interpreted as the unofficial credit market, or 'curb', market (see Taylor, 1983 and Van

Wijnebergen, 1983). There are theoretical and empirical reasons for us preferring the first interpretation to the second, not least the stylised facts relating to the increased international capital flows alluded to in the introduction. Thus, we assume that firms have access to two types of borrowing: domestic bank borrowing and international loans. Rationing of domestic loans to different firms is assumed to depend on the availability of collateral which is related to the firm's capital stock.

The representative firm is assumed to maximize the wealth of its shareholders given by the present discounted value of dividends ( $D_t$ ). The nominal discount rate used in determining the present value is the one which is obtained in the world capital market, denoted  $i_t^*$ , since this is the rate at which shareholders are assumed to be able to borrow or lend as much as they wish.<sup>5</sup> Note that the firm takes both the domestic lending rate  $i_t$  and the world interest rate  $i_t^*$  as determined exogenously in the appropriate market. Moreover, the firm is assumed to be able to raise finance only through borrowing or retained earnings.

Formally, the optimisation problem can be stated as:

$$\text{Max}_{I_t} E_t \left\{ \sum_{s=t+1}^{\infty} \beta_s D_s \right\}, \quad (1)$$

where  $\beta_s = \prod_{l=t+1}^s (1 + i_{l-1}^*)^{-1}$ , subject to the following constraints:

$$D_t = q_t Y_t - p_t I_t + B_t - (1 + i_t) B_{t-1} + A_t - (1 + i_t^*) A_{t-1}, \quad (2)$$

$$K_t = (1 - \delta) K_{t-1} + I_t, \quad (3)$$

$$B_t \leq x_t p_t K_t, \quad (4)$$

where  $E_t\{\cdot\}$  is the expectations operator,  $q_t Y_t$  represents current revenue, where  $q_t$  is the price of output in period  $t$  and  $Y_t$  is output, and where the latter is a function of

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<sup>5</sup>The model assumes that there are two groups of investors in the country: sophisticated investors, who can lend and borrow in the world capital market and who own shares, and unsophisticated investors, who save only in the official banking sector.

the capital stock at the beginning of the period,  $Y_t = f(K_{t-1})$ .<sup>6</sup> The value of current investment is represented by  $p_t I_t$ , where  $p_t$  is the current price of capital goods and  $I_t$  is the quantity of investment made during period  $t$ . New issues of one period debt from the domestic and international market are denoted  $B_t - B_{t-1}$  and  $A_t - A_{t-1}$  respectively, while  $i_t B_{t-1}$  and  $i_t^* A_{t-1}$  are nominal interest payments to the domestic and international capital market, respectively.<sup>7</sup> The exponential rate of depreciation of capital is assumed constant at  $\delta$ .

The first two constraints are standard in models of firm investment. The first constraint is the flow of funds identity for the firm and the second constraint is the equation of motion of the capital stock. The third constraint is specific to D&D; it constrains the supply of domestic bank loans in the domestic market to be a proportion,  $x_t$ , of the value of the firm's capital stock. The capital stock, therefore, represents collateral; banks are willing to lend more to large firms than to small firms.<sup>8</sup>

Taking first-order conditions together yields

$$E_t[q_{t+1}f'(K_t)] = i_t^* p_t + \delta E_t p_{t+1} - (E_t p_{t+1} - p_t) - \frac{p_t(i_t^* - i_t)}{(1 + i_t^*)} x_t. \quad (5)$$

This states that, in equilibrium, the expected marginal revenue product of capital is equal to a modified cost of capital. The modified cost of capital consists of: the financial cost at the rate in the international market  $i_t^* p_t$ ; plus the cost of the fall in the value of the asset  $\delta E_t p_{t+1}$ ; minus the expected capital gain term,  $E_t p_{t+1} - p_t$ ; plus the final term which reflects the reduction in the standard cost of capital relative to the international capital market. This final term shows the cheaper source of finance which is available at rate  $i_t$  but acknowledges that only a proportion  $x_{jt}$  can be financed in this way.

Equation (5) holds for every firm in the economy in the steady-state. D&D show that the same relationship will be observed in the economy as a whole providing that

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<sup>6</sup>Stocks dated  $t$  refer to the end of period  $t$ , equivalent to the beginning of period  $t + 1$ .

<sup>7</sup>In both markets, the model assumes that the nominal interest rate is set at the time the borrowing takes place. Thus, for example, the interest rate applying to official borrowing at the beginning of period  $t$  (the end of period  $t - 1$ , denoted  $B_{t-1}$ ) is determined at the beginning of the period and hence denoted  $i_{t-1}$ .

