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2009

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Abstract

While many developing countries have reformed their financial systems over the last few decades, how an increased level of financial liberalization affects the saving-investment relationship remains unclear. This paper examines the dynamic relationship between the domestic saving and investment rates in India by controlling for the level of financial liberalization. Using data over the period 1950-2005, the results indicate that greater financial liberalization enables more domestic resources to be channeled to investment activities.

Keywords: Saving; Investment; Financial Liberalization; India.

JEL classification: F21; O16; O53

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1. Introduction

The relationship between domestic saving and investment provides some indication about the amount of domestic resources being translated into capital accumulation to fuel long-term growth. If there is no relationship between domestic saving and investment, higher domestic saving does not necessarily lead to higher investment and growth. Changes in domestic saving and investment may be completely independent, however, if capital is internationally mobile. This is because saving accumulated in one country can be easily transferred to and invested in another country. Hence, the study of saving-investment relationship is closely related to the degree of capital mobility. An understanding of capital mobility is important since higher capital mobility may smooth out external shocks to an economy. Conversely, increased capital mobility may also make an economy more vulnerable to financial turbulence.

In a seminal study, Feldstein and Horioka (1980) examine the degree of the association between saving and investment rates across 16 OECD countries. They argue that there should be no relationship between a country's domestic saving and domestic investment in the presence of perfect capital mobility. Extra saving in any country will be channeled to the world capital market to fund other countries with favorable investment climates. Using cross-sectional analysis, they show that on average 85 to 95 percent of the domestic saving are transformed into investment in the domestic economies. Furthermore, the regression coefficient on saving is statistically different from zero but not different from one, suggesting that international capital mobility is rather low, which is counter-intuitive. This observed phenomenon is widely known as the "Feldstein-Horioka puzzle".

Following this controversial finding, the relationship between saving and investment has been the subject of intense research over the past two decades. In brief, there are two main strands of literature that attempt to shed light on the "Feldstein-Horioka puzzle". The first strand of literature, which is in line with the Feldstein-Horioka interpretation, argues that a closer relationship between saving and investment implies greater international capital immobility. Using a cross-sectional framework, the findings of Penati and Dooley (1984) and Dooley *et al.* (1987), among others, suggest a significant relationship between domestic saving and investment rates. In line with these studies, time series analyses in this strand of literature which examine the dynamic saving-investment nexus over time and across different exchange rate and capital control regimes report similar findings (see Jansen and Schulze, 1996).

While it appears that many studies have confirmed the Feldstein-Horioka results of a robust saving-investment relationship, whether the results provide an indication of the degree of capital immobility is still subject to debate. In fact, many economists disagree with the interpretation of the Feldstein-Horioka results for an obvious reason: the increased integration of the global financial markets observed today is inconsistent with the argument of declining capital mobility. Using alternative measures for capital mobility such as purchasing power and interest rate differentials, the studies of Obstfeld (1986) and Baxter and Crucini (1993) have shown that capital mobility is in fact increasing over time.

Based on these findings, the second strand of literature focuses on offering alternative explanations for the saving-investment correlations. For instance, Bayoumi (1990) argues that the saving-investment relationship may be due to the implementation of government policies with the objective of achieving a balanced current account. On the other hand, Wong (1990) argues that the saving-investment relationship is dominated by the existence of the non-traded goods sector, and therefore any interpretation about the degree of capital openness cannot be readily made based on the saving-investment correlation. In an attempt to resolve the 'Feldstein-Horioka' puzzle, Coakley *et al.* (1996) demonstrate that in the presence of current account solvency, saving and investment will be cointegrated irrespective of the degree of capital mobility.

Another line of interpretation is related to the country size. Using a simple neoclassical model, Baxter and Crucini (1993) show that a high correlation can be interpreted as evidence of high capital mobility if country size is considered, given that larger countries tend to have larger effects on the world interest rates. This hypothesis is supported by the empirical findings of Ho (2003) for a panel of 23 OECD countries. More recently, Kasuga (2004) tests the hypothesis that the correlation reflects the change in net worth on investment. The results suggest that the correlations vary significantly according the level of financial development in each country.

While the above studies have made significant contributions to our understanding of the saving-investment relationship, none of them has considered the role of financial liberalization in the saving-investment relationship. Financial liberalization may drive and interact with economic growth to affect saving patterns and investment behavior (McKinnon, 1973; Shaw, 1973). Nevertheless, how financial liberalization impacts on the saving-investment relationship is unclear. On the one hand, greater financial liberalization may encourage domestic savers to save and invest more in domestic markets, and therefore strengthen the saving-investment relationship. On the other, increased financial liberalization may also encourage outflows of funds, resulting in fewer resources available to fund domestic investment projects, and thereby curtail the correlation between saving and investment. Moreover, the effect of financial liberalization on the relationship is further confounded by the theoretically ambiguous effect of financial liberalization on saving, although its effect on investment has generally been found to be positive (Chinn and Ito, 2007).

The present analysis attempts to examine the following issues: 1) has better domestic saving performance led to higher investment in India? 2) how financial liberalization affects the saving-investment relationship? India appears to be an interesting case study for this subject due to its high saving and investment rates over the last few decades. Hence, one may wonder to what extent domestic saving has been used to facilitate the undertaking of investment projects. Another interesting aspect is the recent financial sector reforms undertaken in India, which provides an ideal testing ground for further analysis on the relationship between saving and investment. Finally, the database for India is considered relatively good by developing country standards. The use of annual data covering the period 1950-2005 is sufficiently long to allow for a meaningful time series investigation.

Using the principal component method, we present a composite index of financial liberalization for India. This involves the consideration of 14 financial sector policies, which account for various dimensions of financial sector reforms in the Indian financial system. Section 2 provides an overview of the financial repression and liberalization experience of India. Data and construction of the financial liberalization index are described in Section 3. The econometric methodology used in this study is set out in Section 4. Section 5 presents and analyzes the findings. Finally, the last section concludes the paper.

