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# Financial Development, Financial Fragility, and Growth

# FINANCIAL DEVELOPMENT, FINANCIAL FRAGILITY, AND GROWTH[1]

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ABSTRACT

This paper attempts to reconcile the apparent contradiction between two strands of the literature on the effects of financial intermediation on economic activity. On the one hand, the empirical growth literature finds a positive effect of financial depth as measured by, for instance, private domestic credit and liquid liabilities (e.g., Levine, Loayza, and Beck 2000). On the other hand, the banking and currency crisis literature finds that monetary aggregates, such as domestic credit, are among the best predictors for crises (e.g., Kaminski and Reinhart 1999) and their related economic downturns. This paper starts by illustrating these opposing effects by, first, analyzing the dynamics of output growth and financial intermediation around systemic banking crises and, second, showing that the growth enhancing effects of financial depth are weaker in countries that experienced such crises. After these illustrative exercises, the paper attempts an empirical explanation of the apparently opposing effects of financial intermediation. This explanation is based on a distinction between short-run (transitory) and long-run (trend) changes and effects of domestic credit aggregates. Working with a panel of cross-country and time-series observations, the paper estimates an encompassing model of long- and short-run effects, following, among others, Pesaran, Shin, and Smith (1999)'s Pooled Mean Group Estimator. The main result of the paper is that a positive long-run relationship between financial intermediation and output growth co-exists with a, mostly, negative short-run relationship.

[1] This paper was written while Ranciere worked in the research department of the Central Bank of Chile. We are grateful for the comments and advice from J. Benhabib, J. Cummins, F. Gallego, P. Gourinchas, K. Schmidt-Hebbel, and R. Soto.

# FINANCIAL DEVELOPMENT, FINANCIAL FRAGILITY AND GROWTH

## I. INTRODUCTION

This paper attempts to reconcile the apparent contradiction between two strands of the literature on the effects of financial intermediation on economic activity. On the one hand, the empirical growth literature finds a positive effect of measures of private domestic credit and liquid liabilities on per capita GDP growth. This is interpreted as the growth enhancing effect of financial development (e.g., King and Levine 1993; Levine, Loayza, and Beck 2000). On the other hand, the banking and currency crisis literature finds that monetary aggregates, such as domestic credit, are among the best predictors for crises (e.g., Demirguc-Kunt and Degatriache 1997, 1999; Gourinchas, Landerretche, and Valdes 1999; Kaminski and Reinhart 1999). Since banking crises usually lead to recessions, an expansion of domestic credit would then be associated to growth slowdowns.

A similar divide exists at the theoretical level. According to the endogenous growth literature, financial deepening leads to a more efficient allocation of savings to productive investment projects (see Greenwood-Jovanovic 1990; Bencivenga-Smith 1991). Conversely, the financial crisis literature points to the destabilizing effect of financial liberalization as it leads to overlending. Overlending would occur through a combination of channels, including a limited monitoring capacity of regulatory agencies, the inability of banks to discriminate good projects during investment booms, and the existence of an explicit or implicit insurance against banking failures (Shneider-Tornell, 2000; Aghion, Bacchetta and Banerjee 1999). Not surprisingly, each strand of the literature has produced its own set of policy implications. Thus, researchers that emphasize the findings of the endogenous growth literature advocate financial liberalization and deepening (e.g., Roubini, Sala-y-Martin 1992), while those that concentrate on crises caution against "excesive" financial liberalization (e.g., Balino and Sundarajan 1991; Gavin and Hausman 1995).

This paper seeks to contribute to the debate from an empirical perspective. In section II we examine how the relationship between measures of financial depth and economic growth is affected by the presence of financial crises. For this purpose, we first describe the behavior of financial intermediation and output growth around episodes of banking crises. We then reconsider the evidence on the positive growth effect of financial deepening by analyzing whether this effect is weaker in countries afflicted by financial crises.

In section III the paper attempts an empirical explanation of the apparently contradictory effects of financial intermediation on economic activity. This explanation is based on the distinction between cycle and trend changes of financial intermediation and their corresponding effects on output growth. Working with a panel of cross-country and time-series observations, we estimate an encompassing model of long- and short-run effects. Section IV concludes.

## II. The Relationship between Financial Depth and Growth in the Presence of Financial Crises

In this section we examine how the relationship between measures of financial depth and economic growth is affected by the presence of financial crises. First, we describe the behavior over time of financial intermediation and output growth around banking crises. We do it by using an event-study methodology applied to a panel of countries that have experienced such crises, as identified by Caprio and Klingbiel (1999). Second, we revisit the evidence on the positive growth effect of financial deepening by testing whether this effect is weaker in countries that have experienced banking crises. For this purpose, we follow the GMM cross-country panel-data approach to growth empirics in Levine, Loayza and Beck (2000).[2]

### A. The behavior of financial intermediation and economic activity around episodes of financial crises

Here we describe the behavior of financial intermediation and economic activity in a typical country before and after the start of a banking crisis. We use total liquid liabilities and domestic credit to the private sector, both as ratios to GDP, as the measures of financial intermediation. Economic activity is measured with total and per capita GDP growth rates. We first identify the episodes of banking crises for a large sample of

[2] See also Caselli, Esquivel, and Lefort (1996), Easterly, Loayza, and Montiel (1997), and Beck, Levine, and Loayza (2000).

countries following the Caprio and Klingbiel (1999)'s classification. Then, applying an event-study methodology, we make country experiences comparable by re-scaling calendar time into crisis-centered time for each country. Moreover, to eliminate country-specific effects, we demean each observation with the corresponding country mean. We focus the analysis on the 12-year widow centered on the start of the banking crisis. Figure 1 presents the behavior of the typical country-year observation, which is given by the median across countries in a particular year for each measure of financial intermediation and output growth. Table 1 presents Students' t-tests for the significance of level and correlation changes over the 12-year period.

Both liquid liabilities and private credit rise rapidly before the crisis then drop drastically once it starts. They recover partially in the following years but remain far below their pre-crisis levels. On the other hand, total and per capita GDP growth rates fall in the years prior to the banking crisis, reach the bottom at the onset of the crisis, and recover gradually afterwards. The correlation between the measures of financial intermediation and economic activity depend on the starting and ending year. Prior to the crisis, the correlation between growth and financial intermediation is strongly negative, while it is mildly negative or zero in the aftermath of the crisis. In summary, this first exercise shows that credit booms do precede banking crisis and that the relationship between financial intermediation and growth is negative before and during banking crises.