<sup>8</sup>Note that firms cannot borrow from the domestic market to lend on the international market.

certain aggregation conditions are satisfied and that firm-specific shocks to the proportion of a firm's capital stock financed out of bank loans cancel out across firms. The steady-state relationship can be embedded in a dynamic model that explains aggregate behaviour by assuming that investment is driven by the difference between the actual marginal product of capital and its equilibrium level based on (5). Additional dynamics would be generated by time lags in decision-making, ordering, delivery, installation of new capital, and so on. The dynamic investment equation corresponding to (5) is then given by

$$\frac{I_{jt}}{K_{t-1}} = b_0 + b_1 \frac{I_{t-1}}{K_{t-2}} + b_2 \frac{Y_t}{K_{t-1}} + b_{j3} \left[ \frac{1 + i_t^*}{1 + \pi_t^e} - 1 \right] + b_4 \frac{(i_t^* - i_t)}{(1 + i_t^*)(1 + \pi_t^e)} \frac{B_t}{K_{t-1}}. \quad (6)$$

where the subscript  $j$  refers to firm  $j$ . The term  $Y_t/K_{t-1}$  is interpreted as a proxy for the marginal product of capital and the modified cost of capital is split into two components: the real interest rate in the world capital market and the term capturing financial restraints.

Since we expect investment to depend on the difference between the marginal product and the modified cost of capital, the theoretical model predicts that  $b_2$  should be positive and  $b_3$  negative. The fourth term is present only under financial restraints. A positive  $b_4$  would provide support for the hypothesis that the existence of an alternative market for credit outweighs the credit rationing effect described by McKinnon–Shaw. In such a case, increasing the level of the interest rate ceiling in the domestic market would serve to increase the overall cost of capital (which corresponds to Figure 1 in D&D). On the other hand, a negative  $b_4$  would suggest that the existence of the alternative market is not sufficient to outweigh the McKinnon–Shaw effect, i.e. higher domestic interest will have a positive effect on investment on balance. In this case, the supply of domestic financial savings is elastic with respect to the domestic interest rate so that an increase in the domestic interest rate has a relatively large effect on the domestic supply of investable funds (this corresponds to Figure 2 in D&D).

## 2.2 Operationalising the model in a multi-country analysis

There are three variables in equation (6) that are not directly observed and require modelling assumptions to be made to operationalise the model in a multi-country empirical analysis, the capital stock, the world capital market interest rate and the financial restraints dummy. The construction of the first is based on the perpetual inventory method given by expression (3).<sup>9</sup> The interest rate  $i^*$  used here is the US lending rate. Given the sample of countries we are using, we believe that the US rate is the most appropriate rate to approximate the cost of loans from the world market. The expected inflation series are in turn proxied by the current inflation rate prevailing in each country. The financial restraints dummy is based on nominal interest rate differential  $i^* - i$ . In the theoretical model, the supply of bank loans becomes rationed only if  $i^*$  exceeds  $i$ . This suggests that an observation could be considered as being under conditions of ‘financial restraints’ if  $i^* - i > 0$ . Five variants of Equation (6) are estimated to allow some flexibility in the way that financial restraints are defined and to capture the possible effects of exchange rate risk.<sup>10</sup>

The first model is a "Neo-Classical" investment equation –denoted  $NC$ – which corresponds to a world without financial restraints ( $b_{j4} = 0$ ):

$$\frac{I_{jt}}{K_{jt-1}} = b_{j0} + b_{j1} \frac{I_{jt-1}}{K_{jt-2}} + b_{j2} \frac{Y_{jt}}{K_{jt-1}} + b_{j3} r_t^* + \varepsilon_{jt}, \quad (7a)$$

where the subscript  $j$  refers to country  $j$  and the error term is  $IID(0, \sigma_j^2)$  across time but may be correlated across countries as a result of common real or financial shocks.

The second model –denoted  $FR^A$ – tests the financial restraints hypothesis assuming that all the countries always operate under conditions of financial restraints:

$$\frac{I_{jt}}{K_{jt-1}} = b_{j0} + b_{j1} \frac{I_{jt-1}}{K_{jt-2}} + b_{j2} \frac{Y_{jt}}{K_{jt-1}} + b_{j3} r_t^* + b_{j4} \frac{(i_t^* - i_t)}{(1 + i_t^*)(1 + \pi_t)} \frac{B_{jt}}{K_{jt-1}} + \varepsilon_{jt}. \quad (7b)$$

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<sup>9</sup>The initial capital stock for each country was constructed by using  $K_0 = ((\sum_{t=1970}^{1974} I_t)/5)/\delta$ , where  $\delta$  is the depreciation rate, assumed to be 4%.

<sup>10</sup>Although for tractability reasons, exchange rate risk is not explicitly taken into account in the underlying theoretical model, in reality this may deter domestic firms from borrowing in international markets.

