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Financial development under the shade of globalization and financial institutions: the case of Pakistan

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This study investigates the importance of financial institutions, net capital inflows, and trade openness for financial-sector development in a small developing economy like Pakistan. Two approaches (Johansen test and autoregressive distributive lag approach) were employed for the robustness of long-run relationships among the variables under consideration and found that both techniques provide robust results for long-run relationships, in Pakistan's case. Net capital in inflows has positive impact on financial-sector development in the long run. Trade openness is the main promoter of financial development in both periods. Finally, financial institutions and economic growth also help to improve the development of the financial sector. Further, it examined Rajan and Zingales's [2003a] hypothesis that predicts combined influence of capital account liberalization and trade openness on financial-sector development. However, such relationship does not exist in the case of Pakistan.

In terms of policy implications, our findings suggest that macroeconomic management, which could simultaneously stimulate foreign capital inflows and trade openness, improves quality of financial institutions, and high economic growth in the country would enhance the performance of both capital and financial intermediaries.

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1. Introduction

The performance of an economy is affected by a country's well-functioning financial sector. Financial sector is modeled as lowering the search costs and increasing the level of external finance in the economy. Financial development transfers the incentives of producers toward the good with increasing returns to scale, the intersectoral specialization, and therefore the structure of trade flows is determined by relative level of financial intermediation (Goldsmith [1969]; King and Levine [1993]; Demirgüç-Kunt and Maksimovic [1999]; Jayaratne and Strahan [1996]; Wurgler [2000]; Rajan and Zingales [1998a, 1998b]; Butt et al. [2006]; Akmal, Ahmad, and Butt [2006]). The relationship between financial development and trade has been investigated in two different ways in the literature, which can be identified as the supply side and the demand side.[1] Rajan and Zingales [2003a] focus on the role of supply-side interest groups, especially the vested interests of incumbent industrialists and financial intermediaries. Incumbents, worried about the threat of new entrants, have a lot of reasons to resist the development of the financial sector. These incentives are weakened if the economy becomes more open to foreign competition or to internal flows of capital. According to this model, goods-market openness can improve the supply of external finance, as it aligns the interests of the economically powerful more closely with financial development. In contrast, Svaleryd and Vlachos [2003] emphasize the role of risk diversification. To the extent that openness to trade is correlated with greater risks, such increased exposure to external demand shocks or foreign competition will create new demands for external finance. Firms will need credit to overcome short-run cash-flow problems and adverse shocks. This shows that the effects of trade on finance are likely to work primarily through the demand side.[2] Yongfu Huang and Temple [2005] report that if trade liberalization is followed by high projects of investment and lending booms, there could be a strong relationship between trade openness and finance in the short run, whether or not there is an association in the long run; they also claim that openness has greater link

[1] Financial development can influence the extent of international trade in goods and services through two channels. The first operates through risk-sharing and production specialization. The second channel relies on the ability of the financial sector to divert savings to the private sector. When domestic intermediation is weak and inefficient, firms in export-oriented sectors are burdened by significant liquidity constraints and hence trade less. Financial development can help overcome these constraints by making more external finance available to domestic firms [Carmignani and Chowdhury 2005]. For financial development, see Levine, Loayza, and Beck [2000]; for openness, see Frankel and Romer [1999].

[2] International trade brings about substantial changes in competition, technology, price shocks of intermediary and final goods, and in the long run even in factor endowments and the institutional feature of a society.

with bank-based finance than with stock market development (Beck [2002]; Svaleryd and Vlachos [2000]).[3]

Openness to trade may also influence the demand for external finance through the nature of specialization and sectoral structure, or through the space of innovation and technology transfer—activities that are likely to make intensive use of external finance. Alessandria and Qian [2005] utilize a formal model to argue that capital account liberalization has ambiguous effects on the efficiency of domestic financial intermediaries. The empirical works of Aizenman and Noy [2003], Aizenman [2004], and Chin and Ito [2005] suggest that capital account liberalization is often preceded by goods-market openness, perhaps because trade integration imposes restrictions on capital flows, which may be difficult to sustain. Levine and Renelt [1992] confirm a robust link between openness and share of investment in the gross domestic product (GDP); and if trading economies are also high investment economies, this could promote financial development.[4]

Considering the determinants of financial development, several studies have investigated the impacts of legal, regulatory, and macroeconomic environment on the banking sector and equity markets. La Porta et al. [1998] conclude that origins of the legal code are vital for financial depth because legal systems differ in their treatment of creditors and shareholders, and in contract enforcement. Rajan and Zingales [2003b] point out that government regulation is needed to ensure effective contract enforcement, and transparency in accounting and disclosure. Regulations concerning information disclosure, accounting standards, permissible practice of banks and despite insurance do appear to have material effects on financial development [Mayer and Sussman 2001]. Huybens and Smith [1999] examine theoretically, and Boyd, Levine, and Smith [2001] empirically, the effects of inflation on financial depth. They also conclude that economies with higher inflation rates are likely to have smaller, less active, and less efficient banks and equity markets. Financial development can, in turn, influence/affect the structure and extent of trade. Recently, Beck [2002] investigated this issue in detail. Drawing on the arguments of Kletzer and Bardhan [1987], Beck [2002] develops a model in which countries with well-developed financial markets will tend to have comprehensive advantage in manufacturing.[5] Becker and Greenberg [2004] argue that in countries with good

[3] One mechanism here may be information asymmetries. To the extent that performance in export market is a useful and observable index of a country's productivity, increased openness can reduce information asymmetries among different banks.

