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ABSTRACT *This article contributes to the literature on the finance–growth link by presenting new findings based on a new, larger dataset that is an improvement on earlier studies due to its greater coverage in terms of time periods and countries, as well as the incorporation of additional control variables such as institutional quality and the investment rate. Our results demonstrate that financial development does not have a statistically significant effect on economic growth, a finding that is robust to different model specification and estimation techniques. This suggests that the finance–growth link is not as strong as portrayed in the literature, being dependent on the sample of countries and time periods considered.*

JEL Classification: E44, O16, O43

1. Introduction

There is an extensive literature on the link between financial development and economic growth. However, although many studies have found an association between both variables, the direction of causality remains an issue of debate. Beginning with Bagehot (1873), many scholars believe that the financial sector is a positive force behind economic growth, while others suggest that finance merely follows the dynamics of the real economy (Robinson, 1952). In this study, we re-examine this relationship through a dynamic panel analysis, using a new dataset that includes a larger number of countries than previous studies did, as well as some control variables that were not considered in most longitudinal analyses of the subject. Our analysis focuses exclusively on the financial system, thus excluding the stock market and the insurance industry: the two elements that are also an integral part of the financial sector.

In the introduction to his classic and pioneering study on the financial market, Walter Bagehot compared the size of banking deposits in London with those in New York, Paris

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and the German Empire, emphasizing the English superiority in financial matters. According to him, the possession of large quantities of borrowable money is at the root of a country's economic progress, as exemplified by the case of England. As he put it, "This efficient and instantly-ready organisation gives us [England] an enormous advantage in competition with less advanced countries—less advanced, that is, in this particular respect of credit" (Bagehot, 1873, p. 13). Other economists concurred with this line of argumentation, including Schumpeter (1934), who went further by suggesting that a banker is not merely an intermediary, but an active agent in the process of innovation in industrial activity.

Summarizing the main findings of the literature, Beck (2009) and Levine (2005) argue that the positive impact of financial development on economic growth can be explained by five mechanisms, whose operation reduces the negative effect of information asymmetries among economic agents and the transaction costs involved in their activities. According to them, the financial system (1) provides a means of payment that facilitates a greater number of transactions, (2) concentrates the savings of a large number of investors, (3) makes possible the allocation of resources to their most productive economic use, through the effective evaluation and monitoring of investment projects, (4) improves corporate governance, and (5) contributes to risk reduction and diversification.

Diverse studies have found a relation between the development of the financial system and economic growth, and have suggested a positive influence of the first variable over the second one (Levine, 2005), beginning with Goldsmith's (1969) empirical work and the seminal paper of King & Levine (1993) that comprised a cross-sectional analysis of 77 countries, which included several control variables. The positive link between finance and growth was also highlighted by later works such as Levine *et al.*'s (2000), who introduced a dynamic analysis, and Beck *et al.*'s (2000), who showed that the relation between financial development and economic growth was basically transmitted through an increase in productivity and not through the growth of capital, a result that would seem consistent with the third mechanism we mentioned earlier. According to Rajan & Zingales (1998), the financial system contributes to economic growth by decreasing the costs of external funding to firms.

Nevertheless, the finance and growth link is subject to more complex processes, since the proposed causal mechanisms operate differently according to specific characteristics of the economic environment. Thus, some authors have tried to specify better the conditions under which the positive effect of finance upon growth actually occurs. In this regard, Rousseau & Wachtel (2002) contend that financial development only has a positive influence on growth in a context of moderate to low inflation. Moreover, Rioja & Valev (2004a), working with the same dataset as Beck *et al.* (2000), found that the positive relationship between financial development—measured through the financial system size—and growth only occurs in high- or middle-income countries, suggesting the existence of a minimum threshold of economic development beyond which the financial system has a positive influence on the real economy. In another study, Rioja & Valev (2004b) also suggested that increases in financial development have a positive effect on economic growth only after a certain threshold of financial system development is reached. To account for this finding, their argument is that, in countries with poorly developed financial systems, increases in total credit are typically caused by government intervention through state-owned banks that lend money without paying much attention to the productive consequences of the investment projects they are funding. Consistent with this, La Porta *et al.* (2002) have demonstrated that in financial systems with a greater proportion of state-owned banks, economic growth tends to be smaller.

