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Abstract

This study investigates bidirectional causality between governance and financial development using panel data of 101 countries from 1984 to 2013. The financial development–governance nexus is explored using econometric methods robust to cross-sectional dependence, and the relationship between different levels of development and openness is analyzed. Long-run equation estimates show clear evidence that financial development positively affects governance, and this positive impact is found to be robust to three different measures of governance. Further analysis shows that improving governance quality has positive effects on financial development, while Granger causality tests demonstrate bidirectional causality between financial development and the governance measures. Last, the impact of financial development on governance is dependent on a country's level of development and openness. These findings underscore the crucial role of financial development in bringing about good governance reforms and economic growth that, in turn, can further develop the financial sector. As such, a symbiotic and synergistic relationship can persist between good governance, growth, and financial development. The findings provide significant motivation for policymakers to encourage openness and financial sector development to lift the standard of living, especially in emerging economies.

Keywords: financial development; governance; cross-sectional dependence; economic growth; bidirectional causality; globalization

1 Introduction

The theoretical and empirical literature on the determinants of economic growth and development underscores the importance of financial development and good governance. Efficient financial markets divert resources from unproductive to productive activities, thereby improving overall economic efficiency and increasing economic growth. Schumpeter (1911) first highlighted the role of financial development and financial markets in the growth process, suggesting an efficient banking system is the key to economic growth because of its role in allocating savings to productive investments, thus promoting innovation. By contrast, without participatory, transparent, accountable, and justice-manifesting institutions, including those that guarantee property rights, policymaking can remain paralyzed, constraining countries' abilities to optimize their economic and human development capacity: in short, governance matters. From an institutional perspective, laws and regulations that are effectively enforced by an impartial and efficient governance system can support innovation and investment and create an environment conducive to inclusive economic growth.

The dynamic interaction between financial development and governance remains largely unexplored. Extant studies focus entirely on exploring the role good governance plays in strengthening a country's financial sector. Financial sector development has been found to occur in the presence of an efficient bureaucracy with low levels of corruption and strong law and order (Law & Azman-Saini, 2012; Le, Kim, & Lee, 2015). La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997, 1998) explore the link between financial development and institutions, focusing on whether differences in legal origin can explain capital market development. By contrast, Beck, Demirgüç-Kunt, and Levine (2003) examine the relative importance of the law and finance and initial endowment hypotheses. Their results suggest initial endowments play a more critical role in financial market development than legal origins.

A related strand of literature explores the interrelationship between country-level corporate governance provisions and financial development (Li, Maung, & Wilson, 2018). However, few studies explore the interrelationship between institutional quality and financial development across countries (Banerjee, Bose, & Rath, 2019; Miletkov & Wintoki, 2008). A growing number of studies have explored causality from governance to financial development, but to the best of our knowledge, reverse causality from financial development to governance has not been explored. Several arguments can be made to support causality from financial development to governance. For instance, governance reforms are often costly; hence, well-developed financial markets can be a prerequisite for successful and viable governance reforms (Miletkov & Wintoki, 2008, 2009). It may also be the case that only countries with developed financial markets can afford good governance and build better institutions (Fergusson, 2006). Furthermore, when political power is unequally distributed, and a narrow elite controls political decisions, financial development can be curtailed to restrict political competitors' financial access (Rajan & Zingales, 2003). This suggests improving the level of financial development can lead to a higher degree of political competition that can, in turn, improve the quality of institutions and governance.

A causal relationship from financial development to governance can also be motivated using North's institutional change framework (North, 1971, 1981, 2005). According to this framework, new institutional and governance structures will emerge when the social benefits of change exceed the costs. Thus, any technological shock or change in relative prices alters the cost-benefit possibilities of new institutional and governance arrangements, consequently stimulating the demand for new institutional and governance arrangements or changes in the existing structures. Changes in financial development modify the costs and benefits of particular institutional arrangements (Miletkov & Wintoki, 2008). More specifically, improvements in the level of financial development act as a catalyst for the emergence of higher quality institutions and governance frameworks. Improvements in governance increase the benefits accruing from potential financial arrangements, and after a certain threshold, the benefits from governance reforms will exceed the cost of undertaking those reforms.

Indirectly, lower financial development levels increase economic volatility and uncertainty (Beck, Lundberg, & Majnoni, 2006), which in turn increase the risk of political instability and lead to deterioration in governance quality. Similarly, with higher levels of financial development come greater financial liberalization and more frequent cross-border transactions (Miletkov & Wintoki, 2008). Such frequent interactions create a more informed citizenry, making them more aware of their legal and political rights and stimulating the demand for better governance structures (Khalid, 2017). Further, as Miletkov and Wintoki (2008) highlight, frequent financial transactions create incentives for people to acquire specific skills and education that are more conducive to administering and enforcing contracts, which, in turn, reduces the cost of administering and implementing governance reform. This discussion clearly underscores that financial development is a potential driver of improvements in governance structures, a premise that requires an empirical investigation.

This study builds on cross-country empirical studies that confirm a positive association between financial development and institutional quality. A related strand of literature also explores the link between financial development and legal origins, while a smaller body of work more specifically examines the role of good governance in financial market development. This study adds to this growing body of literature by exploring bidirectional causality between financial development and governance; it tests whether the relationship between financial development and governance varies by level of development and level of openness, which has yet to be considered in the literature.

In this study, bidirectional causality is tested using panel data of 101 countries from 1984 to 2013. Governance quality is measured using the International Country Risk Guide's political risk index (*GOV*) and its two subcomponents, the Investment Profile index (*IP*) and Government Stability index (*GS*). The financial development (*FD*) data are collected from Svirydzenka (2016). The financial development–governance nexus is then explored using econometric methods robust to cross-sectional dependence to identify the models to be cointegrated.

When the time dimension in panel datasets is substantially lower than the number of cross-sections, it is critical to take into account cross-sectional dependence (Sarafidis & Wansbeek, 2012; Shafiullah et al., 2019). Moreover, in a globalized world, governance (institutional quality) in one country affects others, especially its neighbors (Hosseini & Kaneko, 2013; Stiglitz Joseph, 2010), and changes in governance and institutional structures are driven by common global factors (Khalid, 2016). Similarly, convergence (as well as spillover, contagion, etc.) in financial development has been hypothesized and empirically identified across many economies (Apergis, Christou, & Miller, 2012; Bahadir & Valev, 2015; Dekle & Pundit, 2016). These highlight the hypothetical and empirical possibility of cross-sectional dependence in our panel dataset. As such, it is essential to implement econometric methods that are robust to cross-sectional dependence to obtain unbiased and efficient empirical estimates that will allow us to make valid inferences regarding the relationship between financial development and governance. Moreover, testing for cointegration in the specified models is important, as its presence precludes any question of endogeneity because the estimates of a cointegrated system are "superconsistent" (Stock, 1987). Cointegration and causality analysis has been used extensively in the literature related to financial development (see, e.g., Ahamada & Coulibaly, 2013; Coulibaly, 2015; Fromentin, 2017). Thus, our estimation strategy—controlling for cross-sectional dependence in a cointegration-causality analysis framework—enables us to uncover the causal relationship between governance and financial development and identify the direction of causality.

Long-run equation estimates show clear evidence that financial development positively affects governance. The positive impact of financial development on governance is found to be robust for three different measures of governance, while further analysis shows that governance and investment profile index have positive effects on financial development. The Granger causality tests demonstrate bidirectional causality between financial development and each governance measure. Last, we find the impact of financial development on governance depends on the level of economic development and the country's economic openness.