2. Financial Repression and Liberalization: The Indian Experience

There was little financial repression in the financial system of India during the 1950s and 1960s. However, the government gradually imposed more controls on the financial system by raising statutory liquidity and cash reserve requirements over the 1970s and 1980s. Revenue from financial repression was estimated to be 22.4 percent of total central government revenue during the period 1980-85. Furthermore, several interest rate controls were implemented in the late 1980s (see Ang, 2009a, b).

A series of comprehensive financial sector reform policies were undertaken in 1991 as part of the broader economic reform. It was aimed at changing the entire orientation of India’s financial development strategy from its position of a financially repressed system to that of a more liberal, market-type system. It was also hoped that greater benefits of international risk sharing can be reaped through increased financial liberalization. This could help minimize the fluctuations in macroeconomic aggregates.

Since then, interest rates were gradually liberalized and statutory liquidity requirements significantly reduced so that markets could play a greater role in price determination and resource allocation. The equity market was formally liberalized in 1992, although the first country fund was set up earlier in 1986, which allowed foreign investors to access the domestic equity market directly. There has also been a change in the capital account regime from a restricted one to a more open one. The regulatory framework was strengthened significantly in 1992. In addition, entry restrictions were deregulated in 1993, resulting in the establishment of more private and foreign banks. Regulations on portfolio and direct investment were eased since then. The exchange rate was unified in 1993-94 and most restrictions on current account transactions were eliminated in 1994.

However, despite the liberalization programs launched in the early 1990s, the Indian financial system has continued to operate within the context of repressionist policies. For example, significant directed credit programs in favor of certain priority sectors still prevail in the banking system. The bank nationalization program in 1969 has enabled the Reserve Bank of India to effectively implement its credit allocation policy. Although the government divested part of its equity position in some public banks in the 1990s, the banking sector has remained predominantly state-owned. With regard to capital controls, transactions related to capital outflows have remained heavily regulated in India. As such, it appears that repressionist measures coexist with a set of liberalization policies aimed at promoting free allocation of resources.

3. Data and Construction of Variables

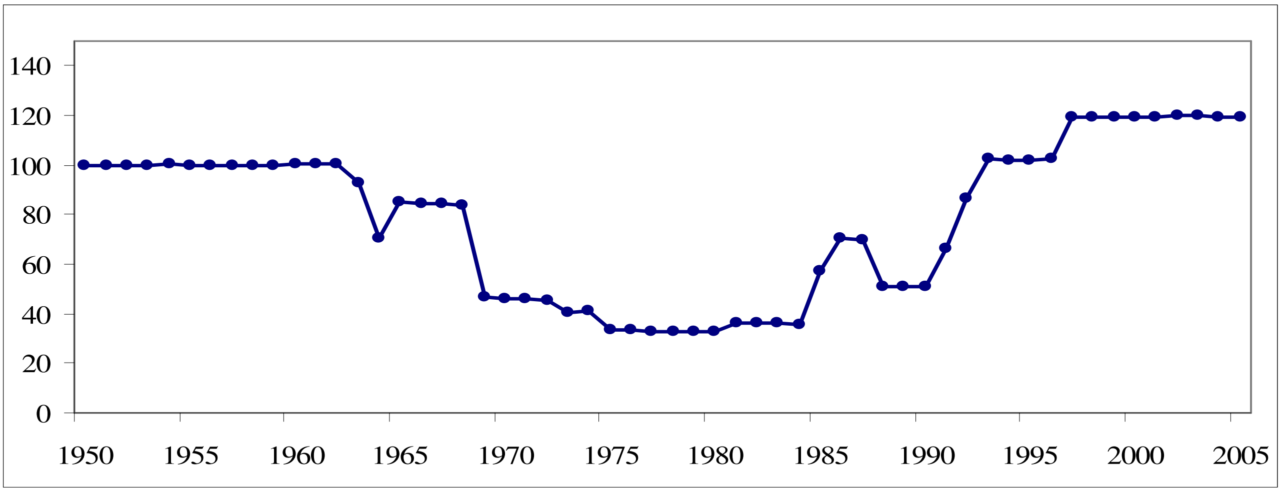
To measure the extent of financial liberalization, it is necessary to take a number of financial sector policies into consideration. However, using all these inputs in the estimation may pose some econometric problems since the underlying policy variables may be highly correlated. To overcome this problem, this study proposes to construct an index to summarize all information contained in each type of financial sector policies.

According to McKinnon (1973) and Shaw (1973), distortions in the financial systems, such as loans issued at an artificially low interest rate, directed credit programs and high reserve requirements are both unwise and unnecessary. These can reduce saving, retard capital accumulation and prevent efficient resource allocation. Therefore, they call for financial liberalization which can be achieved through eliminating or significantly alleviating the financial system distortions. Following this line of argument, we first collect nine series of financial sector policies. Six of them are interest rate controls, including a fixed lending dummy, a minimum lending rate, a maximum lending rate, a fixed deposit dummy, a minimum deposit rate and a maximum deposit rate. These policy controls are translated into dummy variables which take the value of 0 if a control is present and 1 otherwise. The remaining three policies are directed credit programs, the cash reserve ratio and the statutory liquidity ratio. The extent of directed credit programs is measured by the share of directed credit lending in total lending. The other two variables are direct measures expressed in percentages. While constructing the index, we consider, for instance, an increase in the ratio of *non*-directed credit lending in total lending to reflect greater liberalization in the financial system.

