PANEL COINTEGRATION ANALYSIS OF THE FINANCE-INVESTMENT LINK IN OECD COUNTRIES

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Abstract

Applying a panel cointegration analysis to data for 19 OECD countries from 1970 to 1999, the paper provides evidence that financial development is significantly related to investment. Specifically, adding indicators of financial development to a standard investment model (Jorgensen, 1967), we find these indicators to be highly significant. This is interesting, as the previous empirical evidence regarding the finance-investment link has been mixed, especially when the sample was restricted to relatively high-income countries. Also, including financial development indicators yields more plausible estimates for the other determinants of investment, output and the user cost of capital, as compared to those in many other previous studies. For example, the coefficient for output is estimated to be close to one and the coefficient of the user cost of capital is found to be significantly negatively related to investment, as predicted by theory.

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I. Introduction

There is a longstanding debate about whether financial development is conducive to growth and capital accumulation. Recent studies suggest that financial intermediation affects growth through various channels, including its effects on saving and investment rates and thus the rate of capital accumulation, as well as the efficiency of capital allocation (Levine, 1997, 1999, Pagano, 1993, Levine and Renelt, 1992; King and Levine, 1993). Typically, these studies select measures of per capita GDP growth, saving, investment, capital accumulation and productivity and regress them on a number of determinants, often selected without making reference to any specific benchmark model. This can make the interpretation of results difficult. An alternative is to start from a well-established theoretical model to select the control variables. For example, according to the survey by Chirinko (1993), by far the most frequently used specification for the analysis of investment spending has been the neoclassical model pioneered by Jorgenson (1967). The present paper takes this simple neoclassical investment model as a benchmark model for the empirical analysis. It then adds indicators of financial development and examines whether they contain any further explanatory power[1]. Thus, the present paper focuses on just one channel through which finance may influence growth, *i.e.* through its impact on the *level* of investment. In this context, one can distinguish between at least two channels, that is economic growth that follows either from increased total factor productivity or from increases in the country's human or physical capital stocks. By focusing on the correlation between financial development and investment levels, the present paper highlights the latter channel.

The present paper differs from the aforementioned and other papers in this area in terms of this focus, as well as in two other aspects, *i.e.* the sample and techniques used. First, most previous research have mixed developing and developed countries despite their differences. The present sample consists of the sub-group of OECD countries. These form a relatively homogeneous group of countries, so that the pooling of data is more likely to be justified. Note that so far the empirical evidence on the finance-investment link for this sample has been mixed, with several studies failing to identify a significant role of finance for economic development. For example, King and Levine (1993) found that excluding OECD countries from their full sample of developed and developing countries did not affect the significance of

1. Thus, as regards the methodology, the present paper is similar to Benhabib and Spiegel (2000). They use a standard neoclassical growth model to derive an empirical base-growth specification and then add financial development indicators to that equation. Estimating both benchmark and augmented growth regressions, they find some of their financial development

the relationship between financial development and investment. This is consistent with the view that the links are significant for non-OECD, but not for OECD countries. Fernandez and Galetovic (1994) confirmed the weak relationship of financial development and investment in a direct test of this relationship for OECD countries. By contrast, De Gregorio and Guidotti (1995) and Leahy *et. al.* (2001) identified a significant link. Our paper uses a similar, but larger and more recent, data set as Leahy *et. al.*. Second, and even more importantly, we use relatively recent panel cointegration techniques that exploit more fully the information available in the variation over time of the variables than most previous studies have done. The standard approach in the empirical literature on finance and growth is cross-section analysis (Levine, 1997; Wachtel, 2001),2 focusing on long-run average relationships and ignoring the time-series variation in the data. But the variation of variables over time may contain additional information, which may be particularly valuable in situations where the cross-section variation in the data is relatively limited (as is the case for our sample of relatively more homogenous OECD countries). In this context, it is useful to recall the results by Arestis and Demetriades (1996, 1997) and Arestis *et. al*. (2001). Specifically, they point out that time series analysis rather than cross-section analysis of growth determinants may yield noticeably different results regarding the role of different financial development measures. While recent cross-section studies often suggest that stock market capitalisation is a more important determinant of investment and growth than private credit, Arestis *et. al.*, using time series analysis and data for five industrialised countries, find much smaller coefficient estimates for stock market capitalisation than in those cross-section studies. They conclude that the role of stock market capitalisation is probably overestimated in the latter. Against this background the present paper uses a panel cointegration approach, which exploits both cross-section as well as time series variation in the data. Specifically, we use the panel cointegration tests suggested by Kao (1999) and Pedroni (1995). Compared to individual time series tests, these panel cointegration tests have higher power. To estimate the long-term relationships we then use four different panel estimators (OLS, bias-corrected OLS, fully modified OLS and dynamic OLS).

Section II presents the conceptual framework, the simple neoclassical investment model by Jorgensen (1967), as well as an overview of previous empirical results using this model.

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measures to be highly significant in the latter. They conclude that there is a strong link between financial development and growth. Our approach is similar, except that it uses a standard investment instead of growth model as the benchmark model.

2. Another common approach is to aggregate data over a time period of five years or ten years in a panel framework. The temporal aggregation may bias the results and induce an apparent relationship even if no relationship is present. Instrumental variables and GMM estimators do not necessarily solve this problem.

Section III contains the results. It presents the data in the first sub-section and the results of panel unit root tests in its second sub-section. The third sub-section presents the results of panel cointegration tests. They provide strong evidence for the existence of a long-run relationship between financial development and investment. The third sub-section also discusses the estimated parameters. They suggest that the relationship is much stronger when financial development is measured by liquid liabilities or private credit issued by deposit money banks than stock market capitalisation or total value traded. Section IV concludes the paper.