The remainder of this paper proceeds as follows. Section 2 presents a theoretical discussion of the relationship between financial development and governance and develops the hypotheses to be tested. Section 3 introduces the data and empirical methods. Section 4 presents and discusses the results, while Section 5 reports the results of robustness checks. Section 6 offers the conclusions.

2 Financial Development and Governance: An Overview

The role of financial development in a country's economic growth and development has been studied extensively in both the theoretical and empirical literature. Extant studies have identified five channels through which financial development may affect economic growth and development. First, in most theoretical models, the financial system's allocative efficiency is highlighted as a determinant of growth (see, e.g., Bencivenga & Smith, 1991; Greenwood & Jovanovic, 1990; Pagano, 1993; Wu, Hou, & Cheng, 2010). Second, emphasis is placed on the role of financial markets in providing opportunities to hedge against risk by allowing portfolio diversification and increasing liquidity, thereby stimulating economic growth (Levine, 1991; Saint-Paul, 1992). Third, financial development is seen as a mechanism that provides an exit option for agents and improves the efficiency of financial intermediation (Arestis, Demetriades, & Luintel, 2001; Rousseau & Wachtel, 2000). Fourth, as Greenwood and Smith (1997) suggest, efficient financial markets foster technological progress and act as a catalyst for entrepreneurship. Last, efficient financial markets restructure the incentives for corporate control that impact economic growth (Demirgüç-Kunt & Levine, 1996; Jensen & Murphy, 1990).

A growing strand of literature explores the relationship between financial development and governance, including the quality of institutions and legal systems. Most notably, studies have shown that a legal and regulatory system that guarantees property rights protection and contract enforcement is critical for financial development. For instance, La Porta et al. (1997, 1998) focus on whether differences in legal origins can explain capital market developments. Their findings show that poor shareholder rights are associated with less developed equity markets, especially in countries with French civil law. Common law countries enjoy relatively high levels of shareholder rights due to more highly developed equity markets. Similarly, better creditor rights also lead to the development of financial intermediaries. Beck et al. (2003) explore the relative importance of the law and finance and the initial endowment hypotheses. Their results suggest initial endowments play a more critical role in financial market development than legal origins.

Rajan and Zingales (2003) highlight the importance of political forces in shaping policies that influence financial market development. They argue that in countries where political power is unequally distributed and political decisions are controlled by a narrow elite, financial development could be curtailed to restrict political competitors' access to finance. This argument also suggests an increased level of financial development can lead to a higher degree of political competition that in turn can improve the quality of institutions and governance. The link between political institutions and financial development has been further investigated through analyses of the impact of democracy and regime change on financial development (Girma & Shortland, 2008; Huang, 2010). Girma and Shortland (2008) assess how financial development is affected by democratic characteristics and regime change; their findings reveal that the quality of democracy and political stability are substantial factors that drive financial development. They analyze this effect on a disaggregated level and find that, for the most part, political stability and improved democratic processes benefit the banking sector. In fully democratic regimes, there is a swift rise in stock market capitalization (Girma & Shortland, 2008). Huang (2010) also demonstrates that a democratic transition is typically followed by increased financial development and improvement in institutional quality leads to improved financial development, at least in the short run. This relationship holds particularly well for lower income countries, countries that are ethnically divided, and countries with French legal origins.

Roe and Siegel (2011) also highlight the political context of improving financial development, arguing that political stability determines a country's capacity and willingness to reform and improve the institutions and regulations that protect investments. Political instability threatens the proper functioning of institutions and, as a result, may lead to underdeveloped financial markets. This argument is supported by empirical evidence that variations in political stability have a consistent and significant effect on debt and stock market development. Haber, North, and Weingast (2008) examine the role of politics in financial development in the United States and Mexico from 1790 to 1914 and argue that government has strong incentives to behave opportunistically and use financial repression for its benefit. However, institutions that foster political competition reduce the likelihood that governments will behave opportunistically and result in a larger, more competitive, and more efficient banking system.

Mishkin (2009) argues that globalization is a key factor for stimulating institutional reforms in developing countries. Sound institutions are essential for promoting financial development because they establish and maintain strong property rights, an effective legal system, and efficient financial regulation. Therefore, institutional quality plays an important role in mediating the effect of globalization on financial development. Law (2009) demonstrates that trade openness and financial openness appear to have positive impacts on financial development in developing countries. He further analyzes whether the impacts result from fostering competition or upgrading institutional quality and indicates the institutional quality channel outperforms the competition channel.

Apart from formal institutions and enforcement of property and contractual rights, a related strand of the literature analyzes the relationship between informal institutions—in particular, social capital based on trust—and financial development. Social capital is represented by the shared norms that facilitate cooperation between two or more individuals (Coleman, 1988; Fukuyama, 1999; Ostrom, 2000). Shared norms involve developing trust while disincentivizing cheating. Social capital is highly significant in developing markets as financial contracts foster a high level of trust among members of a society (Guiso, Sapienza, & Zingales, 2004). Calderón, Chong, and Galindo (2001) evaluate the relationship between social capital and financial development and find trust has an economically significant and positive effect on financial intermediaries' size and activity, commercial banks' efficiency, and the level of development of stock and bond markets. Furthermore, Calderón et al. (2001) underscore the complementarity between trust and formal institutions in a society where the rule of law is disregarded and vice versa.

This discussion illustrates the importance of governance and institutional quality in improving financial development. However, there may also be reverse causality from financial development to governance and the quality of institutions, an issue that has largely remained unexplored in extant literature. There are several channels through which financial development can affect the quality of governance. Financial development in a country reduces borrowing constraints and increases access to finance for most of the population, which, in turn, increases economic and political competition and lead to governance improvements. Similarly, financial development is often accompanied by financial and trade liberalization, which allow for the free flow of money, goods, and services. As Khalid (2017) notes, trade liberalization eventually leads to improvements in institutional quality and governance. Moreover, financial liberalization increases pressure on governments to improve their institutional structure to prevent the outflow of finance.

Therefore, the aim of this study is to disentangle the relationship between governance and financial development and clearly establish the direction of causality. Moreover, we explore non-linearities in the relationship by analyzing the underlying relationship for different income and globalization levels.

3 Model, Data, and Methodology

3.1 Model and data

Based on recent literature, the determinants of governance are specified in Eq. (1):

$$GOV_{it} = \beta_0 + \beta_1 KOG_{it} + \beta_2 GPC_{it} + \beta_3 FD_{it} + u_{it} \tag{1}$$

Here, governance (*GOV*) for country i in period *t* is explained by globalization as measured by the KOF overall globalization index (*KOG*), economic development as proxied by real GDP per capita (*GPC*), and financial development (*FD*) using the financial development index introduced by Svirydzenka (2016). The quality of governance in Eq. (1) is measured using a composite index derived from the political risk index in the International Country Risk Guide (ICRG).

To further check the robustness of the results estimated from model (1), we substitute the sub-indices from ICRG's political risk index—investment profile (*IP*) and government stability (*GS*)—for *GOV* and estimate the following two equations:

$$IP_{it} = \beta_0 + \beta_1 KOG_{it} + \beta_2 GPC_{it} + \beta_3 FD_{it} + u_{it} \tag{2}$$

$$GS_{it} = \beta_0 + \beta_1 KOG_{it} + \beta_2 GPC_{it} + \beta_3 FD_{it} + u_{it} \tag{3}$$

Our panel dataset covers 101 countries from 1984 to 2013, and Table 1 provides the descriptive statistics of our model variables. As can be seen, *GOV* has a mean of about 24 and ranges from 5 to 38. The standard deviation of 6.4475 produces a coefficient of variation of about 26.5 percent. The mean of *IP* is approximately 7.5, with highest and lowest values of 12 and 0, respectively. The standard deviation is 2.44, giving us a coefficient of variation of 32.6 percent. In contrast, *GS* ranges from 1 to 12, with a mean of 7.65 and a coefficient of variation of 27.6 percent. Judging from the coefficients of variation, *IP* is the most volatile governance measure, followed by *GS* and *GOV*.