**B. Revisiting the evidence on the growth effects of financial deepening**

Working with a large cross-section of countries, King and Levine (1993a, 1993b) find a positive relationship between initial financial intermediation depth and subsequent long-run growth performance. In these and related studies, the long-run growth rate is estimated as the average rate over periods of time as long as 25-30 years. King and Levine use initial measures of financial intermediation (rather than, say, period averages) to be able to conclude that more developed financial systems lead to higher growth. Levine, Loayza, and Beck (2000) address directly the issue of joint endogeneity of financial development through the use of instrumental variables in their growth regressions. They use the countries' legal origin as the "external" instrument for financial depth in their cross-sectional regressions and the lagged observations of all

**Fig 1a: Financial Intermediation**

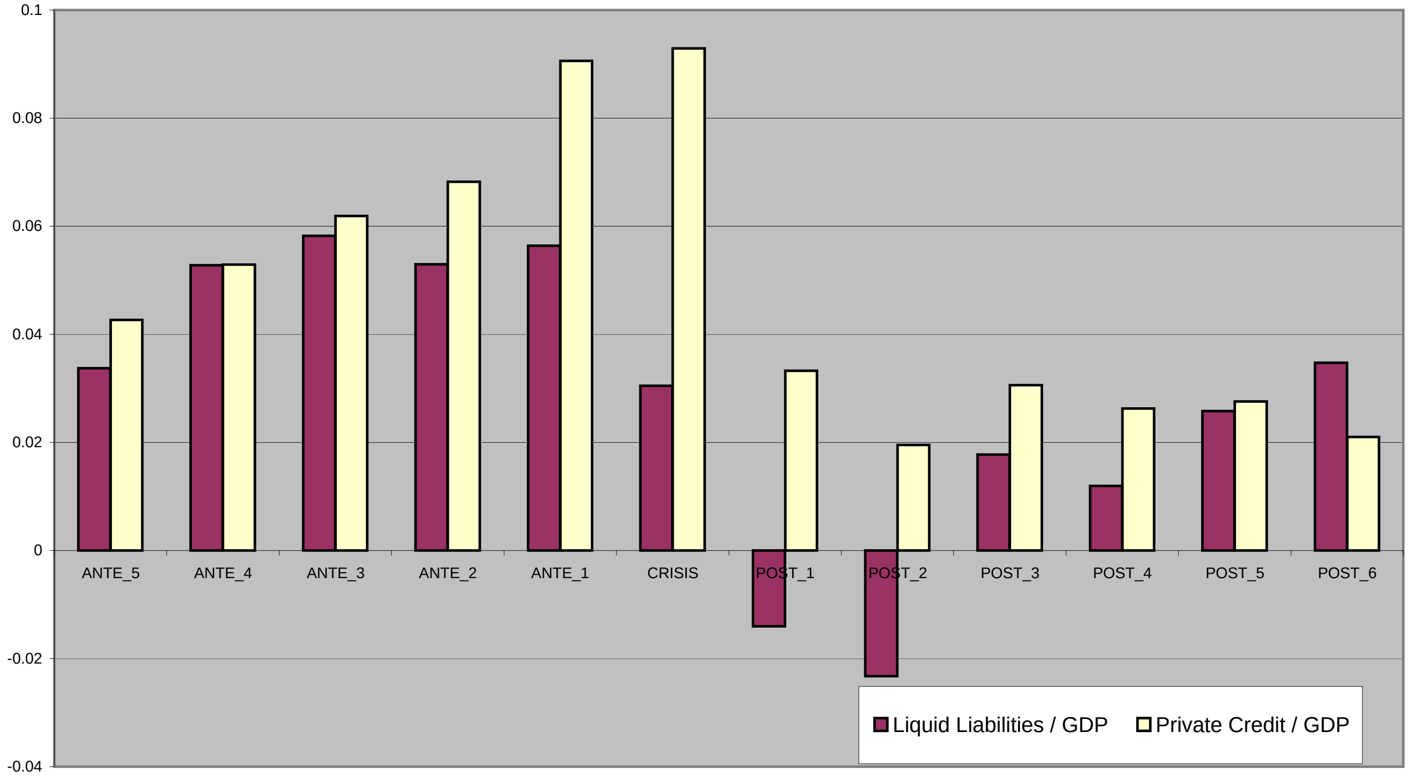
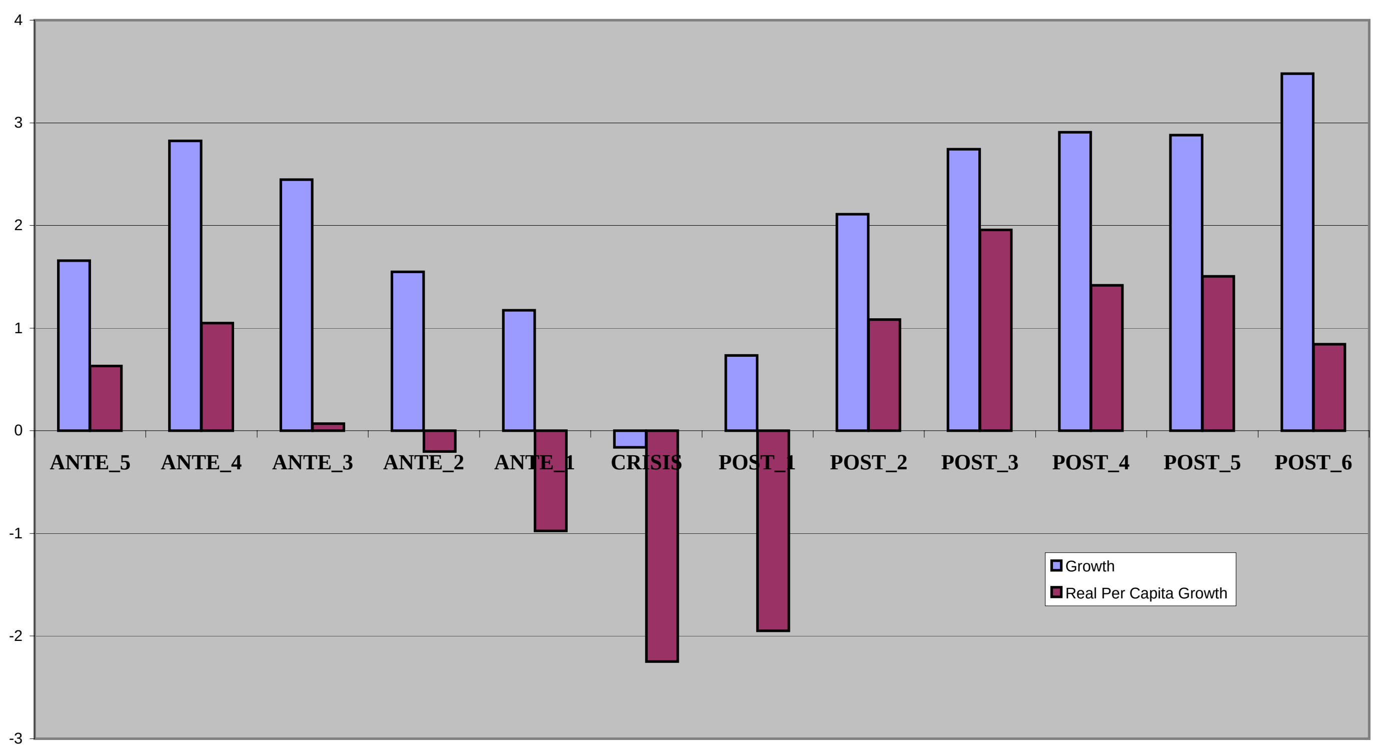



**Fig 1b: Growth**



explanatory variables as "internal" instruments in their pooled (cross-country and time-series) regressions. The data panels used by Levine et al. consist of about 74 countries and, for each of them, non-overlapping five-year averages covering the period 1960-95. They use five-year averages, rather than annual observations, to smooth out transitory or business-cycle fluctuations. Confirming previous results, Levine et al. find robust evidence that financial development and depth lead to an improved growth performance.

