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# Does Foreign Aid Promote Growth? Exploring the Role of Financial Liberalization

*James B. Ang**

## Abstract

This paper examines the impact of foreign aid on the process of economic development in India by controlling for the degree of financial liberalization. A composite index is constructed using the method of principal component analysis to capture the joint influence of various financial sector policies. The results show that while foreign aid exerts a direct negative influence on output expansion, its indirect effect via financial liberalization is positive. Therefore, an important implication of the findings in this paper is that adequate liberalization in the financial system of the host country is a crucial requirement for effective foreign aid. Our results are robust to a number of control variables and estimation techniques.

## 1. Introduction

Although many developing countries have received a substantial amount of foreign aid over the last few decades, there is no consensus regarding its impact on growth (Radelet, 2006). Several recent studies, typified by the work of Burnside and Dollar (2000) and Collier and Dollar (2002), argue that aid assists growth but only in good policy environments. Others suggest that there is a nonlinear effect in the aid–growth relationship due to diminishing returns to aid (see Dalgaard and Hansen, 2001; Hansen and Tarp, 2001; Lensink and White, 2001; Dalgaard et al., 2004, among others). Although the current debate is focused on these two leading views (Alvi et al., 2008), there is also a growing literature showing that the aid–growth nexus can at best be characterized as fragile and ambiguous (see e.g. Clemens et al., 2004; Easterly et al., 2004).

Amidst animated debate on the effectiveness of aid, the present study seeks to examine a different but related issue: the role of financial liberalization in the aid–growth relationship. This issue is examined in the context of India, a large and rapidly growing developing economy that has undergone significant financial sector reforms. From the 1950s to the 1980s, India was the largest recipient of foreign aid. However, its economic growth rates were virtually stagnant during this period (Becker, 2007). The financial sector reforms initiated since the 1990s, along with the rapid economic growth experienced by India since then, suggest that there may be a close association between financial liberalization and the aid–growth nexus. This interesting observation forms the basis for further analysis.

The early literature initiated by McKinnon (1973) and Shaw (1973) suggests that financial liberalization encourages investment and therefore exerts a positive effect on long-term growth. Following the seminal empirical work of King and Levine (1993), the

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relationship between finance and growth has been a subject of considerable academic interest and intense policy debate. The bulk of cross-country evidence appears to suggest that financial development has a positive impact on economic growth (see Ang, 2008, for a survey of the recent literature), although case studies indicate that the direction of causality is less unambiguous (see e.g. Ang and McKibbin, 2007).

These two strands of literature (i.e. the aid–growth and finance–growth links) have recently been combined under an integrated framework by Nkusu and Sayek (2004). They argue that financial development may exert an indirect positive effect on the aid–growth relationship through the conduct of interest and exchange rate management, where the effectiveness of these policies depends on the absorptive capacity of the local financial markets. Significant inflows of foreign aid will put upward pressure on the real exchange rate that can be translated into higher prices. The presence of a broad and deep financial system provides the necessary instruments that could effectively sterilize these undesirable impacts. In other words, foreign aid functions effectively when aid flows are better managed in the context of more developed financial systems. Therefore, it appears plausible that one of the underlying reasons that aid is less effective than expected at spurring development is the failure of financial systems to ensure an efficient allocation of aid resources.

Unlike Nkusu and Sayek (2004), the focus of the present study is on financial liberalization rather than financial development. We emphasize the former since the depth of a financial system is directly shaped by financial sector policies. The presence of a more liberalized financial system also effectively reduces barriers and restrictions on interest and exchange rate controls, providing the monetary authorities with greater flexibility to conduct monetary and exchange rate management (Caporale and Williams, 2001; Kletzer and Kohli, 2001). The main contributions of this study include: (1) Empirically testing the relationship between aid and growth by providing further evidence from a large and rapidly growing developing country. Not only could this enhance our understanding of the aid–growth relationship, but also fill the gap in the extant literature, which is dominated by cross-country analysis. (2) Contributing to the debate on the effectiveness of foreign aid in the Indian economy. There are very few studies on the aid–growth relationship for India, despite its status as one of the largest recipients of foreign assistance. (3) Complementing the literature on the effectiveness of foreign aid by assessing the impact of financial liberalization on economic development. This policy factor has been neglected somewhat in the analysis of the aid–growth nexus. The following section briefly reviews the financial repression and liberalization experience of India.

## 2. Financial Sector Reforms in India

There was little repression in the Indian financial system in the 1950s. However, the Reserve Bank of India gradually imposed more restrictions by introducing interest rate controls in the 1960s. The statutory liquidity ratio was raised from 25% in 1966 to 38% in 1989. The cash reserve rate increased considerably from 3% to 15% during the same period. These high liquidity and reserve requirements enabled the Bank to purchase government securities at low cost. The extent of directed credit programs has also increased significantly since the nationalization of the 14 largest private banks in 1969. A number of priority lending rates were set at levels well below those that would prevail in the free market. This process culminated in the late 1980s when directed lending was more than 40% of the total.

