Growth Volatility and Financial Repression: Time Series Evidence from India

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2009

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Abstract

The main objective of this paper is to explore the determinants of private consumption volatility in India. While considerable effort has been expended on the examining the relationship between growth and volatility, we focus on financial repression and private consumption volatility in India. Using annual time series data, the results show that the implementation of financial repressionist policies are strongly associated with lower consumption volatility in India. The results remain robust after controlling for a wide range of macroeconomic shocks and variables. Additional analysis which involves examining each component of private consumption provides further evidence to support this finding. The presence of a threshold effect suggests that the benefits of financial openness in dampening consumption volatility can only be reaped when India becomes sufficiently liberalized.

Keywords: Growth volatility; financial repression; India.

JEL classification: E44; E58; O53

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1. Motivation

A number of developing countries have undergone significant financial sector reforms over the last two decades, leading to a widely observed increase in the degree of financial integration globally. Accompanying this development, many of these economies have experienced rapid growth, but they have also been subject to high macroeconomic volatility that arises mainly from fluctuations in consumption. Macroeconomic volatility is a fundamental development concern for developing countries since it retards output growth and affects future consumption. As Loayza *et al.* (2007) put forward, macroeconomic volatility may exert large welfare costs to developing countries since it represents 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.

Economic theory generally predicts that greater financial openness helps reduce macroeconomic volatility. Using a simple model of global financial portfolio diversification, Obstfeld (1994) demonstrates that a substantial increase in national welfare can be gained through increasing consumption growth when an economy opens up its financial sector. This model highlights the role of financial integration in sharing country risks and assumes that risky returns are not perfectly correlated across countries. Sutherland (1996) considers a two-country intertemporal general equilibrium model and shows that increasing financial openness reduces consumption volatility given that more integrated financial markets provide greater opportunities for consumption smoothing.

More recently, Levchenko (2005) develops a model which is ideal for developing countries. The model departs from the traditional representative agent model of Obstfeld (1994) and assumes the presence of domestic frictions with unequal access to international financial markets. Under this framework, the model predicts that when risks are idiosyncratic, financial opening allows some agents to insure against these risks in the international financial markets. This reduces the amount of domestic risk sharing and increases volatility of consumption. With the assumption of agent heterogeneity, 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.

The above discussions suggest that the impact of financial openness on volatility of consumption is theoretically ambiguous, and therefore it is ultimately an empirical issue. However, empirical evidence documenting the relationship between macroeconomic volatility and financial openness is rather scant. Moreover, the available studies so far have not been able to establish any systematic relationship between volatility and financial factors. Using data for a sample of 74 countries over the period 1960-97, Easterly *et al.* (2001) find that financial development helps reduce growth volatility but the relationship appears to be nonlinear, implying that very high level 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 instabilities. Similarly, 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.

In an important 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. After controlling for a number of variables typically employed in growth regressions, the results indicate that financial liberalization is strongly associated with lower consumption growth volatility. Although a negative link between financial liberalization and consumption volatility has been documented by Bekaert *et al.* (2006), earlier studies by Razin and Rose (1994), Kose *et al.* (2003) and Buch *et al.* (2005) have failed to find any robust relationship between financial openness and macroeconomic volatility. Hence, it appears that empirical studies have not been able to establish an unambiguous relationship between financial openness and macroeconomic volatility.

While the above studies have made significant contributions to the understanding of the effects of financial openness on volatility, so far there has been no case studies evidence documented. Case studies are particularly useful in disentangling the complexity of the financial environments and economic histories of each individual country. As Kose *et al.* (2006) argue, the lack of consensus with regards to the relationship between financial openness and macroeconomic volatility is probably due to the structural differences between countries included in the cross-country analyses. The objective of this paper is to complement the existing cross-country studies, and enrich the literature by providing further evidence on how financial sector policies affect the evolution of consumption 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 more substantially than developed countries. Moreover, India's recent financial sector reforms provide an ideal testing ground for further analysis on the relationship between financial openness and consumption volatility.