[4] Using stock market for emerging economies, Li et al. [2004] find that goods-market openness is associated with greater marketwide variations, but not greater firm-specific variations.

[5] Recent works by Ju and Wei [2005] and Wynne [2005] also develop connections between financial development and trade specialization.

financial sector, the short-run elasticity of exports is considerably higher, and the allocation of exports across importing countries is much more sensitive to relative exchange rates. Rajan and Zingales [2003a] find that industries which are more in need of outside financing grow more quickly when the financial sector is more developed.

The relationship between institutions and trade openness has been receiving considerable attention. The growing literature focuses on how quality of institutions can affect comparative advantage and international specialization [Becker and Greenberg 2004]. Grossman and Helpman [2004] argue how institutional problems limit the amount of outsourcing to low-wage countries. Levchenko [2003] models how institutions generate comparative advantage, and find that net exports in industries that depend on external finance are higher in countries with good financial sector.[6] There has also been debate on the precise channel through which a country's institutional inheritance affects financial-sector development (Berglof and Von Thadden [1999]; Coffee [2000]; La Porta et al. [1999]; Rajan and Zingales [1998a, 1998b]; Stulz and Williamson [2001]). Moreover, Beck, Demirgüç-Kunt, and Levine [2003a] extend the emphasis of empirical analysis to the settler-mortality hypothesis and find that the initial endowments hypothesis explains more of the cross-variation in financial intermediary and stock market capitalization. Beck, Demirgüç-Kunt, and Levine [2003b] focus their attention on the historical determinants of financial development and do not investigate any of the intermediate linkages. It could be, for example, that the relationship between initial endowments and subsequent financial development reflects factors other than the development of institutions conducive to financial development. While a recent study conducted by Law and Demetriades [2005] found that simultaneous openness to both trade and capital inflows has a positive influence on financial development, in tandem with the [institutions] hypothesis, the quality of a country's institutions has a separate influence on financial development.[7] This study also provides a direct test of both openness and institutional development hypothesis using appropriate specified financial equations.

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[6] Svaleryd and Vlachos use the industry measure introduced by Rajan and Zingales, which was initially shown to capture how much an industry's growth rate depends on financial development. Hence, Svaleryd and Vlachos essentially establish that the faster growth in finance-dependant industries located in financially well-developed countries actually results in higher exports (rather than domestic consumption).

[7] Svaleryd and Vlachos [2000] suggest that development of institutions is necessary for risk diversifications, e.g., financial markets, to reduce barriers to trade. In contrast, Roderick [1998] provides evidence that openness to trade also increases the permanent degree of income volatility in an economy.

The contribution of institutions and openness to improve financial-sector development in economic growth in Pakistan has not been investigated with respect to concerned direction. This study is a first attempt in this particular direction in economic development. It aims to investigate the impact of institutions and openness on financial-sector development in an autoregressive distributive lag (ARDL) framework for Pakistan by utilizing quarterly data over the period 1980-2004. The residual-based cointegration tests are considered inefficient and can lead to contradictory results, especially when there are more than $I(1)$ variables under consideration. This ARDL modeling approach is superior to other methods for analysing the long-run relationships when the variables are in mixed order of integration—i.e., $I(0)$ and $I(1)$—and when there is structural break in the time-series data. The rest of the paper is designed as follows: section 2 explains the model, methodology and data collection procedure, section 3 presents the empirical results. Finally, section 4 presents the conclusion and policy implications.

2. Empirical model and methodology

Based on the model specified by Law and Demetriades [2005] in log-linear form for financial development:

$$LFD = \beta_\circ + \beta_1 LRGDPC + \beta_2 INS + \beta_3 LCIF + \beta_4 LTR + \mu_t \tag{1}$$

where FD is utilized as an indicator of financial development proxied by liquid liability as share of GDP; $RGDPC$ is real per capita GDP; INS is financial institutions; CIF is the net capital inflow, which is proxy for capital account openness; and TR is measure of trade openness (Exports + Imports/GDP). To directly investigate the hypothesis generated by Rajan and Zingales [2003a], an interaction term between the last two variables is included in the model as follows:

$$\begin{aligned} LFD = {} & \alpha_\circ + \alpha_1 LRGDPC + \alpha_2 INS + \alpha_3 LCIF \\ & + \alpha_4 LTR + \alpha_5 L(CIF \times TR) + \eta_t \end{aligned} \tag{2}$$

Equations (1) and (2) provide a basis for the empirical models in the present study. If α_5 is found to be positive and statistically significant, then this would imply that a combination of capital inflows and trade openness exerts an independent influence on financial-sector development, over and above any separate influence each of these two variables may independently and separately have on financial development. Thus, $\alpha_5 > 0$ supports the Rajan and Zingales [2003a] hypothesis.