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The major phase of financial liberalization was undertaken in 1991 as part of the broader economic reform in response to the balance-of-payments crisis of 1990–91. The objective was to provide a greater role for markets in price determination and resource allocation. Consequently, interest rates were gradually liberalized, and the reserve and liquidity ratios were reduced significantly. However, despite this liberalization, the Indian financial system has continued to operate within the context of repressionist policies through the provision of subsidized credit to certain priority sectors. Liberalization of the directed credit programs is only limited to deregulation of priority lending rates, whilst significant controls on the volume of directed lending remain in place. Furthermore, the Bank has tightened supervision and regulation in recent years to ensure that these priority sector requirements are met.

Measuring the extent of financial liberalization is not an easy task. To do this, we follow the approach of Ang (2010a) by constructing a composite index. In particular, nine series for these repressionist policies are collected. 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 1 if a control is present and 0 otherwise. The remaining three policies are the cash (statutory) reserve ratio, the statutory liquidity ratio, and directed credit programs.

Using these nine policy variables, a summary measure of financial repression, which represents the joint impact of the various financial policies, is developed using the method of principal component analysis. In principle, this composite measure is able to capture most of the information from the original dataset that consists of nine policy variables. Given its conciseness, this approach sufficiently deals with the problems of multicollinearity and over-parametrization. The inverse of this measure can be interpreted as the extent of financial liberalization (see e.g. Ang and McKibbin, 2007; Ang, 2010a).

Table 1 presents the results for the financial liberalization index obtained from principal component analysis. The eigenvalues indicate that the first principal component explains about 47.2% of the standardized variance, the second principal component explains another 21.6%, 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.

The six largest principal components are extracted, and they are able to capture 96.6% of the information from the original dataset. 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 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 47.2% of the total variation of the policy variables, has a weight of 47.2/96.6, and so on. The resulting index is presented in Figure 1. It is evident that the trend towards financial repression has been reversed since the early 1990s. The leveling-off observed in the series coincides with the increase in the extent of directed credit programs in recent years.

## 3. Empirical Approach

The importance of financial liberalization as a precondition for the growth-enhancing effects of aid can be illustrated with the AK model of Rebelo (1991) in which total factor productivity ($A$) is a function of aid, financial liberalization, and their interaction.

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*Table 1. Principal Component Analysis for the Financial Liberalization Index*

| | *Principal component* | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | *1* | *2* | *3* | *4* | *5* | *6* | *7* | *8* | *9* |
| Eigenvalue | 4.247 | 1.943 | 1.101 | 0.600 | 0.463 | 0.338 | 0.182 | 0.089 | 0.035 |
| % of variance | 0.472 | 0.216 | 0.122 | 0.067 | 0.051 | 0.038 | 0.020 | 0.010 | 0.004 |
| Cumulative % | 0.472 | 0.688 | 0.810 | 0.877 | 0.928 | 0.966 | 0.986 | 0.996 | 1.000 |
| | *Eigenvector* | | | | | | | | |
| *Policy variable* | *1* | *2* | *3* | *4* | *5* | *6* | *7* | *8* | *9* |
| $FDR_t$ | 0.397 | 0.140 | 0.422 | 0.165 | 0.221 | 0.232 | 0.084 | 0.626 | 0.341 |
| $DRC_t$ | 0.398 | 0.223 | 0.206 | 0.370 | 0.234 | 0.110 | 0.603 | 0.033 | 0.426 |
| $DRF_t$ | 0.223 | 0.522 | 0.195 | 0.473 | 0.155 | 0.187 | 0.569 | 0.109 | 0.148 |
| $FLR_t$ | 0.172 | 0.277 | 0.699 | 0.218 | 0.581 | 0.043 | 0.118 | 0.000 | 0.068 |
| $LRC_t$ | 0.098 | 0.577 | 0.206 | 0.563 | 0.257 | 0.296 | 0.129 | 0.357 | 0.002 |
| $LRF_t$ | 0.318 | 0.353 | 0.345 | 0.103 | 0.302 | 0.621 | 0.221 | 0.294 | 0.180 |
| $CRR_t$ | 0.356 | 0.319 | 0.228 | 0.163 | 0.514 | 0.241 | 0.435 | 0.356 | 0.234 |
| $SLR_t$ | 0.449 | 0.146 | 0.154 | 0.052 | 0.296 | 0.075 | 0.049 | 0.394 | 0.707 |
| $DCP_t$ | 0.407 | 0.005 | 0.129 | 0.456 | 0.161 | 0.600 | 0.187 | 0.317 | 0.297 |

*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, $CRR_t$ = the cash reserve ratio on time deposits, $SLR_t$ = the statutory liquidity ratio, and $DCP_t$ = directed credit program.