Table 1: Descriptive statistics

	Variable					
Measure	*GOV*	*IP*	*GS*	*KOG*	*GPC*	*FD*
Mean	24.3002	7.4790	7.6508	54.4365	14288.4800	0.3210
Maximum	38.2900	12.0000	12.0000	92.6300	110001.1000	1.0000
Minimum	5.0000	0.0000	1.0000	16.1400	130.4367	0.0000
Standard Deviation	6.4475	2.4400	2.1085	18.2182	18633.9200	0.2392
Observations	2929	2929	2929	2929	2929	2929
Time-series range	1984-2013					
No. of countries	101					

The *KOG* has a mean of about 54 and fluctuates between 93 and 16. With a standard deviation of slightly over 18, the coefficient of variation is about 33.5 percent. The mean *GPC* is a little over US$14,200, with highest and lowest values of about US$110,000 and US$130, respectively. The standard deviation is more than US$18,500, resulting in a coefficient of variation of 130.4 percent. *FD* has a mean of about 0.32, with 1 and 0 being the highest and lowest possible values. Based on the standard deviation of 0.24, the coefficient of variation is 74.5 percent. Thus, *GPC* is the most volatile explanatory variable, followed in order by *FD* and *KOG*.

3.2 Empirical strategy

The literature exploring the inter-relationship between financial development and other factors such as trade openness, quality of governance and economic growth and development mainly rely on traditional panel data methods such as Fixed effects, Random effects of GMM methods (see, e.g., Baltagi, Demetriades, & Law, 2009; Çoban & Topcu, 2013; Law & Azman-Saini, 2012; Law, Tan, & Azman-Saini, 2014; Li, Maung, & Wilson, 2018). These techniques, however, provide a biased view of the relationship due to the presence of identification issues. Moreover, these studies do not take into consideration cross-sectional dependence while exploring the underlying relationship using panel data.

Cross-sectional dependence, also known as common correlation, often results in standard panel unit root test results that under-reject the null hypothesis. This introduces type II error in the subsequent empirical analysis. In addition, testing for cross-sectional dependence is important for two reasons: First, in a globalized and highly integrated world, changes in governance and institutional structures across countries are driven by common global factors (Khalid, 2016), and changes in governance and institutional quality in one country often spill over to others, especially its neighboring countries (Hosseini & Kaneko, 2013; Stiglitz Joseph, 2010). Similarly, convergence (as well as spillover, contagion, etc.) in financial development has been hypothesized and empirically identified across many economies (Apergis et al., 2012; Bahadir & Valev, 2015; Dekle & Pundit, 2016). Second, the number of cross-sections (N=101) in the panel dataset described above is considerably higher than the number of time-periods (t=30). Cross-sectional dependence is often the characteristic of such panels because the error term contains unobserved common factors and shocks (Sarafidis & Wansbeek, 2012; De Hoyos and Sarafidis 2006; Dogan et al. 2017; Shafiullah et al., 2019).

Considering the shortcomings of the traditional time-series approaches such as Fixed and random effects model and GMM, this study, therefore, employs a cointegration test together with a Granger-causality test and test for cross-sectional dependence to uncover the causal relationship between financial development and governance quality. This approach of testing for causality has been widely used in the literature on the determinants and effects of financial development (see, e.g., Ahamada & Coulibaly, 2013; Coulibaly, 2015; Fromentin, 2017; Ahmad, Jabeen, Hayat, Khan, & Qamar, 2020; Ali, Yusop, Kaliappan, & Chin, 2020). Because the causal link between financial development and governance is likely to be bidirectional, traditional methods will provide biased estimates of the relationship. To preclude any question of endogeneity, we test for cointegration in the specified models because the estimates of a cointegrated system are “superconsistent” (Stock, 1987). We also test for cross-sectional dependence in our dataset. We implement several tests based on the Lagrange multiplier, including Breusch and Pagan (1980), Pesaran (2004), and Baltagi, Feng, and Kao (2012), as well as one test based on the Dickey-Fuller procedure (Pesaran. 2004). These four techniques test for cross-sectional dependence in individual variables. In addition, it is often useful to test for cross-sectional dependence in the specified model. To achieve this, the Frees (1995) and Pesaran (2004) procedures can be implemented for models (1-3). If cross-sectional dependence is diagnosed in the model variables, standard unit root tests cannot be used to ascertain their stationarity properties. To that end, Pesaran (2007) derived a variant of the Im, Pesaran, and Shin (2003) test that can account for dependence across cross-sections, a procedure is known as the CIPS test.

If the CIPS test indicates the model variables are nonstationary, we are required to test for any presence of cointegration (or long-run equilibrium). If the panel dataset shows cross-sectional dependence, it is necessary to apply cointegration testing methods that are robust to common correlation. The Westerlund and Edgerton (2008) panel cointegration testing procedure is one of the few methods capable of doing this. In addition, it is known for its robustness to residual heteroscedasticity and autocorrelation, as well as good finite sample performance. This procedure is a Lagrange multiplier-based test and estimates the following two statistics:

$$Z_j(N) := \sqrt{N}\left(\overline{LM}_j(N) - E\left(B_j\right)\right), \quad \text{for} \quad j = \{\phi, \tau\} \tag{4}$$

where

$$\overline{LM}_j(N) \coloneqq \frac{1}{N}\sum_{i=1}^{N} LM_j(i), \qquad \text{for} \quad i = \{1, \dots, N\} \tag{5}$$

and the Lagrange multiplier statistic is given by *LMj*.

The empirical analysis then requires estimating long-run equations based on models (1-3) if they are found to be cointegrated. The presence of cross-sectional dependence in the panel dataset requires implementing a suitable regression method. The Pesaran (2006) common correlated effects (CCE) mean group estimator accounts for cross-sectional dependence (or common correlation) by incorporating unobserved common factors in the estimation process. The CCE approach estimates coefficients for each cross-section and averages them across the panel using appropriate weights. The robustness of the CCE regression results may be verified by estimating models (1-3) using the pooled mean group (PMG) estimator by Pesaran and Smith (1995) and Pesaran et al. (1999).

Finally, cointegration tests are often conducted in conjunction with tests for the direction of Granger causality. This becomes essential if there is cointegration in models (1-3). For a cross-sectional dependent panel, it is necessary to apply an appropriate method—in this case the Dumitrescu and Hurlin (2012) approach. This method is robust to cross-sectional dependence in the data and has a null hypothesis of homogeneous noncausality versus the alternative of heterogeneous causation in the specified direction (such as $X \Rightarrow Y$).

4 Empirical Estimation Results

Tests for cross-sectional dependence are first conducted on the individual variables in models (1-3). The list includes the Breusch and Pagan (1980) (BP) *LM*, Pesaran's (2004) scaled (PS) *LM*, Baltagi et al.'s (2012) bias corrected scaled (BCS) *LM*, and Pesaran's (2004) cross-sectional dependence (*CD*) test statistics. As Table 2 shows, the four test statistics reject the null hypothesis that the variables are dependent across countries (cross-sections) at the one percent level of significance.

