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Finance and consumption volatility: Evidence from India

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A B S T R A C T

The main objective of this paper is to explore the determinants of private consumption growth volatility in India, focusing on the role of financial sector policies. Using data for India over the period 1950–2005, the results show that the implementation of financial repressionist policies is strongly associated with lower consumption volatility. The results remain robust after controlling for a wide range of macroeconomic shocks and variables. The presence of a threshold effect implies that the benefits of financial reforms in reducing consumption volatility can only be reaped when the financial system becomes sufficiently liberalized. The results also indicate that the presence of a more open financial system may serve to dampen fluctuations in private consumption.

1. Introduction and Motivation

A number of developing countries have undergone significant financial sector reforms over the last few decades, leading to a widely observed increase in global financial integration. Along with this development, many of these economies have experienced rapid growth, whilst also experiencing significant macroeconomic volatility due mainly to fluctuations in consumption. Macroeconomic volatility is a fundamental concern for developing countries since it retards output growth and affects future consumption. As Loayza et al. (2007) argue, macroeconomic volatility may create large welfare costs for developing countries since it entails deviation from a smooth consumption path. In a similar vein, Prasad et al. (2003) show that potential welfare gains from reducing consumption volatility are enormous for developing countries.

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The theoretical relationship between finance and volatility is well-documented in the literature; however, it is not necessarily unambiguous. For instance, Sutherland (1996) considers a two-country intertemporal general equilibrium model and shows that increasing financial market integration reduces consumption volatility through providing more opportunities for consumption smoothing. Mendoza (1994), however, shows that changes in the volatility of consumption in response to greater financial openness are rather small. When shocks are larger and more persistent, financial integration may exacerbate output fluctuations. Easterly et al. (2001) argue that while increased financial market integration may provide a mechanism to smooth shocks, it may also expose a country to greater volatility, induced by changes in capital flows that disrupt economic activity. More recently, Levchenko (2005) demonstrates that agents who have access to international markets may benefit under greater financial openness whereas those who do not will experience an increase in their consumption volatility and a fall in welfare.

There is also a lack of consensus regarding the theoretical role of financial liberalization in macroeconomic stabilization. Traditional models of market imperfections often predict that improvements in financial markets can effectively absorb the undesirable effects of macroeconomic shocks. Hence, more efficient financial systems allow individuals to smooth consumption in response to fluctuations in income (see, e.g., Bernanke and Gertler, 1989; Greenwald and Stiglitz, 1993). Allen and Carletti (2008) suggest that financial systems can act as shock absorbers by spreading risks efficiently in the absence of market failures that arise due to panics and financial fragility (see also Allen and Gale, 2004, 2007). However, using a dynamic general equilibrium model, Bacchetta and Caminal (2000) show that credit constraints do not always amplify macroeconomic volatility; they can also dampen output fluctuations, depending on the nature of the shock.

Moreover, Caprio et al. (2001) propose that the process of financial liberalization may increase the volatility of asset prices and interest rates, thereby feeding into consumption volatility. It will also tend to induce macroeconomic instability with an initial surge in aggregate credit when financial institutions compete for market share. In line with this, Aghion et al. (2004) show that unrestricted financial liberalization may induce economic instability in a dynamic open economy model. Blanchard and Simon (2001) argue that while a relaxation of credit constraints following liberalization may lead consumers to choose a smoother consumption path, an improved ability to borrow and lend may also result in greater volatility. Hence, on theoretical grounds it is not clear whether financial sector reforms and deepening would lead to lower consumption volatility.

The above discussions suggest that the impact of finance on the volatility of consumption is theoretically ambiguous, and therefore it is ultimately an empirical issue. Although several studies have recently investigated this topic, so far the literature has not been able to establish any systematic relationship between macroeconomic volatility and financial factors (Levchenko, 2005; Kose et al., 2006). For example, while Razin and Rose (1994), Easterly et al. (2001) and Buch et al. (2005) have failed to document any robust relationship between financial openness and macroeconomic volatility, the empirical analysis of O'Donnell (2001) indicates that an increase in the degree of financial integration results in higher output volatility in non-OECD countries, but a negative relationship is found among OECD countries. Kose et al. (2003) find that financial openness is associated with increased consumption volatility, but only up to a certain level. The benefits of financial integration in terms of consumption smoothing can only be realized after a certain threshold is achieved. In an important recent study, Bekaert et al. (2006) examine the effects of equity market liberalization and capital account openness on consumption growth volatility for 95 countries over the period 1980–2000. Their results indicate that financial liberalization is strongly associated with lower consumption growth volatility.

Empirical findings regarding the impact of financial development on volatility are also mixed. For instance, financial development is found to improve the efficacy of consumption smoothing in the empirical analyses of Herrera and Vincent (2008) and Ahmed and Suardi (2009). The important role of financial development in reducing volatility is, however, challenged by Acemoglu et al. (2003), who find that financial intermediation does not influence macroeconomic volatility once the effect of institutions is taken into account. In line with this, Beck et al. (2006) and Bekaert et al. (2006) find only very weak or no association between financial development and macroeconomic volatility. Easterly et al. (2001) find that while financial development helps reduce growth volatility, the relationship appears to be non-linear, implying that very high levels of financial development may serve to magnify shocks to the economy. Their results also show that financial development tends to increase the likelihood of an economic downturn, which induces economic instability.

Hence, it appears that empirical studies have not been able to establish an unambiguous relationship between macroeconomic volatility and financial openness or financial development. While the above studies have made significant contributions to the understanding of the effects of finance on macroeconomic volatility, so far there has been no case study evidence documented. Case studies are particularly useful in disentangling the complexity of the financial environment and economic history of each country. As Mendoza (1994) and Kose et al. (2006) argue, the lack of consensus regarding the relationship between finance and macroeconomic volatility is probably due to the structural differences between countries included in the cross-country analyses. Given that the nature of the shocks and policy regimes may also differ significantly across countries, failure to account for institutional differences may yield misleading conclusions regarding how finance is related to macroeconomic volatility (see Acemoglu et al., 2003).

The main objective of this paper is to complement the existing cross-country studies, and enrich the literature by providing further evidence on how financial sector reforms affect the evolution of consumption growth volatility, drawing on the experience of one of the largest and fastest growing developing economies in the world. We focus our analysis on India rather than OECD countries since volatility affects developing countries substantially more than developed countries. India's recent financial sector reforms also provide an ideal testing ground for further analysis of the relationship between finance and volatility. A key departure here from the existing literature is that we focus on the role of financial sector reforms, given that the depth of a financial system is directly shaped by financial sector policies, and that the impact of financial openness on consumption growth volatility has been well-studied. This study uses two different indicators of financial reforms. The first is based on the approach of Demetriades and Luintel (1997). This indicator provides a measure of the extent of financial repression in the domestic financial system. The second follows the approach advanced by Abiad and Mody (2005), and is a broader measure that considers both domestic and international features.

The paper proceeds as follows. In Section 2, we provide an overview of the financial sector reform experience of India. The empirical model and data are set out in Section 3. Section 4 presents the empirical results, which consistently show that financial liberalization magnifies consumption growth volatility in India. This finding is robust to a number of sensitivity checks. Our results, however, indicate that greater financial openness may help smooth private consumption in India. The last section provides concluding remarks.

2. Financial repression and liberalization: the Indian experience

There was little repression in India's financial system during the 1950s and 1960s. However, the government gradually imposed more controls by raising statutory liquidity and cash reserve requirements over the 1970s and 1980s. An estimated 22.4% of central government revenue was derived from financial repression during the period 1980–1985 (see Giovannini and De Melo, 1993). Furthermore, several interest rate controls were implemented in the late 1980s. During this period, capital markets were underdeveloped and highly fragmented (OECD, 2007). The economy was also subject to a set of industrial licensing requirements that restricted entry and expansion of both domestic and foreign firms (Rodrik and Subramanian, 2005; Aghion et al., 2008). Moreover, the government implemented a set of restrictive trade policies in order to protect domestic industries (Panagariya, 2004; Rodrik and Subramanian, 2005; Madsen et al., 2010).