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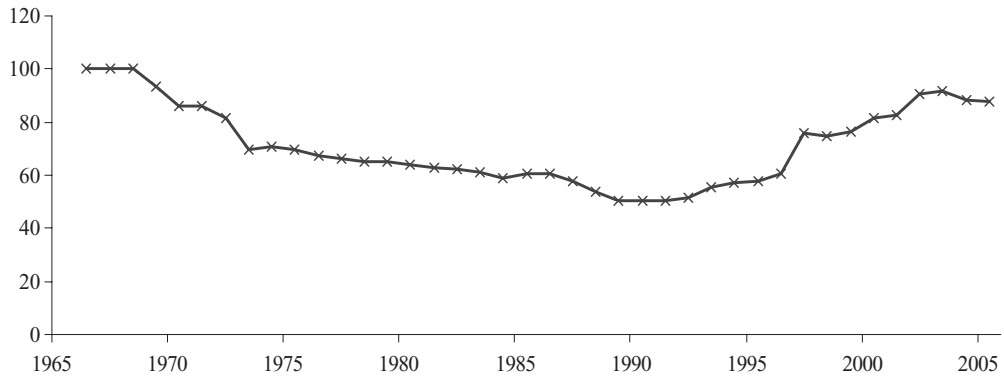


*Figure 1. Financial Liberalization Index (1966–2005)*

*Note*: The first observation is normalized to take the value of 100.

This simple endogenous growth setting can be used to guide our empirical formulation so that economic development ($ED_t$) depends on the capital stock ($KAP_t$), financial liberalization ($FL_t$), foreign aid ($AID_t$), and the interaction term between financial liberalization and foreign aid ($FL_t \times AID_t$), as given in equation (1):

$$\ln ED_t = \beta_0 + \beta_1 \ln KAP_t + \beta_2 \ln FL_t + \beta_3 \ln AID_t + \beta_4 \ln FL_t \times \ln AID_t + \varepsilon_t. \quad (1)$$

In the above equation, $\beta_1$ is expected to carry a positive sign, whereas the signs expected for $\beta_2$ and $\beta_3$ cannot be determined *a priori*. $\beta_4$ is expected to be positive due to the hypothesis that the impact of foreign aid on per capita real GDP is enhanced through the degree of liberalization in the financial system. The model will be estimated using annual data for India over the period 1966–2005. Data sources and construction of variables are explained in the Appendix, Table A1.

The main econometric procedure used to test for the existence of the long-run equilibrium relationship and to provide estimates of this long-run relationship is the ARDL procedure of Pesaran et al. (2001). The approach has several desirable statistical properties, such as precise estimates of long-run parameters and valid *t*-statistics, even in the presence of endogenous explanatory variables. Pesaran and Shin (1999) have shown that the OLS estimators of the short-run parameters are consistent and the ARDL-based estimators of the long-run coefficients are super-consistent in small sample sizes. Hence, valid inferences on the long-run parameters can be made using standard normal asymptotic theory.

The ARDL procedure involves two stages. In the first stage, the existence of the long-run relationship between the variables is tested. The ARDL model for the dynamic output equation can be formulated as:

$$\Delta \ln ED_t = a_0 + b_0 \ln ED_{t-1} + \sum_{j=1}^{k} b_j \ln DET_{j,t-1} + \sum_{i=1}^{p} c_{0i} \Delta \ln ED_{t-i} + \sum_{i=0}^{p} \sum_{j=1}^{k} c_{ji} \Delta \ln DET_{j,t-i} + \varepsilon_t, \quad (2)$$

where $p$ is the lag length, and $DET_t$ is a vector of $k$ determinants of $\ln ED_t$. The null hypothesis of no long-run relationship between $\ln ED_t$ and its determinants is

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*Table 2. Results for Unit-Root Tests*

| | *ADF* | | *PP* | |
|---|---|---|---|---|
| | *Levels* | *First-differenced* | *Levels* | *First-differenced* |
| $\ln ED_t$ | –1.481 | –6.061*** | –0.079 | –6.105*** |
| $\ln KAP_t$ | 2.381 | –2.786* | 3.101 | –3.021** |
| $\ln FL_t$ | –0.785 | –3.285** | –0.774 | –4.748*** |
| $\ln AID_t$ | –2.840 | –5.277*** | –2.724 | –6.164*** |
| $\ln FL_t \times \ln AID_t$ | –1.641 | –6.006*** | –1.916 | –6.027*** |

*Notes*: For ADF, AIC is used to select the lag length and the maximum number of lags is set at five. For PP, Barlett-Kernel is used as the spectral estimation method. The bandwidth is selected using the Newey–West method. *, ** and *** indicate 10%, 5%, and 1% levels of significance, respectively.

$H_0 : b_0 = b_1 = \ldots = b_k = 0$. We first estimate equation (2) by the OLS estimator and then calculate the $F$-statistic. 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 second stage of the procedure is to derive the long-run and short-run estimates using the underlying ARDL model.