Table 2: Test for cross-sectional dependence, individual variables

	Test Statistic			
Variable	BP (*LM*)	PS (*LM*)	BCS (*LM*)	CD
GOV	47797.00***	425.3485***	423.5449***	176.9885***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
IP	59687.60***	543.6644***	541.8609***	211.4611***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
GS	51248.87***	459.6960***	457.8924***	205.3841***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
KOG	113175.4***	1075.888***	1074.147***	331.8246***

	(0.0000)	(0.0000)	(0.0000)	(0.0000)
GPC	90544.52***	850.7022***	848.9609***	215.3841***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
FD	56471.69***	511.6649***	509.8614***	125.3584***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)

Lagrange multiplier and cross-sectional dependence are abbreviated as *LM* and *CD*, respectively. Parentheses include p-values. Null hypothesis: Cross-sectional dependence (common correlation) is not present in the variable. *** When p-value < 0.01, reject H_0.

We then continue to test models (1-3) for cross-sectional dependence using the Frees (1995) and Pesaran (2004) tests and provide the results in Table 3. Both procedures test the null hypothesis that the model is not dependent across the cross-sections. The estimated test statistics from Frees (1995) are greater than the one percent critical value for models (1-3), rejecting the null hypothesis. For the Pesaran (2004) test, the estimated test statistics have p-values that are less than 0.0100 for all three models, indicating the null hypothesis is also rejected for the Pesaran (2004) test. Thus, the presence of cross-sectional dependence is verified in both the individual variables and specified models (1-3).

Table 3: Cross-sectional dependence test results, models

	Frees (1995) [a]		Pesaran (2004) [b]	
Model	Test statistic	1% critical value	Test statistic	p-value
(1)	13.888***	0.1660	68.613***	0.0000
(2)	14.693***	0.1660	70.562***	0.0000
(3)	14.334***	0.1660	128.369***	0.0000

H_0: Model is not dependent (not correlated) across the cross-sections.

[a] *** Reject H_0 when test statistic > 1% critical value. [b] *** Reject H_0 if the p-value < 0.0100.

Faced with cross-sectional dependence in our panel dataset, we test the variables for models (1-3) for unit roots using Pesaran's (2007) CIPS test. The test statistic and respective p-values are shown in Table 4. The p-values associated with the test statistics exceed 0.010 when the variables are measured as levels and fall below 0.010 when the variables are measured as first differences. This implies the levels forms of the variables are non-stationary whereas their first differences are stationary. As such, we can conclude that the order of integration of all variables is 1. Therefore, cointegration tests may be applied to models (1-3). Additionally, we perform the Lluís Carrion-i-Silvestre, Del Barrio-Castro, and López-Bazo (2005) test for structural breaks in the panel variables. The estimated results, reported in Table A1, indicate the variables are free of any structural changes/breaks. This supports the reliability and validity of the CIPS findings that the variables are integrated of order one.

Table 4: Pesaran (2007) CIPS test results

Variable	Test Statistic ($Z(\bar{t})$)	p-value
GOV	-0.729	0.233
Δ *GOV*	-6.380***	0.000
IP	-0.845	0.199
Δ *IP*	-7.207***	0.000
GS	1.575	0.942
Δ *GS*	-9.907***	0.000
KOG	-0.370	0.356
Δ *KOG*	-11.465***	0.000
GPC	9.475	1.000
Δ *GPC*	-2.395***	0.008
FD	3.002	0.999
Δ *FD*	-6.255***	0.000

Deterministic term: constant and trend. Lag order selection: minimization of Akaike Information Criteria (AIC). H_0: Series is non-stationary while cross-sectional dependence is controlled for. *** When p-value < 0.010, reject H_0.

Table 5 provides the estimated test statistics and corresponding p-values for models (1-3) and shows that the p-values for the test statistics are lower than 0.01. As such, the null hypothesis of a lack of cointegration in models (1-3) is rejected at the 1 percent significance level. The number of unobserved common factors is one for models (1) and (3) and two for model (2). Cointegration can thus be observed to exist in the estimated models.

Table 5: Westerlund and Edgerton (2008) cointegration test results

Model	$Z_\tau(N)$	p-value	$Z_\phi(N)$	p-value	No. of common factors
(1)	-12.908***	0.000	-15.950***	0.000	1
(2)	-5.628***	0.000	-3.253***	0.001	2

(3)	-6.463***	0.000	-7.244***	0.000	1

H_0: No cointegration present in the model after controlling for cross-sectional dependence. *** When p-value < 0.010, reject H_0.

As models (1-3) are cointegrated, we continue estimating the long-run equations based on these models. Because our dataset is cross-sectionally dependent, we implement the CCE mean group and PMG estimators. The CCE and PMG estimators' results are provided in Panels A & B, respectively, of Table 6. The coefficient of *KOG*—the proxy for globalization—is negative and significant at the 5 percent level in the CCE estimator. By contrast, the PMG estimator shows a positive and insignificant coefficient on *KOG*. All else being equal, a 1 unit increase in the globalization measure results in a 0.101 unit decrease in the governance indicator. This result is consistent with the framework proposed by Blouin, Ghosal, and Mukand (2012), as well as with the skeptical view of globalization's impact on governance held by Rodrik and Subramanian (2009), Stiglitz (2010), and Krugman (2009). They argue that globalization increases the risk of sudden capital flight and provides governments with the wrong incentives, resulting in "undisciplined" governments and leading to (mis)governance. Similarly, Rodrik (1998) highlights that the risk of capital flight that arises due to global market integration can be mitigated with a large government sector. However, a large public sector can also increase the government's exploitative power, resulting in poor governance quality.

Table 6: Long-run equations (1-3), CCE and PMG estimates

Panel A: CCE estimates			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	-0.692	0.945	0.709
	(0.814)	(0.781)	(0.702)
KOG	-0.101**	-0.066*	-0.064**
	(0.048)	(0.056)	(0.047)
GPC	0.005**	4.26×10^{-5}	0.001**
	(0.002)	(0.982)	(0.040)
FD	**6.376***	**5.914***	**4.234***
	(0.068)	**(0.058)**	**(0.063)**
$\sqrt{MSE}$	1.184	0.421	0.685

Panel B: PMG estimates			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	8.592**	6.365**	-5.325**
	(0.050)	(0.007)	(0.020)
KOG	0.068	0.013	0.110**
	(0.377)	(0.725)	(0.015)
GPC	0.014**	-0.001**	0.006**
	(0.001)	(0.030)	(0.006)
FD	**9.836***	**5.336***	**8.477****
	(0.079)	**(0.071)**	**(0.020)**
$\sqrt{MSE}$	1.521	0.405	1.090

The coefficients are cross-sectional averages. Parentheses indicate *p*-values. Null hypothesis: insignificant effect of coefficient. Mean squared error (σ^2) is abbreviated as MSE. ** Reject H_0 when the *p*-value < 0.0500. * Reject H_0 when the *p*-value < 0.1000.

GPC—per capita income—is positive and significant at the 5 percent level in both the CCE and PMG estimators. A US$1 increase in per capita GDP (base year = 2010) results in an average increase of between 0.005 and 0.014 in the governance indicator (*GOV*). This finding is intuitive and consistent with economic theory. As economic activity increases, investments to improve formal governance increase, and reliance on informal mechanisms decreases. As Dixit (2011) highlights, this is mainly because cooperation can often be efficiently sustained through personal ties and repeated interactions when an economy is small and localized. However, as economic development occurs, it increases the complexity and scale of trade, which may affect the efficiency of formal governance mechanisms (Dixit, 2003; Greif, 1994; J. S. Li, 2003), creating stronger incentives for public investments in governance institutions. In addition, undertaking comprehensive governance reform is often challenging for developing economies that lack administrative expertise and capital resources in terms of expenses and technical know-how (Rodrik, 2008). Therefore, higher GDP per capita means more revenue for government expenditures and other activities. This, in turn, can allow the government to function more efficiently, as well as achieve economies of scale and scope.

