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FINANCE AND INEQUALITY:
THE CASE OF INDIA

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Abstract

Although theory emphasizes the role of financial market frictions in explaining income inequality, there is little empirical research exploring how financial development and financial sector reforms influence the evolution of income inequality. This paper examines how finance impacts on income inequality in India using annual time series data for over half a century. The results indicate that while financial development helps reduce income inequality, financial liberalization seems to have exacerbated income inequality in India. Our results are robust to the use of different measures for financial development and financial liberalization.

Keywords: Financial development; financial liberalization, income inequality; India.
JEL classification: G28; O16; O53

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1. Introduction

Although the relationship between financial development and economic growth has been extensively studied in the literature (see, e.g., King and Levine, 1993; Demetriades and Hussein, 1996; Arestis and Demetriades, 1997; Levine et al., 2000; Bell and Rousseau, 2001; Luintel et al., 2008), little is known about how finance impacts on income inequality. The importance of the finance-inequality relationship has recently been highlighted in an insightful survey article by Claessens and Perotti (2007). They indicate that while financial development can help reduce income inequality, financial liberalization captured by established interests may do the opposite.

The theoretical predictions of the effects of finance on income inequality are controversial. Rajan and Zingales (2003b) argue that improvements in the formal financial sector primarily benefit the rich. Greenwood and Jovanovic (1990) predict a non-linear relationship between financial development and income inequality, where it is hypothesized that income inequality first increases with the degree of sophistication in the financial systems, then stabilizes and eventually declines. Others propose that the presence of financial market imperfections deters the poor from borrowing adequately to invest in human and physical capital, implying that financial development helps alleviate income inequality (Banerjee and Newman, 1993; Galor and Zeira, 1993; Aghion and Bolton, 1997; Mookherjee and Ray, 2003, 2006). Given that theories provide ambiguous predictions regarding the effects of finance on the distribution of income, an alternative is to approach the issue at the empirical level. This could facilitate our understanding of the relationship between finance and inequality, and therefore provide some useful guide for assessing the validity of each theoretical model.

However, despite the important role of financial market frictions in the theories of poverty and income inequality, researchers so far have not adequately addressed whether financial development, and in particular financial sector policies, affect income inequality (Demirgüç-Kunt and Levine, 2007). In this connection, there are two novel studies that focus on examining the effect of financial development on income inequality. Using data for 83 countries over the period 1960-1995, Clarke et al. (2006) examine the effect of financial development on the level of the Gini coefficient – a measure of deviations from perfect income equality. Their results show that financial deepening is associated with lower income inequality. The finding of a non-linear effect of financial development is not robust. A more recent study by Beck et al. (2007) attempts to assess the impact of financial development on the changes of income distribution and income for the poor. Their main findings indicate that financial development is associated with a lower growth rate of the Gini coefficient and a higher growth rate of income for the poor. While these two studies have established that financial development helps reduce income inequality, studies examining the direct impact of financial liberalization on income inequality are particularly scant (Demirgüç-Kunt and Levine, 2007). The limited indirect empirical evidence, based on the survey by Arestis and Caner (2004), seems to suggest that financial liberalization has ambiguous effects on the poor and income distribution.

This paper attempts to contribute to this rather under-researched area by considering an important case study – that of India for the period 1951-2004. Specifically, we analyze the distributional impact of financial development and financial liberalization on the Gini coefficient.[1] The paper aims to complement the above studies, and enrich the literature by providing further evidence on how development of financial systems and implementation of financial sector policies affect the evolution of income inequality, drawing on the experience of one of the most rapidly growing developing economies that has undergone significant financial sector reforms. We focus on India rather than a larger set of countries given that the effects of financial development and financial liberalization may be heterogeneous across countries at different stages of economic development. Moreover, case studies are particularly useful in disentangling the complexity of the financial environments and economic histories of each country. By analyzing case studies, the econometric findings of this project can be related to the prevailing institutional structure (Bell and Rousseau, 2001), and therefore inform academic as well as policy debate.

The main contributions of this study include: 1) empirically testing the effect of financial development on income inequality by providing further evidence from a large and fast growing developing country. Not only could this enhance our understanding of the finance-inequality 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 financial liberalization on the Indian economy. Although various financial restructuring programs have been launched since the early 1990s, there is little empirical evidence informing policy makers the effects of these reforms; and 3) complementing the literature by

[1] Although financial crises may also have significant effects on income inequality, as demonstrated by Pritchett et al. (2000) in the case of Indonesia during the Asian financial crisis of 1997-98, this issue is not addressed in our paper since the economy of India was largely unaffected by the 1997-98 crisis.

assessing the impact of financial liberalization on income inequality. This policy factor has been somewhat neglected in the analysis of the finance-inequality nexus.

Our results show that income inequality decreases as the financial system deepens and broadens, consistent with the general findings in the literature. However, liberalization of the financial systems appears to have a harmful effect on income distribution, a finding that tends to support the political economy argument where captured financial sector reforms benefit only small elites.

The remainder of the paper is structured as follows. The next section describes the financial repression and liberalization experience of India. Three composite indices for financial liberalization are constructed to measure the joint influence of a number of policies implemented in the Indian financial system. Section 3 briefly reviews the theoretical literature on the relationship between finance and inequality. The econometric techniques employed in this study are explained in Section 4. The results are presented and analyzed in Section 5. Finally, we summarize the main findings and conclude.