It is arguable that in most cases, using low-frequency data (such as averages over five or more years) allows the researcher to concentrate on long-run effects. However, in cases of prolonged or deep recessions, such as those associated with financial crises, even averages over long periods may be contaminated by cycle effects. Developing this argument, DeGregorio and Guidotti (1995) present evidence that while in cross-sectional regressions involving a worldwide sample of countries financial intermediation is positively linked with growth, in panel regressions for only Latin American countries, the relationship is negative. They suggest that their results for Latin America may reflect the lasting impact of the repeated financial crises (and associated overlending) that the region has suffered. However, DeGregorio and Guidotti do not offer direct evidence on the role of financial crises in distorting the financial intermediation and growth relationship. Moreover, it is possible that their contrasting results between the worldwide and Latin American samples are actually due to the use of cross-sectional vs.panel-data estimators.

We now analyze how the presence of financial crises modifies the estimated link between measures of financial intermediation and economic growth. For this purpose, we work with the same data and methodology as in Levine, Loayza, and Beck (2000) but allow for, respectively, a banking-crisis and a Latin America effect.

***Data and Methodology***

We work with a pooled (cross-country, time-series) data set consisting of 74 countries and, for each of them, at most 7 non-overlapping five-year periods over 1960-95. This is an unbalanced panel data set. See Annex 1 for the description of the data and country sample.

We estimate a growth regression using panel data. As standard in the literature, the regression equation is dynamic given that it includes the initial level of per capita output as an explanatory variable. Apart from the measure of financial intermediation, the regression equation considers a set of control variables, including initial per capita output, average secondary school attainment of the adult population, the average ratio of government consumption to GDP, the average inflation rate, and the average black market premium on foreign exchange.  To deal with time effects, the regression considers period-specific dummy variables.  Finally, the regression equation allows for unobserved country-specific effects and the joint endogeneity of the explanatory variables.  The method of estimation is the generalized method of moments (GMM) for dynamic models using panel data, as developed by Arellano and Bond (1991) and Arellano and Bover (1995).

The regression equation to be estimated is the following,

$$y_{i,t} - y_{i,t-1} = (\alpha - 1) y_{i,t-1} + \beta' X_{i,t} + \delta FD_{i,t} + \mu_t + \eta_i + \varepsilon_{i,t}$$

where, $y$ is the logarithm of real per capita output, $X$ is a set of control variables, $FD$ is an indicator of financial depth, $\mu_t$ is a time-specific effect, $\eta_i$ is an unobserved country-specific effect, and $\varepsilon$ is the error term. The subscripts $i,t$ represent country and time-period, respectively.  We assess the banking-crisis and the Latin-America effects by introducing a slope dummy on the financial depth indicator.

We relax the assumption of strong exogeneity of the explanatory variables by allowing them to be correlated with current and previous realizations of the error term $\varepsilon$. However, we assume that future realizations of the error term do not affect current values of the explanatory variables.  Furthermore, we assume that the error term $\varepsilon$ is serially uncorrelated.  We allow the unobserved country-specific effect $\eta_i$ to be correlated with the explanatory variables.  However, following a stationarity condition, we assume that *changes* in the explanatory variables are uncorrelated with the country-specific effect.  As Arellano and Bond (1991) and Arellano and Bover (1995) show, this set of assumptions generate moment conditions that allow estimation of the parameters of interest.  The instruments corresponding to these moment conditions are appropriately lagged values of the levels and differences of the explanatory and dependent variables.  Since typically the moment conditions overidentify the regression model, they also allow for specification testing through a Sargan-type test.

### *Results*

Tables 2 and 3 report the growth regression results. We study how the effect of financial intermediation on growth varies in the presence of financial turmoil by including a slope dummy for countries that have suffered a banking crisis (Table 2). Furthermore, in order to reconsider De Gregorio and Guidotti's findings, we also assess the effect of a slope dummy for Latin American countries (Table 3). In each case, we work with two indicators of financial intermediation, namely, the ratio of liquid liabilities to GDP and the ratio of domestic credit to the private sector to GDP.

The results in Table 2 confirm the positive growth effect of larger financial depth. This effect is significantly positive for the samples of non-crisis and crisis countries. However, as the size and significance of the slope dummy coefficient reveals, the positive growth effect is statistically smaller for crisis than for non-crisis countries. This is true for both indicators of financial intermediation (i.e., liquid liabilities and private domestic credit). In Table 3, we reconsider De Gregorio and Guidotti's results. We agree with them that the growth effect of financial deepening is smaller in Latin American countries than in the rest. However, we find that even for Latin American countries an expansion of financial intermediation, as measured in the frequencies of five-year averages, leads to higher growth rates. Qualitatively, the results obtained with the two types of slope dummies are similar; this is not surprising given that most Latin countries have suffered from banking crisis. In summary, the estimated growth effect of financial deepening is smaller, but still positive, in countries that have faced financial crisis, and particularly those in Latin America.

## III. SHORT- AND LONG-RUN GROWTH EFFECTS OF FINANCIAL DEEPENING

In this section we attempt an empirical explanation of the apparently contradictory effects of financial intermediation on economic activity. This explanation is based on the distinction between cycle and trend changes of financial intermediation and their corresponding effects on output growth. Instead of averaging the data to isolate trend effects, we estimate both long- and short-run effects using annual data in a panel containing a large sample of countries. Our method can be summarized as a panel, error-correction model, where long- and short-run effects are estimated jointly from a general autoregressive distributed-lag (ARDL) model.

We propose this panel error-correction method as an alternative to the traditional method of time averaging for the following reasons. First, while averaging clearly induces a loss of information, it is not obvious that averaging over fixed-length intervals effectively eliminates business-cycle fluctuations. Second, averaging eliminates information that may be used to estimate a more flexible model that allows for some parameter heterogeneity across countries. Third, and most importantly for our purposes, averaging hides the dynamic relationship between financial intermediation and economic activity, particularly the presence of opposite effects at different time frequencies.[3]

### A. Methodology

Empirical estimation poses two issues. The first is the need to separate and estimate short- and long-run effects without being able to decompose directly trend and transitory components of growth, financial intermediation, and the other explanatory variables. We treat this issue below in the context of single-country estimation. The second issue is the likely possibility that the parameters in the relationship between financial intermediation and economic activity be different across countries. It can be argued that country heterogeneity is particularly relevant in short-run relationships, given that countries are affected by overlending and financial crises to widely different degrees. On the other hand, we can expect that long-run relationships would be more homogeneous across countries. We discuss below the issue of heterogeneity in the context of multi-country estimation.