This study uses two different indicators of financial repression. The first indicator is developed 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 indicator follows the approach advanced by Abiad and Mody (2005), which is a broader measure that considers both the domestic and international aspects. We attempt to measure financial repression rather than financial openness since the available data on financial sector policies such as loans to priority sectors, statutory reserve ratio and liquidity requirement reflect the strength of policies designed to repress the financial system. Nevertheless, to compare the findings of this paper with other studies, the inverse of the summary measure for financial repression can be interpreted as financial openness (see Ang and McKibbin, 2007).

The paper proceeds as follows. In section 2, we provide an overview of the financial sector reforms experience of India. Empirical model and data are set out in Section 3. Section 4 presents the empirical results, which consistently show that financial repression helps smooth consumption in India. This finding is robust to a number of sensitivity checks. The last section provides concluding remarks.

2. Financial Repression and Openness: The Indian Experience

There was little financial repression in the financial system of India during the 1950s and 1960s. However, the government gradually imposed more controls on the financial system by raising statutory liquidity and cash reserve requirements over the 1970s and 1980s. Revenue from financial repression was estimated to be 22.4 percent of total central government revenue during the period 1980-85 (see Giovannini and De Melo, 1993). Furthermore, several interest rate controls were implemented in the late 1980s.

A series of comprehensive financial sector reform policies were undertaken in 1991 as part of the broader economic reform (Sen and Vaidya, 1999). It was aimed at changing the entire orientation of India's financial development strategy from its position of a financially repressed system to that of a more open, market-type system. It was also hoped that greater benefits of international risk sharing can be reaped through increased financial openness. This could help minimize the fluctuations in macroeconomic aggregates such as consumption and output.

Since then, interest rates were gradually liberalized and statutory liquidity requirements significantly reduced so that markets could play a greater role in price determination and resource allocation. The equity market was formally liberalized in 1992, although the first country fund was set up earlier in 1986, which allowed foreign investors to access the domestic equity market directly (see Bekaert *et al.*, 2005). There has also been a change in the capital account regime from a restricted one to a more open one. The regulatory framework was strengthened significantly in 1992. In addition, entry restrictions were deregulated in 1993, resulting in the establishment of more private and foreign banks. Regulations on portfolio and direct investment were eased since then. The exchange rate was unified in 1993-94 and most restrictions on current account transactions were eliminated in 1994 (see Williamson and Mahar, 1998).

However, despite the liberalization programs launched in the early 1990s, the Indian financial system has continued to operate within the context of repressionist policies. For example, significant directed credit programs in favour of certain priority sectors still prevail in the banking system. The bank nationalization program in 1969 has enabled the Reserve Bank of India to effectively implement its credit allocation policy. Although the government divested part of its equity position in some public banks in the 1990s, the banking sector has remained predominantly state-owned. 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.

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:

$$VOC_t = \beta_0 + \beta_1 PRI_t + \beta_2 VOG_t + \beta_3 PCF_t + \beta_4 FR_t + \varepsilon_t \qquad (1)$$

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. PCF_t is real per capita claims on private sector. Both variables are measured at constant prices using 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

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

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 consider 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, partial liberalized, partial 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 Organisation in India. Following the standard practice, all variables are measured in natural logarithms. Figure 1 shows that volatility in consumption has been subject to much variation over time. Both measures of consumption volatility exhibit very similar pattern 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 percent and 6.4 percent, respectively. An examination of the