The third model –denoted  $FR^D$ – also accommodates the possible effect of financial restraint but the financial restraints term is now interacted with  $D_{jt}$ , a dummy variable that equals 1 when an observation is considered as being under condition of financial restraints (as defined above) and 0 otherwise:

$$\frac{I_{jt}}{K_{jt-1}} = b_{j0} + b_{j1} \frac{I_{jt-1}}{K_{jt-2}} + b_{j2} \frac{Y_{jt}}{K_{jt-1}} + b_{j3} r_t^* + \tilde{b}_{j4} D_{jt} \frac{(i_t^* - i_t)}{(1 + i_t^*)(1 + \pi_t)} \frac{B_{jt}}{K_{jt-1}} + \varepsilon_{jt}. \quad (7c)$$

The fourth model –denoted  $FR^A(\text{unrestricted})$  "unbundles" the financial restraints term into its two components, the real interest rate differential and the inflation rate differential:

$$\frac{I_{jt}}{K_{jt-1}} = b_{j0} + b_{j1} \frac{I_{jt-1}}{K_{jt-2}} + b_{j2} \frac{Y_{jt}}{K_{jt-1}} + b_{j3} r_t^* + \acute{b}_{j4} \frac{(r_t^* - r_{jt})}{(1 + i_t^*)(1 + \pi_t)} \frac{B_{jt}}{K_{jt-1}} + b_{j5} \frac{(\pi_t^* - \pi_{jt})}{(1 + i_t^*)(1 + \pi_t)} \frac{B_{jt}}{K_{jt-1}} + \varepsilon_{jt}. \quad (7d)$$

The fifth model –denoted  $FR-ER$ – introduces a measure of exchange rate uncertainty to capture the risk associated with international borrowing by domestic firms, which may have a negative effect on investment:

$$\begin{aligned} \frac{I_{jt}}{K_{jt-1}} = & b_{j0} + b_{j1} \frac{I_{jt-1}}{K_{jt-2}} + b_{j2} \frac{Y_{jt}}{K_{jt-1}} + b_{j3} r_t^* + b_{j4} \frac{(i_t^* - i_t)}{(1 + i_t^*)(1 + \pi_t)} \frac{B_{jt}}{K_{jt-1}} \\ & + \tilde{b}_{j5} SDEX_{jt} \frac{B_{jt}}{K_{jt-1}} + b_{6jt} SDEX_{jt} + \varepsilon_{jt}, \end{aligned} \quad (7e)$$

where  $SDEX_{jt}$  is the 3-year moving average of the standard deviation of the domestic exchange rate vis-à-vis the US dollar.

### 3 Econometric methodology and empirical results

The empirical analysis consists of three steps. In the first step, we test for non-stationarity in the data using the testing procedures developed by Bai and Ng (2004), labelled by them as PANIC. The basic idea consists of modelling the panel series as

the sum of a set of common factors and idiosyncratic components. Both the factors and the idiosyncratic components can be  $I(1)$  or stationary, so that dependence can be modelled not only through the disturbance terms but also through the common factors. Bai and Ng propose to test the factors and the idiosyncratic components separately. This feature makes it possible to ascertain if nonstationarity comes from a pervasive or an idiosyncratic source.<sup>11</sup> In the second step, we investigate the existence of a cointegrating relationship for all the models. To this end, the panel procedure recently developed by Gengenbach *et al.* (2006) is applied.

1. A preliminary PANIC analysis on each variable  $X_{i,t}$  and  $Y_{i,t}$  to extract common factors is conducted. Tests for unit roots are performed on both the common factors and the idiosyncratic components using the Bai and Ng (2004) procedure.

2. a. If  $I(1)$  common factors and  $I(0)$  idiosyncratic components are detected, then a situation of cross-member cointegration is found and consequently the nonstationarity in the panel is entirely due to a reduced number of common stochastic trends. Cointegration between  $Y_{i,t}$  and  $X_{i,t}$  can only occur if the common factors for  $Y_{i,t}$  cointegrate with those of  $X_{i,t}$ . The null of no cointegration between the estimated factors can be tested using the Johansen (1988) trace test as suggested by Gengenbach *et al.* (2006).

2. b. If  $I(1)$  common factors and  $I(1)$  idiosyncratic components are detected, then defactored series are used. In particular,  $Y_{i,t}$  and  $X_{i,t}$  are defactored separately. Testing for no cointegration between the defactored data can be conducted using standard panel tests for no cointegration such as those of Pedroni (1999) and Pedroni (2004). Cointegration between  $Y_{i,t}$  and  $X_{i,t}$  is found only when the tests for both the common factors and the idiosyncratic components reject the null of no cointegration.<sup>12</sup>

In the third step, we estimate the long-relationship among the variables of interest in the five models under consideration using the continuous-update fully modified (CUP-FM) estimator developed Bai and Kao (2006). These authors discuss the limiting distributions of various panel OLS and FM estimators and argue for the use of

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<sup>11</sup>Other testing procedures based on factor structure generally test the unit root only in the defactored data. See for instance Moon and Perron (2004).

<sup>12</sup>The framework used by Gengenbach *et al.* (2006) leads to panel statistics for the null of no cointegration that have the same distribution as panel unit root tests and hence are not affected by the number of regressors.

CUP-FM estimators.<sup>13</sup>

Our panel dataset contains 20 countries over the period 1972-2000. The countries were chosen because of data availability. A detailed description of the countries involved, measurement of variables and data sources is given in the Appendix. To support the cross-sectional dependence hypothesis, the CD test developed by Pesaran (2004) is applied to our data. The test proposed is:

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left( \sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{ij} \right), \quad (8)$$

where

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^T e_{it}e_{jt}}{(\sum_{t=1}^T e_{it}^2)^{1/2}(\sum_{t=1}^T e_{jt}^2)^{1/2}}$$

denote the sample estimate of the pair-wise correlation of the residuals  $e_{it}$  from the regression of any variable of interest on an intercept, a linear trend and a lagged dependent variable for each country  $i$ . CD test results are reported in Table 1. Clear evidence of cross-sectional dependence is found since the null hypothesis of no cross-correlation is strongly rejected.