2. Financial Sector Reforms in India

There was little financial repression imposed on the Indian financial system in the 1950s. However, the Reserve Bank of India gradually imposed more controls over the financial system by introducing interest rate controls in the 1960s. The statutory liquidity ratio was raised from 25 percent in 1966 to 38 percent in 1989. The cash reserve rate increased considerably from 3 to 15 percent 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 fourteen 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 percent of total lending. Revenue from financial repression was estimated to be 22.4 percent of total central government revenue during the period 1980-85 (see Giovannini and De Melo, 1993).

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 restructure the entire orientation of India’s financial development strategy from its position as a financially repressed system to one that was more open in order to provide a greater role for markets in price determination and resource allocation.

Consequently, interest rates were gradually liberalized, and reserve and liquidity ratios were reduced significantly. The equity market was formally liberalized in 1992, although the first country fund was set up earlier in 1986, allowing 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 relaxed in 1993, resulting in the establishment of more private and foreign banks. Regulations on portfolio and direct investment have since been eased. The exchange rate was unified in 1993-94 and most restrictions on current account transactions were eliminated in 1994.

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. 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.[2] 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. As regards 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.

This study uses the financial liberalization measure advanced by Demetriades and Luintel (1996, 1997). Their approach considers nine indicators of financial repressionist 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 1 if a control is present and 0 otherwise. The remaining three policies are directed credit programs, the cash reserve ratio and the statutory liquidity ratio. The extent of directed

[2] In a study that examines the influence of institutional quality on government ownership of banks, Andrianova et al. (2008) report that India has the highest ratio of state-owned or state-controlled bank assets to total commercial bank assets in a sample of 83 countries.

credit programs is measured by the share of directed credit lending in total lending.[3, 4] The other two variables are direct measures expressed in percentages.

Since we want to summarize the financial sector policies to obtain an overall measure of financial liberalization, the method of principal component analysis seems to be a natural choice. It is a systematic and sophisticated way of examining the patterns of relationship among the variables, with the objective of summarizing the information content of several observed variables into a handful of representative principal components. The method involves computing the linear combinations of the original variables so that the resulting principal components can capture a large proportion of the variance in the original variables. This can therefore serve the same purpose as the full set of original variables, but in a much more succinct manner. Given that the principal components are uncorrelated to each other and their conciseness, this approach sufficiently deals with the problems of multicollinearity and over-parameterization. The inverse of this measure can be interpreted as the extent of financial liberalization (see, e.g., Ang and McKibbin, 2007; Ang, 2008).

To provide a sensitivity check we also consider two alternative measures of financial liberalization. Firstly, the approach of Demetriades and Luintel (1996, 1997) is modified to allow for the policy changes that took place after the liberalization since their work only covers the period to 1991, prior to the reform programs. The modification involves taking 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. We use dummy variables to represent policy changes in these dimensions.[5]

In constructing the third summary measure of financial liberalization, we follow the approach of Abiad and Mody (2005). In particular, six policy dimensions are considered as the inputs to construct the measure: 1) credit controls and reserve requirements; 2) interest

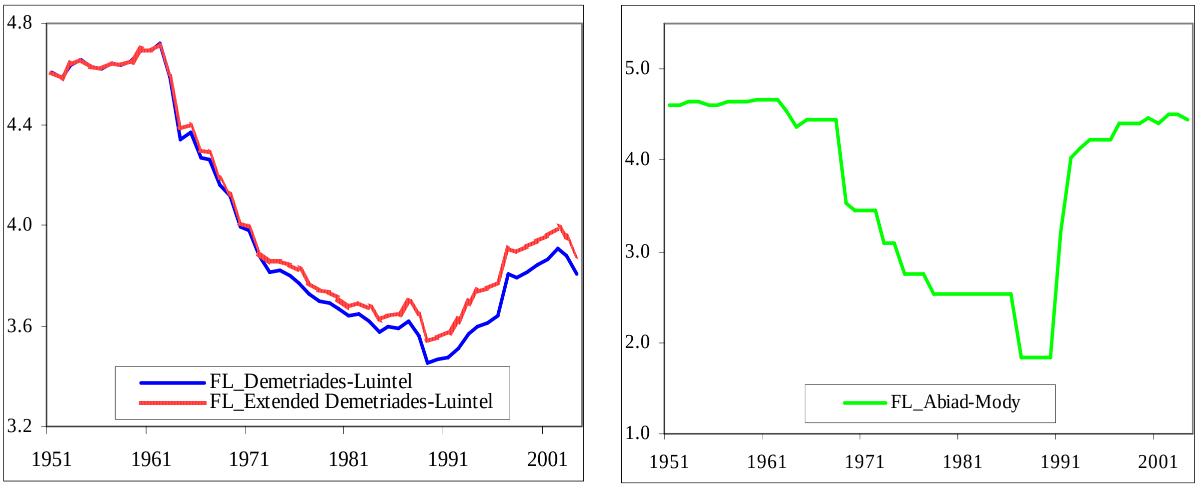
[3] Ideally, one should use a *de jure* (rather than a *de facto*) measure reflecting the strength of directed credit policies designed to repress the financial system. Unfortunately, such information is not available on a consistent and reliable basis for India.

[4] Demetriades and Luintel (1996, 1997) set the measure to zero when the directed credit programs are not implemented, and to 1, 2 and 3 when the programs cover up to 20%, 21-40% and over 40%, respectively, of total bank lending. However, we use the share of directed credit in total credit in order to allow for more variation in the series, particularly in the 1950s and 1960s where the ratio was always below 20%.