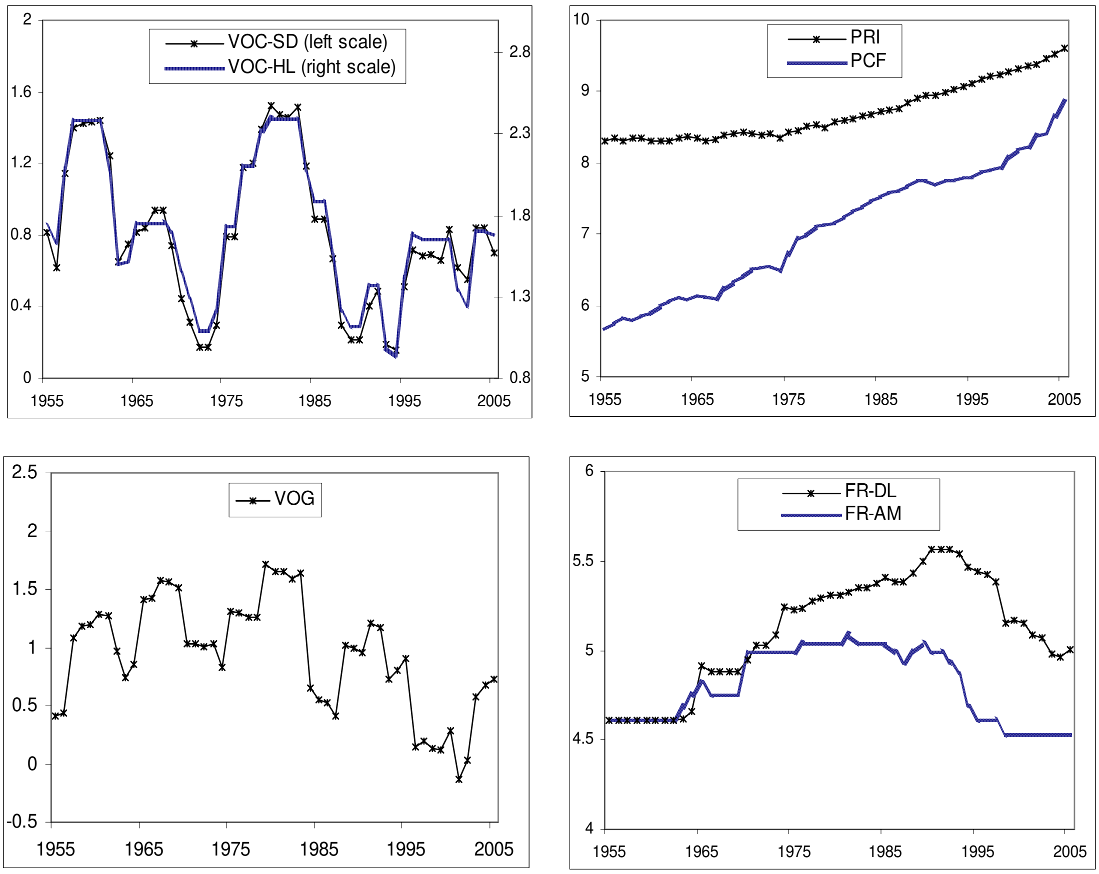
[2] A similar approach has also been used by Ang and McKibbin (2007) and Ang (2008a, b, 2009) in their studies of the relationship between financial development and economic growth in Malaysia.

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

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 be 100, both indicators of financial repression show that the trend towards financial repression has been reversed since 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 reforms. It therefore necessarily reflects a lower extent of financial repression compared to the first measure that focuses exclusively on credit and interest controls.

Figure 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

We now undertake a formal analysis of the relationship between private consumption 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 the conventional 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:

$$\Delta \boldsymbol{x}_t = \boldsymbol{\mu} + \pi \boldsymbol{x}_{t-1} + \lambda \sum_{j=1}^{p-1} \boldsymbol{\gamma}_j \Delta \boldsymbol{x}_{t-j} + \boldsymbol{\varepsilon}_t \tag{2}$$

where $x_t = [VOC_t, PRI_t, VOG_t, PCF_t, FR_t]'$ and $\boldsymbol{\varepsilon}_t \sim \mathrm{IN}(\mathbf{0}, \mathbf{\Omega})$. $\mathbf{\Omega}$ 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.

4.1 Integration and cointegration analysis

The integration properties of the underlying variables are examined using two standard unit root tests - the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. However, the presence of a structural break in a series may bias the results toward non-rejection of the null hypothesis of a unit root when there is none. We therefore also implement unit root tests with an endogenous break to examine whether the series appear to be stationary. For this purpose, we perform the unit root procedure of Zivot and Andrews (1992), which tests the null of a unit root against the alternative of trend stationarity with an unknown break in the series. The results reported in Table 1 clearly show that all variables appear to be integrated at order one, or $I(1)$, at the 5% level of significance. 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.