Table 2 reports the results of the Bai and Ng (2004), Pesaran (2007) and Moon and Perron (2004) panel unit root tests. The CIPS test of Pesaran (2007) and the  $t_a^*$  and  $t_b^*$  tests of Moon and Perron (2004) are included in the analysis because they have greater power than the Bai and Ng test in small samples.<sup>14</sup> In applying the Bai and Ng procedure to test for unit roots, we consider the common factors and the idiosyncratic components separately. The number of common factors is determined using the IC2 criterion developed by Bai and Ng (2002) and one common factor is selected.<sup>15</sup> Where there is only one common factor, Bai and Ng (2004) suggest using a standard Augmented Dickey-Fuller (ADF) test to test stationarity<sup>16</sup>:

<sup>13</sup>For more details see Bai and Kao (2006).

<sup>14</sup>See Appendix C for the results of the Monte Carlo simulations that confirm this. For further details on the CIPS test and the  $t_a^*$  and  $t_b^*$  statistics see Pesaran (2007) and Moon and Perron (2004) respectively.

<sup>15</sup>One common factor is also selected when the other information criteria proposed by Bai and Ng (2002) are considered.

<sup>16</sup>See Bai and Ng (2004, p.1133).

$$\Delta \hat{F}_t = c + \delta_0 \hat{F}_{t-1} + \delta_1 \Delta F_{t-1} + \cdots + \delta_p \hat{F}_{t-p} + v_t \quad (9)$$

where  $F_t$  indicates an  $r \times 1$  vector of common factors. The ADF tests results for the extracted common factor provide evidence of a unit root in all the variables. To test the stationarity of the idiosyncratic component, Bai and Ng (2004) propose pooling individual ADF  $t$ -statistics with de-factored estimated components  $e_{it}$  in the model with no deterministic trend

$$\Delta \hat{e}_{i,t} = d_{i,0} \hat{e}_{i,t-1} + d_{i,1} \Delta \hat{e}_{i,t-1} + \cdots + d_{i,p} \hat{e}_{i,t-p} + \mu_{i,t}. \quad (10)$$

The pooled tests are based on Fisher-type statistics defined as in Maddala and Wu (1999) and in Choi (2001). Let  $P_{\hat{e}}^c(i)$  be the  $p$ -value of the ADF  $t$ -statistics for the  $i$ -th cross-section unit,  $ADF_{\hat{e}}^c(i)$ , then the standardised Choi-type statistics is:

$$Z_{\hat{e}}^c = \frac{-2 \sum_{i=1}^n \log P_{\hat{e}}^c(i) - 2N}{\sqrt{4N}} \quad (11)$$

The previous statistic converges for  $(N, T \rightarrow \infty)$  to a standard normal distribution. In our analysis, we use the Fisher-type statistic defined as in Choi (2001). The pooled  $p$ -value inverse normal tests do not reject the null hypothesis of a unit root for all the variables, providing strong evidence of nonstationarity. Similarly, the results obtained with the tests developed by Pesaran (2007) and Moon and Perron (2004), which are more powerful in small samples, show that the null hypothesis of a unit root cannot be rejected for all the variables.

Since the panel no cointegration hypothesis can be rejected only if the tests for both the common factors and the idiosyncratic components reject the null of no cointegration (see Gengenbach *et al.*, 2006, pp. 698-99), we apply the panel cointegration tests proposed by Pedroni (1999, 2004) to the defactored data and the Johansen (1988) trace test to the common factor components. The results are reported in Table 3. For the panel tests, we use two statistics proposed by Pedroni (1999, 2004). The first statistic is a panel version of a non-parametric statistic that is analogous to the familiar

Phillips and Perron rho-statistic,  $Z_\rho$ . The second is a parametric statistic which is analogous to the familiar ADF  $t$ -statistic,  $Z_t$ . These tests assume the null hypothesis of no cointegration against the alternative that all units (countries) share a common cointegrating vector. The results of these tests provide evidence of a common cointegrating vector for the whole panel. With regard to the common factor components, a cointegrating relationship is found with the Johansen (1998) trace test in all the models.