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 1990–1991 (Sen and Vaidya, 1999; Rodrik and Subramanian, 2005). The objective was to provide a greater role for markets in price determination and resource allocation. It was also hoped that greater benefits of international risk sharing could be reaped through increased financial openness. This could help minimize the fluctuations in macroeconomic aggregates such as consumption and output. Consequently, interest rates were gradually liberalized, and the 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 (see Bekaert et al., 2005). Capital account restrictions were also partially relaxed. The regulatory framework was significantly strengthened in 1994 (OECD, 2007).

Considerable progress has also been made toward opening up the financial sector to encourage more participation by foreign investors. Up to 74% of foreign ownership was allowed and foreign banks were permitted to open a specified number of new branches every year (Panagariya, 2004). The easing of regulations on portfolio and direct investment, along with the deregulation of entry restrictions in 1993, has resulted in the establishment of a significant number of new private and foreign banks (OECD, 2007). Alongside reforms in the financial sector, industrial licensing requirements were fully withdrawn in nearly all industries and the cap on foreign equity investment in almost all industries was raised to 51% (Rodrik and Subramanian, 2005; Madsen et al., 2010). While most goods can be imported without a license or other restrictions following the 1991 reforms, consumer goods remained under licensing until 2001 (Panagariya, 2004). The 1990s' reforms were also accompanied by significant reductions in tariff rates and the lifting of the exchange controls that restricted imports (Rodrik and Subramanian, 2005). The exchange rate was unified in 1993–1994 and most restrictions on current account transactions were eliminated in 1994 (see Williamson and Mahar, 1998).

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. The bank nationalization program in 1969, which brought the total number of branches under government control to 84% (Banerjee et al., 2004), 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. The significant presence of public sector banks have been associated with high intermediation costs and lower intermediating activity in the Indian banking sector (Banerjee et al., 2004). In terms of 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 the free allocation of resources.

3. Empirical model and data

The model specification attempts to examine how financial repression affects consumption volatility in India. In particular, the following empirical framework is adopted for the present study:

$$VOC_t = \beta_0 + \beta_1 PRI_t + \beta_2 VOG_t + \beta_3 PCF_t + \beta_4 FR_t + \epsilon_t \tag{1}$$

Consumption volatility (VOC_t) is measured by the rolling standard deviation of growth rate of real private consumption per capita (denoted as VOC_t^{SD}). We use a window of five years, so that the standard deviation reported for year t is the estimated standard deviation over the period $t-4$ to t. Given that the first available observation is 1950, the first observation for the standard deviation of the growth rate is therefore 1955. Following Bekaert et al. (2006), we also consider an alternative measure of consumption volatility using the high-low range over a period of five years (denoted as VOC_t^{HL}).[1]

We include an income variable (PRI_t) to control for the level of economic development. Since the focus of our analysis is on volatility of private consumption, the relevant income measure is private income rather than GDP. We use per capita real claims on the private sector (PCF_t) to capture the availability of credit. This consideration is important since Aghion et al. (2004) and Mendoza et al. (2007) have shown that the benefit of financial liberalization in macroeconomic stabilization may

[1] We have also considered using a window of seven years for both measures of consumption volatility. However, the results do not vary significantly.

depend on the extent of credit constraints. Both variables are measured at constant prices using the private consumption implicit deflator. VOG_t refers to the standard deviations of the growth rate in real per capita GDP over 5-year overlapping periods.

FR_t is a measure of the extent of financial repression in the preceding period. We use lagged measure of financial repression so that we focus our analysis on how the established level of financial repression affects volatility subsequently. This helps mitigate the concerns of endogeneity bias. Moreover, economic agents may take some time to react to changes in financial sector policies, implying that the use of a beginning period variable is more appropriate. To measure the extent of financial repression, we employ two different summary measures developed by Demetriades and Luintel (1997) and Abiad and Mody (2005) independently, denoted as (FR_t^{DL}) and (FR_t^{AB}), respectively.

The approach of Demetriades and Luintel (1997) considers nine series for the 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, cash (statutory) reserve ratio, and statutory liquidity ratio. The first variable is set to zero when directed credit programs are not implemented, and to 1, 2, 3 when the programs cover up to 20%, 21–40% and over 40%, respectively, of total bank lending. The other two variables are direct measures, which can be expressed in percentages. Thus, except for directed credit programs in which a *de facto* measure is used in absence of *de jure* information, all series are *de jure* measures reflecting the strength of policies designed to repress the financial system in India. Using these nine policy variables, a summary measure of financial repression is developed using the method of principal component analysis.[2]

In constructing the second summary measure of financial repression, 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 rate restraints; 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. We include an additional dimension by also considering the effect of equity market reforms due to Bekaert et al. (2005). 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 seven components is used to obtain an overall measure of financial repression.[3] The second approach 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 Report on Currency and Finance of the Reserve Bank of India and National Accounts Statistics of the Central Statistical Organization in India. Following the standard practice, all variables are measured in natural logarithms. Fig. 1 shows that volatility in consumption has been subject to much variation over time. Both measures of consumption volatility exhibit very similar patterns of change. While consumption volatility increased sharply in the 1960s and the 1980s, both VOC_t^{SD} and VOC_t^{HL} saw a significant decline in the early 1990s, and a subsequent rebound in the years after. Both PRI_t and PCF_t increase steadily over the years, with an average growth rate of 2.6% and 6.4%, respectively. An examination of the changes in the pattern of output volatility over time reveals that VOG_t has been very volatile, but generally on a declining trend.

By normalizing the first observation to 100, both indicators of financial repression show that the trend toward financial repression has been reversed since the liberalization and opening up of the financial system in the early 1990s. These two measures of financial repression show increasing disparity since the 1970s given that the second measure captures more dimensions of financial sector reform. It therefore necessarily reflects a lower extent of financial repression compared to the first measure that focuses exclusively on credit and interest controls.

[2] A similar approach has also been used by Ang and McKibbin (2007), Ang (2009, 2010a,b, forthcoming) and Madsen et al. (2010) in constructing indices of financial liberalization for several countries that have undergone significant financial reforms, including India, Korea and Malaysia.

[3] We have also explored using the first principal component but the results do not vary significantly.