#### *Single-country estimation*

As said above, we face the challenge to estimate long- and short-run relationships without being able to observe the long- and short-run components of the variables involved. Over the last decade or so, a booming cointegration literature has focused on the estimation of long-run relationships among I(1) variables (Johanssen 1995, Phillips and Hansen 1990). From this literature, two common misconceptions have been derived. The first one is that long-run relationships exist *only* in the context of cointegration of

[3] Similar arguments are made by Attanasio, Scorcu, and Picci (2000) in their cross-country study on the dynamic relationship between saving, investment, and growth.

integrated variables. The second one is that standard methods of estimation and inference are incorrect. Pesaran and Smith (1995), Pesaran (1997) and Pesaran and Shin (1999) have argued against both misconceptions, showing how small modifications to standard methods can render consistent and efficient estimates of the parameters in a long-run relationship between both integrated and stationary variables. Furthermore, the methods proposed by Pesaran and co-authors avoid the need for pre-testing and order-of-integration conformability given that they are valid whether or not the variables of interest are I(0) or I(1). The main requirements for the validity of this methodology are that, first, there exist a long-run relationship among the variables of interest and, second, the dynamic specification of the model be augmented such that the regressors are strictly exogenous and the resulting residual is not serially correlated. For reasons that will become apparent shortly, Pesaran and co-authors call their method "an autoregressive distributed lag (ARDL) approach" to long-run modelling.

As an illustration, consider the following simple bivariate model:

$$y_t = a + by_{t-1} + cX_{t-1} + \nu_t \qquad (1)$$

$$X_t = \gamma + \rho X_{t-1} + \varepsilon_t \qquad (2)$$

where $y_t$, the decision variable, is the per capita GDP growth rate in year $t$; and $X$, the forcing variable, represents a set of growth determinants including financial depth and control variables. Furthermore, assume that the residuals (or shocks) have the following distributional properties:

$$\begin{pmatrix} \nu_t \\ \varepsilon_t \end{pmatrix} \; iid\left(0, \Sigma\right), \qquad \Sigma = \begin{pmatrix} \sigma_{\nu\nu} & \sigma_{\nu\varepsilon} \\ \sigma_{\nu\varepsilon} & \sigma_{\varepsilon\varepsilon} \end{pmatrix} \qquad (3)$$

The first point to note is that $X$ does not depend on past values of $y$ (beyond its dependence on previous values of $X$). If a more general process for $X$ were allowed, the long-run relationship between the two variables would not be unique. That is, both variables would be endogenous and additional identification assumptions would be needed to discern between various long-run relationships.[4] Since multiple long-run relationships are beyond the scope of this paper, we restrict the dynamic process for $X$ to be purely autoregressive.

The second point to note is that the existence of a long-run relationship requires the process for *y* to be stable, which in this simple example entails that $|b|<1$. Notice that once we have restricted the process of *X* to be purely autoregressive, the existence of a long-run relationship does not rely on whether *X* is I (0) or I(1); that is, there is no restriction on whether $\rho=1$. Pesaran, Shin, and Smith (2000) present a test for the null hypothesis that there is no long-run relationship when it is not known *a priori* whether *X* is I(0) or I(1). The test consists on examining the null that $b=1$ against the alternative that $|b|<1$.

In order to be able to derive the long-run relationship between *y* and *X*, we must obtain a dynamic regression equation in which, first, the regression residual is serially uncorrelated and, second, the regressors, *X*, are *strictly* exogenous (that is, independent of the residuals at all leads and lags.) Given the assumptions on the distributional properties of the residuals $\nu$ and $\varepsilon$ (equation 3), the requisite that the residuals be serially uncorrelated is met in our simple example. If this were not the case, we would need to augment the lag order in (1) and (2) until the residuals become serially independent (Pesaran and Shin 1999). The second pre-requisite to derive a long-run relationship is, however, not met in our simple example –*X* is not *strictly* exogenous given that the non-zero correlation between the shocks entails a contemporaneous feedback between *y* and *X*. As explained by Pesaran and Shin (1999), the way to control for this contemporaneous feedback is also to augment the dynamic specification in (6). The purpose of augmenting the regression equation is to replace the (correlated) residual $\nu$ with a linear predictor based on leads and lags of *X* and a new residual that by construction is independent of *X*. In our simple example, we model the contemporaneous correlation between $\nu_t$ and $\varepsilon_t$ by a linear regression of $\nu_t$ on $\varepsilon_t$ as follows,

$$\nu_t = \left(\frac{\sigma_{\nu\varepsilon}}{\sigma_{\varepsilon\varepsilon}}\right)\varepsilon_t + \eta_t \qquad (4)$$

where $(\sigma_{\nu\varepsilon}/\sigma_{\varepsilon\varepsilon})$ represents the population coefficient of the regression, and $\eta_t$ is distributed independently from $\varepsilon_t$.

[4] See Hsiao (1997) and Pesaran and Shin (1999).

Substitute the above expression for $v_t$ into equation (1). Then, using the AR model for $X$, express $\varepsilon_t$ in terms of $X_t$ and $X_{t-1}$. The ensuing regression equation is an auto-regressive distributed lag model (ARDL) for $y$ from which a long-run relationship can be derived. The resulting ARDL (1,1) for $y$ is given by,

$$y_t = \left(a - \gamma \frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}\right) + b y_{t-1} + \left(\frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}\right) X_t + \left(c - \rho \frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}\right) X_{t-1} + \eta_t \tag{5}$$

Note that the original process for $y$ (equation 1) is now augmented by the inclusion of the additional regressor $X_t$.

The error-correction model (ECM) implied by the ARDL (1,1) given above can be expressed as,

$$\Delta y_t = -(1-b)\left[ y_{t-1} - \left(\frac{a - \gamma \frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}}{1-b}\right) - \left(\frac{c + \frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}(1-\rho)}{1-b}\right) X_{t-1} \right] + \left(\frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}\right) \Delta X_t + \eta_t \tag{6}$$

where the expression in brackets is the error-correction term and *(1-b)* is the speed of adjustment.

Therefore, the long-run (steady-state) relationship implied by the dynamic system in equations (1)-(4) is given by,

$$y^* = \left(\frac{a - \gamma \frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}}{1-b}\right) + \left(\frac{c + \frac{\sigma_{v\varepsilon}}{\sigma_{\varepsilon\varepsilon}}(1-\rho)}{1-b}\right) X^* + \eta^* \tag{7}$$

or, $y^* = \alpha + \beta x^* + \eta^*$.

The presentation of this simple empirical model serves to highlight the assumptions and properties of the ARDL method proposed by Pesaran and Smith (1995), Pesaran (1997), and Pesaran and Shin (1999) for the estimation of a long-run relationship. The advantage of the method is that standard estimation and inference can be used regardless of whether the regressors are stationary or integrated. The main assumption is that there exist a single long-run relationship between the endogenous and forcing variables.[5] The pre-requisites for consistent and efficient estimation are that the shocks in the dynamic specification be serially uncorrelated and that the forcing variables be strictly exogenous. As we illustrated, the pre-requisites can be met by augmenting sufficiently the lag order of the dynamic regression equation. The resulting equation will generally be an ARDL(p, q) model of sufficiently large lag order.