In the time-series data, realization is used to draw inference about the underlying stochastic process. Therefore, to draw inference from the time-series analysis, stationary tests become essential. A stationary test, which has been widely popular over the past several years, is the unit root test. In this study, the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests were applied to estimate unit root.

2.1. ADF unit root test

In this study, the ADF test is applied to estimate the unit root. The ADF test to check the stationarity series is based on the following equation:

$$\Delta y_t = \beta_1 + \beta_2 t + \delta y_{t-1} + \alpha_i \sum_{t=1}^{m} \Delta y_{t-1} + \varepsilon_t \tag{3}$$

where ε_t is a pure white noise error term and

$$\Delta y_{t-1} = (y_{t-1} - y_{t-2}),\ \Delta y_{t-2} = (y_{t-2} - y_{t-3}) \text{ etc.}$$

This test determines whether the estimates of δ are equal to zero. Dickey and Fuller [1979] provided cumulative distribution of the ADF statistics. If the calculated ratio (value) of the coefficient δ is less than τ critical value from Fuller table, then y is said to be stationary.[8]

2.2. Phillips-Perron test

The distribution theory that supports the Dickey-Fuller tests assumes that the errors are statistically independent and have constant variance. Phillips and Perron [1988] developed the generalization of the Dickey-Fuller procedure that allows for fairly mid-assumptions concerning the distribution of the errors. Thus the Phillips-Perron test allows the disturbance to be weakly dependent and heterogeneously distributed. In this case the regression equations are as follows:

$$y_t = \alpha_1^* + \alpha_2^* y_{t-1} + \mu_t$$

$$y_t = \alpha_1 + \alpha_2 y_{t-1} + \alpha_3 (t - n/2) + \mu_t$$

n = number of observations

$\mu_t = E(\mu_t) = 0\ldots$ but there is no requirement that the disturbance term is serially uncorrelated or homogenous.

The hypothesis in this case is $\alpha_2^* = 1, \alpha_2 = 1$ and $\alpha_3 = 0$.

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[8] 't' ratio of coefficient is always with negative signs.

Data on liquid liabilities are obtained from the various issues of the State Bank of Pakistan's Quarterly Statistical Bulletins and International Financial Statistics [2006]. Quarterly extended data on GDP are obtained from a working paper by Kemal and Arby [2004] and converted to real GDP per capita based on 1980 constant prices. The institution data sets employed in the present study are generated by the authors. A higher value of index implies better institutional quality and vice versa. The proxy for net capital inflows is utilized to investigate whether capital inflows have any impact on financial development, namely private capital inflows.[9] The data on net capital inflows are available in the World Development Indicators [2006]. Trade openness is proxied by the sum of imports and exports divided by GDP; the data of this variable is available in the International Financial Statistics [2006]. Finally, the interaction term between capital inflows and trade openness can be quantified since this has a positive impact on financial development as highlighted in the theory. Table 1 provides descriptive statistics and correlation matrix of variables.

Table 1. Descriptive statistics and correlation matrix

Variables	*Liquid liability*	*Net capital inflows*	*Trade openness*	*Institutions*	*Real GDP per capita*
Mean (million Rs)	1.5760	9.024	-0.064	104.217	-25.728
Median	1.6250	9.221	-0.028	103.881	-92.915
Maximum	2.7345	11.213	1.266	113.242	11468.400
Minimum	0.4110	5.937	-1.198	99.356	-8815.320
Standard deviation	0.6980	1.427	0.688	2.775	1550.870
Skewness	0.0790	-0.484	-0.006	0.494	2.415
Kurtosis	1.6060	2.132	1.728	2.763	44.042
Jarque-Bera	7.7070	6.618	6.338	4.049	6688.890
Probability	0.0210	0.037	0.042	0.132	0.000
Observations	96.0000	96.000	96.000	96.000	96.000
Liquid liability	1.0000				
Net capital inflows (million US dollars)	0.9450	1.000			
Trade openness	0.9780	0.956	1.000		
Institutions	0.9640	0.908	0.949	1.000	
Real GDP per capita (Rs)	0.0340	-0.009	-0.021	0.023	1.000

[9] Net private capital inflows consist of private debt and nondebt flows. Private debt flows include commercial bank lending, bonds, and other private credits; nondebt private flows are FDI and portfolio equity investment. Data are in current US dollars.

Correlations are described as net capital inflows and trade openness, and institutions are positively and strongly correlated with liquid liability (financial development) while real GDP per capita is weakly associated with financial-sector development. Net capital inflow is also highly correlated with trade openness and institutions, while institutions and trade associate highly and positively with each other. Finally, real GDP per capita is correlated weakly and negatively with capital inflows and trade openness, but positively with institutions.

This study utilizes the two alternative approaches for the stable long-run relationship among the variables.