II. Regression specifications

The base-investment model used here is the standard neoclassical investment model (Jorgensen, 1967), some details of which are presented in this paragraph for easy reference. The model describes the determination of the equilibrium capital stock in an environment free of tax, uncertainty, adjustment costs and market imperfections. Combining this with an assumption about the constancy of the long-run relationship between the capital stock and gross investment, an expression for gross investment as a function of output and the user costs of capital can be obtained as follows. Given a constant depreciation rate delta, capital accumulates according to the following identity: $\boldsymbol{K}_{i,t} = (1-\delta_i)\boldsymbol{K}_{i,t-1} + \boldsymbol{I}_{i,t}$, where i is a country index and t a time index, with i running from 1 to N and t from 1 to T. Assuming a constant growth rate of the capital stock in the steady state, gross investment can be expressed as a constant fraction of the capital stock: $\boldsymbol{I}_{i,t} = (\boldsymbol{g}_i + \delta_i)\boldsymbol{K}_{i,t-1}$, where g_i is the (time-invariant) growth rate for the capital stock in country *i*. With a Cobb-Douglas production function, profit maximising yields the steady state identity: $\boldsymbol{c}_i\boldsymbol{K}_{i,t-1} = \alpha_i\boldsymbol{Y}_i\boldsymbol{P}_i$, where $\boldsymbol{Y}_i\boldsymbol{P}_i$ is nominal income, α_i is the share of capital in total income and $\boldsymbol{c}_i$ is the user cost of capital. Taking logs on both sides of the last equation, the following long-run relationship between gross investment, output and the cost of capital is obtained as $\boldsymbol{i}_i = \alpha_i' + \boldsymbol{y}_i - (\boldsymbol{c}_i - \boldsymbol{p}_i)$, where $\alpha_i' = \log(\boldsymbol{g}_i + \delta_i) + \log(\alpha_i)$ and $\boldsymbol{c}_i - \boldsymbol{p}_i$ stands for the user cost in real terms. This suggests that one could estimate the relationship between gross investment, output and the real interest rate (as a proxy for the user cost in real terms) using cointegration analysis. Thus the equation to be estimated, in the following referred to as the benchmark investment equation, is as follows:

$$i_{i,t} = \beta_{1,i} y_{i,t} + \beta_{2,i}(c_{i,t} - p_{i,t}) + u_{i,t} \qquad (1)$$

where $u_{i,t} = \alpha_i^{'} + \varepsilon_{i,t}$, $\alpha_i^{'}$ is an individual fixed effect and $\varepsilon_{i,t}$ is a random error term.

Empirically, according to the survey of the neoclassical investment model by Chirinko (1993), output $y_{i,t}$ is the dominant determinant of investment in regression equations such as (1). Subsequent empirical work confirmed this. The output measure was generally highly significant and robust. Even, the estimated coefficient was sometimes implausibly high (*e.g.* Leahy *et. al.*, 2001), so that some authors imposed a unit coefficient on the long-run finance-investment relationship *a priori* (Bassanini *et. al.* 2000, Ashworth and Davis, 2001). By contrast, empirical studies have found the other determinant in (1), the real user cost $(c_{i,t} - p_{i,t})$, to have at best a very modest effect (Chirinko, 1993). While this statement may have to be qualified in the case of studies for individual countries,3 recent panel data studies confirmed the difficulty in identifying a significant coefficient for the user costs of capital measure (*e.g.* Schich and Pelgrin, 2001).

It is likely that this reflects the omission of other variables conducive to investment, such as financial development. It has become generally accepted today that there exist market imperfections, such as asymmetric information in the relation between investor and saver, so that the Modigliani-Miller theorem, which states that in an ideal market, investment decisions are independent of financial considerations, is irrelevant in actual practise. Market imperfections create a role for the financial sector in reducing the special transaction costs arising from those imperfections. Specifically, the financial sector accumulates special knowledge in evaluating and monitoring investment projects and thus develops a comparative advantage in evaluating risks and designing financial contracts in a way that they ensure incentive compatibility between both investors and savers. For example, banks may gain information advantages from lasting relations with customers by learning from past experience. Financial institutions also serve in the monitoring of investments in order to reduce the risk that investment agents (managers) mismanage resources. For example, stock markets may act as markets for control and thereby function as a disciplining device that reduces costs associated with the delegation of management to agents. All these considerations suggest that more developed financial sectors reduce the costs arising from

3. For example, Ashworth and Davis (2001), using time series analysis, identify a significant role for the user cost in the case of one country (Germany) out of their sample of seven (G-7) countries.

market imperfections. Thus, a positive relationship between financial development and investment is expected in the following modified investment equation:

$$i_{i,t} = \beta_{1,i} y_{i,t} + \beta_{2,i} (c_{i,t} - p_{i,t}) + \beta_{3,i} f_{i,t} + u_{i,t} \qquad (2)$$

where $f_{i,t}$ is a proxy of financial development and all other variables are the same as in (1). We consider specifications (1) and (2) alternatively in our empirical analysis.

III. Data and empirical results

III.1 Data

The data consist of investment, output, two alternative measures of the user cost of capital and three measures of financial development for a panel of 19 OECD countries from 1970 to 1999.4 Gross investment is measured here by the (log) level of real private business sector fixed capital formation. The output variable is the log of real private gross domestic product. The user costs of capital measures are an adjusted real long-term interest rate and the user cost of capital measure calculated by the OECD. The adjusted real long-term interest rate is the *log*(1+*r*), where *r* is the product of a long-term real interest rate (derived from government securities) and the ratio of a deflator of private non-residential fixed capital formation to an output price deflator (to adjust for relative price changes between capital goods and output). This expression would form part of a measure of the user cost of capital (Jorgenson, 1967). Even though an incomplete measure as depreciation rates, rates of investment credit and tax depreciation allowances are not included for lack of sufficient time series in formation across countries, it nevertheless captures most of the time series variation in the user cost of capital (*i.e.* the variation in the real interest rates and price deflators). Alternatively, we used the user cost of capital measure taken from the OECD's *Analytical Database* in the expression *log*(1+*ucc*/*PIB*), where *ucc* is the user cost of capital measure and *PIB* the deflator of private non-residential fixed capital formation. The results were broadly similar, which suggest that the adjusted real long-term interest rate indeed captures part of the time series dynamics of a more complete measure of the user costs of capital. The main text focuses on the results using