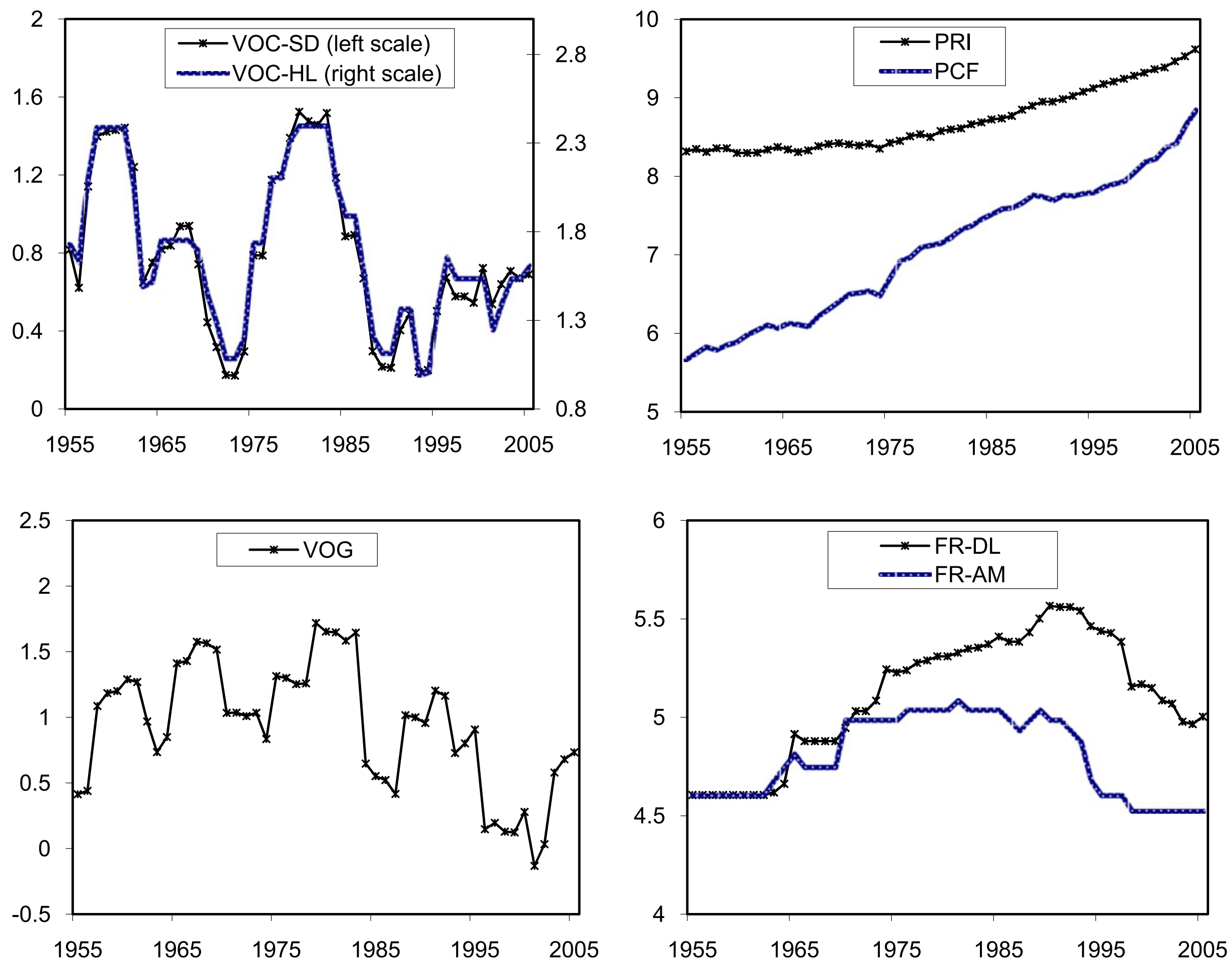

Fig. 1. Evolution of key variables used in the analysis (in natural logarithms). Notes: VOC-SD = ln(5-year standard deviation of growth rate of per capita real private consumption); VOC-HL = ln(high-low range of growth rate of per capita real private consumption over the 5-year period); PRI = ln(per capita real private income); PCF = ln(per capita real claims on private sector); VOG = ln(5-year standard deviation of growth rate of per capita real GDP); FR-DL = ln(financial repression index based on Demetriades and Luintel's approach); and FR-AB = ln(financial repression index based on Abiad and Mody's approach with modifications).

4. Empirical estimation and results

4.1. The VARs methodology

We now undertake a formal analysis of the relationship between finance and volatility and financial repression using the appropriate time series techniques. We begin the analysis by maintaining the assumption that the data generating process for the relationship between the underlying variables is a vector autoregressive (VAR) model at levels. The use of VARs methodology is appropriate in this case given that some of the underlying variables may be endogenous.

The testing procedure involves three steps. First, we perform an integration analysis for each variable using unit root tests. The second step is to test for cointegration using the Johansen techniques for the VARs constructed in levels. If cointegration is detected, the third step is to estimate the long-run relationship. Given that cointegrated variables must have an error-correction representation, the following vector error-correction model (VECM) is adopted:

$$\Delta x_t = \mu + \pi x_{t-1} + \lambda \sum_{j=1}^{p-1} \gamma_j \Delta x_{t-j} + \epsilon_t \tag{2}$$

where $x_t = [\text{VOC}_t, \text{PRI}_t, \text{VOG}_t, \text{PCF}_t, \text{FR}_t]'$ and $\epsilon_t \sim \text{IN}(0, \Omega)$. Ω is the variance-covariance matrix of the residuals. The rank of π is equal to the number of cointegrating vectors. The cointegration tests draw upon the procedure developed by Johansen (1988), which can be performed using the VECM formulated in Eq. (2). By normalizing VOC_t, the cointegrating vector can be interpreted as the long-run equation for the consumption volatility equation.

4.2. Integration and cointegration analysis

Integration properties of variables are conventionally examined using the standard Augmented Dickey–Fuller and Phillips–Perron tests. However, the presence of structural breaks in the series may bias the results toward non-rejection of the null hypothesis of a unit root when there is none. This consideration is of particular importance since the financial system in India has been subject to some drastic changes due to the repressionist or liberalizing policies pursued. Therefore, in principle, it is possible that the financial repression index may be a stationary variable with some breaks. To confirm this conjecture, we implement the standard unit root tests as well as tests with one and two endogenous breaks. For the latter, we perform the unit root procedure of Zivot and Andrews (1992), which tests the null of a unit root against the alternative of trend stationarity with an unknown break in the series. In addition, the endogenous one and two-break unit root procedures of Lee and Strazicich (2003) that allow for breaks under both the null and alternative hypothesis are also used. The results reported in Table 1 clearly show that all variables appear to be integrated at order one, or $I(1)$, at the 1% level of significance, regardless of whether structural breaks are allowed for. Given that all underlying variables share common integration properties, we can now proceed to testing for the presence of a long-run cointegrated relationship between the variables.

It is well-known that the Johansen approach may be sensitive to the choice of lag length. We therefore conduct a series of nested likelihood ratio tests on first-differenced VARs to determine the optimal lag length prior to performing cointegration tests. Given the sample size, we have considered a maximum lag length of five. The optimal lag length is found to be one in all models. Thus, we have followed this lag structure in the remaining analyses. Cointegration tests are then performed for the VARs at levels. In Table 2, both the results of Johansen trace and maximum eigenvalue tests unanimously point to the same conclusion, that there is only one cointegrating vector at the 5% level of significance.

However, it is possible that given the finite sample size used in this study, the Johansen test statistics may be biased (Cheung and Lai, 1993). Hence, we follow the approach of Reinsel and Ahn (1992), who suggest multiplying the Johansen statistics with the scale factor $(N - pk)/N$, where N is the number of observations, and p and k are the order of the VARs and the dimensions, respectively. This procedure corrects for small sample bias so that proper inference can be made. The results are by and large consistent with the standard Johansen cointegration tests.

4.3. The long-run estimates

Following the results of the cointegration tests, we proceed to deriving the long-run estimates. As we can see from Table 3, all equations perform relatively well on the basis of statistical significance and diagnostic checks. The consumption volatility equation is well determined with all variables showing plausible signs and magnitudes. In particular, an increase in the level of private income is associated with lower consumption volatility. Except for Model C, output volatility is found to be positively correlated with consumption volatility. While the financial system may have great potential to be an effective shock absorber, real private credit is found to have an amplifying effect on consumption volatility. This implies that the ease of credit availability may trigger significant fluctuations in consumption patterns, and therefore it is critical to monitor credit expansion cautiously.

Importantly, the coefficients of financial repression are very precisely estimated with a negative sign. The results are not sensitive to the use of different indicators of consumption volatility and summary measures of financial repression. Thus, consistent with the views of Blanchard and Simon (2001), Caprio et al. (2001) and Aghion et al. (2004), financial repression may have a mitigating effect on consumption volatility. While this finding may seem somewhat surprising for some readers, it is highly plausible in the context of India. This is because financial liberalization is more likely to play a positive role in reducing volatility only in countries with well-developed financial systems, relatively good quality of institutions (Kose et al., 2003; Prasad et al., 2003), and equal access to financial markets

Table 1
Unit root tests.