## 4. Empirical Findings

We begin our empirical analysis by assessing the integration properties of the underlying variables. Two standard unit-root tests were used to assess the order of integration of the underlying variables—the augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests. The results reported in Table 2 show that all variables appear to be integrated at order one, or I(1), at the 1% level of significance. This allows legitimate use of the ARDL bounds test since the procedure requires all underlying variables to be integrated at an order less than two. Results based on alternative unit-root tests that consider the presence of structural breaks in the series yield the same conclusion (see Ang, 2009a, for details).

Next, to perform the ARDL bounds test, we estimate equation (2) with one and two lags for the model. We do not consider a lag length greater than two in order to conserve the degrees of freedom, given the small sample used in this study. Table 3 gives the $F$-statistics for the ARDL bounds tests, as well as the Akaike and Schwarz Bayesian Information Criteria (denoted by AIC and SBC, respectively).

The test for the presence of a long-run relationship results in an $F$-statistic of 3.415 when one lag is chosen. This statistic is lower than the 10% upper bound value of 3.520, suggesting that no evidence of cointegration is found when one lag is considered. However, the results indicate the null hypothesis that there exists no per capita output equation is rejected at the 1% significance level for the model when two lags are chosen.[1]

The null of no cointegrated relationship between the variables cannot be rejected when $\ln KAP_t$, $\ln FL_t$, $\ln AID_t$, or $\ln FL_t \times \ln AID_t$ are alternatively chosen to be the dependent variables. Hence, the results suggest these variables can be interpreted as

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*Table 3. ARDL Bounds Tests*

| | | $p = 1$ | $p = 2$ |
|---|---|---|---|
| I. | $F$-statistic | 3.415 | 7.481*** |
| II. | Model section criteria: *AIC* | –4.469 | –4.858 |
| | *SBC* | –3.822 | –3.988 |
| III. | Diagnostic checks: $\chi^2_{NORMAL}$ | 6.681** | 0.151 |
| | | (0.035) | (0.927) |
| | $\chi^2_{SERIAL}$ | 3.675* | 1.827 |
| | | (0.055) | (0.176) |
| | $\chi^2_{WHITE}$ | 11.858 | 20.439 |
| | | (0.617) | (0.368) |
| | $\chi^2_{ARCH}$ | 0.018 | 1.562 |
| | | (0.892) | (0.211) |

*Notes*: $p$ is the optimal lag length for the ARDL model. The test statistics are compared against the critical values reported in Pesaran et al. (2001). For the case with five variables, an intercept and no trend, the 10%, 5%, and 1% critical value bounds for the ARDL bounds test are (2.450, 3.520), (2.860, 4.010), and (3.740, 5.060), respectively. $\chi^2_{NORMAL}$ refers to the Jarque–Bera statistic of the test for normal residuals, $\chi^2_{SERIAL}$ is the Breusch–Godfrey LM test statistic for no first-order serial correlation, $\chi^2_{WHITE}$ denotes White's test statistic to test for homoskedastic errors, and $\chi^2_{ARCH}$ is Engle's test statistic for no autoregressive conditional heteroskedasticity. Figures in parentheses indicate $p$-values. *, **, and *** indicate 10%, 5%, and 1% levels of significance, respectively.

the long-run forcing variables explaining ln $ED_t$.[2] Alternative cointegration test results reported in Ang (2009a) also point to the same conclusion that the underlying variables form a robust long-run relationship.

In line with the results of the bounds test, both AIC and SBC prefer a richer dynamic specification of two lags. Furthermore, the choice of two lags is not subject to any econometric problems, based on the results of the diagnostics checks reported in panel III. Thus, we find that using two lags is more appropriate in this case and have chosen to follow this lag structure in the remaining analyses.

Panel I of Table 4 provides estimates of the long-run relationship, whereas panel II gives the results of the short-run dynamics. Capital stock enters the long-run equation significantly at the 1% level with the expected positive sign. Specifically, the long-run elasticity of per capita real output with respect to capital stock is found to be 0.634. The results suggest that capital stock has played a vital role in the process of economic development in India, a finding consistent with the growth literature that emphasizes the importance of capital deepening.

Every 1% increase in the composite index of financial liberalization is associated with a 4.145% increase in per capita real GDP. The results imply that liberalization of the financial system in India has a favorable effect on economic development. The finding of this positive influence provides some support for the financial liberalization thesis of McKinnon (1973) and Shaw (1973), which argues in favor of removing all financial restraints in order to foster economic growth. Our results are also broadly consistent with Ang (2010b) and Madsen et al. (2010), who find that financial liberalization leads to higher innovative or productivity growth in Korea and India, respectively.