The consideration of only these policy dimensions, however, is rather restricted since they focus exclusively on interest and credit controls in the banking system. Given that the financial system in India has been liberalized beyond relaxing interest and credit controls, as discussed earlier, this necessitates the consideration of more dimensions of financial sector policies in the analysis. We therefore follow the approach of Ang (2009a) by also take into consideration privatization in the financial sector, entry barriers in the banking sector, government regulations on banking operations, equity market liberalization and restrictions on international capital flows. The inclusion of these five additional elements in the construction of the financial liberalization index is particularly relevant to the present study. For instance, policy changes on capital controls and prudential regulations may have direct impact on the saving and investment behavior, and therefore have implications on how closely saving and investment are correlated. We use dummy variables to represent policy changes in these dimensions. The data are directly obtained or compiled from the Annual Report and Report on Currency and Finance of the Reserve Bank of India.

Finally, using the above-mentioned 14 policy variables, a summary measure of financial liberalization, which represents the joint impact of various financial sector policies, is developed using the method of principal component analysis (see, e.g., Ang and McKibbin, 2007; Ang, 2008). The results of the principal component analysis are presented in the appendix.

Figure 1*:* Financial Liberalization Index (1950-2005).

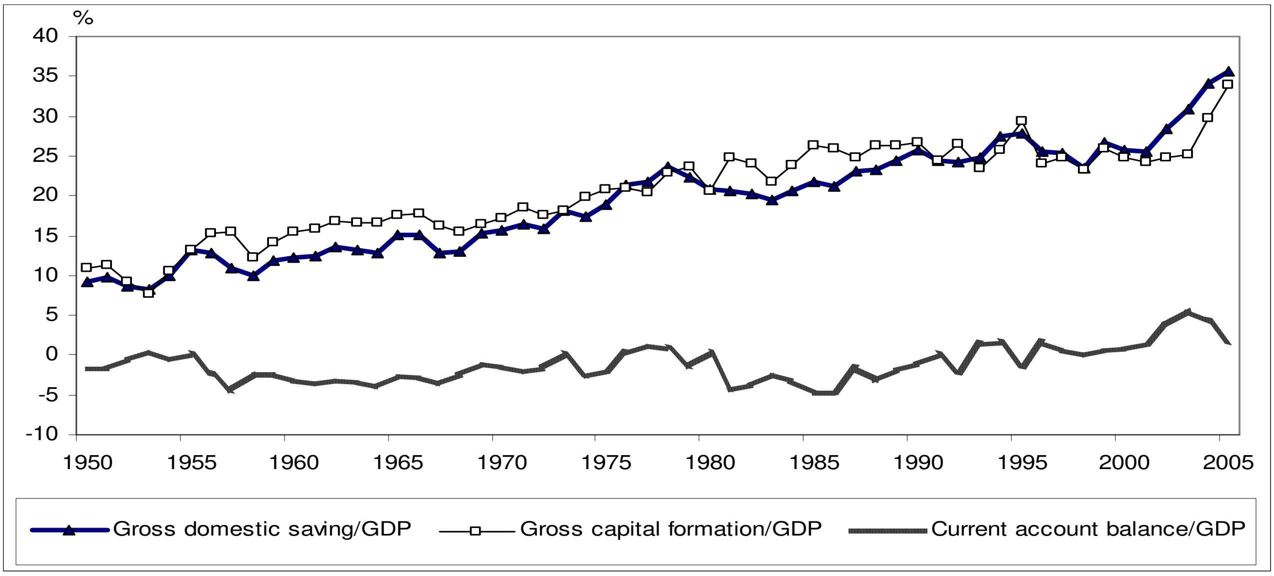


Notes: Index = 100 in 1950.

The resulting composite financial liberalization index displayed in Figure 1 coincides rather well with the actual policy changes that took place in India during the sample period, as discussed earlier in Section 2. The first observation of the index is normalized to 100. However, an index number of 100 does not necessarily refer to a fully liberalized financial system. Relative to other periods, there was very little repression in the financial system during the 1950s. From the beginning of the 1960s the government gradually tightened their controls over the financial system by raising statutory liquidity and cash reserve requirements. The tightening continued throughout the 1970s and 1980s. A major reform in the Indian financial system occurred in 1991 when the central bank launched a series of liberalization programs. The implementation of these liberalization measures is reflected by an upward swing in the series since the early 1990s.

Turing to the saving and investment dynamics, Figure 2 presents the trends of saving and investment rates over the period 1950-2005. It is evident that India has been quite successful in mobilizing saving. The average share of gross domestic saving to GDP rose from an average of just 10 percent in 1950-59 to 30 percent in 2000-05. Gross domestic investment was primarily funded by domestic saving and supplemented by foreign saving. It increased from 12 percent of GDP to 27 percent of GDP during the same period. Before 1990, domestic saving was insufficient to fund expansion in investment activities, during which some of the investment activities were financed by foreign capital. Thus, the current account was in deficits most of the times. The current account surplus recorded since then was mainly due to the rapid increase in saving. Coincidently, a series of financial liberalization policies have been undertaken since 1991. On the whole, both saving and investment rates tend to move quite closely together over time. This suggests that the presence of a cointegrated relationship between them is quite likely. We confirm this by performing two cointegration tests in Section 5.

Figure 2: Trends of saving, investment and current account balance (% of GDP)



Source: National Accounts Statistics, Government of India.

4. Econometric Methodology

The objective of our empirical estimation is to provide estimates of the long-run relationship and the short-run dynamics for the saving and investment relationship in India. We first employ the ARDL bounds procedure of Pesaran *et al.* (2001) and the ECM test of Banerjee *et al.* (1998) to test for the presence of a cointegrated relationship. The former involves a standard *F*-test whereas the latter is a simple *t*-test. Accordingly, the underlying error-correction model can be formulated as:

$$\Delta \ln(I/Y)_t = \alpha_0 + \beta_1 \ln(I/Y)_{t-1} + \beta_2 \ln(S/Y)_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta \ln(I/Y)_{t-i} + \sum_{i=0}^{p} \delta_i \Delta \ln(S/Y)_{t-i} + \varepsilon_t \tag{1}$$

where $(I/Y)_t$ refers to the ratio of gross capital formation to GDP at time *t*, $(S/Y)_t$ is the ratio of gross domestic saving to GDP at time *t*, and *p* is the lag length.