[5] For the first and second measures of financial liberalization, we extract six principal components, which are able to account for 97% and 95% of the total variation in the policy variables, respectively. These components are then summarized into just one composite measure using eigenvalues as the weights. We have also tried using just one and all principal components. However, our empirical results remain insensitive to the number of principal components extracted due to their high correlation structure.

rate restraint; 3) entry barriers in the banking sector; 4) government regulations of operations; 5) privatization in the financial sector; and 6) restrictions on international capital flows. Along each dimension, a score of zero, one, two or three is assigned, indicating fully liberalized, partially liberalized, partially repressed, and fully repressed, respectively. The aggregation of these six components is used to obtain an overall measure of financial repression.[6] Similar to the second approach, this provides a more broad-based measure of financial sector reforms since it considers several other dimensions in addition to credit and interest controls. All data series are directly obtained or compiled from the Annual Report and the Report on Currency and Finance of the Reserve Bank of India.

Figure 1: Financial Liberalization Composite Indices, 1951-2004 (in natural logarithms)



The resulting three composite financial liberalization indices displayed in Figures 1a and 1b coincide rather well with the actual policy changes that took place in India during the sample period, as discussed earlier. In Figure 1a, the two measures of financial liberalization show increasing disparity since the early 1970s given that the second measure captures more dimensions of financial sector reforms. It therefore necessarily reflects a greater degree of financial liberalization compared to the first measure that focuses exclusively on credit and interest controls. The financial liberalization series depicted in Figure 1b shows a rather different pattern of development, due largely to the use of a different coding procedure. Specifically, equal weights have been assigned to each component of financial sector reforms.

[6] We have also explored using the principal component analysis. Correlation analysis shows that this simple arithmetic mean is significantly and highly correlated with the first principal component and weighted average of all six principal components, suggesting that our results would not vary significantly with the use of any of these measures.

On the whole, 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. Conceptual Issues on Finance and Inequality

Developing countries are often characterized by the presence of credit constraints due to market imperfections such as asymmetric information and moral hazard problems. These credit constraints may intensify income inequality since the poor may not have equal access to credit due to the lack of collateral and established relationships with financial institutions. The relaxation of credit constraints achieved through an improvement in the financial systems enables efficient allocation of resources and thereby reduces income inequality. In the models developed by Banerjee and Newman (1993), Galor and Zeira (1993), Aghion and Bolton (1997) and Mookherjee and Ray (2003, 2006), among others, only rich agents can borrow enough to invest in human capital and high-yield investment projects due to credit market imperfections. Their models imply that borrowing constraints triggered by market failures could result in greater income inequality, and consequently, financial development helps alleviate income inequality.

On the other hand, Rajan and Zingales (2003a) argue that in the presence of weak institutional environments, *de jure* political representation is dominated by *de facto* political influence. This allows established interests to influence access to finance, implying that higher financial development induced by captured direct controls is likely to hurt the poor. Rajan and Zingales (2003b) further argue that development of financial systems is more likely to benefit the rich and well connected since they have sufficient wealth for collateral (dubbed "the tyranny of collateral"). The rich may also be able to prevent small firms from accessing external finance and reduce the ability of the poor to improve their economic well-being. Thus, the poor are often excluded from finance and are therefore unable to invest sufficiently in human and physical capital.

However, financial development and income inequality can also be characterized by a hump-shaped relationship. In an influential paper, Greenwood and Jovanovic (1990) present a theoretical model that predicts a non-monotonic relationship between the two variables. They postulate that access to finance involves a fixed transaction cost that poor households cannot afford. Such a market imperfection therefore causes deterioration in their relative position in the distribution of income. However, as the economy becomes more developed, the transaction costs of using financial services decline, allowing the majority to access finance so that financial deepening narrows the income gap between the rich and the poor. Hence, it appears that the above theoretical models offer quite different perspectives about the relationship between financial development and income inequality.

How financial liberalization impacts on income inequality is also theoretically ambiguous. Arestis and Caner (2004) propose that there are three main channels through which financial liberalization can influence poverty and income inequality. The first, known as the economic growth channel, posits that financial liberalization affects income inequality through increasing the rate of economic growth based on the financial liberalization thesis of McKinnon (1973) and Shaw (1973). However, this depends on the empirical links between financial liberalization and economic growth as well as economic growth and income distribution, which are not necessarily unambiguous. The second, the financial crisis channel, works through changes in macroeconomic volatility triggered by crises following financial liberalization. The poor are likely to be more vulnerable to these negative shocks. Finally, the last channel proposes that improved access to credit and financial services due to financial liberalization can have a profound effect on income distribution.