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*Table 4. ARDL Estimates of the Long-Run Relationship and the Short-Run Dynamics*

| *I. The long-run relationship* (*Dep. var.* = ln $ED_t$) | *Coefficient* | *Std. error* | *p-Value* |
|---|---|---|---|
| *Intercept* | –14.623** | 6.848 | 0.043 |
| ln $KAP_t$ | 0.634*** | 0.044 | 0.000 |
| ln $FL_t$ | 4.145** | 1.549 | 0.013 |
| ln $AID_t$ | –2.539** | 0.974 | 0.015 |
| ln $FL_t \times$ ln $AID_t$ | 0.595** | 0.231 | 0.016 |
| *II. The short-run dynamics* (*Dep. var.* = Δ ln $ED_t$) | *Coefficient* | *Std. error* | *p-Value* |
| *Intercept* | –0.022 | 0.020 | 0.288 |
| $ECT_{t-1}$ | –0.381*** | 0.090 | 0.000 |
| Δ ln $KAP_t$ | 1.525*** | 0.408 | 0.001 |
| Δ ln $FL_t$ | 0.631* | 0.342 | 0.076 |
| Δ ln $AID_t$ | –0.457** | 0.217 | 0.045 |
| Δ(ln $FL_t \times$ ln $AID_t$) | 0.108** | 0.050 | 0.039 |
| Δ ln $GDP_{t-1}$ | –0.276* | 0.155 | 0.086 |
| Δ ln $GDP_{t-2}$ | –0.098 | 0.144 | 0.502 |
| Δ ln $AID_{t-2}$ | –0.051 | 0.045 | 0.268 |
| Δ(ln $FL_{t-2} \times$ ln $AID_{t-2}$) | 0.014 | 0.010 | 0.178 |
| *III. Diagnostic checks* | *Test statistic* | | *p-Value* |
| $\chi^2_{NORMAL}$ | 5.704* | | 0.058 |
| $\chi^2_{SERIAL}$ | 0.011 | | 0.919 |
| $\chi^2_{WHITE}$ | 5.627 | | 0.776 |
| $\chi^2_{ARCH}$ | 0.046 | | 0.831 |

*Notes*: The resulting lag structure chosen using AIC for the underlying ARDL model is (2, 1, 2, 2, 2). *, **, and *** indicate 10%, 5%, and 1% levels of significance, respectively.

A key feature of the financial policy environment in India is the presence of significant directed credit programs. A recent study by Ang (2009b) shows that private investment activity in India is indeed negatively associated with the presence of these directed loans. Owing to the nationalization of banks in 1969, the allocation of credit has been mainly performed by government banks, which are often deemed less efficient and subject to political interference. These credit allocation programs, which function mainly as a set of transfer programs, have been found to have little impact on agricultural growth and productive capacity (Hanson, 2001). Given that government intervention in credit allocation has not created many new sources of entrepreneurship, it is not surprising that these programs have been found to have negative impacts on economic expansion. Moreover, although the allocation of credit under the direction of the central bank has benefited some farmers and small traders by allowing them to have adequate access to finance, this may have also discouraged household saving and hence reduced funds available for investment. As such, our results advise in favor of a policy of deregulation in the financial system in the form of reducing the requirements for direct lending in order to boost economic development.

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Moreover, high reserve requirements before the liberalization in the early 1990s have traditionally provided the Reserve Bank of India with funds to purchase government securities at low cost, leaving insufficient funds to finance risky private investment projects. As shown by Ang (2009b), a significant reduction in the ratios of reserve and liquidity requirements following the liberalization has greatly expanded the amount of loanable funds, which has contributed to an economic boom during the 1990s. Therefore, it appears that lowering these requirements can provide significantly more loanable funds, enabling the undertaking of more investment activity to fuel economic development. Our results therefore point to the importance of eliminating these distortionary policies so that aid funds received from foreign donors can be allocated more efficiently, in a financial system that is less subject to interference, in order to boost economic activity.

Our model conjectures that foreign aid has both direct and indirect effects on output expansion. Holding the indirect effect constant, the results show that foreign aid is found to have a direct negative effect on economic development in India, with a negative long-run elasticity of 2.539. Thus, our results do not lend any support to the view that aid will promote growth. The results, instead, suggest that resources from foreign aid have been misused and misallocated and therefore were unlikely to exert any positive impact on growth in India. Our results corroborate the cross-country findings of Knack (2001) and Nkusu and Sayek (2004), who have shown that foreign aid has a dampening effect on economic growth. However, the results stand in sharp contrast to the earlier findings of Chenery and Strout (1966) and Papanek (1973). Our finding is also consistent with the theoretical models developed by Kimbrough (1986) and Gong and Zou (2001), which predict that both investment and output growth rates will move in opposite directions in response to an increase in foreign aid receipts.