The coefficient of *FD* is also found to be positive. However, it is the only variable in the equation that is significant at the 10 percent level in both the CCE and PMG estimators. Ceteris paribus, when *FD* increases by 0.1 units, *GOV* increases between 0.64 and 0.98 units. This is consistent with the argument that only countries with high levels of financial development may be able to support good governance since financial development improves the government's access to funds/revenue, resulting in more efficient operation. Similarly, our result supports Rajan and Zingales's (2003) political economy argument, namely, that when political power is unequally distributed, and political decisions are controlled by a narrow elite, financial development can be curtailed to restrict political competitors' access to finance. Therefore, improvements in financial development can lead to higher levels of political competition that in turn can improve the quality of institutions and governance.

For model (2), we find the coefficient of *KOG* is significant in the CCE estimator at the 10 percent level, and while its PMG counterpart is insignificant. The significant effect of *KOG* on *IP* is negative and slightly lower than that in model (1). The coefficient of *GPC* is found to be positive and insignificant from the CCE but negative and significant (at 5 percent level) from the PMG. This negative effect of *GPC* in model (2) is a contrast to its positive effect in model (1) The coefficient on *FD* are significant when $\alpha = 10\%$ in both the CCE and PMG estimators. The effect of *FD* on *IP* is positive, and its magnitude is almost identical to that in model (1). For model (3), all three coefficients are statistically significant in both the CCE and PMG estimators. The coefficient of *KOG* is significant at the 5 percent level (in both estimators) but have opposing signs. The magnitude of the negative effect (in the CCE estimator) is lower than that of its counterpart in model (1). *GPC* has a positive effect on *GS* and is significant at the 5 percent level. Ceteris paribus, the magnitude of *GPC*'s effect on *GS* in model (3) is between some one-half (PMG) and one-fifth (CCE) its effect on *GOV* in model (1). Like its counterparts in models (1-2), *FD* positively affects *GS,* and this effect is significant at the 10 and 5 percent levels under the CCE and PMG estimators, respectively. However, the coefficient of *FD* in model (3) is substantially lower than its counterpart in model (1). Ultimately, we find *FD* positively affects governance, and this effect is robust to different measures of governance as well as different estimators.

Table 7 provides the estimated test statistics and corresponding p-values from the Dumitrescu and Hurlin (2012) Granger causality tests performed on the variables from models (1-3). As seen in the rightmost column, the p-values are lower than 0.01000 for all causal directions except $GPC \Leftrightarrow GOV$, $IP \Rightarrow GPC$, $GPC \Leftrightarrow GS$, and $FD \Rightarrow GS$. The p-values between $GPC \Leftrightarrow GOV$ are greater than 0.010 but smaller than 0.050. Thus, there is feedback between these two variables at the 5 percent significance level. The estimated p-value for the direction $IP \Rightarrow GPC$ is also less than 0.05000, implying the presence of causality at the 5 percent significance level. For both $GPC \Leftrightarrow GS$ directions, the p-values are greater than 0.10000, demonstrating a lack of causality between these two variables. The p-value for $FD \Rightarrow GS$ is greater than 0.05000 but less than 0.10000, indicating that causality runs in that direction at the 10 percent significance level. In all remaining directions, Granger causality exists at the 1 percent significance level.

Table 7: Dumitrescu and Hurlin (2012) Granger causality test results

Causality direction	Test statistics		p-value
	$\bar{W}$	$\bar{Z}$	
$KOG \Rightarrow GOV$	3.31046***	4.58847***	0.00000
$GOV \Rightarrow KOG$	3.54068***	5.53972***	0.00000
$GPC \Rightarrow GOV$	2.79991**	2.47884**	0.01320
$GOV \Rightarrow GPC$	2.74914**	2.26906**	0.02330
$FD \Rightarrow GOV$	2.92181***	2.98255***	0.00290
$GOV \Rightarrow FD$	3.50866***	5.40743***	0.00000
$KOG \Rightarrow IP$	4.86751***	11.0222***	0.00000
$IP \Rightarrow KOG$	3.66082***	6.03616***	0.00000
$GPC \Rightarrow IP$	4.12621***	7.95913***	0.00000
$IP \Rightarrow GPC$	2.80016**	2.47989**	0.01310
$FD \Rightarrow IP$	4.38912***	9.04548***	0.00000
$IP \Rightarrow FD$	3.88427***	6.95944***	0.00000
$KOG \Rightarrow GS$	2.85649***	2.71265***	0.00670
$GS \Rightarrow KOG$	2.99336***	3.27819***	0.00100
$GPC \Rightarrow GS$	2.56021	1.48839	0.13660
$GS \Rightarrow GPC$	2.59318	1.62463	0.10420
$FD \Rightarrow GS$	2.61741*	1.72477*	0.08460
$GS \Rightarrow FD$	2.96031***	3.14161***	0.00170

$GPC \Rightarrow KOG$	3.28160***	4.54996***	0.00000
$KOG \Rightarrow GPC$	4.31049***	8.84042***	0.00000
$FD \Rightarrow KOG$	3.17164***	4.01483***	0.00000
$KOG \Rightarrow FD$	5.64154***	14.2205***	0.00000
$FD \Rightarrow GPC$	5.26401***	12.6605***	0.00000
$GPC \Rightarrow FD$	4.99966***	11.5683***	0.00000

Selected lag order: 2. Null hypothesis: Lack of causality in the direction. * Reject H_0 when the p-value < 0.10000. ** Reject H_0 when the p-value < 0.05000. *** Reject H_0 when the p-value < 0.01000.

The analysis shows causality in most directions, especially from *FD* to the governance measures. As such, we are curious about the effect of governance on financial development. Accordingly, we estimate equations with *FD* as the dependent variable and each of the three governance measures in turn as an independent variable, along with *KOG* and *GPC*. The CCE and PMG estimates from those equations are provided in Panels A & B of Table 8 and show that the effect of *KOG* and *GPC* on *FD* is positive and significant in all three equations. In addition, the magnitudes of these two coefficients are virtually identical across all three equations. The coefficients of *GOV*, *IP,* and *GS* are positive in their respective equations under both the CCE and PMG estimators. However, only the coefficients of *GOV* and *IP* are significant under the CCE approach at the 10 and 5 percent significance levels, respectively. Ceteris paribus, the effect of a 1 unit increases in *GOV*, *IP,* and *GS* on *FD* are between 0.001 and 0.003 units, respectively, proving that the governance indicators have a positive effect on financial development.

Table 8: Impact of governance on financial development, CCE and PMG estimates

Panel A: CCE estimates			
	Regressand		
Regressor	*FD*	*FD*	*FD*
Constant	-0.046	-0.015	0.020
	(0.389)	(0.751)	(0.665)
KOG	0.002**	0.003**	0.002**
	(0.005)	(0.000)	(0.010)
GPC	2.14×10^{-5}*	2.13×10^{-5}*	2.45×10^{-5}**
	(0.050)	(0.082)	(0.005)
GOV	**0.001***	-	-
	(0.089)		

IP	-	**0.003**** **(0.044)**	-
GS	-	-	6.96×10^{-5} (0.951)
$\sqrt{MSE}$	0.021	0.021	0.023
Panel B: PMG estimates			
	Regressand		
Regressor	*FD*	*FD*	*FD*
Constant	-0.126** (0.013)	-0.124** (0.010)	-0.102** (0.045)
KOG	0.003** (0.000)	0.004** (0.000)	0.003** (0.000)
GPC	3.81×10^{-5}** (0.000)	2.94×10^{-5}** (0.002)	3.78×10^{-5}** (0.000)
GOV	**0.001**** **(0.035)**	-	-
IP	-	**0.002*** **(0.054)**	-
GS	-	-	**0.003**** **(0.040)**
$\sqrt{MSE}$	0.032	0.031	0.033

The coefficients are cross-sectional averages. Parentheses indicate *p*-values. Null hypothesis: insignificant effect of coefficient. Mean squared error (σ^2) is abbreviated as MSE. ** Reject H_0 when the *p*-value < 0.0500. * Reject H_0 when the *p*-value < 0.1000.

