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FINANCIAL LIBERALISATION AND MONETARY POLICY : THE INDIAN EVIDENCE*

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ABSTRACT

It is generally hypothesised that the efficacy of monetary policy is seriously jeopardised by financial liberalisation. However, the empirical evidence to test this proposition is somewhat scanty, especially for LDCs. In this paper, we use an SVAR type model to derive certain causal measures, which can prove useful as indicators of monetary policy potency. Bootstrapping experiments are then conducted to test whether liberalisation has weakened the role of monetary policy in India. We find an attenuation of the relationship between the monetary base and money supply as well as that between money and nominal income. Our results can thus be interpreted as supportive of the hypothesis.

Keywords **: Liberalisation-monetary policy – SVAR- causal measures-Bootstrapping.**

JEL Classification : **E32, C32**

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1.Introduction

The world of finance, throughout its history, has typified the well-known adage " Nothing is permanent, except change". There have, however, been periods when change has been gradual and evolutionary, and others when fundamental and even cataclysmic " shape-shifting" (a term due to Kane (1984)) has occurred within a relatively short span of time. The last three decades have been witness to just such an episodic "space-shifting" in financial markets, institutions and technology on a global scale, with profound consequences for both monetary theory and policy.

These recent changes have usually been subsumed under the generic term ***financial liberalisation***, which is intended to encompass two distinct but interconnected phenomena viz. ***financial deregulation*** and ***financial innovation.*** The latter two phenomena may be viewed as complementary, mutually reinforcing, almost feeding on each other.

Largely as a consequence of the Keynesian legacy, the post World War II period was characterised by extensive financial regulations in most countries. Following Blundell-Wignall et al (1990), these may be classified into two broad categories as follows:

(i) " rate – quantity" (R-Q) regulations on bank deposits and loans (e.g. interest ceilings on deposits and loans, aggregate and sectoral credit controls, capital requirements etc.) and

(ii) "powers" (P) regulations on the scope of activities of categories of financial institutions (e.g. underwriting of securities, combining of banking and insurance, participation in money markets etc.).

Financial deregulation may then be understood to refer to the extent and pace of dismantling of the R-Q and P regulations (see OECD (1989) for a perspective on this process in the OECD countries).

There are pronounced semantic difficulties in the definition of financial innovations (Podolski (1986)). Most analysts prefer to take a pragmatic approach and use the term flexibly to include not only new financial instruments (e.g. swaps, options, junk bonds, CDs etc.) and institutions (e.g. financial futures, money market mutual funds(MMMFs), pension funds etc.) but also adoption of new practices by existing institutions (such as home equity loans, floating rate loans, ATMs, NOWs etc.)[1]

This paper aims to study the impact of financial innovations on the conduct of monetary policy in India. The next Section summarises the received theoretical literature in this area. The framework of the study is outlined in Section 3, whereas Section 4 sets out the econometric

[1] Further the term *financial innovation* is elastic enough to include not only domestic developments, but also developments in foreign exchange and international capital markets.

methodology. Section 5 is devoted to the interpretation of the results of our econometric exercise, with Section 6 tying up the conclusions.

2.Monetary Policy in a Liberalised Context

The process of financial change briefly noted above has impinged on the U.S. and U.K. financial systems in the 1970s, on most other advanced countries in the 1980s and on the LDCs only in the last decade. The change has been unevenly distributed across countries, and even though the general direction of movement has been unambiguously forward, the pace has varied between countries owing to intrinsic structural differences rooted in contestability of markets, structure of competition, industrial concentration and "financial literacy" in general.

Prior to the onset of financial liberalisation, the prevailing paradigm for monetary policy rested on the famous triad of ***instruments-intermediate targets/indicators-objectives.***[2] The guiding principles behind the triad were essentially threefold:

(i) The intermediate targets (usually simple-sum or Divisia monetary aggregates) were "controllable" via the instruments (either a short-term interest rate or the monetary base) within tolerable margins of error.

(ii) The monetary targets bore a stable relationship with macroeconomic aggregates (such as output, inflation and long-term interest rates) so that the intermediate targets served both as early warning signals of portending changes in the macro-aggregates, as well as guideposts for the intended trajectories of these aggregates.

(iii) The flexible exchange rate regime currently in operation, implied a certain independence for the pursuit of national monetary policies (in the case of the EU, this statement has to be suitably qualified).

Financial liberalisation seems to have irreversibly jeopardised all three of the above premises. Firstly, as noted by Tobin (1983), the leverage exerted by the monetary authorities on non-financial variables, was precisely because money bore an exogenously fixed nominal interest rate, inducing portfolio substitution between "money" and "non-money" assets, in response to changes in interest rate levels. The process of financial liberalisation implies a greater role for market forces in the pricing of bank deposits, whose demand thus becomes more dependent on the ***spread*** (between nominal rates on money and near money assets) than on the ***level*** of nominal rates. Since monetary authorities are much

[2] This framework still survives in many countries albeit in a somewhat battered form.

better at influencing short-term rate levels than the spread, this factor seriously erodes their ability to control monetary aggregates.[3] A similar argument can be developed in the context of monetary base control, using the classic Brunner-Meltzer money supply model (Jordan (1984)).

Secondly, the link between monetary aggregates and important macroeconomic magnitudes (especially nominal income) has been rendered tenuous (in the wake of financial liberalisation) due to a host of factors such as

(i) the blurring of the distinction between money and near-money.

(ii) the breakdown of the money demand function (Akhtar (1983), Cotula (1984) etc.)[4]

(iii) the easing of credit and liquidity constraints (owing to the emergence of variable rate lending and large-scale "liability management"- see Goodhart (1986, 1989)) and

(iv) the rising role of arbitrageurs in financial markets, which has introduced volatility in the yield curve (Brown & Manasse (1989)).

In the IS-LM framework, the LM curve becomes both steep as well as stochastic (Tobin (1983)).

Finally, the international dimension of financial liberalisation is reflected in a greater integration of global capital markets in recent years. International capital flows, always on the lookout for profitable portfolio opportunities, are quick to respond to domestic interest changes, setting up a tendency for real interest differentials between countries, to become insignificant. This implies, of course, that the pursuit of domestic monetary policy is seriously circumscribed by the unpredictable responses of global capital flows.

The upshot of the previous discussion appears to be a vastly reduced potency of monetary policy consequent to financial liberalisation. This conclusion is not true in its entirety. Several transmission channels will still remain open, and a few new ones will emerge. A change in interest rates could still affect aggregate consumption via its impact on permanent income, as well as through intertemporal substitution (Bayoumi & Koujianou (1989)), though this channel is likely to be a sluggish one. More importantly, as stressed by Goodhart (1989), the increased elasticity of global capital inflows to domestic monetary policies, implies that the reduced effects on domestic demand are compensated by a greater impact on exchange rates. The latter has, as a matter of fact, emerged as a major channel of monetary policy in several OECD countries, in recent years. A major advantage of the exchange rate channel is claimed to be its direct

[3] Currency, no doubt, is an exception to this phenomenon, but most likely, an increasingly unimportant one.