In general, the receipts of aid can be transferred to the private sector directly, through tax reductions, or public spending. The findings of Swaroop et al. (2000), however, suggest that nearly all external assistance was retained by the central government whilst only a handful of the funds were transferred to state governments in India. Funds administered by the central government were largely spent on unnecessary investment activities that would not have otherwise been undertaken. These projects tend to generate low or negative rates of return and produce little spillovers into other sectors. Moreover, the extensive use of foreign aid to finance centralized government activities may have crowded out private investment due to competition from public enterprises. More red tape and greater uncertainty in the economy may have also contributed to a significant reduction in the incentives to invest in the private sector. Hence, foreign aid is very likely to have resulted in lower aggregate demand in India.

With regard to the indirect effect of foreign aid, the interaction term is found to be statistically significant and has the expected sign (positive). Therefore, it can be inferred that the impact of foreign aid on the Indian economy is strengthened by the level of financial liberalization. This is obvious when we obtain the derivative of $\ln ED_t$ with respect to $\ln AID_t$. The results suggest that aid will have a detrimental effect on economic development unless the financial system in India is liberalized beyond a certain level so that efficient allocation of aid resources can take place. In this case, the threshold from which the effect of aid becomes beneficial is 71.307 (or 4.267 in natural logarithms), a level which has been achieved since 1997 (see Figure 1). Hence, both the direct and indirect effects of financial liberalization on economic development are found to be positive. On the whole, the results tend to support the argument that

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financial liberalization greatly facilitates the management of aid flows, as emphasized by Nkusu and Sayek (2004).

From a policy standpoint, the key finding in this paper gives rise to an important question: if financial liberalization has so far enabled the beneficial effects of development aid to be realized, should the government further liberalize the financial system? There is no unique correct policy stance as this largely depends on the costs and benefits associated with further liberalization. Nevertheless, the government should always adopt a prudent stance towards any further liberalization given that unrestricted financial liberalization may induce instability in the financial system. As noted by Rajan and Zingales (2003), financial liberalization is more likely to work well in environments with strong regulatory capacity. Although the legal system in India was originally based on the British model that emphasizes protection of property rights, India ended up with a much less effective institutional framework since the system was modified in a way that benefited the small number of Europeans that settled in and ran the economy (see Ang, 2010a, and the references therein). This therefore highlights the importance of strengthening the institutional framework before any further liberalization is undertaken so that financial policies can be more effectively carried out to strengthen the effectiveness of aid on growth.

Turning to the short-run dynamics, the regression results for the short-run model reported in panel II of Table 4 show several salient features. In first-differenced form, financial liberalization, aid, and their interaction have signs consistent with those reported in the long-run model, although these effects are much smaller than their long-run counterparts. The coefficient on $ECT_{t-1}$, which measures the speed of adjustment back to the long-run equilibrium value, are statistically significant at the 1% level and correctly signed; i.e. negative. This implies that an error-correction mechanism exists in the per capita output function so that the deviation from long-run equilibrium has a significant impact on the growth rate of per capita output. The equation adjusts at about 38% every year to restore equilibrium when there is a shock to the steady-state relationship. In order to assess the sensitivity of the results, we subject the estimation to a number of robustness checks in the next section.

## 5. Robustness Checks

### *Diagnostic Tests*

The results reported in panel III of Table 4 show that the regression specification fits remarkably well and passes the diagnostic tests against nonnormal residuals, serial correlation, heteroskedasticity, and autoregressive conditional heteroskedasticity at the 5% level of significance. Structural stability of the equations is examined using the cumulative sum (CUSUM) and CUSUM of squares tests on the recursive residuals. The former is able to detect systematic changes in the regression coefficients, whereas the latter is able to detect sudden changes from the constancy of the regression coefficients. The results (not reported) show that the test statistics are within the 5% confidence interval band, suggesting the estimated model is rather stable over time. Moreover, it is also found that the predicted series of per capita real GDP is able to track the actual series very closely over time, providing some confidence that the model is well-fitted.[3]

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*Table 5. Alternative Estimates*

| *Dep. var.* = ln $ED_t$ | *FM-UECM* | *FM-OLS* | *DOLS* | *VECM* |
|---|---|---|---|---|
| *Intercept* | –5.897** | –6.303*** | –30.381*** | –46.251 |
| ln $KAP_t$ | 0.584*** | 0.577*** | 0.679*** | 0.747*** |
| ln $FL_t$ | 2.185*** | 2.286*** | 7.783*** | 11.401*** |
| ln $AID_t$ | –1.293*** | –1.367*** | –4.782*** | –6.954*** |
| ln $FL_t \times$ ln $AID_t$ | 0.298*** | 0.314*** | 1.121*** | 1.639*** |

*Note*: Two lags are used in all estimations. ** and *** indicate 5% and 1% level of significance, respectively.

### *Alternative Estimators*

While the ARDL approach is used for the main results of this paper, to provide a sensitivity check of the results we also consider four other estimators, namely the FM-UECM estimator, the FM-OLS procedure, the DOLS approach, and the vector error-correction model (VECM).[4] Since our focus is on the long-run results, the short-run dynamics generated by each estimator are not reported here for brevity. In general, these approaches give very similar results compared to those estimated using the ARDL procedure (see Table 5). All variables enter the long-run equation significantly at the conventional levels. The main theme is that the interaction term continues to be highly significant, highlighting the importance of the complementary role of financial sector reforms in absorbing the benefits of foreign aid.