4. Empirical Approach

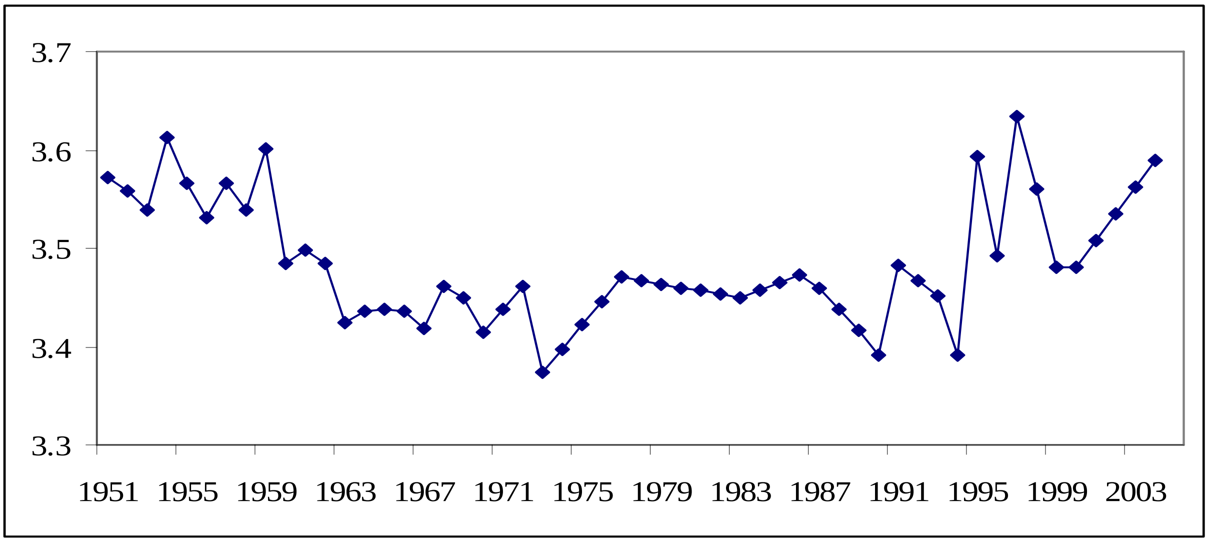
Our empirical model postulates that income inequality ($Gini_t$) depends on per capita growth rate of real GDP (GRO_t),[7] inflation rate (INF_t), trade openness (TO_t), and a variable that captures the effect of finance, as given in Eq. (1). Inflation rate is measured by the growth rate of the GDP deflator, and trade openness is the share of exports plus imports in GDP. For the finance variable, our focus is on the level of financial development and financial liberalization, although we also take into consideration their growth and non-linear effects. Besides, we also pay attention to stock market volatility and banking sector efficiency. The model will be estimated using annual data for India over the period 1951-2004. All data series are directly obtained or compiled from the Report on Currency and Finance of the Reserve Bank of India and National Accounts Statistics of the Central Statistical Organisation in India.

$$Gini_t = \beta_0 + \beta_1 GRO_t + \beta_2 INF_t + \beta_3 TO_t + \beta_4 Finance_t + \varepsilon_t \qquad (1)$$

[7] We use the growth rate of per capita real GDP as a control variable rather than the level of per capita real GDP since the influence of finance on income inequality operates through growth (Demirgüç-Kunt and Levine, 2007). Moreover, the use of a level income variable in our specification may lead to some econometric problems since per capita real GDP is highly correlated with the measures of financial development.

Figure 2 shows the evolution of the Gini coefficient in India over time.[8] It is evident that the series is subject to much variation over time. The Gini coefficient was on a falling trend in the 1950s and 1960s. However, this trend was reversed since the early 1990s, coinciding with implementation of the financial sector reforms in India.

Figure 2: Evolution of the Gini coefficient (on Ln scale)



The dynamic adjustment of the Gini coefficient can be characterized by a conditional error-correction model (ECM), which can be used to test for the existence of a long-run relationship using the ARDL bounds test developed by Pesaran et al. (2001) and the ECM test of Banerjee et al. (1998). The former involves a standard *F*-test whereas the latter is a simple *t*-test. Accordingly, the underlying ECM can be formulated as:

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

where p is the lag length and DET_t is a vector of k determinants of $Gini_t$. The above equation can be estimated by OLS since Pesaran and Shin (1998) 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-

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[8] The Gini coefficient is the ratio of the area between the Lorenz curve (which plots share of population against income share received) to the area below the diagonal. The value ranges from 0 to 1, where 0 means perfect income equality and 1 implies perfect income inequality. We use the Gini data from Deininger and Squire (1996) and Dollar and Kraay (2002). These data are updated with more recent data points available from ADB (2007) and UNDP (2007).

run parameters can be made using standard normal asymptotic theory. The main advantage of this approach is that it can be applied to the model regardless of whether the underlying variables are $I(0)$ or $I(1)$.

The testing procedure involves two stages. In the first stage, the existence of the long-run relationship between the variables is tested. Specifically, two separate statistics are employed to test for the existence of a long-run relationship in Eq. (2): 1) an F-test for the joint significance of coefficients on lagged levels terms of the conditional ECM ($H_0 : b_0 = b_1 = ... = b_k = 0$), and 2) a t-test for the significance of the coefficient associated with $Gini_{t-1}$ ($H_0 : b_0 = 0$). 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 second stage of the procedure is to derive the long-run estimates using the underlying ARDL model.

5. Estimated Results

5.1 Financial development and income inequality

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) test and Phillips-Perron (PP) test. The results, which are not reported to conserve space but are available upon request, show that all variables appear to be $I(0)$ or $I(1)$, suggesting that there is no variable integrated at an order greater than one. This allows legitimate use of the ARDL bounds and ECM tests since these procedures require all underlying variables to be integrated at an order less than two. Next, to perform the cointegration tests, we estimate Eq. (2) with only one lag in order to conserve the degrees of freedom, given the small sample used in this study (54 annual observations). The results reported in panel I of Table 1 strongly suggest that a long-run relationship is present for each of the models estimated.

However, the relationship between financial development and income inequality may be driven by reverse causality since lower income inequality may result in greater political pressures to create a more market-driven type of financial system in order to ensure efficient allocation of resources (Beck et al., 2007). Moreover, banks may also prefer to open branches in richer areas (Rajan, 2006). To address the concern of endogeneity bias, we have attempted to use financial development as the dependent variable. However, no evidence of cointegration is found when the measures of financial development are used as the dependent variables. Hence, the results suggest that financial development can be interpreted as one of the long-run forcing variables explaining the evolution in the Gini coefficient where a reverse causation is absent.