Having found evidence of cointegration in each of the models, we first estimate these models using the FMOLS estimator proposed by Pedroni (2004) under the assumption of cross-sectional independence. Table 4 reports the estimation results. The term proxying the marginal product of capital ( $b_2$ ) is always positive and strongly significant, as predicted by the theory. The coefficient of the world interest rate ( $b_3$ ) is negative and significant in all the models, which is consistent with the interpretation that the world interest rate captures an important component of the cost of capital, irrespective of the extent to which the models incorporate financial restraints. The coefficients on the various financial restraints terms where they appear are positive but rarely significant. In Model  $FR^A$ , which contains the unbundled financial restraints term that is not interacted with the dummy variable,  $b_4$  is positive and highly insignificant. In Model  $FR^D$ , where the unbundled term is interacted with the dummy aimed at capturing the presence of financial restraints,  $\tilde{b}_4$  is again positive and of a similar magnitude as in Model  $FR^A$  and remains highly insignificant. In Model  $FR^A$  (*unrestricted*), which unbundles the financial restraints term, the real interest rate component  $\acute{b}_4$ , which captures the real interest rate differential is positive and insignificant. Interestingly, the inflation rate component  $b_5$  is positive and significant at the 5% level. Its sign suggests that a low domestic inflation rate relative to the world inflation rate has a positive effect on domestic investment (this effect varies with the volume of domestic lending relative to the capital stock). Conversely, when domestic inflation exceeds world inflation, domestic investment decreases (this effect also varies with the volume of loans relative to the capital stock). This is broadly in line with the traditional McKinnon-Shaw effect which suggests that high inflation has a negative effect on investment because it

depresses the supply of investable funds. However, the mechanism here is a different one. The inflation component of the financial restraints term captures the part of the low nominal interest rate that is due to low inflation. If domestic inflation is lower than world inflation, domestic nominal interest rates are low relative to the world capital market and this reduces the cost of capital associated with domestic loans. Model *FR-ER*, which includes the two exchange rate uncertainty variables, suggests that both terms capturing exchange rate uncertainty are negative as expected, although only one of the two -  $\tilde{b}_5$  - is significant at the 5% level while the other one -  $b_6$  - is insignificant. Thus, there is some evidence that exchange rate uncertainty depresses domestic investment.

However, one may argue that the assumption of cross-sectional independence is unrealistic in a world characterised by growing real and financial inter-linkages. In order to check for cross-sectional dependence in the estimates, we compute the long-run cross-sectional correlation matrix of the residuals obtained for each model.<sup>17</sup> The results show that the correlations for model *NC* lie between 0.25 and 0.89, with an overall average of 0.47, for model *FR<sup>A</sup>* between 0.23 and 0.90, with an overall average of 0.46, for model *FR<sup>D</sup>* between 0.23 and 0.88, with an overall average of 0.48, for model *FR<sup>A</sup>* (unrestricted) between 0.24 and 0.89, with an overall average of 0.53, and for model *FR - ER* between 0.24 and 0.88, with an overall average of 0.51. Overall, these results clearly show that the cross-sectional independence assumption is violated for all the models.

Since evidence of cross-sectional dependence is found, we use the CUP-FM estimator of Bai and Kao (2006), which allows for cross-sectional dependence through common factors. Table 5 reports the estimation results. Allowing for cross-country effects impacts on both the magnitude and significance of various coefficients and alters the economic interpretation of some of the results. The term proxying the marginal product of capital ( $b_2$ ) is once again always positive and strongly significant but its coefficient is much larger compared to the estimates obtained assuming cross-sectional

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<sup>17</sup>Tables reporting cross-correlations are not provided here for brevity. These tables are available upon request.

independence. The coefficient of the world interest rate ( $b_3$ ) remains negative and significant in all the models, but once again the estimated coefficients are much larger - hovering around  $-0.25$  compared to  $-0.08$  in Table 4 - suggesting that domestic investment appears to be much more responsive to world capital markets if one allows for cross-country effects. Remarkably, all the financial restraints terms remain positive but are now statistically significant at the 5% level, which now suggests that financial restraints do play an important role in determining investment. In Model  $FR^A$ , the unbundled financial restraints term -  $b_4$  - is positive and significant with a coefficient that has more or less the same size as the one on the world interest rate. The positive coefficient suggests that depressing the domestic interest rate through financial restraints results in additional domestic investment, in contrast to the McKinnon-Shaw prediction. In Model  $FR^D$ , which interacts the financial restraints term with the financial restraints dummy, the coefficient on financial restraints ( $\tilde{b}_4$ ) is more than twice the size of the world interest rate coefficient. This suggests that countries in which financial restraints were present are, in fact, the ones that may have benefited from low domestic interest rates. Model  $FR^A$  (*unrestricted*), which unbundles the interest rate differential into its two components does, however, provide some comfort to supporters of the McKinnon-Shaw hypothesis in that it continues to show, as in Table 4, the positive effects of low inflation on investment. Nevertheless, the effect of the real interest rate differential is now positive and significant at the 5% level, suggesting that depressing the real interest rate to below world levels has a positive effect on domestic investment. The positive effect of low inflation - or negative effect of high inflation - suggests that to some extent McKinnon and Shaw are right to emphasise the damage caused by high inflation. However, in our case this is not so much because of the reduced supply of funds but rather because of the higher cost of capital, since high inflation - in the absence of interest rate ceilings that were common before our sample period and were emphasised by McKinnon and Shaw - normally results in higher nominal interest rates. On balance, as is shown in Model  $FR^D$ , the aggregate effect of financial restraints on domestic investment is positive, although the effect of the inflation rate seems to be broadly along the lines suggested by McKinnon-Shaw.