### *Controlling for Other Effects*

Having examined the robustness of the econometric results, we now turn to presenting the results with additional control variables. We derive the results using the ARDL estimator. Since we are mainly interested in the long-run relationship, only the long-run results are reported to conserve space. Specifically, we control for the presence of a nonlinear effect ($AID_t \times AID_t$) due to diminishing returns to aid (Dalgaard and Hansen, 2001; Lensink and White, 2001), the interaction between aid and other financial or institutional factors, including quality of institutions ($INS_t$), banking sector development ($BANK_t$), stock market development ($STOCK_t$), and share market volatility ($VOL_t$). The results are reported as columns (1) to (5) in Table 6. The construction of these variables is explained in the Appendix, Table A1.

It is evident that there is no support for the presence of a nonlinear effect, implying that a threshold effect does not exist in the relationship between aid and per capita real GDP for India. Apart from the measure of stock market volatility, all other control variables and their interaction terms are found to be statistically insignificant. For instance, the effectiveness of aid on growth in India does not depend on the presence of a good institutional framework such as strong intellectual property rights protection—a finding consistent with Easterly et al. (2004) and Alvi et al. (2008). Moreover, the measures of financial development are found to have no effect on per capita GDP. This is probably due to the inclusion of the financial liberalization index in the specification, as studies have shown that financial liberalization is an important determinant of financial development (e.g. see Ang and McKibbin, 2007).

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*Table 6. Controlling for Nonlinear and Other Interaction Effects*

| *Dep. var.* = ln $ED_t$ | *(1)* | *(2)* | *(3)* | *(4)* | *(5)* |
|---|---|---|---|---|---|
| ln $KAP_t$ | 0.583*** | 0.642*** | 0.691*** | 0.491*** | 0.532*** |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| ln $FL_t$ | 2.513** | 2.623** | 0.463*** | 1.057*** | 1.218*** |
| | (0.024) | (0.025) | (0.003) | (0.000) | (0.000) |
| ln $AID_t$ | –1.847*** | –1.512** | –0.272*** | –0.558*** | –0.457*** |
| | (0.006) | (0.042) | (0.004) | (0.000) | (0.000) |
| ln $FL_t \times$ ln $AID_t$ | 0.348** | 0.351** | 0.055*** | 0.131*** | 0.154*** |
| | (0.034) | (0.046) | (0.004) | (0.000) | (0.000) |
| ln $AID_t \times$ ln $AID_t$ | –0.025 | | | | |
| | (0.155) | | | | |
| ln $INS_t$ | | 0.135 | | | |
| | | (0.918) | | | |
| ln $INS_t \times$ ln $AID_t$ | | 0.036 | | | |
| | | (0.853) | | | |
| ln $BANK_t$ | | | –0.415 | | |
| | | | (0.251) | | |
| ln $BANK_t \times$ ln $AID_t$ | | | –0.033 | | |
| | | | (0.556) | | |
| ln $STOCK_t$ | | | | 0.018 | |
| | | | | (0.839) | |
| ln $STOCK_t \times$ ln $AID_t$ | | | | –0.006 | |
| | | | | (0.661) | |
| ln $VOL_t$ | | | | | –0.542* |
| | | | | | (0.061) |
| ln $VOL_t \times$ ln $AID_t$ | | | | | –0.087* |
| | | | | | (0.051) |
| $F$-statistic (ARDL bounds test) | 6.315*** | 4.515*** | 5.193*** | 5.162*** | 4.519*** |

*Notes*: These additional interaction terms are found to be either I(0) or I(1), allowing legitimate use of the ARDL bounds tests. For the case with six variables (column (1)), an intercept and no trend, the 10%, 5%, and 1% critical value bounds for the ARDL bounds test are (2.26, 3.35), (2.62, 3.79), and (3.41, 4.68), respectively. For the case with seven variables (columns (2) to (5)), an intercept and no trend, the 10%, 5%, and 1% critical value bounds for the ARDL bounds test are (2.12, 3.23), (2.45, 3.61), and (3.15, 4.43), respectively. Figures in parentheses indicate $p$-values. *, **, and *** indicate 10%, 5%, and 1% levels of significance, respectively.

Importantly, the inclusion of additional controls does not alter our main findings. In other words, financial liberalization and its interaction with aid continue to have a positive influence, whereas aid continues to have a direct negative effect on output expansion. These effects are found to be statistically significant at the conventional levels. It should also be highlighted that the evidence of cointegration remains robust to the inclusion of these control variables. Therefore, we conclude that our principal results are not sensitive to the inclusion of a number of control variables.

## 6. Conclusions

Notwithstanding the growing concern over the importance of foreign aid in India, few studies have attempted to assess the impact of aid on the country's growth.