[4] This feature is somewhat mitigated if Divisia indices, rather than simple-sum aggregates are used in the definition of money (Gabb & Mullineux (1995)).

effect on input prices and inflation, but on the flip side, it makes investment in the manufacturing sector tradeables, unduly dependent on the vacillations in domestic monetary policy (thus impeding the long-term growth prospects of the economy). Finally, the credibility of the Central Bank will have a critical bearing on the extent to which it can influence long-term interest rates via short-term rates, in a liberalised environment.

The above arguments have been developed in the context of short-term interest rates as the operating target. But parallel arguments apply even under monetary base control. The authorities, in this case, can undeniably exert leverage on the level of interest rates via their ability to control the base. But with the demand for loans being inelastic to the ***level*** of interest rates, loans can be funded through wholesale deposits. Thus with low reserve ratios there can be major fluctuations in monetary and credit aggregates, relative to any given base. Even if required reserve ratios are high, a rise in the general level of interest rates widens bank spreads, simply leading to a switch in the avenues of financing rather than to any real effects on borrowing.

3.Framework of the Study

The primary focus of the study is to analyse the likely impact of financial liberalisation on the effectiveness of monetary policy in India. Most of the existing empirical studies on various aspects of financial liberalisation have been located in the context of developed countries. The LDC case differs fundamentally in that financial liberalisation is often "government driven" rather than an autonomous evolution in response to market forces. Additionally, the financial innovations, very often, are virtually transplanted from abroad, with little adaptation to domestic conditions. Together, these features imply that the financial liberalisation process in LDCs lacks spontaneity, is somewhat artificial and often premature, and hence may interface with policy in ways quite distinct from the pattern recorded in advanced countries. Thus, our study, it is hoped, would be of more than mere passing interest, as yet one more country case-study.

Empirical modelling in this area is severely handicapped by the difficulties of finding appropriate proxies for financial innovations and their diffusion. The difficulty is particularly pronounced for countries where innovations have been evolving in a steady gradual stream. The Indian case, by contrast, presents two well-delineated epochs – the period prior to 1991 when the financial sector was heavily regulated and the post-1991 period in which substantial liberalisation has taken place.[5] Thus important insights into the financial liberalisation-monetary policy nexus can be gained by a comparative analysis of these two epochs, which we term as Period I (Jan 1977 to Dec 1991) and Period II (Jan 1992 to Dec 2000). The choice of these periods is econometrically substantiated in the next section.

It is not the purpose of this paper to go into the details of the financial liberalisation programme in India. These have been carefully documented in two recent books (Joshi & Little (1996) and Sen & Vaidya (1997)) as also in the Annual Reports of the RBI (Reserve Bank of India – the central bank of the country) especially over the years 1992-1997. It is important, however, to remember, that in several respects the process of financial liberalisation in India is far from complete. Firstly, on the international front, capital inflows are still regulated and exchange controls persist in various guises. Secondly, there are important R-Q regulations on banks and non-bank financial institutions (e.g. committed loans and SLR requirements). Finally, it is only very recently (2000), that a beginning is being attempted with the relaxation of P regulations (the emergence of MMMFs and banking-insurance combines on a limited scale). Even this brief description should suffice to make it clear that the potency of monetary policy in India need not necessarily have been eroded in the wake of financial liberalisation. Substantial liquidity constraints on aggregate spending may still persist, and global inflows are not substantial enough to interfere importantly with the conduct of monetary policy. The role of monetary policy then becomes an open issue, in the settlement of which, empirical econometric methods can have an important role to play. Further, the scope of our inquiry could be meaningfully focussed on the following two aspects:

(i) the relationship between the operating target and the monetary aggregates and

(ii) the relationship between monetary aggregates and indicators of macroeconomic activity.

A comparative analysis of the above two relationships over Periods I and II should then yield useful insights into the issues of our concern.[6] So far as the operating targets are concerned, the RBI has traditionally shown a predilection for the monetary base (and perhaps rightly so, in view of the shallowness of the market for Treasury bills and dated government securities, which renders transmission from short-term to long-term rates

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[5] This is not to deny that some important innovations had already taken place well before 1991, or that regulatory elements are completely absent in recent times but, as a general description, few observers would quarrel with this scheme.

[6] Three facets of the relationships are singled out for analysis viz. the strength of causation, the magnitude of the impacts and the lags.

rather tenuous). The monetary aggregate chosen for our analysis is M3, the broad money aggregate- both because the RBI continues to rely on it as a monetary policy intermediate target, and also because the narrow money concept M1 loses much of its relevance in the context of financial liberalisation (Goodhart (1986)). Additionally, M1 does not fit well into the current framework of monetary analysis, based on the balance sheet identity for banks, as spelt out, for example in Baltensperger (1980). A reliable Divisia aggregate is not available in the Indian context over a substantially long period, otherwise it would have been the natural choice in a study of this kind.

In assessing the impact of monetary policy , one is interested in the effects of monetary policy on real output as well as prices . Since the lags involved may often be less than one year , an annual model has severe limitations . We therefore resort to a monthly model, which also yields substantial high-frequency data suitable to the techniques that we propose to apply. However, the decision to work with monthly data is not without its own set of problems. Because data on Indian GDP is unavailable except on an annual (or semi-annual) basis, we have to proxy the output variable by the IIP (Index of Industrial Production) on which monthly data is available. This is a significant limitation, increasingly so in recent years when services are becoming predominant in national output. Nevertheless there is no reason to believe that the impact of money on the non-industrial sector is likely to differ substantially from the impact on the industrial sector. Thus, at least, as a first approximation, our analysis is not devoid of meaningful interpretations.

The following combinations of series were studied:

(i) B and M

(ii) M, I and P

Here B is the monetary base, M is the broad money measure M3, I is the index of industrial production and P is the wholesale price index. Data on all these variables (upto December 1999) is available from the ***RBI Report on Currency and Finance*** (annual publication), supplemented for the year 2000 by the monthly ***RBI Bulletins***.

4.Methodological Considerations

There are three distinct aspects to our study. The first stage pertains to the identification of monetary policy shocks and shocks to the non-policy variables in the system. In the second stage, frequency domain methods are invoked to study the interrelationship between the identified shocks.

The final stage is concerned with bootstrapping some of the estimates derived at the second stage.