	Augmented Dickey–Fuller test		Phillips–Perron test		Zivot–Andrews test (one break)		Lee–Strazicich test (one break)		Lee–Strazicich test (two breaks)	
	Levels	1st-diff.	Levels	1st-diff.	Levels	1st-diff.	Levels	1st-diff.	Levels	1st-diff.
VOC_t^{SD}	−2.74	−5.54***	−2.41	−5.58***	−3.59 (1977)	−6.20*** (1981)	−2.77 (1984)	−5.96*** (1980)	−2.95 (1963, 1984)	−6.22*** (1972, 1980)
VOC_t^{HL}	−2.98	−5.08***	−2.40	−5.05***	−3.89 (1977)	−5.64*** (1981)	−3.03 (1984)	−5.63*** (1983)	−3.24 (1976, 1983)	−5.91*** (1983, 1995)
PRI_t	−1.46	−5.71***	0.73	−5.82***	−1.72 (1998)	−5.95*** (1960)	−0.75 (1973)	−8.14*** (1965)	−0.88 (1957, 1973)	−8.22*** (1965, 1971)
VOG_t	−2.39	−6.97***	−2.55	−6.97***	−4.19 (1984)	−7.15*** (1998)	−2.17 (1983)	−7.12*** (1999)	−2.64 (1983, 1995)	−7.25*** (1962, 1999)
PCF_t	−1.60	−5.60***	−1.47	−5.61***	−1.92 (1998)	−6.28*** (1979)	−2.06 (2000)	−5.13*** (1978)	−2.25 (1958, 1999)	−5.56*** (1978, 1986)
FR_t^{DL}	−1.60	−5.71***	−1.59	−5.78***	−2.35 (1997)	−7.21*** (1964)	−1.35 (1965)	−6.03*** (1990)	−1.64 (1965, 1997)	−6.42*** (1977, 1990)
FR_t^{AB}	−0.99	−5.55***	−0.95	−5.55***	−3.21 (1998)	−6.25*** (1982)	−1.33 (1969)	−6.08*** (1985)	−1.49 (1969, 1997)	−6.67*** (1962, 1998)

AIC is used to select the lag length and the maximum number of lags is set to five for the Augmented Dickey–Fuller test. Barlett–Kernel is used as the spectral estimation method for the Phillips–Perron tests. The bandwidth is selected using the Newey–West method. Results for tests with structural breaks are based on the "crash" model, which allows for an exogenous shift in the mean of the series. The 10%, 5% and 1% critical values for the Zivot–Andrews test are −4.58, −4.80 and −5.34, respectively whereas those for the Lee–Strazicich tests are −3.22, −3.57 and −4.24 for one break, respectively, and −3.50, −3.84 and −4.55, respectively, for two breaks. The break dates are indicated in the parentheses.
*** indicates 1% level of significance.

Table 2
Johansen cointegration tests.

	Trace statistic (λ_{trace})				
	$r=0$	$r \le 1$	$r \le 2$	$r \le 3$	$r \le 4$
Model A: $(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	85.88** [77.46]	43.35 [40.91]	24.10 [21.74]	9.92 [8.95]	1.42 [1.28]
Model B: $(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	82.86** [74.74]	41.26 [37.22]	19.38 [17.48]	7.44 [6.71]	1.29 [1.17]
Model C: $(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	85.97** [77.54]	47.00 [42.39]	25.24 [22.77]	10.72 [9.67]	1.93 [1.74]
Model D: $(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	82.27** [74.21]	43.06 [38.84]	21.42 [19.32]	8.94 [8.06]	1.25 [1.13]
5% critical values	71.44	49.64	31.88	18.11	8.19
	Maximum eigenvalue statistic (λ_{max})				
	$r=0$	$r=1$	$r=2$	$r=3$	$r=4$
Model A: $(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	40.52** [36.55]	21.25 [19.16]	14.19 [12.79]	8.50 [7.66]	1.42 [1.28]
Model B: $(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	41.60** [37.52]	21.88 [19.74]	11.94 [10.77]	6.14 [5.54]	1.29 [1.17]
Model C: $(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	38.97** [35.15]	21.75 [19.62]	14.52 [13.10]	8.80 [7.93]	1.93 [1.74]
Model D: $(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	39.21** [35.37]	21.64 [19.52]	12.48 [11.25]	7.69 [6.93]	1.25 [1.13]
5% critical values	34.03	27.80	21.49	15.02	8.19

r is the number of cointegrated vector; the optimal lag length is chosen to be one for all models based on likelihood ratio tests; critical values for the tests follow MacKinnon et al. (1999); figures in the brackets indicate the modified Johansen statistics.
** indicates 5% level of significance.

(Aghion et al., 1999; Levchenko, 2005). For most developing countries that do not meet these criteria, such as India, the opposite effect is likely to prevail. We will argue below that the lack of these preconditions may have contributed to our finding of an adverse role for financial liberalization in stabilizing consumption.[4]

First, the theoretical model of Aghion et al. (2004) shows that financial liberalization may destabilize the economy for countries at an intermediate level of financial development. India's ratio of private credit to GDP was 24% in 1991, and was ranked 92nd among 152 countries for which data are available. To the extent that the level of financial development in India can be characterized as intermediate at the point in time when the main phase of liberalization began, financial liberalization is likely to induce instability.

Second, Rajan and Zingales (2003) propose that the process of financial liberalization is likely to be harmful for countries with a weak institutional environment.[5] The legal system in India was originally based on the British model that emphasizes protection of property rights. However, India ended up 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).

Third, in the model developed by Aghion et al. (1999), financial market imperfections and unequal access to finance can produce permanent macroeconomic volatility. Improving access to finance is an important step for macroeconomic stabilization. Prior to the financial deregulation, directed credit programs, repressionist interest rate policies and bank branch regulation have significantly benefited farmers and small traders by improving their access to formal finance (see Ang, 2010a for more details). The resulting reduction in direct lending, higher borrowing costs, and withdrawal of bank branches in rural areas following the liberalizing in the early 1990s have significantly deprived the poor of opportunities to smooth consumption in times of income volatility.

[4] The literature also suggests that a number of financial crises have occurred following financial liberalization programs, resulting in significant volatility in output and consumption (see, e.g., Kaminsky and Reinhart, 1999). Such an explanation is not applicable to India since its economy has not been through a boom-bust credit cycle following the reforms.

[5] Bekaert et al. (2006) find that for countries with poor investor protection, financial liberalization tends to increase consumption volatility considerably.

Table 3
Cointegrating vectors.

	Dep. $= \text{VOC}_t^{SD}$		Dep. $= \text{VOC}_t^{HL}$	
	Model A	Model B	Model C	Model D
Intercept	46.791	111.037	48.963	118.195
PRI_t	−5.713*** (−6.907)	−12.343*** (−7.452)	−5.927*** (−6.759)	−13.129*** (−6.642)
VOG_t	0.436** (2.104)	1.135*** (4.723)	0.338 (1.575)	1.172*** (4.127)
PCF_t	3.144*** (7.133)	5.724*** (8.169)	3.165*** (6.777)	6.130*** (7.342)
FR_t^{DL}	−3.645*** (−7.432)		−3.590*** (−6.859)	
FR_t^{AB}		−9.229*** (−8.364)		−9.710*** (−7.414)
χ^2_{NORMAL}	2.052 [0.841]	7.381 [0.194]	2.899 [0.715]	7.672 [0.175]
χ^2_{SERIAL}	30.344 [0.212]	20.826 [0.702]	33.953 [0.109]	21.701 [0.653]
χ^2_{WHITE}	157.154 [0.978]	165.030 [0.942]	163.915 [0.951]	165.411 [0.939]

The normalized variable is VOC_t; figures in round brackets (.) are *t*-statistics.
χ^2_{NORMAL} refers to the Jarque-Bera statistic of the test for normal residuals.
χ^2_{SERIAL} is the Lagrange multiplier test statistics for no first order serial correlation.
χ^2_{WHITE} denotes the White's test statistic to test for homoskedastic errors; figures in square brackets [.] are *p*-values.
** and ***indicate 5% and 1% levels of significance, respectively.