On the other hand, Loayza & Ranciere (2002) warned of the need to differentiate between short- and long-term effects of finance on growth, showing with panel data techniques that short-term negative effects can be combined with long-term positive effects. More recently, Rousseau & Wachtel (2005, 2011), updating the original panel results of Levine *et al.* (2000), found that the statistically significant association between finance and growth was valid for the period 1960–1989, but disappeared when later years were included in the sample. However, when they introduced statistical controls for the occurrences of banking crises in specific countries, the relationship between finance and growth still held over the whole sample period (1960–2004), suggesting that the weakening of the link was due to the fact that the number of financial crises has increased since the late 1980s (Rousseau & Wachtel, 2011). In addition, using rolling regression techniques (i.e. changing the sample of countries according to specific criteria), they found that the finance–growth link is stronger in middle-income countries or countries with moderate development of their financial system.

To determine the direction of causality, or more precisely the statistical feedback of temporary precedence, the relationship between finance and growth has been analysed with time-series methodology. The results have been mixed and dependent upon the countries and periods of time considered. For example, and taking the case of just one country (Tunisia), Boulila & Trabelsi (2004a) indicate that in a sample sub-period (1963–1987) the direction of causality goes from the real economy to the financial sector, whereas there is bidirectional causality in the total sample period (1962–1998). Moreover, Calderón & Liu (2003) found that the causal effect of financial development on growth is stronger than the reverse causal effect of the same variables. On the other hand, Christopoulos & Tsionas (2004), with an empirical base of 10 developing countries, showed evidence of long-term Granger causality from finance to economic growth, but not of reverse causality. Among the studies that have detected Granger causality flowing from finance to growth, we may mention Chang & Caudill (2005), who found that finance Granger causes growth in the case of Taiwan, and Rousseau & Wachtel (1998), whose study concludes with similar results analysing five industrialized countries over the period 1870–1929.

Yet, there is also evidence contradicting this hypothesis. Shan *et al.* (2001), using time series of nine OECD countries plus China, do not find that finance Granger causes economic growth, but rather find reverse causality in three cases and bidirectionality in five. In a later study, Shan (2005), estimating impulse-response functions for 11 countries, concludes that there is little evidence that finance influences economic growth. With a sample from countries of the Middle East and Mediterranean Africa, Boulila & Trabelsi (2004b) point out that, in most cases, the direction of causality goes from the real economic growth to the financial sector. Al-Awad & Harb (2005) obtained analogous results with 10 Middle Eastern countries, suggesting that there is a short-term direction of causality from economic growth to finance. On the other hand, analysing the case of Kenya, Odhiambo (2008) posits that the direction of causality between finance and growth depends on the indicator used as a proxy for financial development, but overall the financial sector would seem to follow what happens in the real economy, not otherwise. Likewise, in a study of Latin American countries for the period 1961–2005, Blanco (2009) finds that finance does not have a causal effect on growth, but that economic growth leads financial development. In a similar vein, and using a sample of 63 countries and a new technique to evaluate Granger causality, Hurlin & Venet (2008) do not find any evidence of Granger causality from finance to growth, but of reverse causality.

As can be appreciated from the aforementioned studies, the time-series evidence is in general not conclusive regarding the causal direction from finance to economic growth. In contrast, analyses of dynamic panels have shown a positive relation between both variables, although in these cases causality is exclusively inferred from theory. It is therefore important to see what happens with a panel based on a larger sample than previous studies, a wider observation window and a greater number of control variables as well as an improved methodology, to accurately identify the contribution of financial development to economic growth. But before presenting our methodological approach, we will first discuss two methodological limitations pervading most empirical studies on this subject (ours included).

In the first place, to operationalize financial development by measures for the credit size of the banking system involves a problem. In most comprehensive datasets, this indicator does not adequately differentiate between credit to productive firms and credit for consumption, and this latter category is not a minor proportion of total credit in some countries. To differentiate between these two categories is important, since both may contribute to economic growth through different channels: credit to firms exemplifies the classic mechanism initially postulated by Bagehot, whereas credit for consumption could affect the rate of growth by means of an increase in demand (at least in the short run). If the proportions of these types of credit were the same for all countries, there would not be a problem in neglecting this distinction, but this is not the case, and valuable information about the potential contribution of these two distinct mechanisms to economic growth is therefore lost.

In the second place, empirical studies do not take into account the possible spillover effects of highly developed financial systems on smaller or less developed countries. In developing countries, it is not unusual for foreign firms to get funding at lower interest rates through the financial system of their home countries. In this way, the financial system in Country A could, in addition to its own positive domestic effect, contribute to economic growth in Country B. The effect of these global financial flows has not been adequately quantified and taken into account in the studies on the finance–growth link.