***Multi-country estimation***

Our empirical samples below are characterized by time-series (T) and cross-section (N) dimensions of relatively large size. In such conditions, there are a number of alternative methods for multi-country estimation, which allow for different degrees of parameter heterogeneity across countries. At one extreme, the fully heterogeneous-coefficient model imposes no cross-country parameter restrictions and can be estimated on a country-by-country basis -- provided the time-series dimension of the data is sufficiently large. When, in addition, the cross-country dimension is large, the mean of long- and short-run coefficients across countries can be estimated consistently by the unweighted average of the individual country coefficients. This is the “mean group” (MG) estimator introduced by Pesaran, Smith, and Im (1996). At the other extreme, the fully homogeneous-coefficient model requires that all slope and intercept coefficients be equal across countries. This is the simple “pooled” estimator.

In between the two extremes, there are a variety of estimators. The “dynamic fixed effects” estimator restricts all slope coefficients to be equal across countries but allows for different country intercepts. The “pooled mean group” (PMG) estimator, introduced by Pesaran, Shin, and Smith (1999), restricts the long-run coefficients to be the same across countries but allows the short-run coefficients (including the speed of adjustment) to be country specific. The PMG estimator also generates consistent estimates of the mean of short-run coefficients across countries by taking the unweighted average of the individual country coefficients (provided that the cross-sectional dimension is large).

[5] It is worth noting that this assumption underlies implicitly the various single-equation based estimators of long-run relationships commonly found in the cointegration literature. Without such assumption, these estimators would at best identify some linear combination of all the long-run relationships present in the data.

The choice among these estimators faces a general trade-off between consistency and efficiency. Estimators that impose cross-country constraints dominate the heterogeneous estimators in terms of efficiency if the restrictions are valid. If they are false, however, the restricted estimators are inconsistent. In particular, imposing invalid parameter homogeneity in dynamic models typically leads to downward-biased estimates of the speed of adjustment (Robertson and Symons 1992, Pesaran and Smith 1995).

For our purposes, the pooled mean group estimator offers the best available compromise in the search for consistency and efficiency. This estimator is particularly useful when the long run is given by conditions expected to be homogeneous across countries while the short-run adjustment depends on country characteristics such as financial development, institutional quality, and relative price flexibility. Furthermore, the PMG estimator is sufficiently flexible to allow for long-run coefficient homogeneity over only a subset of variables and/or countries.

In view of these considerations, we use the PMG method to estimate a long-run relationship that is common across countries while allowing for unrestricted country heterogeneity in the adjustment dynamics. The interested reader is referred to Pesaran, Shin, and Smith (1999) where the PMG estimator is developed and compared with the MG estimator. Briefly, the PMG estimator proceeds as follows. The estimation of the long-run coefficients is done jointly across countries through a (concentrated) maximum likelihood procedure. Then the estimation of short-run coefficients (including the speed of adjustment), country-specific intercepts, and country-specific error variances is done on a country-by-country basis, also through maximum likelihood and using the estimates of the long-run coefficients previously obtained.[6]

[6] The comparison of the asymptotic properties of PMG and MG estimates can be put also in terms of the general trade-off between consistency and efficiency noted in the text. If the long-run coefficients are in fact equal across countries, then the PMG estimates will be consistent and efficient, whereas the MG estimates will only be consistent. If, on the other hand, the long-run coefficients are not equal across countries, then the PMG estimates will be inconsistent, whereas the MG estimator will still provide a consistent estimate of the mean of long-run coefficients across countries. The long-run homogeneity restrictions can be tested using Hausman or likelihood ratio tests to compare the PMG and MG estimates of the long run coefficients. In turn, comparison of the small sample properties of these estimators relies on their sensitivity to outliers. In small samples (low T and N), the MG estimator, being an unweighted average, is excessively sensitive to the inclusion of outlying country estimates (for instance those obtained with small T). The PMG estimator performs better in this regard because it produces estimates that are similar to *weighted* averages of the respective country-specific estimates, where the weights are given according to their precision (that is, the inverse of their corresponding variance-covariance matrix).

An important assumption for the consistency of our PMG estimates is the independence of the regression residuals across countries. In practice, non-zero error covariances usually arise from *omitted* common factors that influence the countries' ARDL processes. We seek to eliminate these common factors and, thus, ensure the independence condition by allowing for time-specific effects in the estimated regression; this is equivalent to a regression in which each variable enters as deviations with respect to the cross-sectional mean in a particular year.

**B. Data and Results**

The sample consists of 48 countries with annual data for the period 1960-97. Given the procedure's requirements on the time-series dimension of the data, we include only countries that have at least 20 consecutive observations. The measures of financial depth are liquid liabilities and private domestic credit, both as ratios to GDP. The control variables are government consumption, the inflation rate, and the volume of trade.

As outlined in the previous section, the consistency and efficiency of the PMG estimates relies on several specification conditions. The first are that the regression residuals be serially uncorrelated and that the explanatory variables can be treated as exogenous. We seek to fulfill these conditions by including at least one lag of each explanatory variable and as many as 4 lags of the growth rate, the dependent variable. The second are that both country-specific effects and cross-country common factors be eliminated. We do this by including, respectively, country-specific and year-specific intercepts. The third condition refers to the existence of a long-run relationship (dynamic stability) and requires that the coefficient on the error-correction term be negative. We report the estimates for this coefficient and its corresponding standard error, and we find evidence supporting the dynamic stability of the model. The fourth condition is that the long-run parameters be the same across countries. We present Hausman tests which allow us to conclude that the homogeneity of long-run parameters cannot be rejected.

Tables 4 and 5 present the estimation of long- and short-run parameters linking per capita GDP growth, financial intermediation, and other growth determinants. In Table 4 the measure of financial intermediation is the ratio of liquid liabilities to GDP, while in Table 5 it is the ratio of private domestic credit to GDP. In both tables, we present the results obtained using the pooled mean group (PMG) estimator, which we prefer given its gains in consistency and efficiency over other panel error-correction estimators. For comparison purposes, we also present the results obtained with the mean group (MG) and the dynamic fixed-effects (DFE) estimators.

Focusing on the PMG estimator, the main results are the following. In the long-run, per capita GDP growth is positively and significantly related with financial depth. However, in the short-run the average relationship between them is negative. Thus, the sign of the relationship depends on whether their movements are cyclical or permanent. This is true whether the measure of financial depth is liquid liabilities or private domestic credit.

## IV. CONCLUSIONS

The results in this paper can be summarized as follows.

- The dynamic relationship between economic growth and financial intermediation is negative around financial crises. Furthermore, the positive link between "long-run" economic growth and financial deepening is smaller in countries that have suffered banking crises than in the rest.
- Using recent econometric methods for the estimation of dynamic models using panel data, we find that a positive long-run relationship between financial intermediation and output growth co-exists with a, mostly, negative short-run relationship. We propose this result as an empirical explanation for the apparent contradiction between the crisis literature and the endogenous-growth literature on the effects of financial deepening.