4.4. *Financial repression, financial openness and consumption volatility*

Although our main focus is on the impact of domestic financial liberalization or repression, we also examine the effect of financial openness in view of the fact that recent studies have paid considerable attention to its role in consumption smoothing. To examine the role of this financial aspect, we incorporate a measure of financial openness (FO_t) into the base specification. We follow the standard practice in the literature by using the ratio of gross capital flows (sum of capital inflows and outflows) to GDP (see, e.g., Kose et al., 2003). The estimation involves a shorter sample period since data on capital flows are only available from 1970.

The results reported in Table 4 show that financial openness is significantly associated with lower consumption growth volatility when financial repression is measured using the approach of Demetriades and Luintel (1997), indicating that the presence of a more open financial system may be associated with an increased ability to smooth consumption shocks.[6] Our results are broadly consistent with Bekaert et al. (2006). Importantly, our main finding that financial repression reduces consumption growth volatility remains robust to the inclusion of financial openness in the specification. We continue to find very strong evidence of one cointegrated relationship (results not reported) and no econometric specification issues.

4.5. *Robustness*

4.5.1. *Alternative estimators*

Since the finite sample properties of VECM are unknown (Bewley et al., 1994), we propose two single equation approaches to obtain the long-run estimates: the fully-modified unrestricted error-correction model (FM-UECM) and the dynamic ordinary least squares (DOLS) estimator.[7] The FM-UECM estimator of Inder (1993) involves estimating the long-run parameters by incorporating adequate dynamics into the specification to avoid omitted lagged variable bias, as given in Eq. (3).

[6] Since the approach of Abiad and Mody (2005) includes a *de jure* component of restrictions on international capital flows, we have tried to exclude this dimension in the estimations for Models B and D. However, this does not improve the results.

[7] Although our unit root tests reveal that the financial repression measures are $I(1)$ variables even after allowing for structural breaks, we are not absolutely certain that there is no mean reversion component induced by the presence of boundary conditions, given that these measures are bounded by construction. The FM-UECM and DOLS estimators are particularly suitable in this context since they do not require all underlying variables to be integrated at the same order.

Table 4
Financial repression, financial openness and consumption volatility.

	Dep. = VOC_t^{SD}		Dep. = VOC_t^{HL}	
	Model A	Model B	Model A	Model B
Intercept	12.092	97.722	3.341	106.379
PRI_t	−2.684*** (−4.607)	−10.371*** (−11.969)	−1.736*** (−3.122)	−11.158*** (−9.788)
VOG_t	0.415*** (4.950)	1.104*** (11.367)	0.388*** (4.896)	1.126*** (8.939)
PCF_t	2.221*** (9.881)	4.605*** (15.960)	1.879*** (8.799)	4.840*** (12.736)
FR_t^{DL}	−1.123*** (−5.417)		−0.475** (−2.389)	
FR_t^{AB}		−8.438*** (−14.949)		−8.935*** (−12.133)
FO_t	−1.006*** (−3.626)	−0.217 (−0.678)	−1.399*** (−5.069)	−0.101 (−0.237)
χ^2_{NORMAL}	2.341 [0.886]	4.085 [0.665]	3.138 [0.791]	4.045 [0.671]
χ^2_{SERIAL}	34.434 [0.543]	18.760 [0.992]	31.624 [0.677]	16.838 [0.997]
χ^2_{WHITE}	298.655 [0.413]	290.792 [0.542]	306.361 [0.298]	289.265 [0.567]

The normalized variable is VOC_t; figures in round brackets (.) are *t*-statistics.
χ^2_{NORMAL} refers to the Jarque-Bera statistic of the test for normal residuals.
χ^2_{SERIAL} is the Lagrange multiplier test statistics for no first order serial correlation.
χ^2_{WHITE} denotes the White's test statistic to test for homoskedastic errors; figures in square brackets [.] are *p*-values.
** and *** indicate 5% and 1% levels of significance, respectively.

$$VOC_t = \alpha_0 + \sum_{j=1}^{k} \beta_j DET_{j,t} + \sum_{i=0}^{p} \gamma_i \Delta VOC_{t-i} + \sum_{i=0}^{p} \sum_{j=1}^{k} \delta_j \Delta DET_{j,t-i} + \epsilon_t \qquad (3)$$

where DET_t is a vector of *k* determinants of VOC_t. However, this approach may not be asymptotically optimal given that it takes no account of the possible endogeneity of the income variable. In view of this, we follow Bewley (1979) by using the instrumental variable technique to correct the standard errors so that valid inference can be drawn. Specifically, lagged level variables are used as the instruments for the first-different current terms to correct for endogeneity bias.

Next, the short-run effects are removed by defining $VOC_t^* = VOC_t - \hat{\alpha}_0 - \sum_{j=1}^{k} \hat{\beta}_j DET_{j,t} - \sum_{i=0}^{p} \hat{\gamma}_i \Delta VOC_{t-i} - \sum_{i=0}^{p} \sum_{j=1}^{k} \hat{\delta}_{ji} \Delta DET_{j,t-i}$. The fully-modified estimator is then obtained by employing the Phillips-Hansen non-parametric corrections to the regression of VOC_t^* on a constant and $DET_{j,t}$. The resulting estimator thus adequately deals with omitted lag variables bias. Inder (1993) demonstrates that it is asymptotically optimal, even in the presence of endogenous explanatory variables.

The DOLS procedure of Stock and Watson (1993) is asymptotically equivalent to the maximum likelihood estimator of Johansen (1988), and it has been shown to perform well in finite samples. The estimation involves regressing one of the $I(1)$ variables on the remaining $I(1)$ variables, the $I(0)$ variables, leads and lags of the first difference of the $I(1)$ variables, and a constant, as shown in Eq. (4). By doing so, it corrects for potential endogeneity problems and small sample bias, and provides estimates of the cointegrating vector which are asymptotically efficient. The long-run model for VOC_t can be obtained from the reduced form solution by setting all short-run dynamic terms to zero.

$$VOC_t = \alpha_0 + \sum_{j=1}^{k} \beta_j DET_{j,t} + \sum_{i=-p}^{p} \sum_{j=1}^{k} \delta_{ji} \Delta DET_{j,t-i} + \epsilon_t \qquad (4)$$

The results reported in Table 5 indicate that although the magnitudes of the coefficients become smaller, the qualitative aspect of the results remains largely unaltered. This observation is not unusual since the VECM estimator tends to produce larger estimates. Consistent with our previous findings reported in Table 3, financial repression is found to have an important role to play in smoothing consumption volatility. The coefficients associated with the financial repression measures are found to

Table 5
Financial repression and consumption volatility: alternative estimates.