Two separate statistics are employed to test for the existence of a long-run relationship in Eq. (1): 1) an *F*-test for the joint significance of coefficients on lagged levels terms ($H_0 : \beta_1 = \beta_2 = 0$), and 2) a *t*-test for the significance of the coefficient associated with $\ln(I/Y)_{t-1}$ ($H_0 : \beta_1 = 0$). The test for cointegration is provided by two asymptotic critical value bounds when the independent variables are either *I*(0) or *I*(1). The lower bound assumes all the independent variables are *I*(0), and the upper bound assumes they are *I*(1). If the test statistics exceed their respective upper critical values, the null is rejected and we can conclude that a long-run relationship exists.

The long-run estimates are derived using two estimators: the fully-modified unrestricted error-correction model (FM-UECM) of Inder (1993) and the dynamic ordinary least squares (DOLS) estimator of Stock and Watson (1993). The FM-UECM approach involves estimating the long-run parameters by incorporating adequate dynamics into the specification to avoid omitted lagged variable bias, as shown in Eq. (2).

$$\ln(I/Y)_t = a_0 + b_1 \ln(S/Y)_t + \sum_{i=0}^{p} c_i \Delta \ln(I/Y)_{t-i} + \sum_{i=0}^{p} d_i \Delta \ln(S/Y)_{t-i} + u_t \tag{2}$$

However, this approach may not be asymptotically optimal given that it takes no account of the possible endogeneity of the underlying variables. In view of this, we follow Bewley (1979) by using the instrumental variable technique to correct the standard errors so that valid inference can be drawn. Specifically, lagged level variables are used as the instruments for the first-different current terms to correct for endogeneity bias. Next, the short-run effects are removed by defining $\ln(I/Y)_t^* = \ln(I/Y)_t - \hat{a}_0 - \hat{b}_1 \ln(S/Y)_t - \sum_{i=0}^{p} \hat{c}_i \Delta \ln(I/Y)_{t-i} - \sum_{i=0}^{p} \hat{d}_i \Delta \ln(S/Y)_{t-i}$. The fully modified estimator is then obtained by employing the Phillips-Hansen non-parametric corrections to the regression of $\ln(I/Y)_t^*$ on a constant and $\ln(S/Y)_t$. The resulting estimator thus adequately deals with omitted lag variables bias, and Inder (1993) has shown that it is asymptotically optimal, even in the presence of endogenous explanatory variables. Furthermore, using Monte Carlo experiments, Caporale and Pittis (2004) show that this estimator possesses the most desirable small sample properties in a class of 28 estimators.

The key advantage of the DOLS procedure of Stock and Watson (1993) is that it allows for the presence of a mix of $I(0)$ and $I(1)$ variables in the cointegrated system. This estimator is asymptotically equivalent to the maximum likelihood estimator of Johansen (1988). Based on Monte Carlo evidence, Stock and Watson (1993) show that DOLS outperforms a number of alternative estimators of long-run parameters. It has also been shown to perform well in finite samples. This feature is particularly appealing given the small sample size used in the present study. The estimation involves regressing one of the $I(1)$ variables on the remaining $I(1)$ variables, the $I(0)$ variables, leads (p) and lags ($-p$) of the first difference of the $I(1)$ variables, and a constant, as shown in Eq. (3). By doing so, it corrects for potential endogeneity problems and small sample bias, and provides estimates of the cointegrating vector which are asymptotically efficient. The long-run model for $\ln(I/Y)_t$ can be obtained from the reduced form solution of Eq (3) by setting all differenced terms of the regressors to be zero, i.e., $c_i = d_i = 0$.

$$\ln(I/Y)_t = a_0 + b_1 \ln(S/Y)_t + \sum_{i=-p}^{p} c_i \Delta \ln(I/Y)_{t-i} + \sum_{i=-p}^{p} d_i \Delta \ln(S/Y)_{t-i} + u_t \qquad (3)$$

Finally, the error-correction term (ECT) can be obtained by taking $\ln(I/Y)_t - a_0 - b_1 \ln(S/Y)_t$ to formulate an error-correction model. The ECT captures the evolution process on the variable of concern by which agents adjust for prediction errors made in the last period. Hendry's (1995) general-to-specific modeling approach is adopted to derive a satisfactory short-run dynamic model. This involves testing down the general model by successively eliminating statistically insignificant regressors and imposing data acceptable restrictions on the parameters to obtain the final parsimonious dynamic equation. In order to test the robustness of the results, all estimations are subject to various diagnostic tests.

5. Empirical Results

5.1 Unit root tests

We begin the analysis by examining the unit root properties of the underlying variables. Following the standard practice, all variables are measured in natural logarithms. The integration properties of the underlying variables are examined using two standard unit root tests - the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. However, the presence of a structural break in a series may bias the results toward non-rejection of the null hypothesis of a unit root when there is none. We therefore also implement unit root tests with an endogenous break to examine whether the series appear to be stationary. For this purpose, we perform the unit root procedure of Zivot and Andrews (1992) (ZA), which tests the null of a unit root against the alternative of trend stationarity with an unknown break in the series.

Table 1: Test results for unit roots

	ADF	*PP*	*ZA*
$\ln(I/Y)_t$	-3.160	-2.973	-4.097
$\Delta\ln(I/Y)_t$	-7.047***	-10.649***	-7.173***
$\ln(S/Y)_t$	-3.093	-3.144	-7.031***
$\Delta\ln(S/Y)_t$	-4.762***	-12.354***	-8.260***
$\ln FL_t \text{x} \ln(S/Y)_t$	-1.765	-1.718	-3.751
$\Delta[\ln FL_t \text{x} \ln(S/Y)_t]$	-6.279***	-7.707***	-7.885***

Notes: For ADF, AIC is used to select the lag length and the maximum number of lags is set to be five. For PP, Barlett-Kernel is used as the spectral estimation method. The bandwidth is selected using the Newey-West method. Only results for the "crash" model, which allows for an exogenous shift in the mean of the series, are reported for the Zivot-Andrews (ZA) tests. We have also considered the "changing growth" model that allows for a shift in the trend and the "change in level and slope" model that admits both changes. These models yield very similar results that do not alter the conclusions, and therefore they are not reported for brevity. *** indicates 1% level of significance.