These findings are in line with extant studies that find governance has a positive effect on financial development (see, for example, Law & Azman-Saini, 2012b; Le et al., 2015). Improved governance quality implies better enforcement of contractual and property rights, an effective legal system, and efficient financial regulation. These improvements create an environment that promotes rapid financial development.

5 Robustness Checks

Table 9 provides the estimates of long-run equations (1-3) for four country groups by income levels in 2013, as classified by the World Bank (Table A6 shows the composition of countries belonging to each group). As shown, the financial development effect on governance is seen only in upper-middle- and high-income economies. Furthermore, the positive impact of financial development on governance indicators is greater for upper-middle-income countries than high-income countries (Panels C and D). In low- and lower-middle-income countries, the effect of financial development is insignificant, with one exception—financial development worsens investment profiles in lower-middle-income economies (Panel B).

The results indicate the relationship between financial development and governance depends on a country's level of development. Financial development leads to improvements in governance in countries with higher levels of income (and development). These results underscore the importance of the level of development in supporting good governance. Countries with lower GDP per capita may not have sufficient resources to support and improve their governance structures (Fosu, Bates, & Hoeffler, 2006). Moreover, the demand for good governance is likely to be lower at lower levels of development, as most transactions are still managed through personal ties and repeated transactions (Law & Azman-Saini, 2012). This is also in accordance with the median voter hypothesis (Milanovic, 2000). In countries with low per capita income, the poor median voter is more concerned about making "ends meet" and less concerned about the quality of institutions. However, at higher levels of economic development (i.e., higher per capita income), the median voter is better off and is now more aware of or concerned about the quality of institutions and governance. Rising income inequality that is often associated with growth—the Kuznets hypothesis—is also cited as a possible reason for worsening governance at low levels of development, while improving it at high levels of development when inequality declines (Chong & Calderón, 2000). As such, the positive effects on governance quality of improvement in the level of financial development only appear to kick in at the later stages of economic development.

Table 9: Long-run equations (1-3) by country groups of income level, CCE estimates

Panel A: Low-income countries (GDP Per capita ≤$995)			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	2.129	-4.102	0.806
	(0.718)	(0.163)	(0.788)

KOG	-0.064	-0.004	-0.005
	(0.304)	(0.962)	(0.903)
GPC	0.038**	0.003	0.015**
	(0.001)	(0.630)	(0.002)
FD	11.674	-9.410	1.417
	(0.446)	(0.120)	(0.863)
$\sqrt{MSE}$	1.313	0.658	0.701
Panel B: Lower middle-income countries (GDP Per capita ≥$996 and ≤$3,945)			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	-0.533	2.298	-1.681
	(0.894)	(0.739)	(0.373)
KOG	-0.041	0.055	-0.053
	(0.627)	(0.573)	(0.215)
GPC	0.004	-0.005	1.94×10^{-4}
	(0.357)	(0.573)	(0.911)
FD	11.978	**-15.754***	3.112
	(0.346)	**(0.057)**	(0.560)
$\sqrt{MSE}$	1.396	0.473	0.752
Panel C: Upper middle-income countries (GDP Per capita ≥$3,946 and ≤$12,195)			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	-4.083	-2.349	-1.188
	(0.732)	(0.683)	(0.572)
KOG	-0.124	-0.040	-0.101**
	(0.315)	(0.443)	(0.037)
GPC	0.002	0.001	0.001
	(0.228)	(0.105)	(0.120)
FD	**13.602***	**6.202****	**6.470***
	(0.089)	**(0.045)**	**(0.056)**
$\sqrt{MSE}$	1.033	0.464	0.800
Panel D: High-income countries (GDP Per capita ≥$12,196)			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	-4.608	-0.551	-3.528
	(0.596)	(0.824)	(0.347)
KOG	0.138	-0.075*	0.099
	(0.308)	(0.055)	(0.183)
GPC	2.59×10^{-4}	1.44×10^{-4}**	-4.52×10^{-5}
	(0.349)	(0.022)	(0.715)
FD	**7.978****	**2.722***	**3.611***
	(0.042)	**(0.053)**	**(0.077)**
$\sqrt{MSE}$	0.785	0.594	0.488

The coefficients are cross-sectional averages. Parentheses indicate *p*-values. Null hypothesis: insignificant effect of coefficient. Mean squared error (σ^2) is abbreviated as MSE. ** Reject H_0 when the *p*-value < 0.0500. * Reject H_0 when the *p*-value < 0.1000.

Table 10 shows the impact of financial development on governance indicators in countries grouped by level of globalization (Table A7 shows the composition of countries belonging to each group). The groups of economies are classified by quartiles of *KOG* values. For *KOG* quartile 1 (Panel A, Table 10), the effects of *FD* on *GOV*, *IP,* and *GS* are positive, but none are statistically significant. For *KOG* quartile 2 (Panel B, Table 10), *FD* has negative coefficients for all three regressands. However, the negative coefficient of *FD* is only significant when the regressand is *IP*. Finally, for quartiles 3 and 4 of *KOG* in Panels C and D, *FD* has a positive and statistically significant impact on all three governance indicators and *GOV* and *GS* on *FD*. This provides further support for our finding that financial development can improve governance. In addition, the findings in Table 10 are virtually identical to those in Table 9, indicating the governance improving effect of financial development is operative only at high levels of income and globalization. One possible explanation is that a country that is more integrated in the world economy is at higher risk of facing capital flight due to poor governance. Hence, higher levels of globalization may work as a catalyst to trigger institutional and governance reforms in response to increased financial development (Mishkin, 2009). Possible channels of such reforms may include spillover of ideas, information, and technology, as well as improving citizens' affluence as the economy commits fully to integrating with the rest of the world (Shahbaz, Shafiullah, & Mahalik, 2019). In addition, a higher level of globalization involves interdependence between nations, culminating in alignment of economic policy and institutional reforms (Waltz, 1999).

We further test the sensitivity of our results by employing two different measures of financial development that are commonly used in the literature. These two measures are deposit money bank assets to (deposit money + central) bank assets (%) and private credit by deposit money banks to GDP (%); both of these are classified as measures of financial depth (Beck, Demirgüç-Kunt, & Levine, 1999; Beck, Demirgüç-Kunt, & Levine, 2010; Čihák, Demirgüç-Kunt, Feyen, & Levine, 2012). These measures have been widely used in the literature as proxies for financial development (see, e.g., Beck, Levine, & Loayza, 2000; Demirgüç-Kunt & Detragiache, 1998; King & Levine, 1993a; King & Levine, 1993b). The results of this exercise are provided in Appendix Tables A2, A3, and A4. These results are qualitatively similar to those reported using Svirydzenka's (2016) financial development index.

For instance, the long-run estimates in Table A2 reveal that both alternative measures of financial development positively affect governance quality, consistent with our earlier findings reported in Table 6. Further, Table A3 confirms the bidirectional causal relationship between financial development and governance quality when we use the additional financial development measures. As such, the results in Table A3 indicate that bank assets to (deposit money + central) bank assets (%) and private credit by deposit money banks to GDP (%) both cause improvements in governance quality. Last, Table A4 confirms that improvements in governance quality can also lead to improvements in the levels of the two alternative financial development measures; these results are robust to using all three measures of governance.