2. Data and Methods

The development of financial systems, in terms of their capability to provide means of payment, implies a static relationship with the aggregate output of the economy, which basically suggests that the better the quality of the means of payment in an economy, the larger its overall output (i.e. a level–level relationship). However, empirical research on the finance–growth link has largely been focused on the dynamic relationship between financial development and growth of aggregate output (i.e. the relationship between level of financial development and growth of real GDP per capita). In this regard, two explanatory mechanisms mentioned in Beck's (2009) and Levine's (2005) literature reviews are the most relevant: (1) the capability of financial systems to mobilize the population's savings, which in turn increase the investment rate, and (2) the possibility of improving the effect of investment on growth by allocating funds to their most efficient use by economic agents. The dynamic nature of these two causal mechanisms calls for a methodological approach such as the generalized method of moments (GMM) system estimation technique for panel data that we use in this paper.

The dependent variable, economic growth, is operationalized by the real GDP per capita rate of growth in 2005 constant dollars (expressed in percentage points). We calculated averages corresponding to 5-year periods, to smooth the typical annual fluctuations due to business cycles. The data source is the Penn World Tables version 6.3 (Heston *et al.*, 2009).

As in most empirical studies in the literature, we operationalized financial development through a measure of the total credit to the private sector from banks and other financial institutions as a percentage of the GDP. The data source is Beck & Demirgüç-Kunt (2009), whose dataset reports yearly values for diverse countries from 1961 to 2007, which we used to calculate 5-year averages. There were no data available in the case of some developing countries (mostly Latin America ones) in this version of the dataset, so we extended the coverage for these countries by using figures of total credit to the private sector from an earlier version of the dataset (Beck *et al.*, 2000). The yearly time series of total credit to the private sector as a percentage of GDP have discontinuities for some countries. In these cases, we replaced the missing values by linear interpolation to calculate 5-year averages. When the missing values corresponded to the first years of the series, they were replaced by figures calculated from the trend line of the initial 5-year period. In all, only 25 observations were the object of these two interpolation procedures. Moreover, when there were missing values for 3 years within the same 5-year period, this period itself was considered as a missing observation.

It can be contended that using only one indicator for financial development could limit the relevance of our results. However, since the main hypothesis of the empirical studies that tried to test the validity of the finance–growth link at the microeconomic level is that having better access to credit enhances growth for individual firms (Demirgüç-Kunt & Maksimovic, 1998; Beck *et al.*, 2005), we believe that total credit is the best indicator to operationalize the financial development variable for the purpose of our study.

We controlled for the effect of diverse variables deemed by the literature as potential determinants of economic growth. Thus, our analysis includes government size operationalized as government spending share of GDP in current prices, a variable that may have a negative impact upon the growth rate (Scully, 1989; Barro & Sala-i-Martin, 1999), although some authors have also raised the possibility of a positive effect contingent upon the type of government spending considered.

Following the Solow-Swan model of economic growth, we control for countries' initial level of economic development, operationalized as the real GDP per capita at the beginning of each period, to take into account the potential convergence effect (the convergence argument hypothesizes that the growth rate of less developed countries is greater than that of the more developed countries). Moreover, and unlike most of the empirical studies on the finance–growth link, we have also controlled for the effect of the investment rate as a share of GDP in current prices. Trade openness, which in diverse studies appears as a determinant of economic growth, is also considered, operationalized as the sum of exports and imports as a percentage of GDP in current prices. For all variables, we calculated 5-year average values from the yearly time series of these indicators, the data source of which is also the Penn World Tables 6.3.

Human capital is also prominently mentioned in the literature as a determinant of economic growth, and we have operationalized it through the average years of secondary schooling in the adult population older than 15 years at the beginning of each 5-year period. The data source is Barro & Lee's (2000) dataset. Another variable that can affect economic growth is the inflation rate, which at high values typically reflects the degree of macroeconomic instability (Bruno & Easterly, 1998). It was operationalized through the variation in the consumer price index, and the data source for this variable is the World Development Indicators of the World Bank (2008). In very few cases, the time series of this source was supplemented with the inflation rate data from an earlier version of the Beck *et al.* (2000) dataset.

We also added the investment rate, a variable that is omitted in some panel analyses of the relationship between financial development and growth. It was operationalized as the share of investment relative to the GDP in current prices. Also, and unlike previous dynamic panel studies of the finance–growth link, we controlled for the effect of the countries' institutional quality through an indicator that evaluates the nature of a country's government system, the Polity 2 variable of the Polity 4 project (Marshall *et al.*, 2010). The indicator ranges from -10 (strongly autocratic) to 10 (strongly democratic). It is expected that the greater the institutional quality (i.e. the existence of institutions that give citizens voice to express their political preferences, the existence of institutional constraints over the executive branch of government, respect for the rule of law and civil liberties, and related aspects), the better the economic climate, which in turn could generate greater economic growth (Rodrik, 2000; Acemoglu *et al.*, 2005; Shirley, 2008). We have estimated 5-year averages of the yearly values of this indicator.