The results also suggest that exchange rate uncertainty is an even more important determinant of investment if one takes into account cross-country effects. Both terms capturing exchange rate uncertainty ( $\tilde{b}_5$  and  $b_6$ ) are now significant at the 5% level and their coefficients are more than twice the absolute size compared to those reported in Table 4.

## 4 Summary and Conclusion

This paper employs recently developed panel data methods to estimate a model of private investment under financial restraints for 20 developing countries using annual data for 1972-2000. Unit root tests for cross-sectionally dependent panels show that the variables are non-stationary. The application of panel cointegration methods reveals a long run relationship among the variables. The nature of this relationship varies depending on whether we take into account cross-country effects. When we allow for cross-sectional dependence, investment displays more sensitivity to world capital market conditions and exchange rate uncertainty. A perhaps even more surprising result is the finding that financial restraints appear to have had a positive overall effect on domestic investment, in contrast to the McKinnon-Shaw prediction. On the other hand, our findings relating to the impact of inflation on investment accord well with the McKinnon-Shaw hypothesis – regardless of whether allowance is made for cross-sectional dependence. An applied econometrician who does not allow for cross-sectional dependence when estimating investment equations across a panel of countries may therefore find more support for the McKinnon-Shaw hypothesis than is warranted by the data.

Our findings, therefore, demonstrate the importance of cross-country effects in estimating investment models. In addition, they suggest that countries that managed to suppress domestic real interest rates without generating high inflation enjoyed higher levels of private investment than those that would have been obtained under liberalised conditions. There is, of course, a limit to the extent that real interest rates can be depressed by applying nominal interest rate ceilings without resorting to inflationary

policies. When low real interest rates are the result of high inflation, private investment it seems does not appear to increase. Thus, while mild financial repression can stimulate private investment, severe repression through high inflation may well have the opposite effect.

Our findings highlight two new avenues for further research. Firstly, they suggest that cross-country studies of private investment and possibly other macroeconomic aggregates need to take into account cross-country effects. Secondly, they suggest that it may be fruitful to re-examine the effects of financial repression on other key macroeconomic aggregates using the kind of techniques we have used in this paper.

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## Appendix A: Description and Sources of Data

$I$  is private fixed capital formation;  $K$  is private capital stock;  $Y$  is real GDP;  $r^*$  is US real lending rate;  $i^*$  is US nominal lending rate;  $r$  is domestic real lending rate;  $i$  is domestic nominal lending rate;  $B$  is claims on private sector by deposit money banks and other financial institutions;  $\pi^*$  is the US inflation rate (computed using the GDP deflator);  $\pi$  is the domestic inflation rate (computed using the GDP deflator);  $SDEX$  is the 3-year moving average of the standard deviation of the domestic exchange rate vis-à-vis the US dollar. The data is from the World Bank Development Indicators (2008). Data on private investment is from Everhart S.S and M.A. Sumlinski (2001). ‘Trends in Private Investment in Developing Countries, Statistics for 1970-2000 and the Impact on Private Investment of Corruption and the Quality of Public Investment.’ Discussion Paper No. 44, International Finance Corporation.

## Appendix B: List of Countries

The panel comprises Argentina, Bolivia, Chile, Costa Rica, Cote d’Ivoire, Dominican Republic, Egypt, El Salvador, Guatemala, India, Kenya, Malawi, Mexico, Morocco, Paraguay, Philippines, Thailand, Trinidad and Tobago, Uruguay and Venezuela.

## Appendix C: Monte Carlo Simulation Results

Size and Size-Adjusted Power Comparisons of Panel Unit Root Tests										
$N = 20$	Size					Power				
$T = 30$	CIPS	$t_a^*$	$t_b^*$	$BN_{ADF_{\hat{F}}}^c$	$BN_{Z_{\hat{e}}}^c$	CIPS	$t_a^*$	$t_b^*$	$BN_{ADF_{\hat{F}}}^c$	$BN_{Z_{\hat{e}}}^c$
	0.047	0.051	0.039	0.038	0.045	0.520	0.922	0.789	0.130	0.270

Notes: The following DGP was considered:  $y_{it} = \alpha_{i0} + z_{it}^0$ ,  $z_{it} = \rho_i z_{it-1}^0 + \tau \sum_{j=1}^K \beta_{ij} f_{it} + \sqrt{K} \varepsilon_{it}$ . with  $(f_{it}, \varepsilon_{it}, \alpha_{it}) \sim i.i.dN(0, I_3)$ ,  $\tau = 1$ , where the common factors and the idiosyncratic components are assumed to be of the same importance (see Gutierrez, 2006),  $\beta_{ij} \sim U[-1, 4]$ ,  $K = 1$ ,  $\rho_i = 1$ , and  $\rho_i \sim [0.9, 1]$  for the size and the size-adjusted power. The results were obtained using 1000 replications.