Table 10: Long-run equations (1-3) by country groups of globalization level, CCE estimates

Panel A: *KOG* quartile 1			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	-1.787	-1.691	0.826
	(0.647)	(0.142)	(0.703)
KOG	-0.086	-0.081**	0.030
	(0.147)	(0.035)	(0.430)
GPC	0.018**	0.006**	0.004
	(0.003)	(0.002)	(0.143)
FD	10.983	2.663	8.288
	(0.433)	(0.750)	(0.249)
$\sqrt{MSE}$	1.354	0.709	0.723
Panel B: *KOG* quartile 2			
	Regressand		
Regressor	*GOV*	*IP*	*GS*
Constant	-0.395	3.334	-1.934
	(0.918)	(0.518)	(0.207)
KOG	0.036	0.061	-0.041
	(0.632)	(0.205)	(0.290)
GPC	0.005**	-0.004	0.002
	(0.048)	(0.293)	(0.118)
FD	12.371	**-9.402***	2.906
	(0.122)	**(0.096)**	(0.507)
$\sqrt{MSE}$	1.293	0.490	0.738
Panel C: *KOG* quartile 3			
	Regressand		

Regressor	GOV	IP	GS
$Constant$	-18.455**	-3.835	-3.805
	(0.004)	(0.327)	(0.352)
KOG	-0.039	0.021	9.35×10^{-4}
	(0.722)	(0.611)	(0.985)
GPC	-1.56×10^{-4}	4.49×10^{-4}	-2.82×10^{-4}
	(0.843)	(0.335)	(0.513)
$\boldsymbol{FD}$	**12.157****	**4.320****	**7.135****
	(0.037)	**(0.048)**	**(0.033)**
$\sqrt{MSE}$	0.999	0.447	0.505

Panel D: KOG quartile 4

	Regressand		
Regressor	GOV	IP	GS
$Constant$	-3.084	0.310	2.140
	(0.640)	(0.929)	(0.596)
KOG	-0.057	-0.096*	0.004
	(0.710)	(0.073)	(0.958)
GPC	3.31×10^{-4}*	2.34×10^{-4}**	1.85×10^{-4}*
	(0.083)	(0.007)	(0.066)
$\boldsymbol{FD}$	**5.037***	**2.748***	**2.629***
	(0.083)	**(0.085)**	**(0.086)**
$\sqrt{MSE}$	0.929	0.525	0.592

The coefficients are cross-sectional averages. Parentheses indicate p-values. Null hypothesis: insignificant effect of coefficient. Mean squared error (σ^2) is abbreviated as MSE. ** Reject H_0 when the p-value < 0.0500. * Reject H_0 when the p-value < 0.1000.

6 Conclusion

This study is the first attempt to conduct a cross-country analysis of the bidirectional causality between financial development and governance. Further, it tests whether the relationship between financial development and governance varies by level of development and government openness, which, to the best of our knowledge, has not been explored in extant literature. The tests for bidirectional causality between financial development and governance are conducted using cross-sectional dependent panel data of 101 countries spanning more than 30 years. Using econometric methods robust to cross-sectional dependence, we find the specified models to be cointegrated. Long-run equation estimates show that financial development positively affects governance and is robust to different measures of governance quality. Further analysis shows governance quality positively affects financial development, which is consistent with extant literature. The Granger causality tests demonstrate bidirectional causality between financial development and the different governance measures. Last, the impact of financial development is dependent on a country's level of economic development and economic openness.

The empirical findings highlight the pivotal role economic development plays through the financial sector in introducing governance reforms. The improvements in governance, in turn, accelerate financial sector development. As such, there can be a symbiotic, as well as synergistic, relationship among good governance, growth, and economic development. Our results have important policy implications, especially for developing countries, as their financial systems tend to be underdeveloped and well below the global finance frontier. First, governance and institutional reforms are not prerequisites for financial development. However, financial development can drive governance reforms because of governments' ease of access to finance for undertaking such reforms.

Second, improving financial development first may prevent a backlash to institutional and governance reforms, as politicians will face more competition and be less likely to reverse these reforms. A strong financial system will limit relationship banking, which favors loans to friends and relatives at the expense of more productive and profitable commercial lending. As a result, competition in the political arena will improve as politicians see a decline in their de-facto political power. Last, to capitalize further on the positive benefits accruing from financial development, countries should focus on improving their economic and financial liberalization in tandem. Economic and financial liberalization, together with financial development, may trigger far-reaching and deep-rooted institutional and governance reforms that can eventually set a low-income country on a high-growth trajectory.

Although our study highlights a robust link between financial development and good governance at the macro level, the analysis can be extended using micro-data or by conducting a survey to provide further insights into the financial development and governance nexus. Another possible extension of our work at the micro-level is to study how the availability of micro-credit improves the level of governance, especially in rural areas. Further, our data spans a period of only 30 years, hence limiting our ability to introduce non-linearities in analyzing the association between financial development and the quality of governance. With the availability of more data, future studies can extend the analysis to consider any non-linearities in the relationship.

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Appendix

Table A1: Lluís Carrion-i-Silvestre et al. (2005) panel structural break test

	Fixed number of breaks (H_0: m=0 vs H_A: m=1) **		Unknown number of breaks (H_0: m=0 vs H_A: m=?) **		
Variable	Test statistic (*supF*)	Critical value (5%)	Test statistic (*WDmax*)	Critical value (5%)	Break date
GOV	6.8965517	11.470000	6.8965517	12.810000	N/A
IP	6.8571429	11.470000	6.8571429	12.810000	N/A
GS	6.8965517	11.470000	6.8965517	12.810000	N/A
KOG	1.7419355	11.470000	1.7419355	12.810000	N/A
GPC	3.4666667	11.470000	3.4666667	12.810000	N/A
FD	6.8965517	11.470000	6.8965517	12.810000	N/A

** Reject the respective H_0: series contains a unit root process without the specified no. of breaks (m), if test statistic > critical value at the 5% level.

Table A2: Long-run equations (1-3), alternative financial development measures, CCE estimates

Regressor	Regressand *GOV*	*IP*	*GS*	*GOV*	*IP*	*GS*
Constant	-3.062	16.393*	-1.044	-2.546	-3.703	0.141
	(0.436)	(0.053)	(0.648)	(0.531)	(0.579)	(0.947)
KOG	0.020	-0.245	-0.037	-0.002	-0.027	-0.045
	(0.898)	(0.297)	(0.671)	(0.989)	(0.828)	(0.572)
GPC	3.36×10^{-4}*	7.22×10^{-5}	1.28×10^{-4}	-3.51×10^{-4}**	2.63×10^{-4}	-1.52×10^{-4}*
	(0.056)	(0.839)	(0.189)	(0.039)	(0.744)	(0.086)
DEPOSIT MONEY BANK ASSETS to (DEPOSIT MONEY + CENTRAL) BANK ASSETS (%)	**0.123*** **(0.076)**	**0.264**** **(0.045)**	**0.076**** **(0.047)**	-	-	-
PRIVATE CREDIT BY DEPOSIT MONEY BANKS to GDP (%)	-	-	-	**0.112**** **(0.046)**	**0.240**** **(0.011)**	**0.063**** **(0.034)**
$\sqrt{MSE}$	1.136	0.280	0.603	1.128	0.000	0.639

The coefficients are cross-sectional averages. Parentheses indicate *p*-values. Null hypothesis: insignificant effect of coefficient. Mean squared error (σ^2) is abbreviated as MSE. ** Reject H_0 when the *p*-value < 0.0500. * Reject H_0 when the *p*-value < 0.1000.