In testing the effect of financial development on the Gini coefficient, we control for per capital GDP growth, trade openness and inflation. These control variables have also been used by Beck et al. (2007). Panel II in Table 1 reports the long-run estimates derived using the underlying ARDL model. The negative sign and the significance of the coefficient for the growth rate of per capita real GDP suggest that income growth helps alleviate income inequality, a result consistent with the general literature on growth and inequality. In line with the results of Barro (2000), openness to trade enters positively and significantly in all equations except column (5).

The rate of inflation enters negatively and significantly in all regressions, suggesting that monetary instability does not seem to hurt income distribution. In principle, inflation may have an adverse effect on real agricultural wages and hence income inequality. However, it may also be associated with a decline in unemployment due to lower real wages and thus benefit the poor. Our results are consistent with the cross-country findings of Cutler and Katz (1991) and Clarke et al. (2006), but stand in sharp contrast to those of Easterly and Fischer (2001) and Beck et al. (2007). However, as highlighted by Easterly and Fischer (2001), the way inflation affects the poor may well differ between economies due to the complication of the tax system, and therefore is an empirical issue.

Financial development is measured by three commonly used indicators in the literature: the ratio of claims on private sector to GDP, the ratio of broad money M3 minus M1 to GDP, and the share of commercial bank assets in the sum of commercial and central bank assets (see, e.g., Ang and McKibbin, 2007). Our results are compatible with similar studies in this area, in particular Honohan (2004a), Clarke et al. (2006) and Beck et al. (2007). In column (1), an increase in the ratio of private credit to GDP has a significant and favorable effect on income inequality (long-run elasticity of -0.066), indicating that financial deepening has an equalizing effect. In columns (2) and (3), we use two alternative indicators of financial development, but this does not alter our main findings.

Table 1: The effects of financial development (Dep. = Ln Gini coefficient)

I. Cointegration tests	(1)	(2)	(3)	(4)	(5)	(6)
ARDL bounds test (Pesaran et al., 2001)	4.817**	4.549**	4.768**	4.218**	3.626*	4.659**
ECM *t*-test (Banerjee et al., 1998)	-4.522**	-4.341**	-4.371**	-4.087**	-3.954*	-4.491**
II. ARDL estimate						
Intercept	3.659*** (0.000)	3.743*** (0.000)	3.729*** (0.000)	3.388*** (0.000)	3.709*** (0.000)	3.673*** (0.000)
Per capita real GDP growth	-0.331* (0.094)	-0.529* (0.082)	-0.535* (0.063)	-0.479 (0.105)	-0.771** (0.028)	-0.412** (0.044)
Ln Trade openness	0.142*** (0.000)	0.153*** (0.000)	0.143*** (0.000)	0.135*** (0.000)	0.061 (0.125)	0.149*** (0.000)
Inflation rate	-0.351*** (0.008)	-0.495*** (0.009)	-0.545*** (0.002)	-0.470** (0.011)	-0.675*** (0.001)	-0.375*** (0.007)
Ln (Private credit / GDP)	-0.066*** (0.000)					
Ln [(M3 – M1) / GDP]		-0.037** (0.021)				
Ln [Comm. Bank assets / (central bank assets + comm. Bank assets)]			-0.123*** (0.009)			
Ln Banking density				-0.037*** (0.000)		
Ln (Stock market capitalization / GDP)					0.019 (0.195)	
Ln (Modern sector / GDP)						-0.267*** (0.000)
III. Diagnostic checks						
LM test for serial correlation	1.076 (0.301)	2.821* (0.093)	1.843 (0.175)	1.772 (0.183)	3.043* (0.081)	1.034 (0.309)
Ramsey's RESET test	0.327 (0.568)	0.079 (0.778)	0.185 (0.667)	0.001 (0.974)	0.001 (0.991)	0.252 (0.615)
Heteroskedasticity test	0.991 (0.319)	0.345 (0.557)	0.686 (0.407)	0.386 (0.534)	0.994 (0.319)	0.411 (0.521)

Notes: The optimal lag structure for the resulting ARDL model was chosen using SBC. The test statistics of the bounds tests are compared against the critical values reported in Pesaran et al. (2001). The estimation allows for an unrestricted intercept and no trend. The 10%, 5% and 1% critical value bounds for the *F*-test are (2.45, 3.52), (2.86, 4.01) and (3.74, 5.06), respectively. The 10%, 5% and 1% critical value bounds for the *t*-test are (-2.57, -3.66), (-2.86, -3.99) and (-3.43, -4.60), respectively. Numbers in parentheses indicate *p*-values. *, ** and *** indicate 10%, 5% and 1% levels of significance, respectively.

Column (4) shows the effects of banking density on income inequality. Banking density is measured by the number of bank offices per population. The use of banking density as an indicator of financial development has a major advantage – it captures the breadth of financial systems, whereas other indicators reflect their depth. This is particularly important since the theories in finance and income inequality focus on the importance of broad access to finance (Demirgüç-Kunt and Levine, 2007). The results indicate that bank branch expansion is associated with lower income inequality, a finding in line with the results of Burgess and Pande (2005) for the Indian experience. Thus, the social banking program launched by the Indian government during the period 1969-1990 appears to have significantly improved the access of the poor to the formal financial sector.