2.3. Johansen cointegration test

When all the variables are nonstationary at their level but stationary in their 1st difference, this means to proceed further to the Johansen cointegration test. Economically speaking, two variables will be cointegrated if they have a long-term relationship. Thus cointegration of two series suggests that there is a long-run integration test; the system approach developed by Johansen [1991, 1995], of course, can also be applied to a set of variables containing possibly a mixture of $I(0)$ and $I(1)$ [Pesaran, Shin, and Smith 1998, 2001]. The general form of the vector error correction model (ECM) is as follows:

$$\Delta y_t = \alpha_\circ + \alpha_1 t - \Pi z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i' \Delta z_{t-1} + \psi w_t + \mu_t \tag{4}$$

where $z_t = (y_t', x_t') \ldots y_t$ is an $m_{yX}1$ vector of endogenous $I(1)$ var x_t is an $m_{xX}1$ vector of exogenous of exogenous $I(1)$ variables.

$\Delta x_t = \alpha_\circ + \sum_{i=1}^{p-1} \Gamma_i' \Delta z_{t-1} + \Psi w_t + \mu_t \ldots w_t$ is q*1vector of exogenous /deterministic variables $I(0)$.

In the model, the disturbance vector of et and wt satisfies the following assumptions:

(a) $\mu_t = (e_t w_t) iid(0, \Sigma)$

Σ = a symmetric positive-definite matrix

(b) μt = (the disturbance term in the combined model) distributed independently of w_t = i.e., $\mathrm{E}(\mu_t w_t) = 0$ $\alpha_\circ$ and $\alpha_1 \ldots$ intercept and trend coefficients, respectively.

Π = long-run multiplier matrix, i.e., Π_y multiplier matrix of order $(m_y + m)$ where $m = (m_x + m_y)\,\Gamma'_{1y} - \Gamma'_{p-1,y}$ = coefficient matrices that capture the short dynamic effects and are of order $m_y^* m\,\Psi_y$= the $m_y^* m$ matrix of coefficients on the I(0) exogenous variables.

2.4. ARDL approach for integration

As for the newly proposed ARDL approach (Pesaran and Shin [1999]; Pesaran, Shin, and Smith [1996, 2001]), recent studies have indicated that the ARDL approach to cointegration is preferable to other conventional cointegration approaches (e.g., Engle and Granger [1987]; Hansen [1995]). One of the reasons is that it is applicable regardless of whether the underlying regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. The statistic underlying this procedure is the familiar Wald or F-statistic in a generalized Dickey-Fuller-type regression, which is used to test the significance of lagged levels of the variables under consideration in a conditional unrestricted equilibrium ECM [Pesaran, Shin, and Smith 2001]. Another reason for using the ARDL approach is that it is more robust and performs better for small sample sizes (such as in this study) than other cointegration techniques.

The ARDL approach involves estimating the conditional error correction version of the ARDL model for the variable under estimation. The augmented ARDL $\left(p, q_1, q_2, \ldots q_k\right)$ is given by the following equation (Pesaran and Pesaran [1997]; Pesaran, Shin, and Smith [2001]):

$$\alpha(L,p)\,y_t = \alpha_\circ + \sum_{i=1}^{k} \beta_i(L,p)\,x_{it} + \lambda' w_t + \varepsilon_t \tag{5}$$

$$\forall t = 1, \ldots\ldots, n$$

where $\alpha(L,p) = 1 - \alpha_1 L - \alpha_2 L^2 - \ldots - \alpha_p L^P$

$$\beta_i(L,q_1) = \beta_{i\circ} + \beta_{i1} L + \beta_{i2} L^2 + \ldots + \beta_{iq_i} L^{q_i} \,\forall_i = 1,2,\ldots,k$$

y_t is an independent variable, $\alpha_\circ$ is the constant term, L is the lag operator such that $Ly_t = y_t - 1, w_t$ is $s \times 1$ vector of deterministic variables such as intercept term, time trends, or exogenous variables with fixed lags.

The long-term elasticities are estimated by

$$\phi_i = \frac{\hat{\beta}_i(1,\hat{q})}{\alpha(1,\hat{p})} = \frac{\hat{\beta}_{i\circ} + \hat{\beta}_{i1} + \ldots + \hat{\beta}_{i\hat{q}}}{1 - \hat{\alpha}_1 - \hat{\alpha}_2 - \ldots - \hat{\alpha}_p} \qquad \forall i = 1,2,\ldots,k \tag{6}$$

where $\hat{p}$, $\hat{q}$ are the selected (estimated) values of $\hat{p}$ and $\hat{q}$, $i = 1, 2, \ldots, k$.

The long-run coefficients are estimated by

$$\pi = \frac{\hat{\lambda}(\hat{p}, \hat{q}_1, \hat{q}_2, \ldots, \hat{q}_k)}{1 - \hat{\alpha}_1 - \hat{\alpha}_2 - \ldots - \hat{\alpha}_p} \tag{7}$$

where $\hat{\lambda}(\hat{p}, \hat{q}_1, \hat{q}_2, \ldots, \hat{q}_k)$ denotes the OLS estimates of λ in equation (5) for the selected ARDL model.