Table 1: Results for unit root tests

	ADF		PP		ZA	
	Levels	1st-diff.	Levels	1st-diff.	Levels	1st-diff.
VOC_t^{SD}	-2.467	-5.449***	-2.251	-5.545***	-3.661	-5.674***
VOC_t^{HL}	-3.065	-5.139***	-2.369	-5.071***	-3.876	-5.297**
PRI_t	-1.458	-5.713***	0.727	-5.821***	-1.721	-5.948***
VOG_t	-2.389	-6.968***	-2.546	-6.968***	-4.199	-7.145***
PCF_t	-1.603	-5.602***	-1.466	-5.611***	-1.921	-6.279***
FR_t^{DL}	-1.601	-5.712***	-1.592	-5.778***	-2.349	-7.206***
FR_t^{AB}	-0.991	-5.551***	-0.946	-5.551***	-3.214	-6.246***

*Notes: *** and ** indicate 1% and 5% level of significance, respectively. For the Augmented Dickey-Fuller (ADF) test, AIC is used to select the lag length and the maximum number of lags is set to be five. For the Phillips-Perron (PP) tests, Barlett-Kernel is used as the spectral estimation method. The bandwidth is selected using the Newey-West method. Only results for the "crash" model, which allows for an exogenous shift in the mean of the series, are reported for the Zivot-Andrews (ZA) tests. We have also considered the "changing growth" model that allows for a shift in the trend and the "change in level and slope" model that admits both changes. These models yield very similar results that do not alter the conclusions, and therefore they are not reported for brevity.*

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.

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:	85.21**	45.30	24.20	10.46	0.94
$(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	[76.86**]	[40.86]	[21.83]	[9.43]	[0.85]
Model B:	82.19**	40.48	18.81	7.25	1.02
$(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	[74.13**]	[36.51]	[16.97]	[6.54]	[0.92]
Model C:	84.29**	46.88	25.22	11.42	1.59
$(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	[76.03**]	[42.28]	[22.75]	[10.30]	[1.43]
Model D:	81.66**	42.37	21.01	9.21	1.03
$(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	[73.65**]	[38.22]	[18.95]	[8.31]	[0.93]
5% critical values	71.44	49.64	31.88	18.11	8.19
	Maximum eigenvalue statistic ($\lambda_{\max}$)				
	$r=0$	$r=1$	$r=2$	$r=3$	$r=4$
Model A:	39.91**	21.10	13.75	9.51	0.94
$(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	[36.00**]	[19.03]	[12.40]	[8.58]	[0.85]
Model B:	41.72**	21.67	11.56	6.23	1.02
$(VOC_t^{SD}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	[37.63**]	[19.55]	[10.43]	[5.62]	[0.92]
Model C:	37.41**	21.66	13.80	9.83	1.59
$(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{DL})$	[33.74*]	[19.54]	[12.45]	[8.87]	[1.43]
Model D:	39.29**	21.36	11.80	8.19	1.03
$(VOC_t^{HL}, PRI_t, VOG_t, PCF_t, FR_t^{AB})$	[35.44**]	[19.27]	[10.64]	[7.39]	[0.93]
5% critical values	34.03	27.80	21.49	15.02	8.19

*Notes: 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 brackets indicate the modified Johansen statistics. * and ** indicate 5% and 10% level of significance, respectively.*

However, it is possible that given the small sample size used in this study (51 annual observations), 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 observation, and p and k are the order of the VAR 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.