A.Identification of Shocks

Much attention has been devoted in recent years to the identification of monetary policy shocks (Strongin (1995), Sims & Zha (1995), Bernanke & Mihov (1998) etc.). The recommended approach is based on an SVAR (structural VAR). The superiority of the SVAR approach over other approaches , such as the "narrative approach" of Romer & Romer (1994), has been demonstrated, among others, by Gordon & Leeper (1994) and Bagliano & Favero (1997), and derives from the incorporation of the potential endogeneity of monetary policy in a more transparent way, in the SVAR approach. This endogeneity, if not properly modelled, can contaminate the estimated response of macroeconomic variables to a monetary policy shock.

It is now customary to include in the SVAR, an international commodity price index (which we call P*) to account for the so-called "liquidity" and "price" puzzles (Strongin (1995), Kim (1999) etc.). Our SVAR thus includes the following five variables

P*, I, P, M and B

The international commodity price index is measured in US$[7]. The omission of an interest rate variable in the model may seem rather unusual. However, in the Indian context, until very recently (i.e. 1998), the nominal interest rate structure was heavily regulated. The impact of the real interest variable on the economy, was primarily felt through the inflation rate.

The econometrics of the SVAR is too well-known to merit a detailed description. Briefly, we begin with the system

$$G_0\begin{bmatrix} Y(t) \\ X(t) \end{bmatrix} = G(L)\begin{bmatrix} Y(t-1) \\ X(t-1) \end{bmatrix} + B\begin{bmatrix} S_Y(t) \\ S_X(t) \end{bmatrix} \qquad (1)$$

where Y(t) is the vector of non-policy variables (P*, I, P), X(t) is a vector of monetary variables (M, B) and $S_Y(t)$, $S_X(t)$ are vectors of structural shocks to the non-policy variables and monetary variables respectively. Further G_0 describes the contemporaneous relations among the variables,

[7] The data on P* is obtained from the *International Financial Statistics* (line 00176 axd.)

G(L) is a finite-order matrix lag polynomial and the matrix B records the impact of shocks on the variables in the system.[8]

Model (1) is called the structural form and the corresponding reduced form is

$$\begin{bmatrix} Y(t) \\ X(t) \end{bmatrix} = A(L)\begin{bmatrix} Y(t-1) \\ X(t-1) \end{bmatrix} + \begin{bmatrix} u_Y(t) \\ u_X(t) \end{bmatrix} \tag{2}$$

where $A(L) = G_0^{-1}G(L)$ and $[u_Y(t), u_X(t)]^T$ is the vector of reduced-form residuals (or innovations), which are related to the structural shocks via

$$G_0\begin{bmatrix} u_Y(t) \\ u_X(t) \end{bmatrix} = B\begin{bmatrix} S_Y(t) \\ S_X(t) \end{bmatrix} \tag{3}$$

For identification of the structural parameters, we make the following assumptions:

(i) $[S_Y(t), S_X(t)]^T$ is a vector of contemporaneously uncorrelated disturbances.

(ii) The non-policy variables do not simultaneously react to the policy variables. This assumption is a reasonable one for monthly data. Within a month firms may not adjust their output and prices to unanticipated changes in monetary policy, due to menu costs etc. (Sims & Zha(1995)).The converse reactions, however, are allowed.

(iii) To capture the operational procedures of monetary policy in India, we posit that

$$u_M(t) = mu_B(t) + S_M(t) \tag{4}$$

$$u_B(t) = \phi\, S_M(t) + S_B(t) \tag{5}$$

where $u_M(t)$, $u_B(t)$ are the innovations and $S_M(t)$, $S_B(t)$ are the structural disturbances to M and B respectively.

(4) is the standard money-multiplier relationship, whereas (5) is an elementary adjustment rule of monetary base control, in which the base b is adjusted by a factor ϕ to shocks in M. The extreme cases of full and nil adjustment correspond to $\phi = -1$ and $\phi = 0$, respectively.

Model (3) now assumes the following appearance

[8] B need not necessarily be a diagonal matrix, non-zero off-diagonal elements reflecting the possibility that an endogenous variable may be affected by more than one shock.

$$\begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 \\ a_{41} & a_{42} & a_{43} & 1 & a_{45} \\ a_{51} & a_{52} & a_{53} & 0 & 1 \end{bmatrix} \begin{bmatrix} u_{P*} \\ u_I \\ u_P \\ u_M \\ u_B \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & \phi & 1 \end{bmatrix} \begin{bmatrix} S_{P*} \\ S_I \\ S_P \\ S_M \\ S_B \end{bmatrix} \quad (6)$$

We impose the additional restriction that $\phi = -1$ (which is tested later). This makes (6) an over-identified system.

The starting point of our analysis is the reduced form (2), which we estimate (by OLS) over the full period of our analysis (Jan 1977 to Dec 2000) with 6 lags. In spite of the evident non-stationarity of the variables, the model was estimated in log-levels, with no imposition of cointegrating relations. This is done to circumvent a long-run identification problem, which may otherwise crop up (Sims et al (1990) and Hendry (1996)). The model is estimated with monthly seasonal dummies and lag selection was done via Sims'(1980) procedure. This yielded a lag of 6. As is to be expected, this benchmark model shows considerable evidence of parameter instability. The traditional literature on structural breaks presumes that the break point is known to the investigator. In the Indian case, we know that the reforms process was initiated politically in June 1991, but it would be naïve to presume that this date also coincides with the structural break of the econometric model. The change might have been anticipated for some time and on the other hand, there is also the possibility that its consequences might take some time to impact on the economy. It thus seems more reasonable to assume an interval around the date of the episodic exogenous event, as the possible interval, within which the structural break has occurred. As a matter of fact, we take two such potential intervals, viz. April 1985 - April 1987, and Jan 1990 – Dec 1992. The choice of the latter has already been explained, whereas the former is considered to explicitly allow for the view held by several economists that liberalisation had already begun earlier under the Rajiv Gandhi government in 1986. The method proposed by Andrews (1993) seems particularly appropriate to the problem at hand. Andrews indictes how the hypothesis for parameter constancy *over a particular interval* may be tested against various alternatives. Let the given interval be expressed as $(\pi 1, \pi 2)$ where the truncation points $\pi 1$ and $\pi 2$ are expressed as fractions of the total sampling interval. Further, define $\lambda = [(1-\pi 1)\pi 2]/[\pi 1(1-\pi 2)]$. Andrews tabulates critical values for combinations of λ and p (number of regressors). The maximum number of regressors considered by Andrews is p=20 ; however this figure is likely to be substantially exceeded in VAR models. However, as noted by Bagliano & Favero (1997), for values of λ in the range (1,2), these critical values seem to be approximately 1.12 times the critical values of the chi-squared distribution. The test statistic in the regression case is simply *the maximum* over all the usual Chow chi-squared statistics, evaluated successively at each of the data points in the interval under consideration. In Table 1 we denote this quantity as Sup W(π). The results of the stability tests are reported in Table 1 for each of the 5 equations of our model. The intervals of our potential structural breaks can be expressed as (0.33, 0.41) and (0.52, 0.60) with $\lambda = 1.4$ in both cases. The value for p is 31 (6 lags each of 5 variables together with a constant). Resorting to the extrapolation noted above, the approximate critical values are 50.54 (5% level) and 58.72 (1% level). Table 1 indicates that in the first interval, evidence for structural instability occurs only in the P* and M equations, whereas the evidence for instability is much stronger in the more recent interval considered. Thus, we are led to suppose that a major structural break has occurred in the period Jan 1990 to Dec 1992. The table also indicates for each variable exhibiting a break, the point where the Chow statistic is a maximum, which may be identified as the actual structural breakpoint. We find that the breakpoints for the various equations are clustered around Dec 1991, and hence we take this as a dividing threshhold. Thus, we identify Period I as Jan 1977 to Dec 1991 and Period II as Jan 1992 to Dec 2000.