All independent variables are expressed as natural logarithms, with the exception of the institutional quality variable, the human capital variable (average years of secondary schooling in the adult population) and the inflation rate (which enters the equation as the log of 1 plus the inflation rate). Dummy variables were used to control for time-period effects. The resulting dataset is an unbalanced panel that includes information for 98 countries covering 9 five-year periods from 1961–1965 to 2001–2005. When the institutional quality variable enters the regression, we lose three countries from the sample.

We use the method of dynamic panels, which deals with the problem of omitted unobserved variables by taking first differences, and also tackles the issues of endogeneity and reverse causality by using lagged realizations of the explanatory variables as instruments in a GMM framework. The problem of endogeneity is, of course, not fully resolved with this method, but the use of these internal instruments aims to achieve a "weak" form of exogeneity (i.e. the instruments may be correlated with past and current values of the error terms, but not with future realizations of the errors).

We start with the following equation:

$$g_{it} = \alpha + \beta.f_{it} + \gamma.c_{it} + \mu_i + \lambda_t + \varepsilon_{it} \tag{1}$$

in which g is the growth rate, f is the financial development variable, c is a vector of explanatory variables, which we treat as endogenous, and i and t are the subscripts for units of analysis and time periods, respectively. The coefficients to be estimated are β and γ (a vector of explanatory variables coefficients), while μ_i is a vector of unobserved individual (country-specific) effects, λ_t is a vector of time-period effects and ϵ_{it} is the error term.

By taking first differences in Equation (1), we eliminate the country-specific effect term as follows:

$$g_{it} - g_{i,t-1} = \beta(f_{it} - f_{i,t-1}) + \gamma(c_{it} - c_{i,t-1}) + (\lambda_t - \lambda_{t-1}) + (\varepsilon_{it} - \varepsilon_{i,t-1}) \qquad (2)$$

This equation can be estimated with GMM using lagged values of the explanatory variables as instruments. These internal instruments may be correlated with past and current error terms but must not be correlated with subsequent error terms, which is expressed in the following moment conditions:

$$E\lfloor f_{i,t-s}(\varepsilon_{i,t} - \varepsilon_{i,t-1}\rfloor = 0 \quad \text{for each } t = 3 \ldots T,\ s \geq 2 \qquad (3)$$

$$E\lfloor c_{i,t-s}(\varepsilon_{i,t} - \varepsilon_{i,t-1}\rfloor = 0 \quad \text{for each } t = 3 \ldots T,\ s \geq 2 \qquad (4)$$

This GMM difference estimator, however, has some econometric problems, among them the loss of information produced by taking first differences. Therefore, Arellano & Bover (1995) and Blundell & Bond (1998) discuss a system estimator that combines the equation in differences estimated with lagged levels of the explanatory variables with an equation in levels estimated with lagged differences of these variables. It must be assumed that the correlation between the levels of the explanatory variables and the specific country effects is the same for all periods. Under this assumption, lagged differences are valid instruments for the levels equation if they are uncorrelated with future realizations of the error terms, hence the following additional moment conditions:

$$E\lfloor(f_{i,t-s} - f_{i,t-s-1})(\varepsilon_{i,t} + \mu_i)\rfloor = 0 \quad \text{for each } t = 3 \ldots T,\ s = 1 \qquad (5)$$

$$E\lfloor(c_{i,t-s} - c_{i,t-s-1})(\varepsilon_{i,t} + \mu_i)\rfloor = 0 \quad \text{for each } t = 3 \ldots T,\ s = 1 \qquad (6)$$

Here, only the most recent differences are used as instruments in the level equation, as the use of additional lags would imply redundant moment conditions (Arellano & Bover, 1995).

This system estimator approach has been widely used in growth regressions. However, many studies using this technique did not take into account a problem discussed by Windmeijer (2005). According to him, the two-stage method that is routinely used to compute the system estimator calculates standard errors in defect, which in turn leads to the assignment of high but incorrect levels of statistical significance to the independent variables. It is therefore necessary to make a numerical correction, whose omission leads to the acceptance of results that are actually invalid, an error common in much of the literature until a few years ago. In the present study, we have used the xtabond2 module for Stata (Roodman, 2006) to implement Windmeijer's correction for the GMM system estimator.