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Table 1 Descriptive Statistics for Countries with Crisis Experience

| | ANTE CRISIS PERIOD | CRISIS PERIOD | T-test P-Value |
|---|---|---|---|
| | t-5 to the starting year of crisis, t | t+1 to t+6 | Ho: ante=crisis |
| Liquid Liabilities /GDP | 0.047578843 | 0.007509945 | 0.07 |
| *OBS* | *48* | *50* | |
| Private Credit/ GDP | 0.066891752 | 0.027435856 | 0.06 |
| *OBS* | *48* | *49* | |
| Real Per Capita Growth | -0.269641648 | 0.780450416 | 0.0157 |
| *OBS* | *56* | *53* | |
| Correlation (Liquid Liabilities, Growth) | -0.1072 | -0.1208 | 0.35 |
| *OBS* | *42* | *40* | |
| Correlation (Private Credit, Growth) | -0.347 | -0.18 | 0.07 |
| *OBS* | *42* | *41* | |

TABLE 2: Financial Intermediation, Crisis Experience and Growth; system estimator

| Regressors | Coefficient | Std Error | Coefficient | Std Error |
|---|---|---|---|---|
| Constant | 0.751883 | *1.0316* | 3.06879 | *0.9624* |
| Log of Initial Income per Capita | -0.204635 | *0.1096* | 0.10722 | *0.1226* |
| Average year of secondary schooling | 0.477162 | *0.1463* | 0.14471 | *0.1519* |
| Liquid Liabilities | 2.086862 | *0.1837* | | |
| Liquid Liabilities*Crisis Experience | -0.379457 | *0.0414* | | |
| Private Credit | | | 1.43412 | *0.0634* |
| Private Credit*Crisis Experience | | | -0.26059 | *0.0411* |
| Government size | -1.187689 | *0.2865* | -1.90475 | *0.2665* |
| Inflation Rate | 0.325441 | *0.3941* | -0.39897 | *0.3056* |
| Black Market Premium | -1.980017 | *0.09* | -1.18752 | *0.0859* |
| Dummy 71-75 | -0.833267 | *0.08* | -0.98195 | *0.0642* |
| Dummy 76-80 | -0.882677 | *0.1251* | -0.96971 | *0.1158* |
| Dummy 81-85 | -3.043068 | *0.1322* | -2.96185 | *0.1672* |
| Dummy 86-90 | -2.074279 | *0.1594* | -2.01945 | *0.1674* |
| Dummy 91-95 | -2.867901 | *0.1776* | -2.77716 | *0.1637* |
| Sargan Test (P-value) | *0.467* | | *0.41* | |
| 2nd Order Serial Correlation (P-Value) | *0.836* | | *0.642* | |
| Number of Countries | 74 | | 74 | |
| Number of Obsevations | 359 | | 359 | |

TABLE 3: Financial Intermediation,Latin America and Growth; system estimator

| Regressors | Coefficient | Std Error | Coefficient | Std Error |
|---|---|---|---|---|
| Constant | 2.074185 | *0.9213* | 5.379823 | *0.9257* |
| Log of Initial Income per Capita | -0.181326 | *0.0955* | -0.036462 | *0.1106* |
| Average year of secondary schooling | 0.592854 | *0.1141* | 0.434511 | *0.1289* |
| Liquid Liabilities | 2.098478 | 0.1586 | | |
| Liquid Liabilities*Latin America | -0.203884 | 0.0498 | | |
| Private Credit | | | 1.557448 | *0.073* |
| Private Credit*Latin America | | | -0.199361 | *0.053* |
| Government size | -1.946623 | *0.1978* | -2.665188 | *0.2506* |
| Inflation Rate | 0.363155 | *0.357* | -0.287723 | *0.2191* |
| Black Market Premium | -1.741312 | *0.0957* | -1.111259 | *0.0933* |
| Dummy 71-75 | -0.923225 | *0.0941* | -1.03786 | *0.129* |
| Dummy 76-80 | -1.070274 | *0.1002* | -1.146228 | *0.1307* |
| Dummy 81-85 | -3.103926 | *0.1268* | -3.131746 | *0.19* |
| Dummy 86-90 | -2.271343 | *0.1176* | -2.261626 | *0.1375* |
| Dummy 91-95 | -3.18211 | *0.1357* | -3.154942 | *0.1465* |
| Sargan Test (P-value) | *0.467* | | *0.461* | |
| 2nd Order Serial Correlation (P-Value) | *0.836* | | *0.655* | |
| Number of Countries | 74 | | 74 | |
| Number of Obsevations | 359 | | 359 | |

**Table 4: ARDL(3,2,1,1,1,1); Dependant Variable: Growth; Financial Indicator: Private Credit/GDP**

*Pooled Mean Group, Mean Group estimators and Dynamic Fixed Effect, controlling for country and time effects*

*Sample: All Countries 1961-1997*

| | **Pooled Mean Group** | | **Mean Group** | | ***Hausman Tests*** | | **Dynamic Fixed Effect** | |
|---|---|---|---|---|---|---|---|---|
| Variabels | Coef. | St.Er. | Coef. | St.Er. | h-test | p-val | Coef. | St.Er. |
| **Long-Run Coefficients** | | | | | | | | |
| PC | 0.604 | 0.286 | -22.88 | 15.34 | 2.34 | 0.27 | 0.642 | 0.858 |
| GDP(-5) | -9.421 | 0.5 | -4.5 | 4.68 | 1.07 | 0.79 | -3.453 | 1.397 |
| GOV | -0.69 | 0.431 | -3.04 | 4.342 | 0.3 | 0.59 | -2.465 | 0.978 |
| TRAD | 1.388 | 0.417 | 3.19 | 1.643 | 1.28 | 0.26 | 3.848 | 1.116 |
| INF | -9.289 | 1.041 | -9.213 | 5.152 | 0 | 0.99 | -3.121 | 0.492 |
| **Error Correction Coefficients** | | | | | | | | |
| Phi | -0.875 | 0.095 | -1.956 | 0.178 | | | -0.857 | 0.074 |
| **Short-Run Coefficients** | | | | | | | | |
| dGR(-1) | 0.047 | 0.064 | 0.615 | 0.128 | | | -0.081 | 0.041 |
| dGR(-2) | -0.062 | 0.042 | 0.199 | 0.071 | | | -0.059 | 0.039 |
| dPC | -27.27 | 14.134 | -4.183 | 10.06 | | | -3.713 | 1.723 |
| dPC(-1) | 9.1 | 8.63 | 16.897 | 15.444 | | | -1.496 | 1.447 |
| dGDP(-5) | -8.438 | 2.372 | -2.212 | 4.4523 | | | -5.171 | 2.89 |
| dGOV | -14.503 | 2.526 | -8.685 | 3.134 | | | -1.916 | 1.11 |
| dTRAD | -0.209 | 1.728 | -4.933 | 2.321 | | | -2.167 | 1.257 |
| dINF | -2.438 | 4.275 | 8.649 | 5.114 | | | -2.48 | 0.079 |
| Inpt | 67.35 | 7.392 | 176.04 | 63.559 | | | | |
| No. Countries | 48 | | 48 | | | | 48 | |
| No.Observations | 1211 | | 1211 | | | | 1211 | |
| Avg RBarSq | 0.65 | | 0.68 | | | | 0.44 | |

**Table 5: ARDL(3,1,1,1,1,1); Dependant Variable: Growth; Financial Indicator: Liquid Liabilities/GDP**

*Pooled Mean Group, Mean Group estimators and Dynamic Fixed Effect, controlling for country and time effects*