4. The sample includes: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, United Kingdom and United States. In cases where less than 20 observations were available for a country, it was eliminated from the regression specification. For example, in the specifications where financial development is measured by data on stock market capitalisation, Finland, New Zealand and Norway (with 15, 13 and 17 observations, respectively) were excluded. The financial development indicators are obtained from the World Bank's financial development database (see Beck, Demirgüç-Kunt and Levine, 1999). As the data for our sample of OECD countries (in that database) are fully available only until 1996, we update the data up to and including 1999, using the methodology described in Beck et. al.. All other data used here are from the OECD.

the former measure and the appendix reports the results for the second measure for comparison.

Measuring financial development is difficult. Ideally, one wants to measure the extent to which the financial sector reduces the transaction costs arising from market imperfections. This points to the services provided by the financial sector and would suggest measures related to the efficiency and competitiveness of the financial sector. However, time series data are generally not available for those measures, but only for size indicators. Therefore, the standard practice in empirical research is to use measures of components of the financial system (relative to GDP). As each of them has some shortcomings, a set of four indicators — liquid liabilities, private credit of deposit money banks, stock market capitalisation and total value traded as a share of GDP — is used here.5 This set is likely to convey a richer picture of financial development than any single measure. We consider both measures related to financial inter-mediation through banks as well as through stock markets, which allows us to contribute to the discussion about the relative roles of different types of financial development measures (see *e.g.* Arestis *et. al.*).

III.2 Panel unit root tests

Before testing for cointegration we need to confirm whether the variables are actually non-stationary. We adopt the approach suggested by Im, Pesaran and Shin (1997, henceforth IPS), essentially the standard ADF-test in a panel context:

$$\Delta y_{i,t} = \mu_i + \beta_i t + \rho_i y_{i,t-1} + \sum_{j=1}^{p} \varphi_{ij} \Delta y_{i,t-j} + \varepsilon_{i,t} \qquad (3)$$

where $\boldsymbol{y_{i,t}}$ stands for each of the variables presented in section III.1. The null hypothesis and the alternative hypothesis are defined as:

$H_0 : \rho_i = 0$ for all i;

$H_1 : \rho_i < 0$ for at least one i.

Instead of pooling and assuming that ρ_i is the same for all countries (Levin and Lin, 1993), the IPS methodology uses separate unit root tests for the N countries. The test statistic (denoted IPS t-bar statistic) is then calculated as the average of the individual ADF statistics,

5. Liquid liabilities consist of currency and interest-bearing liabilities of bank and non-bank financial intermediaries. Private credit of deposit money banks provided to the private sector are the total claims of deposit money banks on the private sector. Stock market capitalisation measures the value of all listed shares and total value traded the value of total shares traded on domestic stock market exchanges. A discussion of advantages and disadvantages of specific financial development measures is provided in Levine, Loayza and Beck (2000) and Leahy et. al. (2001).

$\bar{t}_{N,T} = \left(1/N\right)\sum_{i=1}^{N} t_i$, where t_i is the ADF t-statistic for the OLS estimate in equation (3) for the *i*-th country. This statistic has the following interesting property. Assuming no cross-country correlation among the errors and T to be the same for all countries, the normalised statistic converges in distribution to a standard normal one, $\sqrt{N}\left(\frac{\bar{t}_{N,T} - E\left[\bar{t}_{N,T}\right]}{\sqrt{Var\left[\bar{t}_{N,T}\right]}}\right) \Rightarrow N(0,1)$, where $\Rightarrow$ denotes convergence in distribution $E\left[\bar{t}_{N,T}\right] = \mu$ and $Var\left[\bar{t}_{N,T}\right] = \sigma^2$ are tabulated on the basis of Monte-Carlo simulations (IPS, 1997). Applying this test to the seven variables under consideration, the null of non-stationarity can not be rejected for any variable (table 1).6 This result is robust to the choice of lag orders (up to three were considered here), as well as to the inclusion of time trend. To see how robust the results are the specific aggregation used here, we also test for the existence of a unit root for each variable in each country individually. Using ADF and Phillips-Perron tests, we find that for almost all countries the presence of a unit root for investment and GDP is not rejected. Results are somewhat more mixed in case of the financial development variables and the user cost of capital, though the null hypothesis of a unit root is not rejected in the majority of countries. We conclude that the variables have a unit root, i.e. they are stationary in first differences, and proceed to test for cointegration.

III.3 Panel cointegration tests

Before conducting panel cointegration tests, we conduct individual cointegration tests for the variables in each country, using the trace and max-eigenvalue statistic from the Johansen test. The results are consistent with the existence of one cointegrating relationship in most countries, with the evidence for a cointegration relationship being slightly stronger for the augmented as compared to the base model.7 Nevertheless, it should be borne in mind that these tests have low power in small samples (such as ours), which is one of the original motivations for applying panel cointegration test.

Next, we apply the panel cointegration tests suggested by Pedroni (1995) and Kao (1999). They assume that the slope coefficients are the same across countries ("homogenous

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6. We also applied the unit root test suggested by Harris and Tzavalis (1999). The results were broadly similar, except when a time trend is included. In this case, the null hypothesis of non-stationary was rejected for the financial development variables. However, it has been shown on the basis of Monte-Carlo simulations (IPS, 1997) that the inclusion of a trend makes the tests less powerful.