Another problem in this context is the use of too many instruments, which has been analysed by Roodman (2009). To see if this was really the case in our regression, we used the Hansen test, under the null hypothesis that the instruments are exogenous. For this test, Roodman recommends using a high p-value of 0.25, instead of the conventional level 0.05. However, we also present the results of the system GMM estimator using "collapsed instruments", a technique implemented in Stata by Roodman to limit the proliferation of instruments, which can weaken the usefulness of the Hansen test. In this regard, high p-values of the test, far from being an indicator that the GMM formulation is valid, can paradoxically be a signal that too many instruments are present and, therefore, of the inadequacy of the model to render unbiased coefficients. Hence, it also makes sense to include a model formulation with the least possible number of instruments, which allows us to have a more stringent test of the validity of these instruments. In addition, we have used the Arellano-Bond test to see whether the error terms have second-order autocorrelation (first-order correlation is expected by construction), since a basic assumption for this model specification to be valid is that error terms are not serially correlated.

3. Results and Discussion

In Table 1, we report the correlation coefficients for all variables. Of these bivariate relationships, the investment rate is the variable that shows the greatest correlation with economic growth, a moderate coefficient of 0.31, while inflation has a negative correlation of 0.28, and credit to the private sector a positive correlation of 0.17. Some independent variables show a high correlation between themselves. The strong correlation between human capital (schooling) and development level (log of initial GDP per capita) of 0.77 is noteworthy, as is that between the log of initial GDP per capita and financial development of 0.75. There is also a moderate correlation between institutional quality and initial GDP per capita (0.53) and human capital (0.52), respectively. The credit to the private sector is highly correlated with human capital (0.64), which is not surprising given that this latter variable is, in turn, highly correlated with the level of economic development. Moreover, the investment rate, institutional quality and inflation have a moderate correlation with private credit (0.60, 0.46 and −0.31, respectively).

The finding that income per capita is highly correlated with financial development is not surprising, since developed economies are presumed to be at the technological frontier in all industries, and therefore at the technological frontier in the banking industry. Although the real GDP per capita may not be a perfect indicator of development (it only measures economic prosperity), in the same way as credit to the private sector may not be a perfect indicator of financial development, this result is interesting. The existence of this basic relationship indicates that there are some conditions in developed countries that allowed them to expand the size of their banking systems. Institutional factors may play a role in this regard, as suggested by La Porta *et al.* (1997) in their analysis of the origins of legal systems. However, we are only interested here in whether financial development does lead to economic growth, which is typically measured as the difference in the log of GDP per capita. Ultimately, relatively high rates of growth, sustained over a long period of time, can lead to economic development, hence the relevance of studying the finance–growth link for development purposes.

Bivariate associations, nevertheless, are to be analysed with caution, since they clearly potentially may reflect endogeneity, which we deal with through a dynamic panel formulation in our multivariate regressions. However, before presenting our own results for the period 1961–2005, it may be helpful to look at the panel findings of other studies that have used system GMM estimation procedures, which we reproduce below (Table 2). As it can be observed in Column 1, the coefficient of the financial development variable, the log of private credit as a percentage of GDP, is positive and statistically significant in our replication of Levine *et al.*'s (2000) seminal work. Column 2 shows the coefficient of private credit when the variable black market premium is excluded from the set of control variables, since it is not available in our own analysis or Rousseau & Wachtel's (2011) analysis. Again, this coefficient is positive and significant. However, Rousseau and

Table 1. Correlation coefficients

	Growth	Ln. initial GDP per capita	Ln. gov.	Ln. openness	Ln. investment	Schooling	Ln. priv. credit	Ln. inflation	Institutional quality
Growth	1	0.08	−0.14	0.05	0.31	0.09	0.17	−0.28	0.14
Ln. initial GDP per capita		1	−0.11	0.17	0.58	0.77	0.75	−0.11	0.53
Ln. gov.			1	0.04	−0.10	−0.07	−0.11	0.12	−0.01
Ln. openness				1	0.21	0.12	0.25	−0.24	0.05
Ln. investment					1	0.41	0.60	−0.15	0.39
Schooling						1	0.64	−0.14	0.52
Ln. priv. credit							1	−0.31	0.46
Ln. inflation								1	−0.07
Institutional quality									1

Note: $n = 709$.