Table A3: Dumitrescu and Hurlin (2012) Granger causality test results, alternative financial development measures

Causality direction	Test statistics		p-value
	$\overline{W}$	$\overline{Z}$	
$DEP \Rightarrow GOV$	3.53138***	5.29631***	0.0000
$GOV \Rightarrow DEP$	4.13937***	7.75000***	0.0000
$DEP \Rightarrow GS$	3.78170***	6.30653***	0.0000
$GS \Rightarrow DEP$	3.35926***	4.60165***	0.0000
$DEP \Rightarrow IP$	3.44808***	4.96011***	0.0000
$IP \Rightarrow DEP$	4.38433***	8.73862***	0.0000
$PRI \Rightarrow GOV$	3.36195***	4.60415***	0.0000
$GOV \Rightarrow PRI$	4.11950***	7.65858***	0.0000
$PRI \Rightarrow GS$	3.87948***	6.69082***	0.0000
$GS \Rightarrow PRI$	3.19229***	3.92009***	0.0000
$PRI \Rightarrow IP$	3.38052***	4.67905***	0.0000
$IP \Rightarrow PRI$	4.16866***	7.85678***	0.0000

Selected lag order: 2. Null hypothesis: Lack of causality in the direction. *** Reject H_0 when the p-value < 0.01000. *DEP* stands for deposit money bank assets to (deposit money + central) bank assets (%). *PRI* stands for private credit by deposit money banks to GDP (%).

Table A4: Impact of governance on alternative financial development measures, CCE estimates

	Regressand					
Regressor	*DEP*	*DEP*	*DEP*	*PRI*	*PRI*	*PRI*
Constant	15.188	18.838	5.890	18.209	20.924	12.234
	(0.277)	(0.181)	(0.705)	(0.324)	(0.195)	(0.504)
KOG	-0.661*	-0.290	-0.380	-0.800	-0.450	-0.589
	(0.069)	(0.380)	(0.333)	(0.287)	(0.516)	(0.398)
GPC	-0.001**	-0.002**	-0.001**	-0.001	-0.001*	-0.001*
	(0.018)	(0.003)	(0.045)	(0.138)	(0.081)	(0.070)
GOV	**0.937****	-	-	**0.936***	-	-
	(0.004)			**(0.053)**		
IP	-	**1.522****	-	-	**2.284****	-
		(0.005)			**(0.043)**	
GS	-	-	**1.523****	-	-	**1.777****
			(0.001)			**(0.031)**
$\sqrt{MSE}$	5.547	5.488	5.636	6.681	6.450	6.380

The coefficients are cross-sectional averages. Parentheses indicate p-values. Null hypothesis: insignificant effect of coefficient. Mean squared error (σ^2) is abbreviated as MSE. ** Reject H_0 when the p-value < 0.0500. * Reject H_0 when the p-value < 0.1000.*DEP* stands for deposit money bank assets to (deposit money + central) bank assets (%). *PRI* stands for private credit by deposit money banks to GDP (%).

Table A5: List of countries in the sample

Albania	Ecuador	Luxembourg	Senegal
Algeria	Egypt	Madagascar	Sierra Leone
Argentina	El Salvador	Malawi	Singapore
Australia	Ethiopia	Malaysia	South Africa
Austria	Finland	Mali	South Korea
Bahamas	France	Malta	Spain
Bahrain	Gabon	Mexico	Sudan
Bangladesh	Gambia	Mongolia	Suriname
Belgium	Germany	Morocco	Sweden
Bolivia	Ghana	Mozambique	Switzerland
Botswana	Greece	Myanmar	Thailand
Brazil	Guatemala	Netherlands	Togo
Brunei	Guinea-Bissau	New Zealand	Trinidad & Tobago
Bulgaria	Guyana	Nicaragua	Tunisia
Burkina Faso	Honduras	Niger	Turkey
Cameroon	Iceland	Nigeria	UAE
Canada	India	Norway	Uganda
Chile	Indonesia	Oman	United Kingdom
China	Iran	Pakistan	United States
Colombia	Ireland	Panama	Uruguay
Congo	Israel	Papua New Guinea	Venezuela
Costa Rica	Italy	Paraguay	Vietnam
Cote d'Ivoire	Jamaica	Peru	Zambia
Cyprus	Japan	Philippines	
Denmark	Jordan	Portugal	
Dominican Republic	Kenya	Saudi Arabia	

Table A6: List of Countries According to World Bank Income Classification

Low-income	Lower-middle-income	Upper-middle-income	High-income	
Bangladesh	Bolivia	Albania	Australia	Saudi Arabia
Burkina Faso	Cameroon	Algeria	Austria	Singapore
Ethiopia	Congo	Argentina	Bahamas	South Korea
Gambia	Cote d'Ivoire	Botswana	Bahrain	Spain
Guinea-Bissau	Egypt	Brazil	Belgium	Sweden
Madagascar	El Salvador	Bulgaria	Brunei	Switzerland
Malawi	Ghana	China	Canada	Trinidad & Tobago
Mali	Guatemala	Colombia	Chile	UAE
Mozambique	Guyana	Costa Rica	Cyprus	United Kingdom
Niger	Honduras	Dominican Republic	Denmark	United States
Sierra Leone	India	Ecuador	Finland	Uruguay
Togo	Indonesia	Gabon	France	Venezuela
Uganda	Kenya	Iran	Germany	
	Mongolia	Jamaica	Greece	
	Morocco	Jordan	Iceland	
	Myanmar	Malaysia	Ireland	
	Nicaragua	Mexico	Israel	
	Nigeria	Panama	Italy	
	Pakistan	Peru	Japan	
	Papua New Guinea	South Africa	Luxembourg	
	Paraguay	Suriname	Malta	
	Philippines	Thailand	Netherlands	
	Senegal	Tunisia	New Zealand	
	Sudan	Turkey	Norway	
	Vietnam		Oman	
	Zambia		Portugal	

Table A7: List of Countries According to Quartiles of KOF Globalization Index

1st Quartile	2nd Quartile	3rd Quartile	4th Quartile
Albania	Algeria	Argentina	Australia
Bangladesh	Bahamas	Bahrain	Austria
Burkina Faso	Bolivia	Brazil	Belgium
Cameroon	Botswana	Brunei	Canada
Congo	China	Bulgaria	Denmark
Cote d'Ivoire	Colombia	Chile	Finland
Ethiopia	Dominican Republic	Costa Rica	France
Guinea-Bissau	Ecuador	Cyprus	Germany
India	Egypt	Israel	Greece
Iran	El Salvador	Jamaica	Iceland
Kenya	Gabon	Japan	Ireland
Madagascar	Gambia	Jordan	Italy
Malawi	Ghana	Malta	Luxembourg
Mali	Guatemala	Mexico	Malaysia
Mongolia	Guyana	Oman	Netherlands
Mozambique	Honduras	Panama	New Zealand
Myanmar	Indonesia	Saudi Arabia	Norway
Niger	Morocco	South Africa	Portugal
Pakistan	Nicaragua	South Korea	Singapore
Papua New Guinea	Nigeria	Thailand	Spain
Sierra Leone	Paraguay	Trinidad & Tobago	Sweden
Sudan	Peru	Tunisia	Switzerland
Suriname	Philippines	Turkey	UAE
Togo	Senegal	Uruguay	United Kingdom
Uganda	Zambia	Venezuela	United States
Vietnam			