	Dep. $=$ VOC_t^{SD}				Dep. $=$ VOC_t^{HL}			
	Model A		Model B		Model C		Model D	
	FM-UECM	DOLS	FM-UECM	DOLS	FM-UECM	DOLS	FM-UECM	DOLS
Intercept	12.585*** (0.000)	35.493*** (0.000)	46.961*** (0.000)	67.033*** (0.000)	13.230*** (0.001)	36.731*** (0.000)	36.599*** (0.000)	66.461*** (0.000)
PRI_t	−1.437*** (0.003)	−4.342*** (0.000)	−5.014*** (0.000)	−7.261*** (0.000)	−1.435*** (0.005)	−4.443*** (0.000)	−3.843*** (0.000)	−7.139*** (0.000)
VOG_t	0.461*** (0.000)	0.301* (0.083)	0.556*** (0.000)	0.635*** (0.005)	0.446*** (0.001)	0.220 (0.228)	0.527*** (0.000)	0.547** (0.020)
PCF_t	0.833*** (0.001)	2.266*** (0.000)	2.099*** (0.000)	3.125*** (0.000)	0.791*** (0.003)	2.212*** (0.000)	1.607*** (0.000)	3.010*** (0.000)
FR_t^{DL}	−1.094*** (0.000)	−2.541*** (0.000)			−0.986*** (0.000)	−2.343*** (0.000)		
FR_t^{AB}			−3.726*** (0.000)	−5.340*** (0.000)			−2.776*** (0.000)	−5.072*** (0.000)

Figures in parentheses indicate p-values.
*, ** and *** indicate 10%, 5% and 1% levels of significance, respectively.

be statistically significant at the 1% level. Moreover, the results are not sensitive to the use of different estimators and measures of financial repression.

4.5.2. Controlling for shocks and macroeconomic variables

Having established the key determinants of consumption volatility, we now turn to presenting the results with additional control variables.[8] These further results are derived using both the FM-UECM approach and the DOLS procedure since they produce more plausible sizes of the coefficients. Moreover, since these estimators yield very similar estimates, for brevity only the results using the FM-UECM approach are reported. We place more emphasis on the results derived from this estimator since Caporale and Pittis (2004) have found that it possesses the most desirable small sample properties in a class of 28 estimators. The results presented in Tables 6 and 7 indicate that most of these additional controls are statistically significant with intuitive signs.

Specifically, we control for terms of trade (TOT_t), monetary (MON_t), fiscal (FIS_t) and asset price (AP_t) shocks. We construct the proxies for these macroeconomic shocks using five-year rolling standard deviations of the rate of change in terms of trade, GDP deflator, real public consumption and share price indices, respectively (see Table 6). We also attempt to control for other macroeconomic variables, including trade openness (TO_t), social security (SOC_t), demographic changes (AGE_t), other macroeconomic reforms (MAR_t) and non-linear effects of financial repression (FR_t^2) (see Table 7). We use the standard trade intensity measure, i.e., the sum of exports and imports over GDP, as the proxy for trade openness. The ratio of accumulated provident and pension funds to private income is used as the measure of expected social security benefits.[9] Demographic changes are captured by the ratio of the number of young (with ages 0–14) and old (with ages 65 and above) dependents to working-age population (with ages 15–64). Other macroeconomic reforms considered as reforms in the trade and industrial sectors. As we can see from Tables 6 and 7, the results remain fairly robust against the inclusion of various proxies for shocks and macroeconomic variables. On the whole, our core results regarding the impact of financial repression remain unaltered.

Volatility in consumption may come from shocks in goods markets due to sudden changes in the international terms of trade. Our results that indicate that the terms of trade induce consumption volatility are in line with the cross-country findings of Kose et al. (2003) and Beck et al. (2006). The proxy for monetary shocks is found to be significant at the 1% level across all equations, with long-run elasticities in the range of 0.288–0.309. The Indian economy has been affected by major increases in the general price level. Therefore, fluctuations in the general price level are likely to have an adverse impact on consumption volatility.

Gavin and Perotti (1997) argue that fiscal policy is often pro-cyclical, expanding in booms but contracting in recessions. Thus, they are more likely to amplify rather than dampen macroeconomic volatility. However, contrary to the above argument, we find that public consumption plays a smoothing role, although its effect is found to be significant only in Model C and Model D, and only at the 10% level. Asset prices shocks, proxied by the standard deviations of the rate of change in the share price index, are found to have an amplifying effect on volatility of private consumption. However, this effect is only found to be significant in Model A and Model C, where financial repression is limited to the domestic components.

Trade openness may act as a shock absorber but it may also increase volatility since tradable sectors tend to be more volatile than non-tradable sectors. Our results are consistent with the cross-country findings of Kose et al. (2003) and Bekaert et al. (2006), who have also found a significant positive link between trade openness and private consumption volatility. In terms of institutional settings, the provision of social security benefits is found to have a dampening effect on consumption volatility.

[8] The finding of only one cointegrated relationship based on the Johansen procedure remains robust to the inclusion of these additional variables, which are entered as exogenous variables individually in the VECM estimation. It is worth noting that in order to conserve the degrees of freedom and avoid the problems of muticollinearity, it is not possible to include all these control variables in a single specification.

[9] However, we must bear in mind the caveat that this measure may be inadequate to capture the expected benefits of the social security programs in India. The pension coverage in India is very poor where only about 13% of the work force are currently covered by the Employee Provident Fund (EPF) and the Employment Pension Scheme (EPS).

Table 6
Financial repression and consumption volatility: controlling for shocks.

	Dep. = VOC_t^{SD}		Dep. = VOC_t^{HL}	
	Model A	Model B	Model C	Model D
[1] Controlling for terms of trade shocks				
Intercept	3.136	37.316***	2.975	34.913***
PRI_t	0.062	−3.614***	0.171	−3.238***
VOG_t	0.651***	0.728***	0.665***	0.726***
PCF_t	0.172	1.459***	0.046	1.279***
FR_t^{DL}/FR_t^{AB}	−1.167***	−3.559***	−0.998***	−3.319***
TOT_t	0.576***	0.461***	0.627***	0.528***
[2] Controlling for monetary shocks				
Intercept	10.289***	43.281***	10.819***	41.901***
PRI_t	−1.132***	−4.608***	−1.113***	−4.389***
VOG_t	0.314***	0.412***	0.298***	0.391***
PCF_t	0.804***	2.023***	0.756***	1.922***
FR_t^{DL}/FR_t^{AB}	−1.152***	−3.613***	−1.043***	−3.385***
MON_t	0.288***	0.299***	0.293***	0.309***
[3] Controlling for fiscal shocks				
Intercept	14.255***	50.465***	15.002***	49.091***
PRI_t	−1.521***	−5.327***	−1.502***	−5.081***
VOG_t	0.429***	0.528***	0.421***	0.519***
PCF_t	0.838***	2.193***	0.782***	2.066***
FR_t^{DL}/FR_t^{AB}	−1.215***	−3.971***	−1.123***	−3.738***
FIS_t	−0.201	−0.143	−0.238*	−0.184*
[4] Controlling for asset prices shocks				
Intercept	16.806***	48.351***	18.327***	47.561***
PRI_t	−1.950***	−5.174***	−2.058***	−5.001***
VOG_t	0.498***	0.573***	0.489***	0.546***
PCF_t	0.986***	2.116***	0.976***	2.033***
FR_t^{DL}/FR_t^{AB}	−1.403***	−3.808***	−1.355***	−3.656***
AP_t	0.259**	0.094	0.315**	0.126

The estimates are derived based on the fully-modified unrestricted ECM estimator of Inder (1993).
*, ** and *** indicate 10%, 5% and 1% levels of significance, respectively.

The effect is found to be highly significant in all models. Our results corroborate the cross-country findings of Bekaert et al. (2006). While India has just recently initiated a pension reform program, there is ample scope for more reforms to improve the coverage of the social security programs.