Lag selection for the reduced forms in both periods was done employing the usual Sims' criterion. The optimal lag was 7 for the first period and 6 for the more recent period. In the interests of uniformity, we chose a common lag of 6 for both periods. As emphasised by Spanos (1990) and Hendry (1996), it is important to subject the reduced form VAR to detailed specification searches. The important misspecification tests are

(i) LM test for autocorrelated residuals
(ii) LM test for ARCH effects in the residuals
(iii) The Shenton-Bowman (1977) test for normality of residuals.

We executed these tests using PCFIML 8.0 (Doornik & Hendry (1995)), and the results are reported in Table 2. These tests[9] are essentially based on the residuals from each equation of the reduced form VAR. By and large, the tests do not indicate serious model inadequacy. The only misspecified equation seems to be that for P* (international commodity price index) in both periods – probably reflecting the fact that the Indian macroeconomic variables have low explanatory power for this international variable. There is also some evidence for non-normality and

[9] The logic of the testing procedure is explained at length in Doornik & Hendry (1995), Chapter 10

ARCH effects for the monetary base in Period II, which we have not been able to successfully account for.

In addition, we performed the vector portmanteau test (described in Doornik & Hendry (1995) p.215) to judge overall goodness-of-fit of the reduced forms over the two periods I and II. This statistic, which we call as VP, has the χ^2 [n^2(s-m)] distribution, with n the number of dependent variables in the system, m the lag length of the dependent variables (assumed equal for all the dependent variables) and s the lag length over which the autocorrelations are tested. For our VARs, n = 5, m = 6 and we select s = 8. Our VP for the two periods has values 48.15 and 62.44 respectively , which are well below the critical values for a χ^2 (50), thus reinforcing our confidence in the models estimated.

We rewrite (6) in the form

$$\begin{bmatrix} S_{P*} \\ S_I \\ S_P \\ S_M \\ S_B \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ D_{21} & 1 & 0 & 0 & 0 \\ D_{31} & D_{32} & 1 & 0 & 0 \\ D_{41} & D_{42} & D_{43} & 1 & D_{45} \\ D_{51} & D_{52} & D_{53} & -\phi & D_{55} \end{bmatrix} \begin{bmatrix} u_{P*} \\ u_I \\ u_P \\ u_M \\ u_B \end{bmatrix} \tag{7}$$

Model (7) is estimated separately over Periods I and II, and the resulting models are referred to as SVAR I and SVAR II. The models were estimated using RATS and in both versions the over-identifying restriction ϕ = -1, is tested for significance. The restriction was supported for Period I but not for Period II. This is in tune with the RBI's observed reliance on M3 as an indicator in Period I[10].Thus, in the final set of estimations, SVAR I was estimated with the restriction ϕ = -1, whereas in the other model ϕ was endogenously estimated. Each of the models produces a vector of structural shocks, which serves as the basis for the second stage of our analysis.

The contemporaneous coefficients are presented in Table 3. Most of the coefficients exhibit the signs that we would expect from theoretical reasoning. Thus, for example, we expect an unanticipated increase in world commodity prices P* to have a simultaneous negative effect on industrial production I, and a positive effect on domestic prices P and money demand M. In a similar manner of reasoning, the other expected effects can be derived. Thus, we *expect* D_{21}, D_{32}, and ϕ to be negative, whereas all the other coefficients are expected to be positive (except D_{52} whose sign is uncertain).Our estimated coefficients in Table 3 have wrong signs for D_{32} in both periods, while D_{42} has the wrong sign only in

[10] M1 was also often used as a subsidiary indicator, during this period.

Period I. This is once again, probably, a reflection of the fact that I is a rather weak proxy for national income. All the other coefficients have the expected signs. The coefficients in the last row of the matrix in (7), indicate how the monetary authorities respond to unanticipated changes in the non-policy variables. Thus the base B is reduced, when either P* or P display unanticipated increases, but actually increased when I shows unexpected increases (i.e. such increases are treated as growth or productivity impulses and accommodated, rather than interpreted as signs of impending inflation).

Additional insights into our models can be obtained by the Forecast Error Variance Decomposition (FEVD) of I and P due to money supply shocks. This is reported in Table 4 for forecast horizons of 6, 12 and 24 months. Three features emerge strongly from the Table viz.

(i) the contribution of money supply shocks to P is of a much higher order of magnitude than their contribution to I

(ii) the importance of money supply shocks declines with the forecast horizon (in case of I as well as P) and

(iii) the importance of money supply shocks has decreased in Period II as compared to Period I (for both I and P)

The last feature, in particular, is consistent with the overall conclusions that follow from the subsequent discussion.

B.Causal Analysis

We may recall that the main aim of this paper is to analyse the interrelationships between the monetary base and money supply on the one hand, and the money supply and macroeconomic aggregates on the other, over the two periods I and II. This may be done via the various structural shocks identified in the first stage of our analysis. We thus focus on the following entities obtained from SVARs I and II

(i) S_M and S_B

(ii) S_M, S_I and S_P

The methodology that we now adopt was first enunciated in the much neglected second half of Granger's (1969) celebrated causality paper. Therein Granger had proposed a novel decomposition of the cross-spectrum to derive the *causal coherency* and other related concepts such as the *causal lag* and *causal gain*. Granger's concept of *causal coherency* and the Geweke (1982, 1984) measures of *linear feedback by frequency* are isomorphic. However, the latter are relatively complicated to compute as they require the factorization of the spectral density matrix (Whittle (1969)), whereas for small systems (order upto 4), the Granger measures are derivable in a straightforward fashion. Additionally, instead of a single measure, as in the Geweke case, Granger's method permits the

calculation of at least three distinct measures (mentioned above), all of which yield useful information about the behaviour of the series.[11] Stiassny (1994) has also exhibited interesting connections between such causal measures and the FEVD of traditional VAR analysis.
As the various structural shocks are contemporaneously orthogonal, we can employ the *simple causal model* discussed at length by Granger (1969). In the trivariate case the model is simply

$$\begin{bmatrix} X(t) \\ Y(t) \\ Z(t) \end{bmatrix} = \begin{bmatrix} a_1(L) & b_1(L) & c_1(L) \\ a_2(L) & b_2(L) & c_2(L) \\ a_3(L) & b_3(L) & c_3(L) \end{bmatrix} \begin{bmatrix} X(t) \\ Y(t) \\ Z(t) \end{bmatrix} + \begin{bmatrix} \varepsilon_1(t) \\ \varepsilon_2(t) \\ \varepsilon_3(t) \end{bmatrix} \tag{8}$$