*Sample: All Countries 1961-1997*

| | **Pooled Mean Group** | | **Mean Group** | | ***Hausman Tests*** | | **Dynamic Fixed Effect** | |
|---|---|---|---|---|---|---|---|---|
| Variabels | Coef. | St.Er. | Coef. | St.Er. | h-test | p-val | Coef. | St.Er. |
| **Long-Run Coefficients** | | | | | | | | |
| LLY | 2.124 | 0.453 | -30.162 | 48.46 | 0.45 | 0.5 | 1.223 | 1.6725 |
| GDP(-5) | -8.616 | 0.513 | 62.848 | 64.449 | 1.23 | 0.27 | -3.43 | 1.48 |
| GOV | -0.584 | 0.469 | -24.21 | 18.569 | 1.62 | 0.2 | -2.28 | 1.08 |
| TRAD | 2.716 | 0.379 | 13.99 | 14.012 | 0.65 | 0.42 | 3.859 | 1.103 |
| INF | -10.177 | 0.986 | -10.745 | 25.692 | 0 | 0.98 | -3.173 | 0.485 |
| **Error Correction Coefficients** | | | | | | | | |
| Phi | -0.889 | 0.092 | -1.727 | 0.124 | | | -0.849 | 0.07 |
| **Short-Run Coefficients** | | | | | | | | |
| dGR(-1) | 0.079 | 0.055 | 0.453 | 0.087 | | | 0.076 | 0.038 |
| dGR(-2) | -0.051 | 0.038 | 0.116 | 0.051 | | | -0.057 | 0.038 |
| dLLY | -22.22 | 7.34 | -17.467 | 7.911 | | | -11.781 | 3.147 |
| dGDP(-5) | -7.193 | 2.576 | -6.117 | 3.975 | | | -4.714 | 2.194 |
| dGOV | -12.889 | 2.706 | -7.707 | 2.842 | | | -1.83 | 1.044 |
| dTRAD | -1.827 | 1.805 | -6.494 | 2.324 | | | -2.219 | 1.259 |
| dINF | -1.077 | 3.195 | 10.962 | 3.905 | | | -2.512 | 0.814 |
| Inpt | 57.162 | 5.924 | 68.55 | 36.584 | | | | |
| No. Countries | 49 | | 49 | | | | 49 | |
| No.Observations | 1235 | | 1235 | | | | 1235 | |
| Avg RBarSq | 0.61 | | 0.65 | | | | 0.44 | |

APPENDIX A1: LIST OF SYTEMIC BANKING CRISES*

| Country Name | Start | End | Start | End | Start | End |
|---|---|---|---|---|---|---|
| Argentina | 1980 | 1982 | 1989 | 1990 | 1995 | 1995 |
| Benin | 1988 | 1990 | | | | |
| Burkina Faso | 1988 | 1994 | | | | |
| Bolivia | 1986 | 1987 | 1994 | 2000 | | |
| Brazil | 1990 | 1990 | 1994 | 1996 | | |
| Central African Republic | 1988 | 1999 | | | | |
| Chile | 1976 | 1976 | 1981 | 1983 | | |
| Cote d'Ivoire | 1988 | 1991 | | | | |
| Cameroon | 1987 | 1993 | 1995 | 1998 | | |
| Congo, Rep. | 1992 | 2000 | | | | |
| Colombia | 1982 | 1987 | | | | |
| Czech Republic | 1989 | 1991 | | | | |
| Algeria | 1990 | 1992 | | | | |
| Ecuador | 1996 | 2000 | | | | |
| Egypt, Arab Rep. | 1977 | 1985 | | | | |
| Spain | 1977 | 1985 | | | | |
| Estonia | 1992 | 1995 | | | | |
| Finland | 1991 | 1994 | | | | |
| Ghana | 1982 | 1989 | | | | |
| Guinea | 1985 | 1985 | 1993 | 1994 | | |
| Hungary | 1991 | 1995 | | | | |
| Indonesia | 1987 | 2000 | | | | |
| Israel | 1977 | 1983 | | | | |
| Kenya | 1985 | 1989 | 1992 | 1992 | 1993 | 1995 |
| Korea, Rep. | 1997 | 2000 | | | | |
| Kuwait | 1988 | 1990 | | | | |
| Lebanon | 1988 | 1990 | | | | |
| Sri Lanka | 1989 | 1993 | | | | |
| Lithuania | 1995 | 1996 | | | | |
| Latvia | 1995 | 1996 | | | | |
| Madagascar | 1988 | 1988 | | | | |
| Mexico | 1995 | 2000 | | | | |
| Mali | 1987 | 1989 | | | | |
| Mauritania | 1984 | 1993 | | | | |
| Malaysia | 1997 | 2000 | | | | |
| Niger | 1987 | 1993 | | | | |
| Norway | 1988 | 1998 | | | | |
| Nepal | 1988 | 1988 | | | | |
| Peru | 1983 | 1990 | | | | |
| Philippines | 1998 | 2000 | | | | |
| Paraguay | 1995 | 2000 | | | | |
| Russian Federation | 1995 | 1995 | 1998 | 1998 | | |
| Senegal | 1988 | 1991 | | | | |
| El Salvador | 1989 | 1989 | | | | |
| Slovak Republic | 1991 | 2000 | | | | |
| Slovenia | 1992 | 1994 | | | | |
| Sweden | 1991 | 1994 | | | | |
| Chad | 1992 | 1992 | | | | |
| Thailand | 1997 | 2000 | | | | |
| Ukraine | 1997 | 1997 | | | | |
| Uruguay | 1981 | 1984 | | | | |
| Venezuela | 1994 | 2000 | | | | |
| Congo, Dem. Rep. | 1991 | 1992 | 1994 | 2000 | | |
| Zimbabwe | 1995 | 1995 | | | | |

Source: Caprio and Klingbiel (1999)

* Here are only listed countries for which we get a precise time period for Banking Crises.