7. Without inclusion of a trend (both with and without a constant), the null hypothesis of no cointegrating relationship could be rejected for most countries (and the null hypothesis of at most one could not be rejected). However, when additional lags are included, the evidence becomes weaker in the case of some countries. Nevertheless, the null of no cointegration is still rejected for most countries.

cointegration tests”). To see whether this is consistent with the data, we apply a Wald test comparing the $\beta_i = \beta$ for all i with unrestricted β_i ’s (Pedroni, 2000) and a joint Haussmann test. The Wald test rejects the null hypothesis of identical coefficients at conventional significance levels. However, Pesaran *et. al*. (1999) argue that it typically rejects the null hypothesis in most applied examples and suggest using a joint Haussmann test instead. Applying this test, we can not reject the null of homogeneity in most cases. Specifically, we do not reject homogeneity in the case of the model without financial development (a statistic of 3.17 with p-value 0.17) as well when financial development is proxied either by private credit (2.82 with p-value 0.42), stock market capitalisation (1.21 with p-value 0.75) or total value traded (1.14 with p-value 0.78). While the results are different when financial development is proxied by liquid liabilities (15.06 with p-value 0.00), taking all the evidence together, we conclude that the homogeneity assumption is justified overall.8 This appears to be a reasonable as regards the long-run relationship for a sample of relatively homogenous group such as the OECD countries. It may be due, for example, to arbitrage conditions or common technologies influencing all countries in a similar way.

The two aforementioned tests are based on the null of no cointegration. Kao (1999) presents two types of cointegration tests, the Dickey-Fuller (DF) and the Augmented Dickey-Fuller (ADF) types. Each one is based on the OLS estimated residuals $\hat{u}_{i,t}$ from the estimated equations (1) or (2). The DF-type test consists of estimating ρ in the following form:

$$\hat{u}_{i,t} = \rho \hat{u}_{i,t-1} + v_{i,t} \qquad (4)$$

and testing the null $H_0 : \rho = 1$. In this context, four different statistics are used (see also Kao, 1999). They differ with regard to the assumptions regarding exogeneity of regressors with respect to the errors and possible cointegration with endogenous regressors (see footnote to the table 2 for details). The ADF-type test is based on the following regression:

$$\hat{u}_{i,t} = \rho \hat{u}_{i,t-1} + \sum_{j=1}^{p} \psi_j \Delta \hat{u}_{i,t-j} + v_{i,t} \qquad (5)$$

and the null $H_0 : \rho = 1$. The test suggested by Pedroni (1995) involves a non-parametric correction. He provides two statistics from a pooled Phillips and Perron-type test.

8. We also experimented with heterogeneous panel cointegration tests (Pedroni, 1997, 2000). The results obtained were broadly similar to those described here.

The results of the cointegration tests for the variables in equations (1) and (2) are shown in table 2. Almost all statistics reject the null of no cointegration.9 An exception is the DF_t statistics. However, as shown in Kao (1996), when the cross-section and time series dimensions are comparable and the estimated variance is small (as is the case here), the DF_rho* and DF_t* tests have size and power properties superior to the DF_t statistics. We conclude that the variables are cointegrated. This holds both for the variables in (1) as well as in (2). If anything, the evidence for cointegration is slightly stronger when the financial development variables are included. Specifically, the DF-t* test rejects the null of no cointegration at a higher level of significance in the case of the model augmented by financial development variables than in the case of the base-investment model. The results from the test suggested by Pedroni are similar, also rejecting the null hypothesis of no cointegration.

III.4 Inference

Four different homogenous estimation approaches are used, the standard OLS, BCOLS (bias-corrected OLS), FMOLS (fully modified OLS) and DOLS (dynamic OLS) estimators described in Kao and Chiang (2000). To explain the different estimators, we write the equations (1) and (2) alternatively as a system of the following form:

$$\begin{aligned} \boldsymbol{Y} &= \alpha + \boldsymbol{X^t}\beta + \boldsymbol{u}_1 \\ \Delta\boldsymbol{X} &= \boldsymbol{u}_2 \end{aligned} \qquad (6)$$

where Y is an NT x 1 vector, α is NT x 1 vector of constants, β is an 3 x 1 vector of the slope parameters and X is a 3xNT matrix composed of the regressors (output, real adjusted interest rate and one of the measure of financial development), with N the number of countries and T the number of years. u_1 and u_2 are NT x 1 vectors of residuals. The second equality in (4) states that the independent variables are an integrated process of order one for all i so that their first differences are stationary. Under these specifications, equation (6) specifies a system of cointegrated regressions. In addition, we assume that $\{y_{i,t}, x_{i,t}\}$ are independent across countries and $w_{i,t} = \{u_{i,t}, \varepsilon_{i,t}^{'}\}^{'}$ satisfies the assumptions of Kao and Chiang (2000).

9. An alternative test is that for the null of cointegration, as suggested by McCoskey and Kao (1998a, b) and discussed in Banerjee (1999) and Baltagi and Kao (2000). Applying this type of test gives weaker results in favour of cointegration. However, these tests are quite new and the small size properties not as well as analysed as for the more standard test for the null of no cointegration. This is why we have more confidence in the results of the latter.

The standard OLS estimator, $\hat{\beta}_{OLS}$, is obtained by subtracting individual means for each cross-section and regressing $\boldsymbol{y}_{i,t} - \boldsymbol{y}_{i,.}$ on $\boldsymbol{x}_{it,k} - \boldsymbol{x}_{i.,k}$, where $y_{i,.}$ and $y_{i,.k}$ are the mean of the dependent variable of each cross-section element and the mean of the *k*-th explanatory variable for each cross-section and *k* denotes the *k*-th regressor with *k*=1, 2 and 3, i.e. the OLS estimator of β is $\hat{\beta}_{OLS} = \left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{i,t}-\bar{x}_i)(x_{i,t}-\bar{x}_i)'\right]^{-1}\left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{i,t}-\bar{x}_i)(y_{i,t}-\bar{y}_i)\right]$ where $\bar{x}_i = (1/T)\sum_{t=1}^{T} x_{i,t}$ and $\bar{y}_i = (1/T)\sum_{t=1}^{T} y_{i,t}$.