Table 1: CD Tests Results	
	<i>Statistics</i>
$\frac{I_{jt}}{K_{jt-1}}$	4.696 (0.009)
$\frac{Y_{jt}}{K_{jt-1}}$	7.865 (0.000)
$\frac{(i_t^* - i_{jt})}{(1+i_t^*)(1+\pi_{jt})} \frac{B_{jt}}{K_{jt-1}}$	5.289 (0.000)
$\frac{(r_t^* - r_{jt})}{(1+i_t^*)(1+\pi_{jt})} \frac{B_{jt}}{K_{jt-1}}$	5.077 (0.006)
$\frac{(\pi_t^* - \pi_{jt})}{(1+i_t^*)(1+\pi_{jt})} \frac{B_{jt}}{K_{jt-1}}$	7.631 (0.000)
$SDEX_{jt} \frac{B_{jt}}{K_{jt-1}}$	3.657 (0.034)
$SDEX_{jt}$	4.563 (0.011)

Notes: Pesaran (2004) shows that under the null hypothesis of no cross-sectional dependence  $CD \xrightarrow{d} N(0, 1)$ .  $p$ -values are in parenthesis.

Table 2: Panel Unit Root Tests Results

	$BN_{ADF_{\bar{F}}^c}$	$BN_{Z_{\bar{e}}^c}$	CIPS	$t_a^*$	$t_b^*$
$\frac{I_{jt}}{K_{jt-1}}$	-2.153 (0.180)	-1.037 (0.717)	-1.512 (0.840)	-1.270 (0.102)	-0.107 (0.410)
$\frac{Y_{jt}}{K_{jt-1}}$	-1.742 (0.410)	-1.910 (0.972)	-1.462 (0.851)	-0.291 (0.388)	-0.051 (0.490)
$\frac{(i_t^* - i_{jt})}{(1+i_t^*)(1+\pi_{jt})} \frac{B_{jt}}{K_{jt-1}}$	-1.659 (0.402)	-1.244 (0.694)	-0.438 (0.956)	-1.145 (0.120)	-0.983 (0.176)
$\frac{(r_t^* - r_{jt})}{(1+i_t^*)(1+\pi_{jt})} \frac{B_{jt}}{K_{jt-1}}$	-2.070 (0.195)	-0.639 (0.906)	-2.001 (0.163)	-1.170 (0.112)	-1.042 (0.150)
$\frac{(\pi_t^* - \pi_{jt})}{(1+i_t^*)(1+\pi_{jt})} \frac{B_{jt}}{K_{jt-1}}$	-1.097 (0.510)	-0.708 (0.929)	-1.416 (0.867)	-0.789 (0.210)	-1.070 (0.131)
$SDEX_{jt} \frac{B_{jt}}{K_{jt-1}}$	-2.090 (0.210)	-0.972 (0.780)	-0.363 (0.978)	-0.923 (0.170)	-1.120 (0.134)
$SDEX_{jt}$	-1.071 (0.520)	-0.456 (0.885)	-1.988 (0.167)	-1.119 (0.125)	-1.242 (0.110)

Notes: Sample period 1972-2000. The number of common factors selected using the IC2 criterion is equal to 1. The maximum number of factors is fixed to 4.  $BN_{ADF_{\bar{F}}^c}$  and  $BN_{Z_{\bar{e}}^c}$  denote the Bai and Ng (2004) unit root tests on common factor and idiosyncratic component respectively. The ADF test regression only includes a constant. The number of lags is selected using the Bayesian information criterion (BIC). The maximum number of lags is fixed to 4. CIPS denotes the panel unit root test proposed by Pesaran (2007). The truncated version is applied as suggested by Pesaran (2007). The appropriate lag-length for CIPS is selected using the Akaike information criterion (AIC) with a maximum number of lags equal to 4.  $t_a^*$  and  $t_b^*$  are the two statistics developed by Moon and Perron (2004).  $p$ -values are in parenthesis.

Table 3: Panel Cointegration Tests Results

Pedroni (1999, 2004) idiosyncratic cointegration		
Models	$Z_\rho$ test	$Z_t$ test
$NC$	-18.055 (0.000)	-24.699 (0.000)
$FR^A$	-6.068 (0.000)	-2.775 (0.002)
$FR^D$	-6.254 (0.000)	-4.519 (0.000)
$FR^A(unrestricted)$	-3.691 (0.000)	-6.719 (0.000)
$FR - ER$	-6.923 (0.000)	-7.215 (0.000)
Johansen (1988) factor cointegration		
Models	$H_0 : r =$	Trace Test
$NC$	0	23.029 (0.003)
	1	3.841 (0.531)
$FR^A$	0	31.482 (0.001)
	1	0.220 (0.639)
$FR^D$	0	24.395 (0.002)
	1	0.343 (0.558)
$FR^A(unrestricted)$	0	16.413 (0.036)
	1	0.074 (0.785)
$FR - ER$	0	15.495 (0.048)
	1	0.189 (0.663)

Notes: Sample period 1972-2000. The Pedroni tests include individual effects.  $Z_\rho$  and  $Z_t$  denote the panel coefficient  $\rho$  type and  $t$ -ratio tests. The number of cointegrating vectors in the Johansen (1988) trace test is denoted by  $r$  and the Akaike information criterion (AIC) lag length is 4.  $p$ -values are in parenthesis.