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The paper empirically investigates this relationship by focusing on the complementary role of financial sector reforms. Specifically, we tested the hypothesis that financial liberalization enhances recipient countries' ability to effectively absorb foreign aid, drawing on the experience of a large and rapidly growing developing economy.

Using the ARDL bounds test, the empirical evidence shows a significant long-run relationship between per capita real GDP and all its determinants. The results also reveal that financial liberalization is an important factor determining output growth for India both in the short run and long run, indicating the financial liberalization thesis of McKinnon (1973) and Shaw (1973) is firmly supported by the Indian data. We found that aid inflows per se may retard output expansion. However, India is able to gain significantly more from aid flows with a more liberalized financial system. Our empirical results are robust to several sensitivity checks, including model stability, the use of different estimators, and the inclusion of a number of control variables.

The results have some policy implications. First, any impediments to financial sector reform are likely to harm India's economic prosperity. Second, foreign aid has a direct detrimental effect on the economic performance of India. However, the benefits of aid can be realized through liberalizing the financial sector. Thus, liberalizing the financial system can enhance the effectiveness of aid on growth. Third, the presence of a good policy environment in the form of a stronger intellectual property rights protection framework has no direct implication for the aid–growth nexus in India. Fourth, the beneficial effect of foreign aid is more likely to be realized through the channel of financial liberalization rather than banking sector or stock market development. Finally, there is no evidence to support the presence of a nonlinear effect in the relationship, suggesting that diminishing returns to aid have not occurred in India.

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## Appendix

*Table A1. Construction of Variables and Data Sources*

| *Variable* | *Description* | *Sources* |
|---|---|---|
| $ED_t$ | Economic development is measured per capita GDP at 1993 prices. | *National Accounts Statistics*, Government of India |
| $KAP_t$ | The initial capital stock is computed by taking the initial gross capital formation at constant prices divided by the sum of the depreciation rate (assumed to be 10%) and the average growth rate of gross capital formation at constant prices over the period 1966–2005. Capital stocks are then computed using the standard perpetual inventory model. | *National Accounts Statistics*, Government of India |
| $AID_t$ | Foreign aid is measured by the ratio of utilization of external assistance to nominal GDP. | *National Accounts Statistics*, Government of India |
| $FL_t$ | The composite index of financial liberalization consists of six interest rate control dummies, cash reserve ratio, statutory liquidity ratio, and directed credit programs. The reserve and liquidity ratios are direct measures expressed in percentages. The extent of directed credit program is measured by 0, 1, 2, and 3 when the programs cover 0%, up to 20%, 21% to 40%, and more than 40%, respectively, of total bank loans. The index is constructed using the method of principal component analysis. Details of index construction are described in section 2. | *Annual Reports* and *Report on Currency and Finance*, Reserve Bank of India |
| $INS_t$ | Quality of institutions in this study is measured by the protection of intellectual property rights. The intellectual property rights index covers five dimensions: (1) patentability of various kinds of inventions, (2) membership in international patent arrangements, (3) provisions for loss protection, (4) enforcement mechanisms, and (5) duration of the patent term. Each dimension is assigned a value ranging from zero to one. The unweighted sum of these five values provides an indication of the overall level of intellectual property rights protection, with higher values reflecting greater levels of protection. Missing years are interpolated. | Park and Lippoldt (2005) |
| $BANK_t$ | Following the established practice, banking sector development is measured by bank credit to commercial sector divided by nominal GDP (see e.g. Dellas and Hess, 2002; Beck, 2003). | *Annual Reports* and *Report on Currency and Finance*, Reserve Bank of India |
| $STOCK_t$ | Stock market development is measured by the ratio of share market capitalization to nominal GDP. Data for stock market capitalization before 1976 are constructed using the share price index. | *International Financial Statistics* CD-ROM and Beck et al. (2000) |
| $VOL_t$ | Stock market volatility is measured by the five-year rolling standard deviation of the growth rate of the share price index. | *International Financial Statistics* CD-ROM |

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## Notes

1. The bounds test statistics are also compared against the critical values for small samples provided by Narayan (2005), which are calculated using stochastic simulations specific to the sample size based on 40,000 replications. The 10%, 5%, and 1% critical bounds for a sample size of 40 are (2.660, 3.838), (3.202, 4.544), and (4.428, 6.250), respectively. Hence, our conclusion regarding the evidence of cointegration remains unchanged even if the critical bounds are adjusted for finite samples.
2. The issue of causality is more formally addressed in Ang (2009a).
3. These results are not reported to conserve space, but they are available upon request.
4. The coefficient of the error-correction term in the VECM model is found to be –0.238 with a *t*-statistic of –2.815, providing further evidence in support of cointegration. This estimate suggests that the economy adjusts at the rate of 23.8% per year or that it takes approximately 4.2 years to restore equilibrium when there is a shock to the steady-state relationship.

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