The results reported in Table 1 unanimously suggest that both $\ln(I/Y)_t$ and $\ln FO_t \text{x} \ln(S/Y)_t$ are $I(1)$ variables. The results are significant at the one percent level. Although both ADF and PP suggest that $\ln(S/Y)_t$ contains a unit root, the ZA procedure indicates that the variable is stationary once the presence of a structural break in the series is considered. This points to the importance of considering a cointegration approach that is appropriate for the presence of a mix of $I(0)$ and $I(1)$ variables in the model.

5.2 The base line model

Next, to perform cointegration tests on the basic saving-investment equation without the implications of financial liberalization, we regress the conditional error-correction model in Eq. (1) by allowing up to five lags. To ascertain the existence of a level relationship between the variables, this requires satisfying both the *F*- and the *t*-tests. Table 2 gives the *F*-statistics for the ARDL bounds tests, *t*-statistics for the ECM test, the Akaike's and Schwarz's Bayesian Information Criteria, denoted by AIC and SBC, respectively, and several diagnostic test statistics.

The results reported in Table 2 indicate the null hypothesis that there exists no level investment rate equation is rejected at the five percent significance level with three and four lags, providing strong support for the existence of a long-run relationship between the rates of saving and investment. No evidence of cointegrating is found when saving rate is used as the dependent variables. The evidence of cointegration has also been confirmed using the Johansen trace and eigenvalue tests. The results, which are not reported here to conserve space but available upon request, also show that the null of no cointegration is rejected at the five percent level when three or four lags are assumed. The finding of a cointegrated relationship between saving and investment rates is consistent with the time series findings of Jansen and Schulze (1996) for Norway, Moreno (1997) for the US and Japan, De Vita and Abbott (2002) for the US, Pelagidis and Mastroyiannis (2003) for Greece and Payne (2005) for Mexico.

Table 2: Cointegration tests for the saving-investment relationship

	$p=1$	$p=2$	$p=3$	$p=4$	$p=5$
			A. Bounds test statistics		
F-statistic	3.493	2.676	7.706**	7.106**	4.366
t-statistic	-2.642	-2.211	-3.486**	-3.662**	-2.938*
			B. Model selection criteria		
AIC	-2.249	-2.296	-2.529	-2.565	-2.471
SBC	-2.029	-1.998	-2.154	-2.111	-1.935
			C. Diagnostic checks		
$\chi^2_{NORMAL}(2)$	0.248 (0.883)	4.749* (0.093)	0.761 (0.683)	0.802 (0.669)	0.916 (0.632)
$\chi^2_{SERIAL}(1)$	2.192 (0.138)	9.344*** (0.002)	0.013 (0.911)	0.272 (0.601)	0.023 (0.879)
$\chi^2_{SERIAL}(2)$	3.168 (0.205)	11.608*** (0.003)	0.259 (0.878)	0.546 (0.761)	0.953 (0.621)
$\chi^2_{ARCH}(1)$	2.842* (0.091)	0.053 (0.817)	0.310 (0.577)	0.188 (0.664)	0.174 (0.676)
χ^2_{WHITE}	19.703** (0.049)	22.454** (0.025)	13.391 (0.530)	22.141 (0.346)	22.059 (0.688)

Notes: p is the lag length. The test statistics of the bounds tests are compared against the critical values reported in Pesaran *et al.* (2001). The 10%, 5% and 1% critical value bounds for the F-test are (4.04, 4.78), (4.94, 5.73) and (6.84, 7.84), respectively. The 10%, 5% and 1% critical value bounds for the t-test are (-2.57, -2.91), (-2.86, -3.22) and (-3.43, -3.82), respectively. $\chi^2_{NORMAL}(2)$ refers to the Jarque-Bera statistic of the test for normal residuals, $\chi^2_{SERIAL}(1)$ and $\chi^2_{SERIAL}(2)$ are the Breusch-Godfrey LM test statistics for no first and second order serial relationship, respectively, $\chi^2_{ARCH}(1)$ is the Engle's test statistic for no autoregressive conditional heteroskedasticity and χ^2_{WHITE} denotes the White's test statistic to test for homoskedastic errors, with degrees of freedom equal to the number of slope coefficients. Numbers in parentheses indicate p-values. *, ** and *** indicate 10%, 5% and 1% level of significance, respectively.

In line with the results of cointegration tests, the AIC favors modeling with four lag whereas the SBC points to specifying the model with three lags. The results are not surprising given that the SBC always tends to select a model with less dynamics. Both regressions fit rather well and pass the diagnostic tests against non-normal residuals, serial correlation, autoregressive conditional heteroskedasticity and heteroskedasticity. However, the model with four lags fails functional misspecification at the five percent level. This is probably due to over-parameterization of the ARDL model. Nevertheless, to provide some robustness checks, it seems prudent to choose the lag length to be three and four.

To obtain the long-run model, we first estimate Eq. (2) by OLS. The long-run model for $\ln(I/Y)_t$ is obtained from the reduced form solution of Eq. (2), when all differenced terms of the regressors are set to zero. Based on the FM-UECM procedure, the results reported in panel A of Table 3 show that the elasticity of investment rate with respect to saving rate is found to be about on average 0.71 in the long run, suggesting a positive long-run relationship between saving and investment rates. The estimates obtained based on the DOLS procedure produce very similar results.