Table 2. Earlier analyses of the finance–growth link using dynamic panels (system GMM estimator)

	Levine *et al.* (2000)[1]	Levine *et al.* (2000)[2]	Rousseau & Wachtel (2011)
Private credit	1.41*	2.05*	− 0.001
Period	1961–1995	1961–1995	1960–2004
n	77	77	83
Hansen test (*p*-value)	0.58	0.38	0.50

Notes: All coefficients contemplate the Windmeijer correction. [1] Our replication of Levine *et al.*'s results is based on the dataset of Beck *et al.* (2000) and uses the program xtabond2 for Stata, exactly matching the same replication by Roodman (2009), whose Stata commands we have used. The control variables in this regression are initial GDP per capita, secondary education, trade openness, black market premium, inflation rate, and government expenditure over GDP. [2] Same control variables as (1) minus black market premium. *Significant at $p < 0.05$.

Wachtel, using the same set of control variables as our replication in Column 2, find that the financial development coefficient lacks statistical significance when a longer time period is considered. Our own study is an extension of these analyses by including more countries and taking into account additional control variables, institutional quality and the investment rate, which could better model the determinants of economic growth and isolate the potential contribution of financial development to growth.

First, we present the results of a fixed effects panel (Table 3), which should be interpreted with caution since it is invalidated by endogeneity. The results of all models show that most control variables are statistically significant with coefficients that are consistent with the literature. Trade, investment and human capital have a positive effect on growth, while inflation has a negative one as well as the initial level of GDP per capita, which is consistent with the Solow-Swan theory of the convergence in growth rates. Our main variable of interest, private credit, does not have a significant effect on growth. The addition of the institutional quality variable in Model 3 does not make a difference in this result, and this variable itself has a negative coefficient that lacks statistical significance.

The problem of endogeneity is dealt with using the GMM difference estimator. However, as we indicated earlier, this technique is not devoid of econometric problems, so our presentation here is basically for the purpose of testing the robustness of the results to those obtained with the GMM system estimator. All the models with the difference estimator are valid, as they pass the Hansen and Arellano-Bond autocorrelation tests. Again, the results show no significant effect of the financial development indicator, private credit, on economic growth. Among other independent variables, only the log of the initial GDP per capita and the human capital indicator remain statistically significant.

Next, we present the results of the GMM system estimator in Table 4. In our view, the GMM system estimator is the most appropriate technique to analyse the models at hand, provided, of course, that the model specifications pass the tests of instrument validity and serial correlation. For each of our three basic models, we present three methods of estimation: (1) Windmeijer-corrected standard errors, (2) Windmeijer correction omitted, and (3) collapsed instruments (with Windmeijer-corrected standard error). In the Windmeijer-correction version of Model 1, only the inflation rate has an effect on growth that is statistically significant. However, if we estimate the same model without this numerical correction, other variables acquire statistical significance, in particular private

Table 3. Determinants of economic growth (1961–2005)—fixed effects and GMM difference estimator

	Fixed effects panel			GMM difference estimator		
Variables	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
Constant	51.279	53.066	53.447			
	(5.938)	(6.049)	(6.292)			
Ln. GDP per capita initial	−5.799***	−5.694***	−5.783***	−6.589**	−7.917***	−8.973***
	(0.695)	(0.701)	(0.728)	(2.041)	(1.970)	(2.348)
Ln. gov	−0.529	−0.161	−0.309	1.389	1.133	1.203
	(0.805)	(0.803)	(0.848)	(2.267)	(2.306)	(2.126)
Ln openness	1.810***	1.555**	1.668**	1.371	1.185	1.956
	(0.476)	(0.495)	(0.491)	(1.623)	(1.585)	(1.741)
Ln. priv. cred.	0.372	0.219	0.225	0.299	0.575	0.342
	(0.297)	(0.320)	(0.323)	(1.024)	(0.982)	(0.966)
Ln. inflat	−2.817***	−2.828***	−2.770***	−1.654	−1.776	−1.739
	(0.609)	(0.645)	(0.670)	(1.186)	(1.198)	(1.271)
Sec. schooling	0.794*	0.853**	0.833*	2.921*	2.295*	2.466*
	(0.320)	(0.307)	(0.354)	(1.200)	(0.957)	(1.058)
Ln. invest.		1.425*	1.398*		0.137	0.188
		(0.651)	(0.657)		(1.569)	(1.564)
Instit. quality			−0.030			−0.116
			(0.029)			(0.090)
Adj. R^2	0.308	0.319	0.320			
F	22.71	19.86	18.23			
Pr > F	0.000	0.000	0.000			
N. Inst				49	56	63
Hansen				0.325	0.354	0.400
AB (2)				0.548	0.508	0.809
Obs.	730	730	709	575	575	559
Units	98	98	95	97	97	94

Notes: Fixed and period effects not reported; robust standard errors in parentheses. $***p < 0.001$; $**p < 0.01$; $*p < 0.05$.