Studies have suggested that stock market activity may also predict growth (e.g., Levine and Zervos, 1998; Arestis et al., 2001; Beck and Levine, 2004). Given that the financial development measures we have considered so far are primarily bank-based in nature, we also take into account a market-based measure. The results in column (5) show that stock market development has no statistically significant impact on income inequality. However, the results must be interpreted with caution since our measure of stock market development is a rather noisy indicator of financial development. The series was backdated using the share price index for the period before 1976 due to data unavailability. We are unable to relate the findings to the literature since the impact of stock market development on the Gini coefficient has not been studied in the literature.

We also include a proxy for modern sector development in the estimation due to Clarke et al. (2006). This allows us to examine how the sectoral structure of the economy affects income inequality. The influence of the modern sector is measured by the share of industrial and service sectors value added in total GDP. When the economy moves away from subsistence agriculture activities to advanced service oriented activities, this may be reflected by credit facilities becoming more readily and cheaply available. Thus, in general, countries with more developed financial systems tend to have relatively larger service and industrial sectors, and therefore this measure may provide an indirect indicator of financial development. The results reported in the last column show that development in the non-farm sector has an equalising effect, consistent with the findings of Datt and Ravallion (1998) for India.

Panel III reports some diagnostic statistics. We do not find any evidence of serial correlation, functional misspecification, and heteroskedasticity at the five percent level of significance. In sum, our results suggest that financial development helps in reducing the Gini coefficient through helping the poor to access finance by reducing financial market frictions. Thus, quite consistent with the vast literature showing a positive relationship between financial development and economic growth, our results reveal that financial development is effective in reducing income inequality in India.

5.2 Financial liberalization and income inequality

Evolution of financial market frictions in the financial systems, which can have a significant impact on access to finance and thus income inequality, may be driven by financial sector policies. Therefore, we want to examine the extent to which financial sector policies matter to income inequality. This addresses the concern raised by Demirgüç-Kunt and Levine (2007) that there is surprisingly little research that investigates whether financial sector policies influence the evolution of the distribution of income. We first look at how each type of financial sector policy affects income inequality. This includes examining the three main components of domestic financial sector reforms – directed credit programs, reserve and liquidity requirements, as well as interest rate restraint (see McKinnon, 1973; Shaw, 1973).

As explained earlier, the extent of directed credit controls is measured by the share of direct lending in total lending. Reserve and liquidity requirements are the sum of the cash reserve and statutory liquidity ratios. The index for interest rate restraint is constructed using the method of principal component using the six interest control variables discussed in Section 2. The results reported in Table 2 show that the estimated elasticity of the Gini coefficient with respect to a steady-state increase in the extent of directed credit programs is -0.249, and for reserve and liquidity requirements and interest rate restraint, the elasticities are -0.229 and -0.035, respectively (see columns (1) – (3)).

Taken together, the results in columns (1) to (3) appear to suggest that financial repressionist policies are pro-poor, and thus financial liberalization is likely to aggravate the income inequality problem in India. To confirm this, columns (4) – (6) examine how income inequality responds to the overall financial sector reforms in India using different measures of financial liberalization. The measure used in column (4) encompasses the three financial sector policies considered earlier, where they are summarized into just one single variable. We take the inverse so that this composite variable can be interpreted as financial liberalization.

Table 2: The effects of financial sector policies (Dep. = Ln Gini coefficient)

I. Cointegration tests	(1)	(2)	(3)	(4)	(5)	(6)
ARDL bounds test (Pesaran et al., 2001)	4.024**	3.908*	3.857*	4.661**	4.660**	3.715*
ECM *t*-test (Banerjee et al., 1998)	-4.033**	-4.193**	-3.746*	-4.467**	-4.467**	-4.011**
II. ARDL estimate						
Intercept	3.797*** (0.000)	3.803*** (0.000)	3.709*** (0.000)	3.759*** (0.000)	3.751*** (0.000)	3.621*** (0.000)
Per capita real GDP growth	-0.527* (0.081)	-0.512* (0.098)	-0.276 (0.171)	-0.294 (0.148)	-0.305 (0.134)	-0.574* (0.072)
Ln Trade openness	0.123*** (0.000)	0.106*** (0.000)	0.083*** (0.000)	0.109*** (0.000)	0.105*** (0.000)	0.081*** (0.001)
Inflation rate	-0.502*** (0.008)	-0.465** (0.020)	-0.353*** (0.009)	-0.338** (0.015)	-0.348** (0.012)	-0.556*** (0.004)
Directed credit programs	-0.249*** (0.000)					
Reserve and liquidity requirements		-0.229** (0.028)				
Interest rate restraint			-0.035*** (0.000)			
Ln Financial liberalization index (Demetriades-Luintel)				0.074*** (0.000)		
Ln Financial liberalization index (Extended Demetriades-Luintel)					0.077*** (0.000)	
Ln Financial liberalization index (Abiad-Mody)						0.019** (0.041)
III. Diagnostic checks						
LM test for serial correlation	2.125 (0.145)	1.077 (0.299)	1.578 (0.209)	1.538 (0.215)	2.248 (0.134)	2.092 (0.148)
Ramsey's RESET test	0.041 (0.841)	0.256 (0.613)	0.264 (0.607)	0.051 (0.821)	0.289 (0.591)	0.036 (0.851)
Heteroskedasticity test	0.756 (0.384)	0.506 (0.477)	1.875 (0.171)	1.848 (0.174)	2.871* (0.091)	1.371 (0.242)

Notes: see previous table.