It is likely that our measure of financial repression may also capture changes in other policy environments apart from the financial system. As highlighted by Bekaert et al. (2006), financial sector reforms may occur simultaneously with other reforms undertaken in the economy. This concern is particularly pertinent for India given that the financial sector reforms in the early 1990s were part of a larger package of economic reforms aimed at enhancing growth (see Section 2). To ensure the financial repression index does not pick up other macroeconomic reforms (MAR_t) so that the results are specific to financial sector reforms, we estimate the consumption volatility equation by controlling for their effects. We do this by creating a summary index for macroeconomic policy reforms that focus on trade sector and industrial reforms. This involves considering the first principal component of the ratio of foreign collaborations to total firms, the ratio of foreign firms to total domestic private firms, and one minus the tariff rate.[10] The resulting index shows a clear upward trend since the early 1990s, coinciding rather well with the actual changes in the overall macro policy environment in India. Its coefficients are

[10] Although the ratio of FDI to GDP may be an appropriate measure to capture changes in industrial policies, data for FDI are only available from 1970. Therefore, we consider the number of technical collaborations and the presence of foreign firms as the alternative measures.

Table 7
Controlling for other macroeconomic variables and the non-linear effects.

	Dep. = VOC_t^{SD}		Dep. = VOC_t^{HL}	
	Model A	Model B	Model C	Model D
[1] Controlling for trade openness				
Intercept	29.542***	46.226***	30.506***	45.333***
PRI_t	−3.157***	−4.906***	−3.188***	−4.735***
VOG_t	0.192**	0.283**	0.172*	0.251**
PCF_t	0.740***	1.438***	0.695***	1.317***
FR_t^{DL}/FR_t^{AB}	−0.708***	−2.197***	−0.591***	−1.919***
TO_t	1.648***	1.409***	1.678***	1.488***
[2] Controlling for social security				
Intercept	7.489**	40.186***	7.265**	38.152***
PRI_t	−2.188***	−5.365***	−2.288***	−5.230***
VOG_t	0.404***	0.507***	0.378***	0.482***
PCF_t	1.853***	2.888***	1.958***	2.909***
FR_t^{DL}/FR_t^{AB}	−0.808***	−3.342***	−0.649**	−3.075***
SOC_t	−1.325***	−1.083***	−1.531***	−1.237***
[3] Controlling for demographic changes				
Intercept	26.354***	52.909***	27.588***	52.357***
PRI_t	−3.576***	−6.298***	−3.664***	−6.194***
VOG_t	0.458***	0.571***	0.449***	0.558***
PCF_t	0.481**	1.596***	0.429*	1.497***
FR_t^{DL}/FR_t^{AB}	−0.498**	−2.738***	−0.374	−2.512***
AGE_t	−13.242***	−12.479***	−13.770***	−12.982***
[4] Controlling for other macroeconomic reforms				
Intercept	18.436***	51.547***	20.175***	51.920***
PRI_t	−1.890***	−5.487***	−1.981***	−5.433***
VOG_t	0.625***	0.595***	0.631***	0.596***
PCF_t	0.769***	2.122***	0.713**	2.020***
FR_t^{DL}/FR_t^{AB}	−2.014***	−4.089***	−2.054***	−4.014***
MAR_t	−0.562***	−0.197*	−0.662***	−0.273**
[5] Controlling for non-linear effects				
Intercept	48.754***	282.641***	50.963***	294.352***
PRI_t	−1.959***	−3.318***	−1.979***	−3.138***
VOG_t	0.453***	0.609***	0.440***	0.581***
PCF_t	1.159***	1.206***	1.132***	1.103***
FR_t^{DL}/FR_t^{AB}	−14.254**	−106.681***	−14.718**	−111.765***
$(FR_t^{DL})^2/(FR_t^{AB})^2$	1.272**	10.832***	1.327**	11.386***

The estimates are derived based on the fully-modified unrestricted ECM estimator of Inder (1993).
*, ** and *** indicate 10%, 5% and 1% levels of significance, respectively.

found to be negative and significant across all models, suggesting that reforms in the trade and industrial sectors are associated with lower consumption volatility.

In our empirical analysis, we have detected a significant effect of age dependency. The results suggest that the private sector tends to exhibit less fluctuation in consumption spending with the increase of dependent population relative to working population. The finding that age dependency is associated with lower consumption volatility seems rather intuitive given that dependents tend to have more stable consumption patterns compared to the working population.

Finally, we also include a quadratic term to test for evidence of non-linearity in the data. We find evidence in favor of such a non-linear effect which implies the presence of a threshold effect. That is, while financial repression and consumption volatility is found to have a negative first order relationship, once financial repression crosses a threshold, the link becomes positive. This implies that very high levels of financial repression may serve to magnify consumption volatility. On the other hand, the results also imply that the benefits of financial reforms in reducing consumption volatility can only be realized when India becomes sufficiently liberalized.

5. Conclusions

The present study is motivated by the significant increase in the degree of financial integration and macroeconomic volatility observed across the developing world, and the lack of any previous time series attempts to analyze the relationship between financial reforms and consumption volatility in developing countries. The study contributes to the existing body of literature by investigating the unique experience of India, where its recent financial sector reforms provide an excellent case for further analysis. Specifically, we test how financial repression affects private consumption volatility in India using annual time series data over the period 1950–2005. In this study, financial repression is measured by two summary measures, which consider various types of domestic and international financial sector policies adopted by the Indian government.

Using the Johansen cointegration techniques, the empirical evidence shows a significant long-run relationship between consumption volatility and its determinants. After documenting these basic cointegration results, the long-run estimates are derived using several different estimators. The results are insensitive to the choice of estimators and consistently suggest that financial repression has a significant dampening effect on consumption volatility.

Sensitivity of the results to the inclusion of various other effects, including financial openness, trade openness, age dependency, social security benefits, macroeconomic reforms and non-linearity are examined. Moreover, in order to assess whether the results are affected by how other shocks influence consumption volatility, the volatility of the terms of trade, inflation, government spending, and asset prices are also considered for robustness checks. In all cases, the results regarding how financial repression affects consumption volatility remain robust to the inclusion of these control variables.

While financial repression may not always be desirable, the evidence presented in this paper does provide some support to the argument that some form of financial regulation may help alleviate consumption volatility in developing countries. Our results are largely consistent with the Krugman-Stiglitz view that developing countries should not fully liberalize their financial systems since they do not function efficiently, but rather some forms of control that regulate the operation of the financial markets should be imposed.

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References

Abiad, A., Mody, A., 2005. Financial reform: what shakes it? What shapes it? American Economic Review 95, 66–88.

Acemoglu, D., Johnson, S., Robinson, J., Thaicharoen, Y., 2003. Institutional causes, macroeconomic symptoms: volatility, crises and growth. Journal of Monetary Economics 50, 49–123.

Aghion, P., Bacchetta, P., Banerjee, A.V., 2004. Financial development and the instability of open economies. Journal of Monetary Economics 51, 1077–1106.

Aghion, P., Banerjee, A., Piketty, T., 1999. Dualism and macroeconomic volatility. Quarterly Journal of Economics 114, 1359–1397.

Aghion, P., Burgess, R., Redding, S., Zilibotti, F., 2008. The unequal effects of liberalization: evidence from Dismantling the license Raj in India. American Economic Review 98, 1397–1412.

Ahmed, A.D., Suardi, S., 2009. Macroeconomic volatility, trade and financial liberalization in Africa. World Development 37, 1623–1636.

Allen, F., Carletti, E., 2008. Financial System: Shock Absorber or Amplifier? BIS Working Papers No. 257.

Allen, F., Gale, D., 2004. Financial Intermediaries and markets. Econometrica 72, 1023–1061.

Allen, F., Gale, D., 2007. Understanding Financial Crises. Oxford University Press, Oxford.

Ang, J.B., 2009. Private investment and financial sector policies in India and Malaysia. World Development 37, 1261–1273.

Ang, J.B., 2010a. Finance and inequality: the case of India. Southern Economic Journal 76, 738–761.

Ang, J.B., 2010b. Research, Technological change and financial liberalization in South Korea. Journal of Macroeconomics 32, 457–468.

Ang, J.B., Savings Mobilization, financial development and liberalization: the case of Malaysia. Review of Income and Wealth, forthcoming, doi:10.1111/j.1475-4991.2010.00411.x.