Table 3: Cointegrating vectors

	$Dep. = VOC_t^{SD}$		$Dep. = VOC_t^{HL}$	
	Model A	Model B	Model C	Model D
Intercept	46.998	111.129	49.003	117.825
PRI_t	-5.726*** (-6.742)	-12.339*** (-7.419)	-5.917*** (-6.412)	-13.077*** (-6.567)
VOG_t	0.445*** (2.078)	1.118*** (4.598)	0.362 (1.583)	1.161*** (4.015)
PCF_t	3.209*** (7.127)	5.751*** (8.180)	3.207*** (6.566)	6.143*** (7.309)
FR_t^{DL}	-3.803*** (-7.523)		-3.674*** (-6.747)	
FR_t^{AB}		-9.288*** (-8.391)		-9.740*** (-7.389)
χ^2_{NORMAL}	2.948 [0.708]	6.658 [0.247]	4.747 [0.447]	7.464 [0.188]
χ^2_{SERIAL}	30.435 [0.208]	20.106 [0.741]	34.813* [0.092]	20.488 [0.721]
χ^2_{WHITE}	162.851 [0.816]	171.089 [0.671]	168.811 [0.715]	174.598 [0.599]

*Notes: 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, respectively; χ^2_{WHITE} denotes the White's test statistic to test for homoskedastic errors; figures in square brackets [.] are p-values; * and *** indicate 10% and 1% level of significance, respectively.*

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 economic development 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, private credit flows have been found to have an amplifying effect on consumption volatility. This implies that the ease of credit availability may trigger significant fluctuations in consumption pattern, and therefore it is critical to monitor credit expansion carefully.

Importantly, financial repression is found to have a mitigating effect on consumption volatility. The results are not sensitive to the use of different indicators of consumption volatility and summary measures of financial repression. To this end, our results are consistent with the evidence of Kaminsky and Reinhart (1999), who have highlighted that a number of financial crises have occurred following financial liberalization programs. These crises are often associated with a loss of access to world credit markets and greater fluctuations in output and consumption. Prasad *et al.* (2003) argue that developing countries do not seem to benefit from financial openness through reducing consumption volatility due to the presence of weak institutional setting. In this regard, the presence of a sound institutional and regulatory framework is necessary for India so that any potential benefits of greater financial openness can be reaped.

4.2 Alternative estimators

Since the small 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 dynamic ordinary least squares (DOLS) estimator. 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).

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

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 endogeneous explanatory variables. Furthermore, using Monte Carlo experiments, Caporale and Pittis (2004) show that this estimator possesses the most desirable small sample properties in a class of 28 estimators.

The 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 be zero.

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

The regression specifications reported in Table 4 pass the diagnostic tests against non-normality and heteroskedasticity at the conventional levels. However, the estimated equations fail the serial correlation tests at the 1% level of significance. The presence of serial correlation in the residuals may be due to the use of 5-year overlapping periods to provide a measure for consumption volatility. Therefore, in all regression analyses, we deal with the moving average component in the residuals by adjusting the standard errors following the approach of Newey and West (1987) in order to obtain heteroskedastic and autocorrelation consistent estimates.

The results indicate that although the magnitudes of the coefficients are relatively smaller, the qualitative aspect of the results remains largely unaltered. This finding 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 be statistically significant at the 1% level. The results are not sensitive to the use of different estimators and measures of financial repression.