The results reported in column (4) show that the measure of financial liberalization is significantly and positively associated with income inequality, with a long-run elasticity of 0.074. To provide some sensitivity checks, we also consider two other broader indicators of financial liberalization (see details in Section 2). Financial liberalization continues to enter positively and significantly even when we use different measures in columns (5) and (6). The estimates are found to be 0.077 and 0.019, respectively. Thus, the results unanimously show that financial liberalization appears to have a harmful effect on the distribution of income in India, confirming the earlier findings in columns (1) to (3). In all cases, we continue to find evidence of cointegration (panel I). There is also little evidence of econometric problems (panel III).

Similar to the effect of financial development, changes in the distribution of income may also affect the political economy in shaping financial sector policies so that variations in the Gini coefficient may influence the composite indices of financial liberalization and bias our results. We address this concern by also treating the measures of financial liberalization as the dependent variables but no evidence of cointegration is found (results not reported). This provides some support that financial liberalization can be entered "exogenously" in our specification.

Although in principle financial sector reforms may reduce market frictions and thereby alleviate income inequality, our results seem to suggest that liberalization of the domestic and international financial system has led to an increase in income inequality in India, providing some support to the arguments of Claessens and Perotti (2007). How could these results be interpreted within the specific context of India?

The results reported in Tables 1 and 2 show that financial development and financial liberalization have different effects on income inequality. Although increased financial liberalization is generally associated with deeper financial systems, case study evidence by Demetriades and Luintel (2001) and Ang (2008) have shown that financial sector reforms could retard development of financial systems. This is likely to be the case when banks act like a cartel and behave like a monopolist, a situation that reflects the current banking system in India, which is dominated by several big banks.

In India, directed credit has been extended to the agricultural sector and small and medium enterprises over the last few decades. These programs have significantly benefited farmers and small traders, allowing the poor direct access to financial services. Therefore, reducing the extent of these programs as part of the financial sector reforms is likely to hurt the poor. Similarly, the deregulation of interest rates may increase the costs of borrowing to the poor since this involves higher transaction costs relative to the size of the loan. The resulting higher borrowing costs, along with the reduction of direct lending, can have undesirable effect on income inequality since these policies deter the poor from adequately accessing finance.[9]

Financial liberalization in India did not necessarily lead to a relaxation of credit constraints to poor individuals that result in lower inequality. Before the liberalization, banks were required to open a certain number of branches in rural areas, and this policy was an important factor behind the savings rate increases of the 1970s and 1980s in India. However, this requirement was relaxed in 1991 following the launch of financial reforms. Thus, foreign and private banks would necessarily have a bias in providing consumer credit to richer areas, and access to finance by the poor would fall as banks withdrew branches from rural areas.

As Aghion et al. (2004) have shown, unrestricted financial liberalization may induce instability. While financial repression may not be desirable, the evidence presented in this paper does provide some support to the argument that some form of financial restraint may help in alleviating income inequality in developing countries. However, as noted by Demetriades and Luintel (2001) and Honohan and Stiglitz (2001), financial restraints are more likely to work well in environments with strong regulatory capacity, pinpointing the importance of strengthening the institutional framework.

For instance, Beck et al. (2008) find that bank branch deregulation reduces the Gini coefficient of income inequality in the U.S., a result that contradicts our findings for India. This highlights the fact that the effect of financial deregulation on income inequality may depend on the quality of institutions. As Rajan and Zingales (2003b) propose, the process of financial liberalization is likely to be harmful for countries with a weak institutional environment. Although the legal system in India was originally based on the British model that emphasizes protection of property rights, India ended with a much less effective institutional framework since the legal system was modified in a way that benefited the small number of Europeans that settled in and ran the economy (Mishkin, 2006).

In sum, financial sector reforms may lead to well-connected elites capturing most of the gains from new opportunities (Rajan and Zingales, 2003a, b; Claessens and Perotti, 2007). Therefore, the presence of these established interests may deepen rather than broaden access

[9] Although many farmers in developing countries obtain credit through microfinance, this issue is not formally addressed in this paper due to the lack of reliable time series data. Moreover, the microfinance finance penetration rate, defined as borrowing clients as a percentage of population) stands at only 1.1% for India. This ratio is relatively lower than many other developing countries such as Bangladesh, Indonesia and Thailand where microfinance claims 13.1%, 6.7% and 6.5% of the population as clients, respectively (see Honohan, 2004b).

to finance, resulting in higher income inequality. For example, Das and Mohapatra (2003) have shown that stock market liberalization in emerging markets has mainly benefited high income individuals at the expense of others. Furthermore, the presence of weak institutional environments in many liberalizing markets has allowed insiders to expropriate the interests of minority shareholders (Claessens et al., 2002; Claessens, 2006). Unlike central and eastern Europe where the extent of captured reform and rent seeking is limited (Roland, 2002), in India financial liberalization has primarily benefited the well-connected rich, leaving the poor to fall further behind due to unequal access to finance.

5.3 The effects of financial growth, non-linearity, efficiency and volatility on income inequality

Next, we analyze the results further by examining the growth and non-linear effects of financial development and financial liberalization. Since our earlier results reveal that the relationship between income inequality and finance are not sensitive to the measures of financial development and financial liberalization, we use only the ratio of private credit to GDP and the approach of Demetriades and Luintel (1996, 1997), respectively, in performing these additional tests for brevity. The results in the first two columns in Table 3 show that while the growth rate of financial development is found to be statistically significant at the conventional levels, the growth rate of financial liberalization is not.