The regression results for the equation of $\Delta\ln(I/Y)_t$ reported in panel B of Table 3 provide the short-run dynamics of the investment rate function. All coefficients are statistically significant at the conventional levels. In first-differenced contemporaneous form, the coefficients on $\Delta\ln(S/Y)_t$ are consistent with the long-run results. The magnitudes of the short-run coefficients, however, are found to be slightly smaller than their long-run counterparts. Both estimators yield very similar results.

The coefficients on ECT_{t-1}, which measure the speed of adjustment back to the long-run equilibrium value, are statistically significant at the one percent level and correctly signed, i.e., negative, providing further evidence against no cointegration. This implies that an error-correction mechanism exists so that the deviation from long-run equilibrium has a significant impact on the growth rate of investment rate. Based on the estimates obtained from using the FM-UECM procedure, the magnitudes of the coefficients suggest that investment rate adjusts at the speed of about 36.8 percent and 58.2 percent every year for Model A and Model B, respectively (or it takes about 2.7 years and 1.7 years, respectively) to restore equilibrium when there is a shock on the steady-state relationship. The model with three lags estimated using the DOLS approach appears to be subject to some econometric problems. However, for the model with four lags, the regression specifications fit remarkably well and pass the diagnostic tests against non-normality, serial correlation, autoregressive conditional heteroskedasticity, heteroskedasticity and functional misspecification.

The findings of a robust long-run cointegrated relationship between domestic saving and investment rates suggest that any change in domestic saving will be closely associated with a change in investment. Hence, financial sector policies targeting at mobilizing domestic saving are critical for capital accumulation. This is further examined in the next section. However, it should also be highlighted that over-reliance on domestic saving may limit the growth opportunity of an economy. As such, policy makers should also focus on attracting foreign capital as part of the development policy while mobilizing resources in the domestic economy. Policy makers should ensure these additional foreign resources are channeled to the productive sectors and utilized efficiently for sustained development.

Table 3: Results for the saving-investment relationship

	Model A (p = 3)		Model B (p = 4)	
	FM-UECM	*DOLS*	*FM-UECM*	*DOLS*
A. The long-run relationship (Dep. Var. = $\ln(I/Y)_t$*)*				
Intercept	-0.359***	-0.357***	-0.422***	-0.332***
	(0.000)	(0.000)	(0.000)	(0.000)
$\ln(S/Y)_t$	0.728***	0.721***	0.693***	0.734***
	(0.000)	(0.000)	(0.000)	(0.000)
B. The short-run dynamics (Dep. Var. = $\Delta\ln(I/Y)_t$*)*				
Intercept	-0.006	-0.015	0.012	-0.001
	(0.559)	(0.195)	(0.198)	(0.884)
ECT_{t-1}	-0.368***	-0.323***	-0.582***	-0.579***
	(0.001)	(0.005)	(0.000)	(0.000)
$\Delta\ln(S/Y)_t$	0.671***	0.673***	0.517***	0.532***
	(0.000)	(0.000)	(0.000)	(0.000)
$\Delta\ln(S/Y)_{t-1}$		0.232*		
		(0.053)		
$\Delta\ln(I/Y)_{t-3}$			0.224**	0.234**
			(0.041)	(0.034)
$\Delta\ln(S/Y)_{t-3}$			-0.228*	-0.232*
			(0.067)	(0.063)
$\Delta\ln(S/Y)_{t-4}$			-0.280***	-0.274***
			(0.006)	(0.008)
C. Diagnostic checks				
$\chi^2_{NORMAL}(2)$	3.099	0.723	1.342	0.699
	(0.212)	(0.697)	(0.511)	(0.705)
$\chi^2_{SERIAL}(1)$	0.028	1.546	0.237	0.088
	(0.865)	(0.214)	(0.626)	(0.766)
$\chi^2_{SERIAL}(2)$	1.695	5.889*	0.296	0.383
	(0.428)	(0.053)	(0.862)	(0.826)
$\chi^2_{ARCH}(1)$	3.790*	4.317**	0.001	0.056
	(0.051)	(0.038)	(0.966)	(0.813)
χ^2_{WHITE}	4.216	16.910***	2.861	5.649
	(0.121)	(0.009)	(0.721)	(0.844)

Notes: see notes to Table 2.

5.3 The Saving-Investment Relationship and Financial Liberalization

In order to examine the implications of financial liberalization on the saving-investment relationship, we incorporate an interaction term between saving and financial liberalization into the empirical specification. The expected sign for this interaction term is ambiguous, as discussed earlier. Specifically, the cointegration tests are performed based on the following equation:

$$\Delta \ln(I/Y)_t = \alpha_0 + \beta_1 \ln(I/Y)_{t-1} + \beta_2 \ln(S/Y)_{t-1} + \beta_3 \ln(S/Y)_{t-1} \text{ x} \ln FL_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta \ln(I/Y)_{t-i} + \sum_{i=0}^{p} \delta_i \Delta \ln(S/Y)_{t-i} + \varepsilon_t \quad (4)$$

The results reported in Table 4 show that the null of no cointegration is rejection at the conventional levels only when three or four lags are chosen. The choice of these lags is consistent with the model selection criteria given by the AIC or SBC. The diagnostic results show a pattern very similar to those reported in Table 2. We therefore conclude that there exists a long-run relationship between investment rate, saving rate and the interaction between saving and financial liberalization.