APPENDIX A2 : SYSTEMIC AND BORDERLINE BANKING CRISES IN BLL 74 COUNTRIES PANNEL

| | | systemic banking crises | borderline banking crises | |
|---|---|---|---|---|
| ARG | Argentina | x | | |
| AUS | Australia | | x | |
| AUT | Austria | | | |
| BEL | Belgium | | | |
| BOL | Bolivia | x | | |
| BRA | Brazil | x | | |
| CAF | Central African Republic | x | | |
| CAN | Canada | | x | |
| CHE | Switzerland | | | |
| CHL | Chile | x | | |
| CMR | Cameroon | x | | |
| COL | Colombia | x | | |
| CRI | Costa Rica | | x | |
| CYP | Cyprus | | | |
| DEU | Germany | | | |
| DNK | Denmark | | x | |
| DOM | Dominican Republic | | | |
| DZA | Algeria | x | | |
| ECU | Ecuador | x | | |
| EGY | Egypt, Arab Rep. | x | | |
| ESP | Spain | x | | |
| FIN | Finland | x | | |
| FRA | France | | | |
| GBR | United Kingdom | | | |
| GHA | Ghana | x | | |
| GMB | Gambia, The | | | |
| GRC | Greece | | x | |
| GTM | Guatemala | | x | |
| HND | Honduras | | | |
| HTI | Haiti | | | |
| IDN | Indonesia | x | | |
| IND | India | | x | |
| IRL | Ireland | | | |
| IRN | Iran, Islamic Republic of | | | |
| ISR | Israel | x | | |
| ITA | Italy | | | |
| JAM | Jamaica | | | |
| JPN | Japan | x | | |
| KEN | Kenya | x | | |
| KOR | Korea, Republic of | x | | |
| LKA | Sri Lanka | x | | |
| LSO | Lesotho | | x | |
| MEX | Mexico | x | | |
| MUS | Mauritius | | | |
| MWI | Malawi | | | |
| MYS | Malaysia | x | | |
| NER | Niger | x | | |
| NIC | Nicaragua | | | |
| NLD | Netherlands | | | |
| NOR | Norway | x | | |
| NZL | New Zealand | | x | |
| PAK | Pakistan | | | |
| PAN | Panama | | | |
| PER | Peru | x | | |
| PHL | Philippines | x | | |
| PNG | Papua New Guinea | | x | |
| PRT | Portugal | | | |
| PRY | Paraguay | x | | |
| RWA | Rwanda | | x | |
| SDN | Sudan | | | |
| SEN | Senegal | x | | |
| SLE | Sierra Leone | | | |
| SLV | El Salvador | x | | |
| SWE | Sweden | x | | |
| SYR | Syria | | | |
| TGO | Togo | x | | |
| THA | Thailand | x | | |
| TTO | Trinidad and Tobago | | | |
| URY | Uruguay | x | | |
| USA | United States | | x | |
| VEN | Venezuela | x | | |
| ZAF | South Africa | | | |
| ZAR | Zaire | | | |
| ZWE | Zimbabwe | x | | |
| Total | 74 | 36 | 12 | |

Source: Caprio and Klingbiel (1999)

## Appendix B:Countries in the Sample

Algeria
Argentina
Australia
Austria
Belgium
Bolivia
Brazil
Cameroon
Canada
Central African Republic
Chile
Colombia
Costa Rica
Cyprus
Denmark
Dominican Republic
Ecuador
Egypt
El Salvador
Finland
France
Gambia
Germany
Ghana
Great Britain
Greece
Guatemala
Haiti
Honduras
India
Indonesia
Iran
Ireland
Israel
Italy
Jamaica
Japan
Kenya
Lesotho
Malawi
Malaysia
Mauritius
Mexico
Netherlands
New Zealand
Nicaragua
Niger
Norway
Pakistan
Panama
Papua New Guinea
Paraguay
Peru
Philippines
Portugal
Rwanda
Senegal
Sierra Leone
South Africa
Spain
Sri Lanka
Sudan
Sweden
Switzerland
Syria
Taiwan
Thailand
Togo
Trinidad and Tobago
United States of America
Uruguay
Venezuela
Zaire
Zimbabwe

## Appendix C: Variables and Sources

| Variable | Definition | Original source | Secondary source |
|---|---|---|---|
| | | | |
| Level and growth rate of GDP | Real per capita GDP | World Development Indicators | Loayza et al. (1998) |
| | Real per capita GDP (for initial GDP in cross-section regressions) | Penn World Tables | |
| Government size | Government expenditure as share of GDP | World Development Indicators | Loayza et al. (1998) |
| Openness to trade | Sum of real exports and imports as share of real GDP | World Development Indicators | Loayza et al. (1998) |
| | | | |
| | | | |
| Inflation rate | Log difference of Consumer Price Index | International Financial Statistics (IFS), line 64 | |
| Average years of schooling | Average years of schooling in the population over 25 | Barro and Lee (1996) | |
| Average years of secondary schooling | Average years of secondary schooling in the population over 15 | Barro and Lee (1996) | |
| Black market premium | Ratio of black market exchange rate and official exchange rate minus one | Pick's Currency Yearbook through 1989 ; and World Currency Yearbook. | |
| Liquid Liabilities | {(0.5)*[F(t)/P_e(t) + F(t-1)/P_e(t-1)]}/[GDP(t)/P_a(t)], where F is liquid liabilities (line 55l), GDP is line 99b, P_e is end-of period CPI (line 64) and P_a is the average annual CPI. | IFS | |
| Commercial-Central Bank | DBA(t) / (DBA(t) + CBA(t)), where DBA is assets of deposit money banks (lines 22a-d) and CBA is central bank assets (lines 12 a-d). | IFS | |
| Private Credit | {(0.5)*[F(t)/P_e(t) + F(t-1)/P_e(t-1)]}/[GDP(t)/P_a(t)], where F is credit by deposit money banks and other financial institutions to the private sector (lines 22d + 42d), GDP is line 99b, P_e is end-of period CPI (line 64) and P_a is the average CPI for the year. | IFS | |

Appendix D : 1960-1997 ANNUAL DATA CORRELATION (five year average data correlation in parenthesis)

| | com | lly | pc | growth | inf | gov | school | trade | bmp | initial |
|---|---|---|---|---|---|---|---|---|---|---|
| **com** | 1.00 | | | | | | | | | |
| **lly** | 0.47 (0.51) | 1.00 | | | | | | | | |
| **pc** | 0.55 (0.6) | 0.84 (0.84) | 1.00 | | | | | | | |
| **growth** | **0.21 (0.33)** | **0.15 (0.22)** | **0.14 (0.2)** | 1.00 | | | | | | |
| **inf** | -0.2 (0.6) | -0.2 (-0.26) | -0.1 (-0.26) | **-0.2 (-0.29)** | 1.00 | | | | | |
| **gov** | 0.24 (0.6) | 0.37 (0.21) | 0.27 (0.24) | **-0.0 (-0.04)** | -0.11 | 1.00 | | | | |
| **school** | 0.31 | 0.56 | 0.56 | **0.09 (0.13)** | 0.03 | 0.41 | 1.00 | | | |
| **trade** | 0.26 (0.6) | 0.16 (0.13) | 0.08 (0.09) | **0.05 (0.13)** | -0.16 | 0.48 | 0.05 | 1.00 | | |
| **bmp** | -0.3 (0.6) | -0.1 (-0.03) | -0.2 (0.22) | **-0.1 (-0.2)** | 0.26 | -0.13 | -0.10 | -0.21 | 1.00 | |
| **initial income** | 0.52 (0.6) | 0.62 (0.61) | 0.55 (0.76) | **0.14 (-0.14)** | -0.11 | 0.43 | 0.80 | 0.08 | -0.23 | 1.00 |
| **OBS** | 2656.00 | 2509.00 | 2521.00 | 2612.00 | 2577.00 | 1551.00 | 2484.00 | 2620.00 | 2576.00 | 2766.00 |
| **5 year avg OBS** | *359.00* | *359.00* | *359.00* | *359.00* | *359.00* | *359.00* | *359.00* | *359.00* | *359.00* | *359.00* |

VARIABLES

com = Commercial Banks Assets /(Central Banks + Commercial Banks Assets)
lly = Liquid Liabilities / GDP
pc = Private Credit/ GDP
growth= real per capita Growth
inf = inflation rate
gov= governement expenditures / GDP
school = average year of secondary education
trade = trade openess
bmp=black market premium
initial income = beginning of the period real per capita income