Ang, J.B., McKibbin, W.J., 2007. Financial liberalization, financial sector development and growth: evidence from Malaysia. Journal of Development Economics 84, 215–233.

Bacchetta, P., Caminal, R., 2000. Do capital market imperfections exacerbate output fluctuations? European Economic Review 44, 449–468.

Banerjee, A.V., Cole, S., Duflo, E., 2004. Banking reform in India. In: Bery, S., Bosworth, B., Panagariya, A. (Eds.), India Policy Forum, vol. 1. National Council of Applied Economic Research; Brookings Institution, Washington, DC, pp. 277–332.

Beck, T., Lundberg, M., Majnoni, G., 2006. Financial intermediary development and growth volatility: do intermediaries dampen or magnify shocks? Journal of International Money and Finance 25, 1146–1167.

Bekaert, G., Harvey, C.R., Lundblad, C., 2005. Does financial liberalization spur growth? Journal of Financial Economics 77, 3–55.

Bekaert, G., Harvey, C.R., Lundblad, C., 2006. Growth volatility and financial liberalization. Journal of International Money and Finance 25, 370–403.

Bernanke, B., Gertler, M., 1989. Agency costs, net worth, and business fluctuations. American Economic Review 79, 14–31.

Bewley, R., 1979. The direct estimation of the equilibrium response in a linear dynamic model. Economics Letters 3, 357–361.

Bewley, R., Orden, D., Yang, M., Fisher, L.A., 1994. Comparison of Box-Tiao and Johansen Canonical estimators of cointegrating vectors in VEC(1) models. Journal of Econometrics 64, 3–27.

Blanchard, O., Simon, J., 2001. The long and large decline in U.S. output volatility. Brookings Papers on Economic Activity 1, 135–164.

Buch, C.M., Döpke, J., Pierdzioch, C., 2005. Financial openness and business cycle volatility. Journal of International Money and Finance 24, 744–765.

Caporale, G.M., Pittis, N., 2004. Estimator choice and Fisher's Paradox: a Monte Carlo study. Econometric Reviews 23, 25–52.

Caprio, G., Hanson, J.A., Honohan, P., 2001. Introduction and overview: the case for liberalization and some drawbacks. In: Caprio, G., Honohan, P., Stiglitz, J.E. (Eds.), Financial Liberalization: How Far, How Fast? Cambridge University Press, Cambridge, pp. 3–30.

Cheung, Y.-W., Lai, K.S., 1993. Finite-Sample sizes of Johansen's likelihood ratio tests for cointegration. Oxford Bulletin of Economics and Statistics 55, 313–328.

Demetriades, P.O., Luintel, K.B., 1997. The direct costs of financial repression: evidence from India. Review of Economics and Statistics 79, 311–320.

Easterly, W., Islam, R., Stiglitz, J.E., 2001. Shaken and Stirred: explaining growth volatility. In: B.Pleskovic, N.Stern (Ed.), Annual World Bank Conference on Development Economics. World Bank, Washington, pp.

Gavin, M., Perotti, R., 1997. Fiscal policy in Latin America. In: Bernanke, B., Rotemberg, J. (Eds.), NBER Macroeconomics Annual. MIT Press, Cambridge, pp. 11–61.

Giovannini, A., De Melo, M., 1993. Government revenue from financial repression. American Economic Review 83, 953–963.

Greenwald, B.C., Stiglitz, J.E., 1993. Financial market imperfections and business cycles. Quarterly Journal of Economics 108, 77–114.

Herrera, S., Vincent, B., 2008. Public Expenditure and Consumption Volatility The World Bank Policy Research Working Paper Series No. 4633.

Inder, B., 1993. Estimating long-run relationships in economics: a comparison of different approaches. Journal of Econometrics 57, 53–68.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12, 231–254.

Kaminsky, G.L., Reinhart, C.M., 1999. The Twin crises: the causes of banking and balance-of-payments problems. American Economic Review 89, 473–500.

Kose, M.A., Prasad, E.S., Terrones, M.E., 2003. Financial integration and macroeconomic volatility. IMF Staff Papers 50, 119–142.

Kose, M.A., Prasad, E.S., Terrones, M.E., 2006. How do trade and financial integration affect the relationship between growth and volatility? Journal of International Economics 69, 176–202.

Lee, J., Strazicich, M.C., 2003. Minimum Lagrange multiplier unit root test with two structural breaks. Review of Economics and Statistics 85, 1082–1089.

Levchenko, A.A., 2005. Financial liberalization and consumption volatility in developing countries. IMF Staff Papers 52, 237–259.

Loayza, N.V., Rancière, R., Servén, L., Ventura, J., 2007. Macroeconomic volatility and welfare in developing countries: an introduction. World Bank Economic Review 21, 343–357.

MacKinnon, J.G., Haug, A.A., Michelis, L., 1999. Numerical Distribution Functions of Likelihood Ratio Tests For Cointegration. Journal of Applied Econometrics 14, 563–577.

Madsen, J.B., Saxena, S., Ang, J.B., 2010. The Indian growth Miracle and endogenous growth. Journal of Development Economics 93, 37–48.

Mendoza, E.G., 1994. Robustness of macroeconomic indicators of capital Mobility. In: Leiderman, L., Razin, A. (Eds.), Capital MOBILITY: The Impact on Consumption, Investment and Growth. Cambridge University Press, Cambridge, pp. 83–111.

Mendoza, E.G., Quadrini, V., Ríos-Rull, J.-V., 2007. On the Welfare Implications of Financial Globalization without Financial Development NBER Working Papers No. 13412.

Mishkin, F.S., 2006. The Next Great Globalization: How Disadvantaged Nations Can Harness Their Financial Systems to Get Rich. Princeton University Press, Princeton and Oxford.

ODonnell, B., 2001. Financial Openness and Economic Performance unpublished manuscript.

OECD, 2007. OECD Economic Surveys: India 2007, Issue 14, 1–246 (complete edition: ISBN 9264033513).

Panagariya, A., 2004. India in the 1980s and 1990s: a Triumph of Reforms IMF Working Papers No. 04/43.

Prasad, E., Rogoff, K., Wei, S.-J., Kose, A., 2003. Effects of Financial Globalization on Developing Countries: Some Empirical Evidence IMF Occasional Papers No.: 220.

Rajan, R.G., Zingales, L., 2003. Saving Capitalism from the Capitalists: Unleashing the Power of Financial Markets to Create Wealth and Spread Opportunity. Crown Business, New York.

Razin, A., Rose, A., 1994. Business Cycle Volatility and Openness: An Exploratory Cross-Section Analysis NBER Working Papers No.: 4208.

Reinsel, G.C., Ahn, S.K., 1992. Vector autoregressive models with unit roots and reduced rank structure: estimation, likelihood ratio test, and Forecasting. Journal of Time Series Analysis 13, 353–375.

Rodrik, D., Subramanian, A., 2005. From "Hindu growth" to productivity surge: the mystery of the Indian growth transition. IMF Staff Papers 52, 193–228.

Sen, K., Vaidya, R., 1999. The Process of Financial Liberalization in India. Oxford University Press, Oxford.

Stock, J.H., Watson, M.W., 1993. A Simple estimator of cointegrating vectors in higher order integrated systems. Econometrica 61, 783–820.
Sutherland, A., 1996. Financial market integration and macroeconomic volatility. Scandinavian Journal of Economics 98, 521–531.
Williamson, J., Mahar, M., 1998. A Survey of Financial Liberalization. Essays in International Finance No.: 211. Department of Economics, Princeton University.
Zivot, E., Andrews, D.W., 1992. Further evidence on the great crash, the Oil-price shock, and the unit-root hypothesis. Journal of Business and Economic Statistics 10, 251–270.