Table 4: Financial repression and consumption volatility: alternative estimators

	$Dep. = VOC_t^{SD}$				$Dep. = VOC_t^{HL}$			
	Model A		Model B		Model C		Model D	
	FM-UECM	DOLS	FM-UECM	DOLS	FM-UECM	DOLS	FM-UECM	DOLS
Intercept	12.240*** (0.001)	35.345*** (0.000)	47.217*** (0.000)	67.788*** (0.000)	12.867*** (0.001)	36.199*** (0.000)	38.392*** (0.000)	67.448*** (0.000)
PRI_t	-1.374*** (0.005)	-4.320*** (0.000)	-5.010*** (0.000)	-7.326*** (0.000)	-1.364*** (0.008)	-4.368*** (0.000)	-4.809*** (0.000)	-7.233*** (0.000)
VOG_t	0.446*** (0.000)	0.282 (0.114)	0.543*** (0.000)	0.622** (0.006)	0.445*** (0.001)	0.215 (0.255)	0.539*** (0.000)	0.532** (0.029)
PCF_t	0.830*** (0.001)	2.278*** (0.000)	2.117*** (0.000)	3.181*** (0.000)	0.796*** (0.003)	2.202*** (0.000)	2.027*** (0.000)	3.080*** (0.000)
FR_t^{DL}	-1.126*** (0.000)	-2.558*** (0.000)			-1.042*** (0.000)	-2.346*** (0.000)		
FR_t^{AB}			-3.807*** (0.000)	-5.452*** (0.000)			-3.604*** (0.000)	-5.199*** (0.000)
χ^2_{NORMAL}	2.161 (0.339)	0.512 (0.774)	1.909 (0.384)	4.294 (0.117)	1.814 (0.403)	0.651 (0.722)	2.059 (0.357)	4.059 (0.131)
χ^2_{SERIAL}	29.098*** (0.000)	16.668*** (0.000)	31.770*** (0.000)	31.391*** (0.000)	27.926*** (0.000)	16.652*** (0.000)	31.483*** (0.000)	31.292*** (0.000)
χ^2_{WHITE}	27.517 (0.490)	41.575 (0.318)	37.123 (0.116)	45.997 (0.175)	29.619 (0.382)	36.593 (0.534)	37.636 (0.105)	45.362 (0.192)

*Notes: figures in parentheses indicate p-values. ** and *** indicate 5% and 1% level of significance, respectively.*

4.3 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.[4] We derive the results using both the FM-UECM approach and the DOLS procedure. Since these estimators produce very similar results, for brevity only the results using the former approach are reported. Results of diagnostic tests, which are very similar to those reported in Table 4, are also not reported to conserve space. The results presented in Tables 5 and 6 indicate that the additional controls are statistically significant with their expected signs.

[4] 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 problems of muticollinearity, it is not possible to include all these control variables in a single specification.

Table 5: 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	2.639	24.983***	2.101	33.949***
PRI_t	0.131	-2.305**	0.322	-3.074***
VOG_t	0.648***	0.582***	0.663***	0.724***
PCF_t	0.133	0.915**	0.109	1.229***
FR_t^{DL} / FR_t^{AB}	-1.138***	-2.501***	-1.047***	-3.352***
TOT_t	0.585***	0.396***	0.658***	0.560***
[2] Controlling for monetary shocks				
Intercept	9.581***	42.993***	10.412***	41.467***
PRI_t	-1.056***	-4.542***	-1.037***	-4.303***
VOG_t	0.301***	0.401***	0.301***	0.394***
PCF_t	0.794***	2.016***	0.758***	1.912***
FR_t^{DL} / FR_t^{AB}	-1.176***	-3.657***	-1.092***	-3.433***
MON_t	0.285***	0.301***	0.287***	0.307***
[3] Controlling for fiscal shocks				
Intercept	13.799***	50.548***	14.512***	48.939***
PRI_t	-1.446***	-5.304***	-1.419**	-5.029***
VOG_t	0.412***	0.513***	0.420***	0.518***
PCF_t	0.829***	2.201***	0.782***	2.069***
FR_t^{DL} / FR_t^{AB}	-1.236***	-4.037***	-1.168***	-3.805***
FIS_t	-0.191	-0.138	-0.228*	-0.178*
[4] Controlling for asset prices shocks				
Intercept	16.167***	48.058***	17.787***	47.534***
PRI_t	-1.858***	-5.114***	-1.973***	-4.992***
VOG_t	0.480***	0.556***	0.484***	0.558***
PCF_t	0.979***	2.119***	0.982***	2.027***
FR_t^{DL} / FR_t^{AB}	-1.406***	-3.845***	-1.389***	-3.671***
AP_t	0.235**	0.076	0.297**	0.144

*Notes: the estimates are derived based on the fully-modified unrestricted ECM estimator of Inder (1993). The summary measure for financial repression used in Model A and Model C follows the approach of Demetriades and Luintel (1997) (denoted as FR_t^{DL}) whereas that in Model B and Model D follows the procedure of Abiad and Mody (2005) (denoted as FR_t^{AB}). *, ** and *** indicate 10%, 5% and 1% level of significance, respectively.*

Specifically, we control for terms of trade (TOT_t), monetary (MON_t), fiscal (FIS_t) and asset prices (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 index, respectively. We also attempt to control for other macroeconomic variables, including trade openness (TO_t), social securities (SOC_t), demographic changes

(DEM_t) and non-linear effects of financial repression (FR_t^2). 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.[5] 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). As we can see from Tables 5 and 6, the results remain fairly robust against the inclusion of various proxies for macroeconomic shocks and macroeconomic variables. On the whole, our core results about the effects of financial repression remain unaltered.