Table 4: Cointegration tests for the saving-investment relationship and financial liberalization

	$p=1$	$p=2$	$p=3$	$p=4$	$p=5$
			A. Bounds test statistics		
F-statistic	2.471	1.848	6.583***	6.335**	4.008
t-statistic	-0.365	-2.217	-3.811**	-4.189***	-3.431*
			B. Model selection criteria		
AIC	-2.164	-2.172	-2.481	-2.501	-2.390
SBC	-1.833	-1.726	-1.918	-1.819	-1.587
			C. Diagnostic checks		
$\chi^2_{NORMAL}(2)$	0.519 (0.771)	5.851* (0.054)	0.626 (0.731)	0.129 (0.941)	0.083 (0.959)
$\chi^2_{SERIAL}(1)$	1.921 (0.166)	9.589*** (0.002)	1.649 (0.199)	0.009 (0.925)	0.047 (0.828)
$\chi^2_{SERIAL}(2)$	2.967 (0.227)	10.621*** (0.005)	1.834 (0.399)	0.251 (0.882)	0.561 (0.351)
$\chi^2_{ARCH}(1)$	3.057* (0.081)	0.001 (0.982)	0.002 (0.969)	0.301 (0.583)	0.127 (0.721)
χ^2_{WHITE}	11.849 (0.158)	17.923* (0.083)	11.818 (0.621)	15.803 (0.538)	15.984 (0.718)

Notes: p is the lag length. The test statistics of the bounds tests are compared against the critical values reported in Pesaran *et al.* (2001). The 10%, 5% and 1% critical value bounds for the F-test are (3.17, 4.14), (4.79, 4.85) and (5.15, 6.36), respectively. The 10%, 5% and 1% critical value bounds for the t-test are (-2.57, -3.21), (-2.86, -3.53) and (-3.43, -4.10), respectively. *, ** and *** indicate 10%, 5% and 1% level of significance, respectively.

Table 5 reports the estimated results for this long-run relationship and its short-run dynamics using both the FM-UECM and DOLS procedures. The results are obtained by regressing Eqs. (2) and (3) with an additional interaction term. It is evident that saving enters the equation significantly, with a long-run elasticity in the range of 0.593-0.621. In terms of short-run dynamics, the coefficients on $\Delta \ln(S/Y)_t$ are also highly significant, in the range of 0.487-0.666. On the whole, there are no significant variations in the short-run results, compared to those of the base line model reported in Table 3. The diagnostic checks produce rather satisfactory results.

Table 5: The saving-investment relationship and financial liberalization

	Model C (p = 3)		Model D (p = 4)	
	FM-UECM	*DOLS*	*FM-UECM*	*DOLS*
A. The long-run relationship (Dep. Var. = $\ln(I/Y)_t$*)*				
Intercept	-0.393*** (0.000)	-0.407*** (0.000)	-0.425*** (0.000)	-0.376*** (0.000)
$\ln(S/Y)_t$	0.621*** (0.000)	0.595*** (0.000)	0.615*** (0.004)	0.593*** (0.000)
$\ln(S/Y)_t \times \ln FL_t$	0.021** (0.022)	0.022** (0.036)	0.017* (0.088)	0.026** (0.018)
B. The short-run dynamics (Dep. Var. = $\Delta \ln(I/Y)_t$*)*				
Intercept	-0.005 (0.641)	-0.005 (0.613)	0.007 (0.413)	-0.006 (0.527)
ECT_{t-1}	-0.384*** (0.001)	-0.617*** (0.000)	-0.602*** (0.000)	-0.602*** (0.000)
$\Delta \ln(S/Y)_t$	0.666*** (0.000)	0.487*** (0.000)	0.513*** (0.000)	0.523*** (0.000)
$\Delta \ln(S/Y)_{t-1}$		0.214** (0.027)		
$\Delta \ln(I/Y)_{t-3}$			0.223** (0.041)	0.226** (0.037)
$\Delta \ln(S/Y)_{t-3}$			-0.225* (0.067)	-0.227* (0.065)
$\Delta \ln(S/Y)_{t-4}$			-0.281*** (0.006)	-0.278*** (0.006)
C. Diagnostic checks				
$\chi^2_{NORMAL}(2)$	3.448 (0.178)	1.516 (0.468)	1.738 (0.419)	1.758 (0.415)
$\chi^2_{SERIAL}(1)$	0.019 (0.889)	0.132 (0.716)	0.362 (0.548)	0.302 (0.582)
$\chi^2_{SERIAL}(2)$	1.695 (0.428)	0.285 (0.867)	0.365 (0.833)	0.323 (0.851)
$\chi^2_{ARCH}(1)$	4.255** (0.039)	0.782 (0.376)	0.015 (0.902)	0.042 (0.838)
χ^2_{WHITE}	4.158 (0.125)	0.969 (0.809)	2.872 (0.717)	3.089 (0.686)

Notes: see notes to Table 3.

In the long-run equation, the interaction term is found to be statistically significant at the five percent level and has a positive sign. However, no significant effect of financial liberalization is found in the short run. Therefore, it can be inferred that the impact of saving on investment is strengthened through greater liberalization in the financial system in the long run. This is obvious when we obtain the derivative of $\ln(I/Y)_t$ with respect to $\ln(S/Y)_t$, which gives

$0.621+0.021\ln FL_t$ for the model with three lags estimated using the FM-UECM procedure. Hence, both the direct and indirect effects of saving on investment are found to be positive in the long run.

The saving glut hypothesis of Bernanke (2005) posits that the large increase in investment in the U.S. observed in recent years is mainly due to large influx of funds from the developing world, whose financial development is relatively weak. In Bernanke's view, this excess saving problem can be mitigated through better developed financial systems, which can be achieved through financial sector reforms. Our results lend some support to this view. Moreover, our results are not contradictory to the Feldstein-Horioka interpretation of the saving-investment correlation given that, in principle, a financial system can be liberalized and yet remain closed (so that capital is immobile) or vice versa. For instance, the Japanese and Korean financial systems during most of the years in 1970-1990 were repressed but open.

6. Conclusions

Many developing countries have reformed their financial systems over the last few decades. While an increased level of financial development has generally been observed across the world, the issue concerning how financial liberalization impacts on the saving-investment relationship remains unknown. An understanding of this relationship is important in order to assess the costs and benefits associated with greater financial liberalization.