(Here $a_j(L)$, $b_j(L)$ and $c_j(L)$, j = 1,2,3 are polynomials in the lag operator L, with the constant term zero, the $\varepsilon_j(t)$ are contemporaneously uncorrelated with var $\{\varepsilon_j(t)\} = \sigma_j^2$, j = 1,2,3)
We now denote by a_j the expression $a_j[\exp(-i\omega)]$, with b_j, c_j similarly defined. Further, let $\alpha = (a_1-1)$, $\beta = (b_2-1)$, $\gamma = (c_3-1)$
The *partial cross-spectrum* between X(t) and Y(t) given Z(t) is known to be (at any frequency ω)

$$C_{XY\uparrow Z} = -\frac{\sigma_1^2\sigma_2^2 b_3 a_3}{\Delta} - \frac{\sigma_1^2\sigma_3^2 \beta a_2}{\Delta} - \frac{\sigma_2^2\sigma_3^2 b_1 \alpha}{\Delta} \tag{9}$$

where

$$\Delta = \sigma_1^2|\beta\gamma - c_2 b_3|^2 + \sigma_2^2|c_1 b_3 - b_1\gamma|^2 + \sigma_3^2|b_1 c_2 - c_1\beta|^2 \tag{10}$$

Granger proposed the decomposition of (9) as follows

$$C_{XY\uparrow Z} = C^{(1)} + C^{(2)} + C^{(3)} \tag{11}$$

In (11), $C^{(1)}$ represents the interactions between X and Y *through* Z, $C^{(2)}$ the causal influence of X on Y *given* Z, and $C^{(3)}$ the causal influence of Y on X *given* Z. We now define the three quantities of our interest as follows.

Causal Coherency: The causal coherency from X to Y *given* Z is defined as

[11] We are not making any general claims for the superiority of Granger's method over Geweke's- only maintaining that that the former is more germane to the problem at hand. Geweke's method is possibly better suited for higher order systems. No Monte Carlo evidence on this, however, seems to be available.

$$C(X \to Y \uparrow Z) = \frac{|C^{(2)}|^2}{f_{X\uparrow Z} f_{Y\uparrow Z}} \tag{12}$$

at any frequency ω, where $f_{X\uparrow Z}$ and $f_{Y\uparrow Z}$ are the conditional power spectra of X (given Z) and of Y (given Z).
In turn, the conditional power spectrum of X (given Z) is defined by

$$f_{X\uparrow Z} = \frac{\sigma_1^2\sigma_2^2|b_3|^2 + \sigma_1^2\sigma_3^2|\beta|^2 + \sigma_2^2\sigma_3^2|b_1|^2}{\Delta} \tag{13}$$

with $f_{Y\uparrow Z}$ defined analogously.

<u>Causal lag:</u> The causal phase lag function from X to Y (given Z) is defined as

$$D(X \to Y \uparrow Z) = \tan^{-1}\left\langle \frac{\operatorname{Im} C^{(2)}}{\operatorname{Re} C^{(2)}} \right\rangle \tag{14}$$

for each frequency ω. (The intuitive interpretation of this concept is given in Granger & Hatanaka (1964) p.115-118).

<u>Causal Gain:</u> The causal gain from X to Y (given Z) is defined as

$$G(X \to Y \uparrow Z) = C(X \to Y \uparrow Z)\left\langle \frac{f_{X\uparrow Z}}{f_{Y\uparrow Z}} \right\rangle = \left|\frac{C^{(2)}}{f_{Y\uparrow Z}}\right|^2 \tag{15}$$

for each frequency ω.

The above concepts are notably distinct from the usual concepts of *coherency, phase* and *gain* used in spectral analysis. More specifically, the *causal coherency* measures the strength of causality from one variable to another at each frequency, the *causal lag* is the delay in transmission effects from one variable to another and the *causal gain* is the impact of a unit change in one variable on another. Further, in computing these quantities, we are correcting for the influence of other variables in the system (i.e Z in the above definitions).
We now apply the above definitions to the problem at hand. To gauge the influence of monetary policy shocks on the macro economy, we proceed as follows.The role of X is played by S_M, with Y and Z being in turn, S_I and S_P . Further, by suppressing the role of Z, we can always specialize our definitions to the bivariate case to obtain the quantities C(X→Y) etc. We study the monetary block of our model as a bivariate system, with X assuming the values S_B and Y the values S_M
We study the various causality relationships over three frequency ranges:
R1 : (.042,.056) corresponding to a *long-run* period between 3 to 4 years.

R2 : (.082,.164) corresponding to a *medium-run* period between 1 to 2 years.
R3 : (.164,.656) corresponding to a *short-run* period between 3 months to 1 year.
In each range, we average over the Fourier frequencies in that range. The results of our analysis are presented in Tables 3 to 5.

C. Bootstrapping

No asymptotic theory seems to be available for the distribution of the causal measures discussed above. Hence bootstrapping experiments seem to be indicated before the results in Tables 5 to 7 can be interpreted. This is the third and final stage of our computations.
There are two particular groups of hypotheses of interest. Firstly, whether the specific causal measures are significantly different from zero in each of periods I and II, and secondly, whether they differ significantly over the two periods. Hypothesis testing via bootstrapping in time-series models is a rather complicated field, but the papers by Freedman & Peters (1984), Runkle (1987) and Li & Maddala (1997) furnish several useful guidelines for the practitioner. In dynamic time series models, residual-based bootstrapping is recommended over direct sampling, and hence the first step in our analysis is bootstrapping the residuals in the reduced-form (2). The residuals have to be *rescaled* before the bootsrapping is undertaken (in the manner outlined by Bergstrom (1999)).
We denote the (rescaled) residuals from the reduced form of (2) over Period I as

$$u_t = \{u^t_{P*}, u^t_I, u^t_P, u^t_M, u^t_B\}; (t = 1,...T_1)$$

and those over Period II as

$$v_t = \{v^t_{P*}, v^t_I, v^t_P, v^t_M, v^t_B\}; (t = T_1 + 1,..T_1 + T_2)$$

($T_1$and T_2 are the total number of observations in Periods I and II respectively).
Let these be stacked row-wise in two matrices U and V respectively, so that

$$U = \begin{bmatrix} u_1 \\ u_2 \\ .. \\ .. \\ u_{T_1} \end{bmatrix} \qquad V = \begin{bmatrix} v_{T_1+1} \\ v_{T_1+2} \\ .. \\ .. \\ v_{T_1+T_2} \end{bmatrix}$$