Volatility in consumption may come from shocks in goods market due to sudden changes in international terms of trade. Our results that 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 0.285-0.307. 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 share price index, are found to have an amplifying effect on volatility of consumption in the private sector. 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 output 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), who have found a positive link between trade openness and private consumption volatility. In terms of institutional setting, the provision of social security benefits is found to have

[5] However, caveat must be borne in mind 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 per cent of the work force is currently covered by the Employee Provident Fund (EPF) and the Employment Pension Scheme (EPS).

a dampening effect on consumption volatility. 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 much scope for more reforms to take place to improve the coverage of the social security programs.

Table 6: 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	29.225***	46.475***	30.229***	45.471***
PRI_t	-3.094***	-4.895***	-3.124***	-4.704***
VOG_t	0.177*	0.266**	0.171	0.244**
PCF_t	0.735***	1.439***	0.701***	1.310***
FR_t^{DL} / FR_t^{AB}	-0.739***	-2.256***	-0.648***	-1.975***
TO_t	1.651***	1.432***	1.682***	1.523***
[2] Controlling for social securities				
Intercept	7.405**	40.803***	7.134*	38.396***
PRI_t	-2.105***	-5.351***	-2.195***	-5.193***
VOG_t	0.391***	0.496***	0.378***	0.479***
PCF_t	1.817***	2.871***	1.929***	2.902***
FR_t^{DL} / FR_t^{AB}	-0.855***	-3.442***	-0.717***	-3.166***
SOC_t	-1.274***	-1.033***	-1.482***	-1.213***
[3] Controlling for demographic changes				
Intercept	25.371***	53.182***	26.557***	52.523***
PRI_t	-3.410***	-6.269***	-3.486***	-6.153***
VOG_t	0.449***	0.567***	0.454***	0.567***
PCF_t	0.489**	1.631***	0.452*	1.525***
FR_t^{DL} / FR_t^{AB}	-0.559**	-2.872***	-0.465*	-2.643***
AGE_t	-12.668***	-12.121***	-13.131***	-12.719***
[4] Controlling for non-linear effects				
Intercept	44.979***	293.482***	46.823***	306.987***
PRI_t	-1.864***	-3.244***	-1.871***	-3.021***
VOG_t	0.439***	0.596***	0.439***	0.582***
PCF_t	1.134***	1.185***	1.113***	1.068***
FR_t^{DL} / FR_t^{AB}	-13.001**	-111.367***	-13.361**	-117.331***
$\left(FR_t^{DL}\right)^2$ / $\left(FR_t^{AB}\right)^2$	1.147*	11.317***	1.189*	11.963***

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

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 level of financial repression may serve to magnify consumption volatility. Hence, the results also seem to suggest that the benefits of financial liberalization in reducing consumption volatility can only be realized when India becomes sufficiently open financially.

4.4 Further analysis: the composition of private consumption

Different components of volatility in private consumption may respond to financial repression differently. As Blanchard and Simon (2001) put forward, how financial factors affects consumption volatility is theoretically ambiguous. On the one hand, a more developed financial system, which can be achieved through greater financial openness, enables consumers to achieve consumption smoothing through spending more on services and non-durables. On the other hand, better access to credit also allows consumers to adjust faster toward their desired stock of durables, resulting in more volatility of spending on consumer durables.