The BCOLS is obtained by subtracting an error correction term from the standard OLS estimator, *i.e.* $\hat{\beta}_{BCOLS} = \hat{\beta}_{OLS} - \frac{\hat{\delta}}{T}$. The FMOLS estimator is obtained by modifying the variable $\boldsymbol{y}_{i,t}$ to correct for possible endogeneity, $\boldsymbol{y}^{+}_{i,t} = \boldsymbol{y}_{i,t} - \hat{\Omega}\Delta\boldsymbol{x}_{i,t}$, where $\hat{\Omega}$ is derived from consistent estimates of partitioned elements of the long-run covariance matrix of $\{w_{i,t}\}$. As well, a term, $\hat{\Delta}^{+}_{\varepsilon u}$, is included to correct for serial correlation, with $\hat{\Delta}^{+}_{\varepsilon u} = \hat{\Delta}_{\varepsilon u} - \hat{\Delta}_{\varepsilon}\hat{\Omega}'$ where $\hat{\Delta}_{\varepsilon u}$ and $\hat{\Delta}_{\varepsilon}$ are kernel estimates of their respective elements in the one-sided long-run covariance matrix of $\{w_{i,t}\}$. Thus, the FMOLS estimator is $\hat{\beta}_{FMOLS} = \left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{it}-\bar{x}_i)(x_{it}-\bar{x}_i)'\right]^{-1}\left[\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)\hat{y}^{+}_{it} - T\hat{\Delta}^{+}_{\varepsilon u}\right)\right]$. The DOLS estimator is calculated by estimating the following equation:

$$\mathrm{i}_{i,t} = \alpha + \beta_1 \mathrm{y}_{i,t} + \beta_2(\mathrm{c}_{i,t} - \mathrm{p}_{i,t}) + \beta_3 \mathrm{fd}_{i,t}$$
$$+ \sum_{j=-p}^{q}\gamma^1_j \Delta \mathrm{y}_{i,t+j} + \sum_{j=-p}^{q}\gamma^2_j \Delta(\mathrm{c}_{i,t+j} - \mathrm{p}_{i,t+j}) + \sum_{j=-p}^{q}\gamma^3_j \Delta \mathrm{fd}_{i,t+j} + \mathrm{u}_{i,t} \qquad (7)$$

where *p* and *q* are the lags and leads.

Table 3 reports the empirical results, five of which are singled out for special attention. First, the coefficient for output is always positive and highly significant, consistent with earlier empirical work. Second, the coefficient is almost always close to one, except in the case of the OLS estimator and in one instance when using the BCOLS estimator. However, as pointed out by Chen, McCoskey and Kao (1999), the latter two estimators have a non-negligible bias in small samples such as ours. On that account the FMOLS and DOLS are preferable, with the DOLS exhibiting the least bias in small samples as shown by Kao and Chiang (2000) using Monte Carlo simulations. Often the DOLS estimator has the drawback that its results are sensitive to the choice of number of lags and leads (Kao and Chiang, 2000), however, for our sample we find that most coefficient estimates vary only little when the number of leads and lags are changed (appendix table 2). Thus, we conclude that the estimated output coefficient is insignificantly different from one. This is consistent with theoretical priors. It is an important result, as the coefficient estimates in similar previous investment equations were either implausibly high (*e.g.* Leahy *et. al.*, 2001) or a unit coefficient was imposed *a priori* (Bassanini *et. al.*, 2001). We think that this reflects the choice of estimation techniques, with the present cointegration approach (which focuses on long-term relations) allowing us to obtain more precise estimates of the coefficient in the long-term relations.

Third, the user costs of capital proxy is almost always negative, as predicted by theory. Moreover, it is almost always both negative and significant when including either private credit or stock market capitalisation to measure financial development. The result that the user cost of capital are significantly negative related to investment is important, as previous work has mostly failed to identify a significant role for this variable. The higher values for the cost of capital obtained under the FMOLS and DOLS are consistent with earlier findings by Caballero (1994, 1999) for US data. Applying the Stock and Watson (1993) procedure, the author found a coefficient on the cost of capital close to minus one in a regression for US capital equipment, where the unit coefficient on output is imposed.

Fourth, adding financial development to the base equation, this variable is always significant, irrespective of the indicator and estimator chosen. This provides strong support for the hypothesis that financial development is conducive to investment even in high-income countries, even though such estimates do not establish causality. This confirms earlier work by Leahy *et. al.* (2001). It also expands on earlier work which has failed to identify a significant role for financial development for investment and capital accumulation in OECD countries, using methods that do not fully exploit both the cross-sections and time dimensions.

Fifth, while all three financial development measures - private credit, stock market capitalisation and total value traded - are always significantly positive, the coefficients estimates suggest a much stronger role for private credit than for the two stock markets indicators (stock market capitalisation or total value traded). This result differs from the suggestions in many other studies, using cross-section analysis. Indeed, to further shed light on the issue of the relative importance of these two sources of financing, we conduct a "horse-race" exercise, including both private credit and a stock market measure at the same time in the regression specification. The results confirm a stronger role for private credit rather than stock market measures (table 4, columns 2 and 4). While this result differs from those in several previous empirical studies, it confirms the hypothesis by Arestidis *et. al.* (2001), specifically that standard cross-section analysis tends to overestimate the role of stock market capitalisation.