A bootstrap sample may be denoted as $\{u^*_1,...u^*_{T_1}; v^*_{T_1+1},...v^*_{T_1+T_2}\}$, where the first part of the vector represents a resampling with replacement from the rows of U, and the second part a similar resampling from the rows of V. Since the entire vector of residuals is being resampled, the contemporaneous correlation structure of the residuals is being preserved. However, such a scheme, without modification, is likely to give inconsistent estimates in dynamic time-series models (Liu & Singh (1992)). A suggested improvement is to incorporate the *moving overlapping blocks* method of Kunsch (1989). Let L1 and L2 be the chosen block lengths for U and V respectively, then we get (T_1-L1 +1) blocks in U and (T_2 –L2 + 1) blocks in V. (The k-th block from U is, for example $\{u_k, u_{k+1}, \ldots u_{k+L1-1}\}$ and the m-th block from V is $\{v_m, v_{m+1}, \ldots v_{m+L2-1}\}$). We now select randomly b1 blocks from U and b2 blocks from V, where T_1=L1*b1 and T_2=L2*b2. From the ($T_1 + T_2$) sized sample of residuals, we construct a new artificial data set {Y*(t), X*(t)}, t=1,2....($T_1 + T_2$) in the standard recursive fashion (Runkle (1987)). We now compute all the causal measures over both periods from this artificial data set, and this completes the full cycle of a single bootstrap calculation. The bootstrap experiment is repeated B times.[12]

For our problem T_1=180, T_2=108 and we select L1=L2 =12 with b1=15 and b2=9. The number of bootstrap replications B is set at 5000. The simplest way to proceed in our rather complicated hypothesis testing set-up is to follow Efron & Tibshirani (1993), Ch. 16. Let β^I and β^{II} denote a typical causality measure over Periods I and II respectively. We have two groups of tests to perform

1. $H_0 : \beta^I = 0$ vs $H_1 : \beta^I \neq 0$ and similarly for β^{II}
2. $H_0 : \beta^I = \beta^{II}$ vs $H_1 : \beta^I \neq \beta^{II}$

Both groups of hypotheses are tested by constructing appropriate confidence intervals. Several methods of constructing bootstrap confidence intervals are available (e.g. the bootstrap-t, percentile, BC_a, ABC etc.) The BC_a (bias corrected and accelerated) method has the two major advantages of being transformation respecting (i.e. the endpoints transform correctly if we change the parameter of interest from θ to some function of θ) and it is second-order accurate (i.e. it converges to the true confidence interval at rate (1/n) rather than at rate $(1/n)^{0.5}$ as is the case with the bootstrap-t and percentile methods). The construction of the BC_a method may be briefly explained as follows. Let θ be the parameter of interest, $\theta^{\#}$ its estimate from the original sample and $\theta^*(k)$ its estimate from the k-th bootstrap sample, k = 1,2,...B. Let θ^*_α indicate the 100 α-th percentile of the B bootstrap estimates $\theta^*(1)$, $\theta^*(2)$,... $\theta^*(B)$. Then a BC_a interval of intended coverage (1-2α) is $(\theta^*_{\alpha_1}, \theta^*_{\alpha_2})$, where

[12] Even the *moving overlapping block* bootstrap is not free of problems. Politis & Romano (1994) suggest the so-called *stationary bootstrap* to alleviate some of these problems. This refinement has not been attempted here.

$$\alpha_1 = \Phi\left\{b + \frac{b + z_\alpha}{1 - a(b + z_\alpha)}\right\} \tag{16}$$

$$\alpha_2 = \Phi\left\{b + \frac{b + z_{(1-\alpha)}}{1 - a(b + z_{(1-\alpha)})}\right\} \tag{17}$$

(Φ is the standard normal cumulative distribution function and z_α is the 100α-th percentile point of the standard normal distribution). The bias-correction factor b is obtained as

$$b = \Phi^{-1}\{p\} \tag{18}$$

where p is the proportion of the bootstrap replications $\theta^*(k)$ less than the original estimate $\theta^\#$.

The acceleration a is given by the formula (Efron & Tibshirani (1986))

$$a = \frac{\sum_{k=1}^{n} \{\theta^\#(.) - \theta^\#(k)\}^3}{6\left\{\sum_{k=1}^{n} (\theta^\#(.) - \theta^\#(k))^2\right\}^{1.5}} \tag{19}$$

where $\theta^\#(k)$ is a jackknife estimator of θ from the original sample with the k-th value deleted, and

$$\theta^\#(.) = \frac{1}{n}\left\{\sum_{k=1}^{n} \theta^\#(k)\right\} \tag{20}$$

Reverting to our original problem, to test the first group of hypotheses, we construct a BC_a interval for the particular parameter (β^I or β^{II}) of intended coverage 90% i.e.$\alpha = 0.05$ and see whether the origin is covered by the interval.[13] If so, then the null hypothesis is not rejected at the 10% level of significance. To test the second group of hypotheses, we construct a BC_a interval for (β^I - β^{II}), and then proceed as before. The only difference of course, is that in the first group of hypotheses our bootstrap sample is either $\{u_1^*,...u_{T_1}^*\}$ or $\{v_{T_1+1}^*,...v_{T_1+T_2}^*\}$, whereas in the second group of hypotheses the full sample $\{u_1^*,...u_{T_1}^*; v_{T_1+1}^*,...v_{T_1+T_2}^*\}$, is used.

Table 8 summarises the main conclusions of our analysis.We use the symbols (+), (-) and (=) to denote respectively the possibilities that a causal measure has increased, decreased or remained constant from Period I to II. Note that BC_a intervals need to be computed for

[13] α is taken somewhat larger than the conventional value (.025), as BC_a intervals sometimes tend to be too short and are thus prejudiced against the null.

comparisons, only if a particular measure is significant in both periods. We adopt the obvious convention of denoting by (+) the situation where a measure is insignificant in Period I but significant in Period II, by (-) the converse situation, whereas (=) covers the situation when the measure is insignificant in both periods.

We now turn to an interpretation of the results.

5. Interpretation of Results

In Section 3, we had set out our objective as the testing of whether the following twin relationships had weakened in the wake of financial liberalisation in India, viz.