In addition, we do not find any evidence of a non-linear effect of finance, a finding consistent with Beck et al. (2007). When the linear and squared terms for financial development are added to the base model, these two terms become jointly insignificant, suggesting that a threshold effect of financial development is not present in the relationship (column (3)). Similarly, we do not find any significant non-linear effect of financial liberalization. The coefficient of squared financial liberalization is insignificant although its linear term is weakly significant at the 10 percent level (column (4)).

In order to shed additional light, we test the effects of banking efficiency and volatility in the share market on the degree of economic opportunity. Demirgüç-Kunt and Levine (2007) propose that financial innovation can affect the level and evolution of financial market frictions, and therefore has an impact on income inequality. On the other hand, although in principle macroeconomic volatility may affect economic growth (see Ramey and Ramey, 1995), empirical research has not yet established whether financial volatility has an impact on income distribution.

Table 3: The effects of financial growth, nonlinearity, efficiency and volatility

I. Cointegration tests	(1)	(2)	(3)	(4)	(5)	(6)
ARDL bounds test (Pesaran et al., 2001)	3.432	3.762*	3.975**	3.459*	4.698**	3.969*
ECM *t*-test (Banerjee et al., 1998)	-3.869*	-3.903*	-4.438**	-4.154*	-4.272**	-4.112**
II. ARDL estimate						
Intercept	3.815*** (0.000)	3.701*** (0.000)	3.781*** (0.000)	3.753*** (0.000)	3.765*** (0.000)	3.640*** (0.000)
Per capita real GDP growth	-1.281*** (0.007)	-0.651* (0.089)	-0.344* (0.096)	-0.291 (0.170)	-0.601* (0.052)	-0.793** (0.041)
Ln Trade openness	0.116*** (0.000)	0.089*** (0.006)	0.121*** (0.000)	0.109*** (0.000)	0.112*** (0.000)	0.084** (0.018)
Inflation rate	-1.051*** (0.002)	-0.562** (0.023)	-0.319** (0.037)	-0.266 (0.103)	-0.529*** (0.006)	-0.633** (0.016)
Growth rate of private credit / GDP	-0.485** (0.034)					
Growth rate of financial liberalization index		0.188 (0.329)				
Ln (Private credit / GDP)			0.118 (0.488)			
Squared Ln (Private credit / GDP)			0.047 (0.277)			
Ln Financial liberalization index				0.086* (0.096)		
Squared Ln (Financial liberalization index)				0.015 (0.753)		
Interest spread					-0.694*** (0.000)	
Ln Share price volatility						0.020 (0.337)
III. Diagnostic checks						
LM test for serial correlation	2.901* (0.089)	4.252** (0.039)	0.789 (0.374)	1.189 (0.275)	2.778* (0.096)	4.775** (0.029)
Ramsey's RESET test	0.002 (0.961)	0.239 (0.625)	0.016 (0.898)	0.040 (0.841)	0.549 (0.459)	0.067 (0.795)
Heteroskedasticity test	1.452 (0.228)	0.789 (0.374)	1.287 (0.257)	2.335 (0.127)	1.578 (0.209)	1.175 (0.278)

Notes: the dependent variable is Ln Gini coefficient. A different set of critical values apply to columns (3) and (4) since the estimation involves six variables. The relevant 10%, 5% and 1% critical value bounds for the *F*-test are (2.26, 3.35), (2.62, 3.79) and (3.41, 4.68), respectively. The 10%, 5% and 1% critical value bounds for the *t*-test are (-2.57, -3.86), (-2.86, -4.19) and (-3.43, -4.79), respectively.

The interest rate spread is measured by the difference between the average lending rates and average deposit rates.[10] We use the rolling standard deviation of growth rate of the ratio of share price index to GDP deflator with a five year rolling window as the measure of share price volatility. The results in column (5) show that efficiency in the banking system is effective in reducing income gaps. However, volatility in the stock market, measured by variations in relative share prices, has no statistically significant impact on income inequality (see column (6)).

6. Summary and Conclusions

Since theories provide different predictions about the impact of finance on income inequality, more empirical analysis is necessary to shed light on their relationship. In this paper, we examined the determinants of the Gini coefficient in an autoregressive distributed lag framework, paying particular attention to testing the effects of financial development and financial sector reforms. Employing the ECM cointegration test and the ARDL bounds technique, the empirical evidence showed a significant steady-state relationship between the Gini coefficient and its determinants. After documenting these basic cointegration results, the long-run estimates were derived using the underlying ARDL model.

Our evidence suggests that underdevelopment of financial systems hurts the poor more than the rich, resulting in higher income inequality. Both the level and growth effects of financial development are found to be significant. The results therefore highlight the importance of developing financial systems in order to alleviate income inequality. However, both domestic and international financial sector reforms do not seem to reduce unequal access to finance, but rather tend to aggravate income inequality in India. Our results are not sensitive to the measures of financial development and financial liberalization. In addition, increased banking density and banking efficiency are found to have a favorable effect on income inequality in India. Finally, there is no evidence to support the presence of a non-linear effect in the finance-inequality relationship, providing no support to the Greenwood and Jovanovic (1990) hypothesis.

[10] Honohan (2004a) propose that the interest rate spread is a better measure that reflects the efficiency of the financial system, compared to other rudimentary measures of financial development commonly used in the literature. This measure has also been used by Rousseau (1998) as a measure of technical progress in the banking sector.

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