The error correction model linked to the ARDL $(\hat{p}, \hat{q}_1, \hat{q}_2, \cdots, \hat{q}_k)$ can be obtained by writing equation (5) in terms of lagged levels and the first difference of $y_t, x_{1t}, x_{2t}, \ldots, x_{kt}$ and w_t:

$$\Delta y_t = \Delta\alpha_\circ - \alpha(1, \hat{p}) EC_{t-1} + \sum_{i=1}^{k} \beta_{i^\circ} \Delta_{it} + \lambda' w_t \tag{8}$$

$$-\sum_{j=1}^{\hat{p}-1} \alpha^\bullet j \Delta y_{t-1} - \sum_{i=1}^{k} \sum_{j=1}^{\hat{q}_{t-1}} \beta_{ij} \Delta_{i,t-j} + \varepsilon_t$$

where ECM is the error correction model, which is defined as follows:

$$ECM_t = y_t - \alpha - \sum \hat{\beta}_i x_{it} - \lambda' w_t \tag{9}$$

x_t is the k- dimensional forcing variables, which are not cointegrated among themselves. ε_t is a vector of stochastic error terms, with zero means and constant variance-covariance.

The existence of an error-correction term among a number of cointegrated variables implies that changes in dependent variable are a function of both the levels of disequilibrium in the cointegration relationship (represented by the ECM) and the changes in the other explanatory variables. This tells us that any deviation from the long-run equilibrium will reflect back on the changes in the dependent variable in order to force the movement toward the long-run equilibrium [Masih and Masih 1996].

The ARDL approach involves two steps for estimating long-run relationship [Pesaran, Shin, and Smith 2001]. The first step is to investigate the existence of a long-run relationship among all variables in the equation under estimation. The second step is to estimate the long-run and short-run coefficients of the same equation. We run the second step only if we find a long-run relationship in the first step [Narayan and Smyth 2004]. This study uses a more general formula of ECM with unrestricted intercept and unrestricted trends [Pesaran, Shin, and Smith 2001]:

$$\Delta y_t = c_\circ + c_1 t + \pi_{yy} y_{t-1} + \pi_{yx.x} x_{t-1} + \sum_{i=1}^{p-1} \psi_i' \Delta z_{t-1} + w' \Delta X_t + \mu_t \tag{10}$$

where $c_\circ \neq 0$ and $c_1 \neq 0$. The Wald test (F-statistics) for the null hypothesis $H_0^{\pi_{yy}} : \pi_{yy} = 0$, $H_0^{\pi_{yx.x}} : \pi_{yx.x} = 0'$, and alternative hypothesis $H_1^{\pi_{yy}} : \pi_{yy} \neq 0$, $H_1^{\pi_{yx.x}} : \pi_{yx.x} \neq 0'$. Hence, the joint null hypothesis of the interest in the above equation is given 135by $H_0 = H_0^{\pi_{yy}} \cap H_0^{\pi_{yx.x}}$, and the alternative hypothesis is correspondingly stated as $H_0 = H_1^{\pi_{yy}} \cap H_1^{\pi_{yx.x}}$.

The asymptotic distributions of the F-statistics are nonstandard under the null hypothesis of no cointegration relationship between the examined variables, regardless of whether the variables are purely $I(0)$ or $I(1)$, or mutually cointegrated. Two sets of asymptotic critical values are provided by Pesaran and Pesaran [1997]. The first set assumes that all variables are $I(0)$ while the second set assumes that all variables are $I(1)$. If the computed F-statistics is greater than the upper-bound critical value, then we reject the null hypothesis of no cointegration and conclude that there exists steady state equilibrium among the variables. If the computed F-statistics is less than the lower-bound critical value, then we cannot reject the null of no cointegration. If the computed F-statistics falls within the lower- and upper-bound critical values, then the result is inconclusive; in this case, following Kremers, Ericsson, and Dolado [1992] and Banerjee, Dolado, and Mestre [1998], the error correction term will be a useful way to establish cointegration. The second step is to estimate the long-run coefficient of the same equation and the associated ARDL error coercion models. Finally, the third step is to investigate the short-run behavior of variables in the long-run equilibrium.

3. Empirical interpretations

Before we proceeded with the ARDL bound test, we tested for the stationarity status of all variables to determine their order of integration. This is to ensure that the variables are not $I(2)$ stationary so as to avoid spurious results. According to Ouattara [2004], in the presence of $I(2)$ variables, the computed F-statistics provided by Pesaran, Shin, and Smith [2001] are not valid because the bound test is based on the assumption that the variables are $I(0)$ or $I(1)$. Therefore, the implementation of unit root tests in the ARDL procedure might still be necessary in order to ensure that none of the variables are integrated of order 2 or beyond.

Augmented Dickey-Fuller and Phillips-Perron (PP) tests were used on each variable. The lag length for the ADF tests was selected to ensure that the residuals were white noise.

Both tests provided the same positions of all variables in the model describing that liquid liabilities, trade, intuitions, and real GDP per capita are stationary at level I(0) as shown in Table 2. All variables in the concerned model are stationary at first difference as mentioned in Table 3. The test regressions included both drift and trend in the investigation of order of integration of variables as discussed in equation (3).