Table 4. Determinants of economic growth (1961–2005)—GMM System estimator

	GMM system estimator								
	Model 1			Model 2			Model 3		
Variables	Windmeijer correction	Without correction	Collapsed instruments	Windmeijer correction	Without correction	Collapsed instruments	Windmeijer correction	Without correction	Collapsed instruments
Constant	2.108	2.108	17.048	17.532	17.532	39.671	12.733	12.733	35.719
	(7.229)	(1.222)	(12.645)	(7.030)	(1.741)	(12.789)	(9.382)	(3.033)	(17.872)
Ln. initial GDP	0.177	0.177	−1.664	−1.472*	−1.472***	−3.461**	−0.886	−0.886**	−2.847
per capita	(0.825)	(0.132)	(1.710)	(0.786)	(0.185)	(1.501)	(0.953)	(0.300)	(2.046)
Ln. gov	0.369	0.369*	−0.507	0.142	0.142	0.422	0.266	0.266	0.564
	(0.821)	(0.192)	(1.463)	(0.138)	(0.801)	(1.173)	(0.961)	(0.243)	(1.114)
Ln openness	−0.344	−0.344***	3.294**	−0.169	−0.169	0.279	−0.740	−0.740***	0.296
	(0.540)	(0.116)	(1.437)	(0.659)	(0.106)	(1.131)	(0.558)	(0.171)	(1.100)
Ln. priv. cred.	0.177	0.177*	−0.242	0.295	0.295***	0.161	0.262	0.262**	0.163
	(0.444)	(0.103)	(0.880)	(0.339)	(0.069)	(0.535)	(0.452)	(0.092)	(0.497)
Ln. inflat	−1.741**	−1.741***	−0.522	−1.099	−1.099***	−0.820	−2.423*	−2.423***	−1.579
	(0.837)	(0.147)	(1.094)	(0.840)	(0.123)	(1.179)	(0.980)	(0.300)	(1.076)
Sec. schooling	0.632	0.632*	1.882	1.318**	1.318***	1.765*	0.658	0.658***	1.172
	(0.701)	(0.369)	(1.330)	(0.563)	(0.151)	(0.901)	(0.560)	(0.150)	(1.223)
Ln. invest.				1.530**	1.530***	4.437***	1.364	1.364***	4.638***
				(0.745)	(0.145)	(1.210)	(0.877)	(0.227)	(1.334)
Instit. quality							0.057	0.057***	−0.023
							(0.036)	(0.009)	(0.052)
N. Inst	92	92	20	106	106	22	120	120	24
Hansen	0.314	0.314	0.164	0.760	0.760	0.713	0.966	0.966	0.263
AB (2)	0.571	0.571	0.365	0.413	0.413	0.253	0.377	0.377	0.263
Obs.	615	615	615	615	615	615	597	597	597
Units	98	98	98	98	98	98	95	95	95

Notes: Period effects not reported; standard errors in parentheses. $^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$.

credit. This shows how empirical support for certain hypothesis can be shown using statistical techniques that are later demonstrated to be deficient, as they undervalue standard errors and therefore inflate the statistical significance of some regression coefficients. In the third column of Model 1, we introduce the collapsed-instrument estimation, which represents the most stringent possible selection of instruments (one lag only with collapsed instruments) to facilitate the detection of invalid specifications due to instrument proliferation. In this case, it can be observed that the instruments are invalid, since they do not pass the Hansen test—its *p*-value is less than the recommended threshold of 0.25. If the collapsed-instruments specification does not pass this test, it can also be concluded that the normal-instruments model is also invalid, since its larger *p*-value is merely the result of adding instruments. The utility of the collapsed-instruments technique relies, among other things, on its power to detect that apparently valid models with normal instruments are actually invalid.