To test for robustness of results, we re-estimated the regressions shown in tables 3 for all sub-sample periods obtained by deleting from the original sample (ending in 1999) the most recent year until the sub-sample ends in 1995. Appendix table 3 shows the results for the shortest sub-sample (1970 to 1995). Three observations are singled out for attention. First, the results are not significantly different from those shown in table 3 shown in the main text. Second, even though not significant, an interesting change in the estimated coefficients for the financial development variables can be observed. Specifically, if we use the short sample ending in 1995 as a starting point and increase the sample successively to include more recent data up to 1999, the estimated coefficients for the two stock market indicators variables increase, while the opposite is true for the private credit variable. This can be seen by comparing columns 1 with 2 and 3 with 4 in table 4. This change in the relative importance of the two types of financial development variables is confirmed by including both types of variables jointly.10 This may reflect the effect of the stock market boom, which started in many OECD countries around 1995 and it suggests the usefulness of exploiting the time-series variation in the data when analysing the roles of the two types of financial development measures. Third, the estimated coefficients for the user cost of capital proxies increases (become less negative), as more recent data are included. With the real interest rate remains broadly unchanged during the second half of the 1990s, the increasing contribution of the user cost of capital to explaining investment appears to reflect the recent decrease in the relative price of capital goods. This suggests that it is useful to account for the latter when constructing user cost of capital measures.

10. In principle, given that one of the financial market indicators is sufficient to yield a co-integration relationship, one would have to test for whether more than one co-integration vectors can be found to make the results of Table 4 strictly comparable to those from Table 3. However, the application of a multiple equation test akin to the Johansen-Juselius test in time-series is cumbersome to apply in a panel setting and therefore beyond the scope of this paper.

IV. Conclusions

The present paper uses a standard investment model and augments it to include financial development as a framework for the empirical analysis. It provides empirical evidence for the hypothesis that financial development contributes to growth through its effect on the level of investment, even in the case of countries with relatively high income levels. This is similar to the results presented in Schich and Pelgrin (2000). But it differs from those in other previous empirical studies restricting the sample to high-income-level countries, generally failing to identify a significant contribution of finance to overall growth. Moreover, the estimates for the other investment determinants identified here appear more plausible in two aspects than those in previous studies. First, the coefficient on output is close to one, consistent with a constant capital output ratio in the steady state. Previous papers have either identified coefficients significantly different from one or have imposed a unit coefficient *a priori*. Second, the user cost of capital measure is negative and significant, as expected. Finally, our results contribute to the discussion of the relative roles for growth of either bank or stock market financing. Specifically, they provide some support for the hypothesis by Arestidis *et. al.* (2001) that previous empirical studies have tended to overestimate the role of stock markets as compared to bank credits in explaining investment levels over the long-term by using techniques that ignore the time variation in the data.

Table 1. Panel Unit root tests

	Variables	t-value	p-value
Im, Pesaran and Shin (1995)			
Without time trend	Investment	7.55	0.00
	Gross domestic product	7.52	0.00
	Real adjusted interest rate	7.18	0.00
	Liquid Liabilities	7.39	0.00
	Private credit	7.66	0.00
	Stock market capitalisation	6.49	0.00
	Total Value Traded	6.46	0.00
With time trend	Investment	10.80	0.00
	Gross domestic product	10.72	0.00
	Real adjusted interest rate	10.95	0.00
	Liquid Liabilities	10.40	0.00
	Private credit	10.87	0.00
	Stock market capitalisation	9.37	0.00
	Total Value Traded	9.27	0.00

Table 2. Panel Cointegration tests

	Benchmark model	Inclusion of financial development variables			
	Without financial development	Liquid liabilities	Private credit	Stock market capitalisation	Total Value Traded
Pedroni (1995)[1]					
PC1	– 13.93	– 15.41	– 15.32	– 17.76	– 18.09
PC2	– 13.69	– 15.14	– 15.05	– 17.37	– 17.71
Kao (1999)[2]					
ADF	– 3.52	– 4.00	– 3.91	– 5.97	– 5.03
Kao (1999)[3]					
DF-rho	– 3.55	– 4.18	– 4.13	– 5.33	– 5.48
DF-t	19.33	18.81	18.87	13.99	13.90
DF-rho*	– 6.91	– 7.87	– 7.85	– 8.67	– 9.19
DF-t*	– 2.30	– 2.65	– 2.60	– 3.39	– 3.45

Note: All tests are left-hand side, i.e. large negative values are used to reject the null of no cointegration.
1. PC1 and PC2 are the non-parametric Phillips-Perron tests.
2. The ADF test is analogous to the parametric Augmented Dickey-Fuller test for nonstationary time series.
3. The DF test is analogous to the parametric Dickey-Fuller test for nonstationary time series. The DF-rho and DF-t statistics assume strict exogeneity of the regressors with respect to errors and no autocorrelation. DF-rho* and DF-t* statistics are based upon endogeneous regressors. Note that these tests depend on consistent estimates of the long-run variance-covariance matrix to correct for nuisance parameters once the limiting distribution has been found.

Table 3. Panel estimates of investment equations (data from 1970 to 1999)

	Standard model	Models including financial development indicators			
	Without financial development	Liquid liabilities	Private credit	Stock market capitalisation[2]	Total Value Traded[2]
OLS estimates					
Financial development[1]		0.162	0.115	0.045	0.019
		(0.037)**	(0.022)**	(0.013)**	(0.008)**
Gross domestic product	1.230	1.158	1.158	1.423	1.396
	(0.031)**	(0.034)**	(0.036)**	(0.049)**	(0.045)**
Adjusted real interest rate	– 0.290	– 0.291	– 0.445	– 0.376	– 0.409
	(0.191)	(0.187)	(0.187)**	(0.169)**	(0.137)**
Memorandum item:					
R-square	0.78	0.78	0.79	0.88	0.88
Bias-corrected OLS estimates					
Financial development[1]		0.219	0.145	0.045	0.016
		(0.071)**	(0.045)**	(0.015)**	(0.008)**
Gross domestic product	1.181	1.083	1.048	1.352	1.393
	(0.076)**	(0.078)**	(0.082)**	(0.078)**	(0.080)**
Adjusted real interest rate	– 0.300	– 0.299	– 0.508	– 0.305	– 0.313
	(0.256)	(0.247)	(0.246)**	(0.202)	(0.204)**
Memorandum item:					
R-square	0.78	0.78	0.79	0.88	0.88
FMOLS estimates					
Financial development[1]		0.387	0.257	0.045	0.036
		(0.074)**	(0.046)**	(0.016)**	(0.008)**
Gross domestic product	0.926	0.931	0.933	0.931	0.933
	(0.079)**	(0.081)**	(0.085)**	(0.081)**	(0.084)**
Adjusted real interest rate	– 0.259	– 0.183	– 1.071	– 1.704	– 1.810
	(0.266)	(0.256)	(0.254)**	(0.211)**	(0.213)**
Memorandum item:					
R-square	0.73	0.77	0.77	0.74	0.79
DOLS estimates[3]					
Financial development[1]		0.180	0.166	0.064	0.041
		(0.076)**	(0.050)**	(0.018)**	(0.009)**
Gross domestic product	1.004	0.995	0.998	1.006	1.005
	(0.082)**	(0.084)**	(0.092)**	(0.090)**	(0.093)**
Adjusted real interest rate	0.105	0.090	– 0.505	– 1.227	– 0.969
	(0.275)	(0.266)	(0.274)*	(0.232)**	(0.235)**
Memorandum item:					
R-square	0.54	0.62	0.68	0.72	0.76