Table 2. Unit-root estimation

Variables	*Level ADF test statistics*	*Lags*	*Level PP test statistics*	*Lags*
Liquid liability	-6.519*	0	-6.514*	3
Trade	-6.292*	1	-6.481*	3
Institutions	-6.122*	0	-6.153*	3
Net capital inflows	-2.400	1	-2.155	3
Real per capita GDP	-4.588*	1	-5.125*	3

* At 1 percent significance level.

Table 3. Unit-root estimation

Variables	*1st Difference ADF test statistics*	*Lags*	*1st Difference PP test statistics*	*Lags*
Liquid liability	-5.237*	0	-16.438*	0
Trade	-5.237*	3	-10.929*	0
Institutions	-10.385*	1	-19.160*	3
Net capital inflows	-5.176*	1	-6.1954*	3
Real per capita GDP	-7.0414*	1	-11.734*	3

* At 1 percent significance level.

After establishing that all the individual series under consideration are stationary, the traditional cointegration method is used to estimate the long-run relationship among variables, particularly in liquid liability, net capital inflows as proxy for capital account openness, institutions, trade as share of GDP, and GDP per capita. As mentioned, the Johansen's maximum likelihood approach is applied for the cointegration test. The results from the Johansen cointegration analysis are summarized in Table 4, where both the maximum eigenvalue and the trace-test value examine the null hypothesis of no cointegration against the alternative of cointegration. Starting with the null hypothesis of no cointegration ($R = 0$) among the variables, the trace-test statistic is 128.29, which is above the 1 percent (96.58) and 5 percent critical values (87.31). Hence it rejects the null hypothesis $R \leq 0$ in favor of the general alternative $R = 1$. As shown in Table 4, the null hypothesis of $R \leq 1$ can be rejected at 5 percent significance level, hence its alternative of $R = 2$ is accepted. Consequently, we concluded that there are two cointegrating vectors among liquid liability, net capital inflows as proxy for capital account openness, institutions, globalization of trade openness, and GDP per capita. This confirms the existence of a long-run relationship between concerned variables in the basic specified model. Turning to the maximum eigenvalue test, the null hypothesis of no cointegration ($R = 0$) is rejected at 1 percent significance level in favor of the general alternative, i.e., one cointegrating vector, $R = 1$. The test also rejected the null hypothesis of $R = 1$ in favor of the alternative $R = 2$. It is concluded that there are two cointegrating relationships among the five $I(0)$ or $I(1)$ or mutually cointegrated variables. Therefore, quarterly data from 1981 to 2004 appear to support the proposition that in Pakistan there exists a stable long-run relationship among liquid liability, net capital inflows as proxy for capital account openness, institutions, globalization of trade openness, and GDP per capita.

Table 4. Johansen first information maximum likelihood test for cointegration

Hypotheses	*Likelihood ratio*	*5 percent critical value*	*1 percent critical value*	*Maximum Eigen values*	*5 percent critical value*	*1 percent critical value*
$R = 0$	128.29	87.31	96.58	62.16	37.52	42.36
$R \leq 1$	66.13	62.99	70.05	32.55	31.46	36.65
$R \leq 2$	33.58	42.44	48.45	13.83	25.54	30.34
$R \leq 3$	19.75	25.32	30.45	12.82	18.96	23.65
$R \leq 4$	6.96	12.25	16.26	6.96	12.25	16.26

The main assumption of ARDL is that the included variables in the model are having cointegrated order $I(0)$ or $I(1)$, and it can be implemented if there is a structural break in the economy. This supports the implementation of bounds testing, which is a three-step procedure. In the first step we selected lag order on the basis of Akaike information criterion (AIC) because the computation of F-statistics for cointegration is very sensitive with lag length, so lag order 6 is selected on the lowest value of AIC.

The results of bound-testing approach for cointegration show that the calculated F-statistic is 4.573, which is higher than the upper-bound critical value of 4.06 at 10 percent significance level, implying that the null hypothesis of no cointegration cannot be accepted and that there is indeed a cointegration relationship among the variables in this model.[10] Having found a long-run relationship, we applied the ARDL method to investigate the long-run and short-run elasticities. Long-run coefficients of the variables under investigation are shown in Table 6 in the first column, which is the basic model under investigation.

It is important to note that the signs of all estimated coefficients of explanatory variables are consistent and highly significant according to the theory. All the variables promote the efficiency of financial development through the institutions, especially trade openness. Net capital inflows proxy for capital account openness is associated positively with financial-sector development and significant at 10 percent level, which means that capital account is also an important determinant of financial development. Real GDP per capita impacts the financial sector positively and significantly at the 5 percent level but with low coefficient because of the low rate of saving in the economy. We included the interactive term to check the combined impact of capital account openness and trade openness.

Results reveal that the coefficient of interaction term indicates the existence of Rajan and Zingales's hypothesis in Pakistan but is insignificant, which predicts the combined as well as individual impact of net capital inflows and trade openness on financial development except for the separate influence of each variable on financial-sector development. The main reason is that beneficial effects from net inflows could not be attained due to the low performance of the financial sector. Capital account openness has a positive but insignificant impact on the efficiency of financial development. Therefore, institutions, trade openness, and real GDP per capita are still important determinants of financial-sector development.