(i) the link between the monetary base (B) and the broad money measure (M) and

(ii) the link between money (M) and real output (proxied by I) and prices (P)

So far as the link between the monetary base B and the money supply M is concerned, Table 5 indicates that the causal coherency was significant in the pre-liberalisation period (Period I) in the frequency ranges R1 and R2 (corresponding to the long and medium-term periods respectively). Thus, prior to liberalisation, the monetary authority seemed to be reasonably successful in maintaining control over the money supply via its chosen operating instrument B, except in the very short run. This control seems to have been totally eroded subsequent to liberalisation (as evidenced by the causal coherencies in Period II). From Table 6, we also find that the lag (from B to M) seems to have increased both at the long and short ends, whereas the message from Table 7 is that the size of the money-multiplier (as measured by the causal gain) has increased over the medium and long runs. Overall, the association between the monetary base and money supply has become more uncertain as well as sluggish, but the reponsiveness (in the sense of elasticity) of money supply to the base has increased. The last feature, in particular, is in line with *a priori* theoretical expectation (Akhtar (1983) and Podolski (1986)).

The relationship between money and real output (I) is insignificant in both periods. While this conclusion agrees broadly with that of Kim (1999) for the G-7 countries, it has to be tempered with the realisation that I is a very unsatisfactory proxy for national output in the Indian context, where agriculture still accounts for about a quarter of the GDP, and services for another 40%. (As already noted earlier, our choice of I is dictated by the fact of our methods requiring substantial degrees of freedom and the unavailability of data on the other components of GDP at frequencies higher than the annual). From Table 6, we see that that the lags have increased substantially subsequent to liberalisation, whereas

Table 7 indicates a significant increase in the elasticity of output with respect to money (post-liberalisation). The last observation is consistent with Tobin's (1983) hypothesis that financial liberalisation results in a significant steepening of the LM curve.

From Table 5, significant causal links from money to prices emerge over the long and medium terms in the pre-liberalisation period – post-liberalisation only the long-term link persists. The message from Table 6 suggests additionally that monetary policy might be affecting prices with a long lag (the lag is about half the period of the average cycle length in the frequency range R1 i.e. about 18 to 24 months). Table 6 further indicates that for the long-term, this lag is constant over the two periods, though the lag length seems to have increased significantly in Period II over the medium and short cycles. There is also a pronounced increase in the causal gain at all frequencies (see Table 7), indicating an overall increase in the elasticity of both output and prices with respect to money supply in the wake of financial liberalisation.

6.Conclusions

Financial innovations have been proceeding apace in the last few decades throughout the world , and even the LDCs have not been immune from their influence. The innovations have been accompanied by extensive dismantling of the regulatory framework. These developments have wide-reaching ramifications, the full extent of which has yet to be grasped. Empirical modelling of this phenomenon is only recently beginning to emerge, but largely in the context of the OECD countries. The evidence for the LDCs seems to be largely of an episodic and narrative kind. This paper hopes to break some fresh ground, by analysing the evidence for the Indian case via formal econometric methods.

The emphasis is on examining the potency of monetary policy in the wake of financial liberalisation. In particular, two questions are singled out for analysis :

(i) Has the link between the monetary base and money supply weakened consequent to liberalisation ? and

(ii) Does money influence output and prices, and how has this influence changed with liberalisation?

The basic issue of identifying exogenous monetary policy shocks has received much attention in the literature and we follow the mainstream approach of identifying such shocks via an SVAR. In the second stage of our analysis, the identified shocks are further processed via causality measures, originally suggested by Granger (1969). Since no asymptotic theory seems to be available for these measures, an extensive set of bootstrapping experiments was resorted to.

Our results seem to indicate that monetary policy has been rendered considerably impotent, consequent to financial liberalisation. The broad money aggregate M3 (M in our econometric exercise) is less firmly related to the monetary base, while simultaneously its association with macroeconomic magnitudes has weakened (it was never significant with real output anyway). The frequency domain results are further supported by the FEVD analysis reported in Table 4. By and large, our results are in concordance with what the theoretical writing on the subject leads us to expect – as always, however, certain peculiarities are present, which cannot be accounted for fully by any existing theory.

The study undoubtedly, suffers from several limitations. We have focussed on a single measure of money supply (M3) and our conclusions could be sensitive to that fact. As noted in Section 1, a Divisia index, if available, would be more appropriate for the study. Unfortunately, for India no reliable or officially accepted set of Divisia aggregates, exists. This constitutes an important limitation of the study. Another major limitation is that open economy considerations have been kept in the background, except through the commodity price level P*. Some allowance for the exchange rate channel could be made (as Kim (1999) seems to have done) by measuring P* in domestic currency (Rupees) instead of US$. But in such a case, our identifying restriction that P* is not affected contemporaneously by monetary policy shocks, cannot be sustained. Nevertheless, the limitation is not as serious, as it would be say, in the context of the OECD countries, since total foreign trade as a percentage of GDP is barely 15% in the Indian context, so that the Indian economy is not fully open now, and certainly was not, prior to 1991. A final limitation is imposed by the fact that our methodology is somewhat aggregative, and the detailed working of the transmission mechanism, cannot be exhibited transparently in the study. This can only be achieved by a full-scale computable general-equilibrium model.

However, in spite of the above limitations, the model can hopefully serve as a useful benchmark study. Two broad policy conclusions emerge from our analysis. Firstly, M3 seems to be becoming increasingly irrelevant, as a monetary policy intermediate target. Broader measures of money supply, or Divisia aggregates, might be better at performing this role. Secondly, monetary base control may, at some stage, have to make way for short-term interest rate targeting. This latter development must, however, await significant restructuring of the Indian money market.

TABLE 1

Test for Structural Breaks

			Sup W(π) in respect of equation for				
Break Interval	(π1,π2)	λ	P*	I	P	M	B
1985:4 to 1987:4	(0.33, 0.41)	1.4	61.42**	48.34	37.66	54.17*	32.78
1990:1 to 1992:12	(0.52, 0.60)	1.4	68.93** (1992:3)	52.01* (1991:9)	46.30	56.49* (1991: 10)	70.23** (1992:2)

Notes:

(i) The notation is as per text (Section 4A)

(ii) (*) and (**) denote significance at 5% and 1% levels respectively

(iii) The calculation of the critical values is also explained in the text.