Note: ** significant at 5 per cent level. * significant at 10 per cent level.

1. Financial development is alternatively proxied by liquid liabilities, private credit, stock market capitalisation and total value traded.
2. When measuring financial development by stock market and total value traded, data availability required us to reduce the sample size to 16 countries.
3. The results shown here are those for our preferred lead-lag combination. They are one lead and no lag (1,0) for the model without financial development, (1,0) for liquid liabilities, (2,0) for private credit and (0,2) for stock market capitalisation and total value traded.

Table 4. Robustness of results

	1976-1995	1976-1999	1976-1995	1976-1999
OLS estimates				
Financial development				
Private credit	0.243 (0.056)**	0.204 (0.048)**	0.280 (0.057)**	0.251 (0.049)**
Stock market capitalisation[1]	0.063 (0.014)**	0.063 (0.012)**		
Total value traded			0.025 (0.006)**	0.029 (0.005)**
Gross domestic product	1.233 (0.063)**	1.259 (0.051)**	1.251 (0.063)**	1.250 (0.052)**
Adjusted real interest rate	– 0.399 (0.177)**	– 0.470 (0.159)**	– 0.410 (0.178)**	– 0.483 (0.159)**
R-square	0.86	0.89	0.85	0.90
Bias-corrected OLS estimates				
Financial development				
Private credit	0.284 (0.081)**	0.225 (0.077)**	0.319 (0.085)**	0.275 (0.079)**
Stock market capitalisation[1]	0.062 (0.016)**	0.063 (0.014)**		
Total value traded			0.018 (0.008)**	0.026 (0.007)**
Gross domestic product	1.148 (0.092)**	1.180 (0.080)**	1.201 (0.097)**	1.189 (0.007)**
Adjusted real interest rate	– 0.463 (0.205)**	– 0.480 (0.080)**	– 0.466 (0.097)**	– 0.490 (0.084)**
R-square	0.86	0.90	0.85	0.90
FMOLS estimates				
Financial development				
Private credit	0.427 (0.085)**	0.356 (0.081)**	0.397 (0.089)**	0.293 (0.083)**
Stock market capitalisation[1]	– 0.006 (0.017)	0.013 (0.015)		
Total value traded			0.004 (0.008)	0.018 (0.008)**
Gross domestic product	0.930 (0.096)**	0.932 (0.084)**	0.931 (0.103)**	0.933 (0.088)**
Adjusted real interest rate	– 1.121 (0.215)**	– 1.612 (0.129)**	– 1.115 (0.103)**	– 1.557 (0.195)**
R-square	0.74	0.78	0.75	0.81
DOLS estimates[2]				
Financial development				
Private credit	0.266 (0.095)**	0.240 (0.088)**	0.249 (0.100)**	0.218 (0.091)**
Stock market capitalisation[1]	0.029 (0.018)*	0.036 (0.016)**		
Total value traded			0.015 (0.009)*	0.020 (0.008)**
Gross domestic product	1.005 (0.108)**	1.006 (0.092)**	1.004 (0.115)**	1.005 (0.096)**
Adjusted real interest rate	– 0.594 (0.241)**	– 0.923 (0.211)**	– 0.511 (0.251)**	– 0.866 (0.214)**
R-square	0.71	0.74	0.71	0.74

Note: ** significant at 5 per cent level. * significant at 10 per cent level.
1. When measuring financial development by stock market and total value traded, data availability required us to reduce the sample size to 16 countries.
2. The results shown here are those for our preferred lead-lag combination, which are the same for the four different specifications for each financial development measure. They are one lead and no lag (1,0) for the model without financial development, (1,0) for liquid liabilities, (2,0) for private credit and (0,2) for stock market capitalisation and total value traded.

Appendix table 1: Panel Estimates of investment equations with alternative user cost of capital variable (data from 1970 to 1999)