[10] As can be seen from Table 4, although the results of the F-test changes significantly at lag order 5 and 7, support for cointegration exists. F-test statistic is highly sensitive with the lag order; there is strong evidence for cointegration because our F-value is greater than its critical value even when sixth lag is imposed.

Table 5. Lag length criteria

Lag order	*Akaike information criterion*	*Schwartz criterion*	*Log likelihood*	*F-statistics for* cointegration
4	11.502	14.419	-412.631	5.387**
5	10.469	14.104	-335.909	11.989*
6	9.683	14.047	-271.094	4.573***
7	9.728	14.830	-243.187	5.107**

Note: *(**) and *** represent the significance level at 1 percent (5 percent) and 10 percent while critical values are 5.72 (4.57) and 4.06, respectively.

Table 6. Estimated long-run coefficients using the ARDL approach (Dependent variable: Liquid)

Variables	*Model 1*	*Model 2*
Constant	-7.78*	-7.48*
	(-5.17)	(-4.48)
LCIF	0.047***	0.063
	1.71	1.49
LTRADE	0.566*	0.47**
	6.98	2.01
INS	0.086*	0.08*
	6.05	4.76
RGDPC	0.017**	0.02**
	2.17	2.18
LCIF*LTRADE		0.01
		0.425
R2	0.973	0.9726
Adjusted-R2	0.971	0.9711
F-Statistics	790 (0.00)	626 (0.00)
Durbin-Watson stat	1.4323	1.416

Note: t values are in parentheses while *(**) and *** represent 1 percent (5 percent) and 10 percent levels of significance.

After the investigation of long-run relationships among the variables, to obtain short-run dynamics of these variables, we establish the error correction representation of ARDL model based on equation (1), which can be shown as follows:

$$\Delta LIQUID = \alpha_\circ + \sum_{J=0}^{n} \beta_2 \Delta LCIF + \sum_{j=0}^{n} \beta_2 \Delta INS + \sum_{j=0}^{n} \beta_3 \Delta GDPPC + \sum_{j=0}^{n} \beta_4 \Delta LTRADE + \eta CE_{t-1} + \varepsilon_t \ldots \tag{11}$$

After establishing the long-run relationship between financial development and net capital inflows, financial institutions, GDP per capita, and trade openness in the case of Pakistan as discussed above, Table 7 reports the short-run coefficient estimates obtained from the ECM version of ARDL model. The ECM coefficient shows the rate of adjustment toward the equilibrium, and it should have a statistically significant coefficient with negative sign.

$$\mu_t = -0.016 + 0.032\Delta LCIF - 0.065\Delta INS - 0.000018\Delta GDPPC - 0.189\Delta LTRADE + 0.697CR(-1) \tag{12}$$

Table 7. Error correction representation for the selected ARDL model
(Dependent variable: Liquid)

Variables	*Model 1*	*Model 2*
Constant	0.016	1.418
DLCIF	-0.032	-0.375
DINS	0.065	5.187
DGDPPC	1.83E-05	4.056
DLTRADE	0.189	2.474
CR (-1)	-0.679	-7.384
R-squared = 0.536132	Akaike info criterion = -1.736	
Adjusted R2 =0.509473	Schwartz criterion = -1.572	
Durbin-Watson stat =2.032	F-statistic = 20.1(0.00)	
Short-run diagnostic tests		
Serial Correlation LM Test = 2.283 (0.114)		
W-Heteroskedasticity Test = 1.969(0.174)		
Ramsey RESET Test = 1.008(0.320)		
Jarque-Bera Test = 0.412(0.814)		

Note: ARDL (1, 1, 1, 1 and 1) selected on the basis of AIC.

The error correction term CEt-1, which measures the speed of adjustment to restore equilibrium in the dynamics model, appears with negative sign and is significant at 1 percent level, ensuring that long-run equilibrium can be attained. Banerjee, Dolado, and Bestre [1998] argue that a highly significant error correction term is further proof of the existence of a stable long-run relationship. Indeed, he argued that testing the significance of CEt-1, which is supposed to carry a negative coefficient, is a relatively more efficient way of establishing cointegration. The lag length of the short-run model is selected on the basis of the Akaike information criterion.

The results of short-run dynamic coefficients associated with the long-run relationships obtained from the ECM version of equation (12) are given in Table 7. The signs of the short-run dynamic impacts are maintained in the long run. The equilibrium correction coefficient (CRt-1), with an estimated value of -0.679, which is highly significant at the 1 percent level, has the correct sign and implies a fairly high speed of adjustment to equilibrium level after a shock. Approximately 67.9 percent of disequilibrium from the previous quarter's shock converges back to the long-run equilibrium in the current quarter. Therefore, financial institutions, GDP per capita, and trade are found significant at the 1 percent level and positively associated with financial-sector development, which means that these determinants improve the efficiency of the financial sector. Finally, capital inflows impact the financial sector negatively but insignificantly in the short run.