Table 4: Panel Estimation Results with Cross-Sectional Independence

Pedroni FMOLS					
	$NC$	$FR^A$	$FR^D$	$FR^A(unrestricted)$	$FR - ER$
$b_2$	0.0014 <sup>†</sup> (0.0004)	0.0028 <sup>†</sup> (0.0007)	0.0019 <sup>†</sup> (0.0006)	0.0021 <sup>†</sup> (0.0008)	0.0025 <sup>†</sup> (0.0007)
$b_3$	-0.0947 <sup>†</sup> (0.0169)	-0.0892 <sup>†</sup> (0.0202)	-0.0789 <sup>†</sup> (0.0190)	-0.0734 <sup>†</sup> (0.0210)	-0.0787 <sup>†</sup> (0.0215)
$b_4$	-	0.1570 (0.2135)	-	-	0.1480 (0.2319)
$\tilde{b}_4$	-	-	0.1320 (0.2314)	-	-
$\acute{b}_4$	-	-	-	0.1243 (0.1045)	-
$b_5$	-	-	-	0.2123 <sup>†</sup> (0.1071)	-
$\tilde{b}_5$	-	-	-	-	-0.1529 <sup>†</sup> (0.0768)
$b_6$	-	-	-	-	-0.1787 (0.1483)
$b_1$	0.5083 <sup>†</sup> (0.0923)	0.5183 <sup>†</sup> (0.1023)	0.4712 <sup>†</sup> (0.0893)	0.4892 <sup>†</sup> (0.1153)	0.5032 <sup>†</sup> (0.1093)

Notes: Sample period 1972-2000. The standard errors in parenthesis are computed using a sieve bootstrap procedure (see Fachin, 2004; Chang *et al.*, 2006).<sup>†</sup>denotes significance at the 5% level.

Table 5: Panel Estimation Results with Cross-Sectional Dependence  
Bai and Kao FMOLS

	$NC$	$FR^A$	$FR^D$	$FR^A(unrestricted)$	$FR - ER$
Two Stage					
$b_2$	0.0482 <sup>†</sup> (0.0215)	0.0512 <sup>†</sup> (0.0143)	0.0498 <sup>†</sup> (0.0141)	0.0503 <sup>†</sup> (0.0145)	0.0494 <sup>†</sup> (0.0141)
$b_3$	-0.2165 <sup>†</sup> (0.1024)	-0.2521 <sup>†</sup> (0.0831)	-0.2461 <sup>†</sup> (0.0804)	-0.2435 <sup>†</sup> (0.0811)	-0.2672 <sup>†</sup> (0.0762)
$b_4$	-	0.2323 <sup>†</sup> (0.0754)	-	-	0.2412 <sup>†</sup> (0.0761)
$\tilde{b}_4$	-	-	0.5321 <sup>†</sup> (0.1841)	-	-
$\acute{b}_4$	-	-	-	0.2127 <sup>†</sup> (0.0639)	-
$b_5$	-	-	-	0.2699 <sup>†</sup> (0.0923)	-
$\tilde{b}_5$	-	-	-	-	-0.3131 <sup>†</sup> (0.0982)
$b_6$	-	-	-	-	-0.5017 <sup>†</sup> (0.1218)
$b_1$	0.7378 <sup>†</sup> (0.1513)	0.7810 <sup>†</sup> (0.1934)	0.6951 <sup>†</sup> (0.1653)	0.7352 <sup>†</sup> (0.2397)	0.7012 <sup>†</sup> (0.2476)
Iterative					
$b_2$	0.0461 <sup>†</sup> (0.1083)	0.0152 <sup>†</sup> ( )	0.0489 <sup>†</sup> (0.0139)	0.5021 <sup>†</sup> (0.0141)	0.0154 <sup>†</sup> (3.71)
$b_3$	-0.2191 <sup>†</sup> (0.1023)	-0.2651 <sup>†</sup> (0.0689)	-0.2414 <sup>†</sup> (0.0789)	-0.2419 <sup>†</sup> (0.0792)	-0.2710 <sup>†</sup> (0.0796)
$b_4$	-	0.2231 <sup>†</sup> (0.0739)	-	-	0.2504 <sup>†</sup> (0.0761)
$\tilde{b}_4$	-	-	0.5215 <sup>†</sup> (0.1823)	-	-
$\acute{b}_4$	-	-	-	0.2214 <sup>†</sup> (0.0721)	-
$b_5$	-	-	-	0.2701 <sup>†</sup> (0.0923)	-
$\tilde{b}_5$	-	-	-	-	-0.3529 <sup>†</sup> (0.1123)
$b_6$	-	-	-	-	-0.5271 <sup>†</sup> (0.1391)
$b_1$	0.7630 <sup>†</sup> (0.1456)	0.7832 <sup>†</sup> (0.1835)	0.6783 <sup>†</sup> (0.1735)	0.7414 <sup>†</sup> (0.1998)	0.7234 <sup>†</sup> (0.1956)

Notes: Sample period 1972-2000. The standard errors in parenthesis are computed using a sieve bootstrap procedure (see Fachin, 2004; Chang *et al.*, 2006). <sup>†</sup>denotes significance at the 5% level.