Model 2 adds investment rate to the set of independent variables. This specification passes the Hansen test, despite having more instruments. We can see again that private credit appears as a significant determinant of economic growth in the normal instruments model, but only in the model that does not include the Windmeijer correction, which, of course, is an incorrect result. In contrast, the log of initial GDP per capita has a negative and statistical effect on growth, as predicted by the Solow-Swan theory, in both the normal-instruments and collapsed-instruments specifications. Other statistically significant results are the positive effects of both the investment rate and human capital. The institutional quality variable enters in Model 3, but only becomes statistically significant with a positive effect on economic growth in the estimation that does not include the Windmeijer correction. The inflation rate is also significant with a negative effect on growth, while GDP per capita and the investment rate have the same signs as in Model 2 but lack statistical significance, except in the uncorrected specification. Again, the statistical significance of private credit as a predictor of economic growth only appears in the uncorrected model. It must be pointed out that the high *p*-value of the Hansen in this model, far from being a good sign, can indicate a problem, as very high *p*-values in this test are usually the sign of a problem of instrument proliferation. We can therefore conclude that the normal-instruments specification of Model 3 is invalid. The collapsed-instruments estimation passes the Hansen test. However, the only variable that has statistical significance in the collapsed-instruments version is the investment rate, with a positive effect. If we add further lags to all models to test for robustness, the results of the coefficient are similar to those of the models presented. Using additional lags, though, increases the *p*-value of the Hansen test to values closer to 1.

To sum up, our findings present a compelling case that financial development, operationalized as credit to the private sector, does not have a positive effect upon economic growth over the time period studied. Such a result is robust to different model specifications and estimation procedures. We did not find any consistently robust results for the other independent variables. The investment rate, though, has a significant and positive effect in the system GMM models (except in the normal-instruments specification of Model 3, which is invalid), but lacks significance with the GMM difference estimator. One lesson from this empirical analysis is that, in estimating economic growth, the instrumental variable models are simply invalid, or the findings, when robust, may at best indicate that some variables lack statistical significance (more negative than positive results). Identifying significant variables may be contingent upon the countries in the sample, the periods selected and the model specification of choice.

As Durlauf *et al.* (2005) observed, the econometrics of growth suffers from one fundamental difficulty, the fact that there are only few units of analysis (i.e. countries) available for obtaining statistically significant results. While panel methods increase the number of units of analysis, they also have their own problems. In the case of GMM methods, the basic weakness is that too many instruments may invalidate the models. As a rule of thumb, it is suggested that the number of units of analysis should not be greater than the number or instruments (Roodman, 2009), but as this is seldom possible, the recently introduced collapsed-instruments technique is very useful. Still, as Bazzi & Clemens (2009) observe, there is no test yet available to easily identify which specific instruments are weak or invalid in GMM growth regressions.

4. Conclusion

The finance–growth link has received much attention in the economic growth literature. While the empirical time-series evidence is mixed regarding the direction of causality in this relationship, the approach most often cited as providing the most convincing evidence that finance leads economic growth is the methodology of dynamic panels. We have re-examined this relationship with a new dataset that is an improvement on earlier studies due to its greater coverage in terms of both countries and time periods. In addition, we have controlled for the effect of institutional quality and the investment rate, two variables not considered in earlier studies using dynamic panels.

From a methodological viewpoint, our analysis also represents an improvement on many earlier studies by (1) including the Windmeijer correction for the GMM system estimator, which avoids the incorrect estimation of standard errors and therefore misrepresentation of the statistical significance of the explanatory variables, and (2) explicitly dealing with the problem of instrument proliferation, which appears in many empirical studies on the determinants of economic growth (Bazzi & Clemens, 2009), by incorporating a collapsed-instruments specification.

Our results demonstrate that financial development, operationalized as credit to the private sector, does not have a statistically significant effect on economic growth. This suggests that the finance–growth link is not as firm as portrayed in the literature, a finding consistent with Rousseau & Wachtel's (2011) recent empirical analysis, which also casts doubts on the strength of this link as a general hypothesis, regardless of the sample of countries and time periods considered. As these authors argue, in some countries, financial systems are indeed useful for economic growth, while in others, the same systems are crisis-prone, typically because they generate credit expansion that lacks a sound basis in the real sector and ultimately dampens the prospects for growth. For instance, recent empirical work (Chun, 2011) supports the claim that the larger the size of the financial system, the greater the probability of a banking crisis with its subsequent negative effect on growth (even excluding the recent global financial crisis).

However, the question still remains: what specific mechanisms account for the mixed effects of banking systems on economic growth? Is it a question of irresponsible credit expansion? And if so, why? Is it due to faulty monetary policy? Is it just a matter of deficient banking supervision? Is it a problem caused by fundamental institutional flaws that distort the incentive systems for economic agents? Once institutional quality (as measured by indicators of respect for civil and political rights) is controlled for, a positive effect of finance on growth might be expected, but according to our results this appears not to be the case. Further research is therefore needed to ascertain the specific conditions under which finance leads to economic growth—in other words, when and why financial systems are a blessing rather than simply irrelevant to, or even negative for, growth. So far and pending such clarification, it is suggested that policy makers should not put too much faith in credit created by the financial system as a driver for economic growth.

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