The model has passed all short-run diagnostic tests for nonserial correlation: no heteroskedasticity, no specification problem in functional form, and normality of error term. The regression for the underlying ARDL equation fits very well at R2 = 97.2 percent and also has passed the diagnostic tests against serial correlation, functional form misspecification, heteroskedasticity, and non-normality of error term as shown in Table 6. The cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) plot from recursive estimation of the model also indicates stability of long-run coefficient over the sample period because graphs have not exceeded the critical boundaries of both figures at 5 percent significance level.

4. Conclusions and policy implications

There is a paucity of empirical literature on the specific relationship between financial institutions, capital account liberalization, and trade openness, but none exists in the case of Pakistan. This study investigates the importance of financial institutions, net capital inflows, and trade openness for financial-sector development in a small developing economy like Pakistan. Two approaches

(Johansen test and ARDL approach) were employed for the robustness of long-run relationships among the variables under consideration and found that both techniques provide robust results for long-run relationships in Pakistan's case. Net capital inflows have positive impact on financial-sector development in the long run. Trade openness is the main promoter of financial development in both periods. Finally, financial institutions and economic growth also help to improve the development of the financial sector. Further, it examined Rajan and Zingales's hypothesis [2003a] that predicts the combined influence of capital account liberalization and trade openness on financial-sector development. However, such relationship does not exist in the case of Pakistan.

In terms of policy implications, our findings suggest that macroeconomic management, which could simultaneously stimulate foreign capital inflows and trade openness, improves the quality of financial institutions, and high economic growth in the country would enhance the performance of both capital and financial intermediaries. Therefore, there is a need to design appropriate macroeconomic policies for the financial market to obtain greater benefits from globalization.

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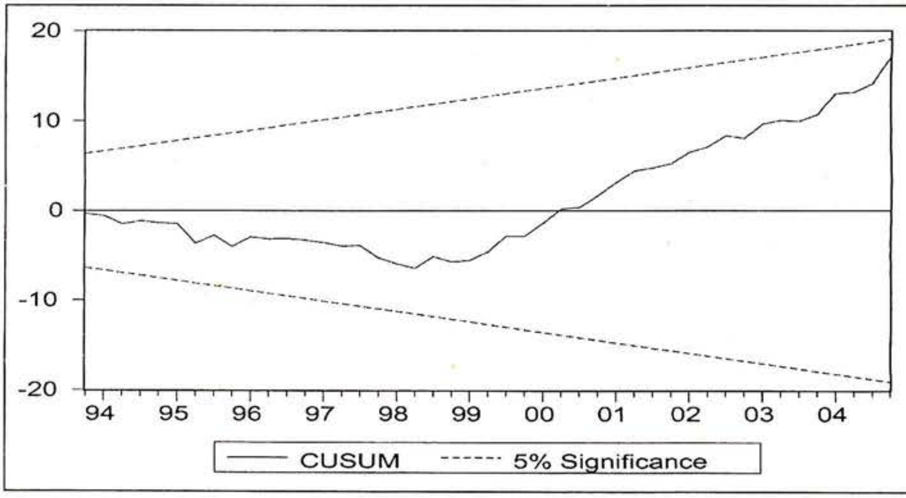
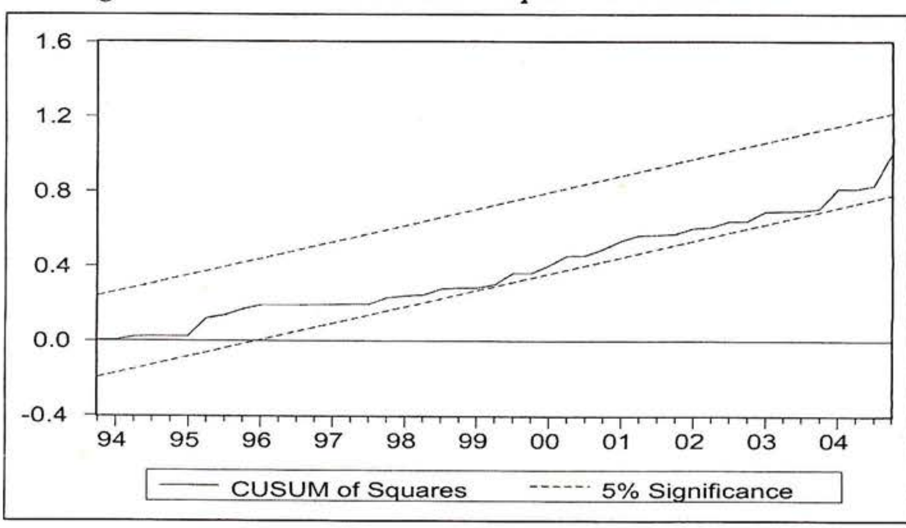
Appendix

Figure 1. Plot of cumulative sum of recursive residuals



The straight lines represent critical bounds at 5 percent significance level.

Figure 2. Plot of cumulative sum of squares of recursive residuals



The straight lines represent critical bounds at 5 percent significance level.