	Standard model	Models including financial development indicators			
	Without financial development	Liquid liabilities	Private credit	Stock market capitalisation[2]	Total Value Traded[2]
OLS estimates					
Financial development[1]		0.165 (0.034)**	0.092 (0.021)**	0.059 (0.012)**	0.023 (0.005)**
Gross domestic product	1.188 (0.025)**	1.115 (0.029)**	1.097 (0.032)**	1.316 (0.048)**	1.332 (0.048)**
User cost of capital[3]	1.159 (0.229)**	1.177 (0.225)**	0.955 (0.230)**	0.351 (0.414)	0.770 (0.506)
Memorandum item:					
R-square	0.80	0.81	0.81	0.89	0.89
Bias-corrected OLS estimates					
Financial development[1]		0.225 (0.067)**	0.118 (0.044)**	0.058 (0.016)**	0.019 (0.008)**
Gross domestic product	1.131 (0.072)**	1.033 (0.073)**	1.013 (0.079)**	1.235 (0.078)**	1.274 (0.083)**
User cost of capital	1.176 (0.456)**	1.210 (0.437)**	0.934 (0.445)**	0.752 (0.639)	1.223 (0.781)
Memorandum item:					
R-square	0.80	0.81	0.81	0.89	0.89
FMOLS estimates					
Financial development[1]		0.370 (0.069)**	0.208 (0.045)**	0.044 (0.016)**	0.039 (0.008)**
Gross domestic product	0.931 (0.074)**	0.936 (0.076)**	0.934 (0.082)**	0.934 (0.078)**	0.936 (0.083)**
User cost of capital	– 1.068 (0.472)**	– 0.930 (0.452)**	– 0.776 (0.460)*	– 1.061 (0.639)*	– 1.125 (0.677)*
Memorandum item:					
R-square	0.72	0.76	0.78	0.79	0.83
DOLS estimates[4]					
Financial development[1]		0.180 (0.072)**	0.141 (0.048)**	0.068 (0.017)**	0.043 (0.009)**
Gross domestic product	1.004 (0.077)**	0.994 (0.078)**	0.998 (0.088)**	1.008 (0.085)**	1.007 (0.091)**
User cost of capital	– 0.078 (0.489)	– 0.267 (0.468)	– 0.182 (0.494)	0.205 (0.759)	0.041 (0.765)
Memorandum item:					
R-square	0.52	0.59	0.69	0.66	0.71

Note: ** significant at 5 per cent level. * significant at 10 per cent level.

1. Liquid liabilities, private credit, stock market capitalisation and total value traded are alternatively included as financial development measure.
2. When measuring financial development by stock market and total value traded data availability required reducing the sample size to 16 countries.
3. The user cost of capital is measured by log(1+ ucc / pib), where ucc are the user costs of capital and pib are the investment prices deflator.
4. The results shown here are those for our preferred lead-lag combination. They are one lead and no lag (1,0) for the model without financial development, (1,0) for liquid liabilities, (2,0) for private credit and (0,2) for stock market capitalisation and total value traded.

Appendix table 2: Robustness of DOLS estimates
(data from 1970 to 1999, using real adjusted interest as user cost of capital proxy)

	Benchmark model	Financial development variables			
	Without financial development	Liquid liabilities	Private credit deposits	Stock market capitalisation[2]	Total Value Traded[2]
Financial development[1]		(0.176;0.213)	(0.165;0.178)	(0.044;0.066)	(0.035;0.047)
Gross domestic product	(0.998;1.004)	(0.987;0.995)	(0.989;1.001)	(1.001;1.009)	(1.002;1.008)
Real adjusted interest rate	(0.074;0.817)	(0.080;0.957)	(– 0.505;0.295)	(– 1.958;– 0.736)	(– 1.324;– 0.507)

1. Financial development is alternatively proxied by liquid liabilities, private credit, stock market capitalisation and total value traded.
2. The maximum number of lags and leads was set to 3.

Appendix table 3. Panel estimates of investment equations (data from 1970 to 1995)

	Benchmark model	Inclusion of financial development variables			
	Without financial development	Liquid liabilities	Private credit	Stock market capitalisation[2]	Total Value Traded[2]
OLS estimates					
Financial development[1]		0.117 (0.042)**	0.087 (0.024)**	0.039 (0.015)**	0.020 (0.006)**
Gross domestic product	1.174 (0.037)**	1.122 (0.041)**	1.094 (0.043)**	1.476 (0.060)**	1.416 (0.055)**
Adjusted real interest rate	– 0.174 (0.189)	– 0.164 (0.189)	– 0.267 (0.190)	– 0.408 (0.189)**	– 0.468 (0.184)**
Memorandum item:					
R-square	0.74	0.74	0.75	0.84	0.84
Bias-corrected OLS estimates					
Financial development[1]		0.168 (0.072)**	0.120 (0.046)**	0.037 (0.017)**	0.012 (0.008)
Gross domestic product	1.125 (0.078)**	1.048 (0.082)**	1.013 (0.087)**	1.416 (0.086)**	1.391 (0.092)**
Adjusted real interest rate	– 0.187 (0.251)**	– 0.169 (0.247)	– 0.315 (0.246)	– 0.410 (0.220)*	– 0.515 (0.232)**
Memorandum item:					
R-square	0.74	0.74	0.75	0.84	0.84
FMOLS estimates					
Financial development[1]		0.398 (0.075)**	0.273 (0.048)**	0.036 (0.018)**	0.030 (0.009)**
Gross domestic product	0.925 (0.081)**	0.931 (0.085)**	0.932 (0.090)**	0.932 (0.090)**	0.931 (0.097)**
Adjusted real interest rate	– 0.180 (0.262)	– 0.107 (0.257)	– 0.980 (0.256)**	– 1.441 (0.233)**	– 1.513 (0.245)**
Memorandum item:					
R-square	0.71	0.72	0.71	0.67	0.71
DOLS estimates[3]					
Financial development[1]		0.169 (0.078)**	0.178 (0.051)**	0.058 (0.020)**	0.037 (0.010)**
Gross domestic product	1.005 (0.085)**	0.995 (0.090)**	0.996 (0.099)**	1.004 (0.103)**	1.004 (0.106)**
Adjusted real interest rate	0.126 (0.273)	0.107 (0.268)	– 0.446 (0.271)*	– 0.975 (0.262)**	– 0.709 (0.271)**
Memorandum item:					
R-square	0.51	0.59	0.68	0.71	0.73

Note: ** significant at 5 per cent level. * significant at 10 per cent level.
1. Financial development is alternatively proxied by liquid liabilities, private credit, stock market capitalisation and total value traded.
2. When measuring financial development by stock market and total value traded, data availability required us to reduce the sample size to 16 countries.
3. The results shown here are those for our preferred lead-lag combination. They are one lead and no lag (1,0) for the model without financial development, (1,0) for liquid liabilities, (2,0) for private credit and (0,2) for stock market capitalisation and total value traded.

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