In this connection, we examine the hypothesis by considering each component of private consumption. This involves analyzing how private consumption in durables, non-durables and services respond to changes in financial sector policies. The results reported in Table 7 clearly show that there is no support for the hypothesis put forward by Blanchard and Simon (2001). Our main finding that financial repression reduces consumption volatility remains robust, irrespective of the types of consumption volatility considered. The coefficients on the financial repression measures are highly significant at the 1% level across all equations. Its effect is found to be largest in durables but smallest in services.

Table 7: Financial repression and the components of consumption volatility

	$Dep. = VOC_t^{SD}$		$Dep. = VOC_t^{HL}$	
	Model A	Model B	Model C	Model D
[1] Durables				
Intercept	12.589***	45.323***	13.455***	43.923***
PRI_t	-1.442***	-4.826***	-1.446***	-4.591***
VOG_t	0.422***	0.508***	0.378***	0.458***
PCF_t	0.908***	2.089***	0.853***	1.963***
FR_t^{DL} / FR_t^{AB}	-1.117***	-3.631***	-1.021***	-3.384***
[2] Non-durables				
Intercept	11.048***	43.432***	12.072***	34.708***
PRI_t	-1.297***	-4.671***	-1.334***	-3.680***
VOG_t	0.439***	0.528***	0.401***	0.481***
PCF_t	0.761***	1.982***	0.729***	1.558***
FR_t^{DL} / FR_t^{AB}	-0.921***	-3.425***	-0.831***	-2.588***
[3] Services				
Intercept	10.933***	43.355***	11.999***	42.715***
PRI_t	-1.274***	-4.656***	-1.317***	-4.516***
VOG_t	0.427***	0.517***	0.393***	0.482***
PCF_t	0.742***	1.969***	0.714***	1.891***
FR_t^{DL} / FR_t^{AB}	-0.909***	-3.416***	-0.825***	-3.228***

*Notes: "durables" consists of furniture and fixtures, electrical appliances, musical instruments, jewels, clothing and footwear, miscellaneous personal goods, rubber and plastic products, etc.; "non-durables" include food and beverages, newspapers and books, petrol and diesel, fireworks, etc.; "services" refers to rent and water charges, medical services, educational fees, entertainment and recreational services, hotels and restaurants, etc.; the estimates are derived based on the fully-modified unrestricted ECM estimator of Inder (1993). *** indicates 1% level of significance.*

5. Conclusions

Many developing countries have reformed their financial systems over the last few decades. While an increased level of financial openness has generally been observed across the world, the debate concerning how financial openness impacts on growth volatility remain contentious. Moreover, the issue is typically discussed within the framework of the relationship between financial development and output volatility; so far little effort has been made to examine the relationship between financial openness and private consumption volatility. An understanding of the way financial openness impacts on macroeconomic volatility is important in order to assess the costs and benefits associated with financial reform policies. We focus on analysing consumption volatility instead of output volatility due to its implications on economic welfare for developing countries.

The present study is motivated by the significant increase in the degree of financial openness and output volatility observed across the developing world, and the lack of any previous time series attempts to analyze the relationship between financial openness 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 1955-2005. In this study, financial repression is measured by two summary measures, which consider various types of domestic and international financial sector policies adopted in the India financial system.

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, we derive the long-run estimates using several different estimators. The results are insensitive to the choice of estimators. The estimated results based on annual data for the period 1955-2005 consistently suggest that financial repression has a significant dampening effect on consumption volatility. We examine the sensitivity of the results to the inclusion of additional control variables, taking into account of various macroeconomic variables, including trade openness, age dependency, social security benefits, and the non-linear effects of financial repression. In order to explore how different sources of volatility influence consumption volatility, we also analyze the volatility of terms of trade, inflation, government spending, and asset prices. The impact of financial repression on consumption volatility is robust to the inclusion of these control variables.

Our study should be seen in the context of a burgeoning literature examining the effects of globalization on growth volatility. While the empirical results presented in this study are intriguing, more analysis is warranted. We hesitate to generalize the findings of this study to other developing countries since the results may be unique to the experience of India due to its own institutional and historical settings. Future studies can look at how financial repression or openness affects consumption volatility in other developing countries using the framework established in this paper.

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