TABLE 2

Diagnostic Tests for Reduced Form

	Equation for	LM Test		ARCH Test		Normality Test
		$\chi^2(s)$	F(s,T-k-s)	$\chi^2(s)$	F(s,T-k-s)	$\chi^2(2)$
SVAR I	P*	10.72*	2.61*	9.92*	3.11*	11.54**
	I	8.31	0.92	6.93	2.09	5.14
	P·	9.17	1.14	6.22	0.73	1.29
	M	8.03	0.84	8.41	1.94	3.68
	B	8.85	1.98	7.95	1.54	4.57
SVAR II	P*	10.46*	2.74*	9.58*	2.66*	7.34*
	I	9.32	1.17	8.42	2.27	5.47
	P	8.50	1.46	7.19	1.84	1.88
	M	6.38	0.67	8.06	1.63	2.51
	B	8.77	1.05	9.61*	2.59*	6.11*

Notes:

(i) The LM test is examining for serial correlation upto the fourth order (i.e. s=4). Similarly the ARCH test is for ARCH effects upto the fourth. k is the number of regressors, which is 42 in our case (6 lags on each of 5 variables plus a constant and 11 seasonal dummies). The number of effective observations (T) is 174 for SVAR I and 102 for SVAR II. Thus for SVAR I, the relevant F statistic is F(4,128) and for SVAR II, it is F(4,60)

(ii) (*) and (**) denote significance at 5% and 1% levels respectively

(iii) The notation of the table closely follows that of Doornik & Hendry (1995)

TABLE 3

Contemporaneous Coefficients in the Structural Models

Coefficient	SVAR I	SVAR II
D_{21}	-0.17	-0.29*
D_{31}	0.06	0.11
D_{32}	0.14	0.02
D_{41}	0.23*	0.71*
D_{42}	-0.12	0.16*
D_{43}	1.37**	1.94*
D_{45}	2.61*	3.21*
D_{51}	-0.17*	-0.11*
D_{52}	0.09	0.22*
D_{53}	-0.73*	-0.49**
ϕ	-1	-0.62*
D_{55}	0.88	0.75

Notes:

(i) (*) and (**) denote significance at 5% and 1% levels respectively

(ii) ϕ is restricted at –1 in SVAR I

(iii) the notation is as in (7) of the text.

TABLE 4

FEVD for Industrial Production (I) and Prices (P) due to Money Supply Shocks

	Forecast Horizon (months)	Industrial Production (I)	Prices (P)
SVAR I	6	20.84 (5.12)	45.21 (26.93)
	12	27.36 (8.71)	42.27 (20.54)
	24	8.53 (6.46)	31.19 (28.44)
SVAR II	6	11.47 (5.83)	20.34 (16.23)
	12	9.13 (6.69)	18.87 (11.31)
	24	6.08 (3.15)	10.49 (6.79)

Notes:
(i) The standard errors (displayed in brackets), are calculated by bootstrapping in the manner suggested by Runkle (1987) and Hamilton (1994)

TABLE 5

Causal Coherency [C(X→Y↑Z) or C(X→Y)]

Shocks	Frequency Range	Period I	BC_a	Period II	BC_a
$X=S_M$	R1	0.001	(-.0022,.0018)	0.084	(-.016,.103)
$Y=S_I$	R2	0.0008	(-.0037,.0015)	0.082	(-.011,.099)
$Z=S_P$	R3	0.0000	---	0.015	(-.024,.032)
$X=S_M$	R1	0.140*	(.091, .236)	0.116*	(.081,.230)
$Y=S_P$	R2	0.329*	(.205, .361)	0.018	(-.007,.031)
$Z=S_I$	R3	0.039	(-.008,.053)	0.003	(-.012,.009)
$X=S_B$	R1	0.561*	(.509, .593)	0.081	(-.013,.135)
$Y= S_M$	R2	0.414*	(.378,.443)	0.056	(-.008,.098)
	R3	0.051	(-.019,.074)	0.012	(-.091,.031)

Notes:

(i) (*) denotes significance at the intended level of significance viz. 10%. On the choice of this wider significance see footnote # 13 of the text.

TABLE 6

Causal Lag [D(X→Y↑Z) or D(X→Y)]

Shocks	Frequency Range	Period I	BC_a	Period II	BC_a
$X=S_M$	R1	0.0016	(-.018,.011)	0.490*	(.443,.513)
$Y=S_I$	R2	0.021	(-.029,.040)	0.417*	(.388,.447)
$Z=S_P$	R3	0.057	(-.038,.073)	0.409*	(.374,.434)
$X=S_M$	R1	0.523*	(.487, .551)	0.522*	(.493,.546)
$Y=S_P$	R2	0.155*	(.106, .193)	0.385*	(.342,.408)
$Z=S_I$	R3	0.041	(-.008,.096)	0.331*	(.309,.359)
$X=S_B$	R1	0.179*	(.120, .207)	0.298*	(.268,.326)
$Y= S_M$	R2	0.171*	(.103,.198)	0.136*	(.107,.162)
	R3	0.187*	(.119,.223)	0.289*	(.266,.313)

Notes:
(i) Same as Table 5

TABLE 7

Causal Gain [$G(X \rightarrow Y \uparrow Z)$ or $G(X \rightarrow Y)$]

Shocks	Frequency Range	Period I	BC_a	Period II	BC_a
$X=S_M$	R1	0.08	(-0.21, 0.13)	0.39*	(0.28, 0.54)
$Y=S_I$	R2	0.19	(-0.03, 0.28)	0.34*	(0.21, 0.46)
$Z=S_P$	R3	0.57*	(0.34, 0.77)	0.76*	(0.62, 0.95)
$X=S_M$	R1	0.19	(-0.05, 0.31)	0.29*	(0.14, 0.48)
$Y=S_P$	R2	0.15	(-0.06, 0.26)	0.38*	(0.18, 0.51)
$Z=S_I$	R3	0.04	(-0.19, 0.12)	0.33*	(0.17, 0.46)
$X=S_B$	R1	0.49*	(0.28, 0.66)	0.73*	(0.55, 0.91)
$Y= S_M$	R2	0.24*	(0.09, 0.47)	0.41*	(0.23, 0.66)
	R3	0.06	(-0.14, 0.19)	0.02	(-0.19, 0.21)

Notes:
(i) Same as Table 5

TABLE 8

Comparison of Measures over Periods I and II

Shocks	Frequency Range	Causal Coherency $C(X \to Y \uparrow Z)$ Or $C(X \to Y)$	Causal Lag $D(X \to Y \uparrow Z)$ Or $D(X \to Y)$	Causal Gain $G(X \to Y \uparrow Z)$ Or $G(X \to Y)$
$X=S_M$ $Y=S_I$ $Z=S_P$	R1	(=)	(+)	(+)
	R2	(=)	(+)	(+)
	R3	(=)	(+)	(+) (-0.29,-0.05)
$X=S_M$ $Y=S_P$ $Z=S_I$	R1	(=) (-0.018, 0.053)	(=) (-0.12, 0.08)	(+)
	R2	(-)	(+) (-0.39, -0.12)	(+)
	R3	(=)	(+)	(+)
$X=S_B$ $Y=S_M$	R1	(-)	(+) (-0.26, -0.04)	(+) (-0.38, -0.11)
	R2	(-)	(=) (-0.11, 0.17)	(+) (-0.31, -0.06)
	R3	(=)	(+) (-0.24, -0.04)	(=)

Notes:

(i) The notation in this table, is explained towards the end of Section 4C. Note that BC_a intervals are calculated only in those cases, where the measure is significant in both Periods I and II. The intended coverage of the BC_a intervals is 90% (see footnote # 13 of the text).

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