In this paper, we examine the relationship between domestic saving and investment rates in a cointegration framework using the Indian data for the period 1950-2005. Employing the ARDL and ECM cointegration techniques, the empirical evidence shows a fairly robust long-run relationship between domestic saving and investment rates, consistent with the prediction of an intertemporal current account model. After documenting these basic cointegration results, we derive the long-run estimates using two different estimators. The qualitative aspects of the results are insensitive to the choice of estimators. The estimated results based on annual data for the period 1959-2005 suggest that saving and investment are strongly related, both in the short run and long run.

To provide more insight into the role of financial factors in the saving-investment dynamics, we also examine the saving-investment nexus by controlling for the level of financial liberalization. We attempt to address the difficult problem of measuring the extent of financial liberalization by using the method of principal component analysis to create an index that represents the overall level of liberalization in the financial system. Our results show that saving and investment rates are still robustly cointegrated, even after controlling for the effect of financial liberalization. The interaction between saving and financial liberalization is found to have a positive significant effect on investment in the long run. This suggests that, in addition to a direct effect, saving also has an indirect effect in stimulating investment through greater financial liberalization.

While the empirical results presented in this study are intriguing, more analysis is warranted. We hesitate to generalize the findings of this study to other developing countries since the results may be unique to the experience of India due to its own institutional and historical settings. Future studies can look at how financial liberalization affects the saving-investment relationship in other countries, which have experienced significant financial sector reforms, using the framework established in this paper.

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Appendix I: Principal component analysis for the financial liberalization index

	Principal component													
	PC 1	PC 2	PC 3	PC 4	PC 5	PC 6	PC 7	PC 8	PC 9	PC 10	PC 11	PC 12	PC 13	PC 14
Eigenvalues	6.57	3.86	1.22	0.74	0.55	0.37	0.23	0.21	0.09	0.06	0.04	0.03	0.02	0.01
% of variance	0.47	0.28	0.09	0.05	0.04	0.03	0.02	0.02	0.01	0.00	0.00	0.00	0.00	0.00
Cumulative %	0.47	0.75	0.83	0.89	0.92	0.95	0.97	0.98	0.99	0.99	1.00	1.00	1.00	1.00
	Eigenvector (loadings)													
Policy Variable	PC 1	PC 2	PC 3	PC 4	PC 5	PC 6	PC 7	PC 8	PC 9	PC 10	PC 11	PC 12	PC 13	PC 14
FDR_t	0.32	0.16	-0.33	0.26	0.25	-0.09	-0.05	-0.12	0.35	0.19	-0.49	-0.07	0.29	-0.36
DRC_t	0.16	0.42	0.19	0.18	0.06	-0.25	0.24	-0.39	-0.49	-0.33	-0.04	0.05	0.29	0.10
DRF_t	-0.03	0.43	-0.25	0.13	0.27	0.38	-0.64	-0.07	-0.12	-0.10	0.15	-0.03	-0.16	0.19
FLR_t	0.16	-0.01	0.71	0.02	0.63	0.04	-0.05	0.15	0.12	0.18	0.04	0.01	0.02	0.05
LRC_t	-0.09	0.32	-0.02	-0.85	0.11	0.18	0.12	-0.10	0.04	0.01	-0.31	0.08	0.03	0.00
LRF_t	0.32	-0.01	-0.19	0.14	0.10	0.66	0.56	0.20	-0.07	-0.05	0.11	0.03	0.03	0.11
CRR_t	0.31	0.07	0.35	0.06	-0.47	0.23	-0.11	-0.38	0.50	-0.15	-0.08	0.10	-0.07	0.22
SLR_t	0.23	0.38	0.16	-0.07	-0.24	0.02	0.04	0.04	-0.11	0.22	0.33	-0.29	-0.27	-0.62
DCP_t	0.33	0.18	-0.14	-0.22	-0.11	-0.26	-0.11	0.34	0.19	0.07	0.44	-0.02	0.51	0.28
PRI_t	0.33	0.21	-0.15	0.03	0.05	-0.40	0.14	0.24	0.03	0.02	-0.18	0.24	-0.65	0.28
ENB_t	-0.28	0.30	0.12	0.21	-0.26	0.09	0.10	0.11	-0.13	0.65	-0.27	-0.17	0.10	0.36
REG_t	-0.30	0.30	0.10	0.18	-0.12	0.07	0.01	0.30	0.12	-0.08	0.03	0.74	0.13	-0.29
EML_t	-0.33	0.17	-0.18	0.06	0.26	-0.13	0.35	-0.45	0.41	0.18	0.45	0.03	-0.11	0.09
ICF_t	-0.31	0.27	0.09	0.12	0.01	-0.05	0.15	0.36	0.32	-0.53	-0.14	-0.50	-0.03	0.03

Notes: FDR_t = fixed deposit dummy, DRC_t = deposit rate ceiling dummy, DRF_t = deposit rate floor dummy, FLR_t = fixed lending dummy, LRC_t = lending rate ceiling, LRF_t = lending rate floor, CRR_t = cash (statutory) reserve ratio on time deposit; SLR_t = statutory liquidity ratio, DCP_t = directed credit program, PRI_t = privatization in the banking sector, ENB_t = entry barriers, REG_t = regulations, EML_t = equity market liberalization, and ICF_t = international capital flows. The above table presents the results for the financial liberalization index obtained from principal component analysis. The eigenvalues indicate that the first principal component explains about 47 percent of the standardized variance, the second principal component explains another 28 percent and so on. The first principal component is computed as a linear combination of the nine policy measures with weights given by the first eigenvector. In this case, the six largest principal components are extracted, and they are able to capture 95 percent of the information from the original data set. The remaining principal components are not considered since their marginal information content is relatively small. The percentages of variance are adjusted to make sure that their absolute values sum up to one. These adjusted values are then used as the weights to compute the index. In this connection, the first principal component, which accounts for 47 percent of the total variation of the policy variables, has a weight of 47/95, and so on.