

Money-income Link in Developing Countries: a Heterogeneous Dynamic Panel Data Approach

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I. INTRODUCTION

The question whether real money causes real output appears to be important for many economists working in the area of macroeconomics and, has been subjected to a variety of modern econometric techniques, producing conflicting results. One often applied method to investigate the empirical relationship between money and real activity is Granger causality analysis [Granger (1969)]. Using this approach, the causality question can be sharply posed as whether past values of money help to predict current values of output. This concept, however, should be clearly distinguished from any richer philosophical notion of causality [cf. Holland (1986)]. Present paper examines the relationship between money (both M1 and M2) and income (Real GDP) for 15 developing countries using a newly developed heterogeneous dynamic panel data approach.¹ Sims (1972) postulated “the hypothesis that causality is unidirectional from money to income agrees with the post war U.S. data, whereas the hypothesis that causality is unidirectional from income to money is rejected”. Since then a voluminous literature has emerged testing the direction of causality.² Some studies have tested the relationship between these variables and the direction of causality for a particular country using time series techniques [e.g., Hsiao (1979) for Canada, Stock and Watson (1989) for U.S. data, Friedman and Kuttner (1992, 1993) for U.S. data, Thoma (1994) for U.S. data, Christiana and Ljungquist (1988) for U.S. data, Davis and Tanner (1997) for U.S. data, Jusoh (1986) for Malaysia, Zubaidi, *et al.* (1996) for Malaysia, Biswas and Saunders (1998) for India, and Bengali, *et al.* (1999) for Pakistan]. Other studies

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¹For more detail, see Hurlin and Venet (2001).

²See Hayo (1999) for a survey on this issue.

have tested the above on a number of countries, for example Krol and Ohanian (1990) used the data for Canada, Germany, Japan and the U.K. Hayo (1999) using data from 14 European Union (EU) countries plus Canada, Japan, and the United States. More recently Hafer and Kutan (2002) used a sample of 20 industrialised and developing countries. This paper contributes to this later strand of the literature, which it extends in three directions. First, it employed a newly developed panel cointegration technique [Larsson, *et al.* (2001)], to examine the long-run relationship between money and income. Second, the study performs panel causality test, recently developed by Hurlin and Venet (2001), to explore the direction of causality between the said variables. Third, the important contribution of the present study is to test whether relationship between money and income is homogeneous or heterogeneous across countries.

Friedman (1961, 1964); Friedman and Schwartz (1963) postulated money or its rate of change tends to “lead” income in some sense. A body of macroeconomic theory, the “Quantity Theory”, explains these empirical observations as reflecting a causal relationship running from money to income. However, it is widely recognised that no degree of positive association between money and income can be itself prove that variation in money causes variation in income. More recent studies, both theoretical and empirical, also have shown money to have little or no direct effect on economic cycles. Rudebusch and Svensson (1999, 2002), for example, conclude that the behaviour of money (real or nominal) has no marginally significant impact on deviations of real output from potential (the output gap) once past movements in the gap and real rates of interest are accounted for. Such findings, on the basis of what Meyer (2001) refer to as the “Consensus macro model”, have achieved an influential position among macroeconomists and policy-makers.

The empirical evidence on the relationship between these variable is overwhelming, however the existing empirical studies do not settle the issue of causality. A variety of modern econometric techniques producing conflicting results. For example, Stock and Watson (1989) use a vector autoregressive (VAR) model that accounts for several important economic variables and find that money exerts a statistically significant effect on real economic activity. Friedman and Kuttner (1992, 1993), on the other hand, show that using the same specification as Stock and Watson but extending their sample through the 1980s obviates the money income link. Friedman and Kuttner’s results indicate that interest rates are relatively more useful in explaining movements in output. Thoma (1994) also reports that changes in money do not have a statistically significant impact on output in the United States.

In this paper we use new data and new econometric procedures that directly confront the potential biases induced by simultaneity and unobserved country-specific effects that have plagued previous empirical work on this issue. There are several key advantages of using panel data over a single time series or cross section

data.³ One is the larger sample size and hence more powerful significance tests; another is the possibility of analysing dynamic properties of the relationship under study and to include country and time specific effects.

Specifically, we use, Im, Pesaran and Shin (hence after IPS) (2002) approach to test for the order of integration in panel data. The standard trace rank test by Johansen (1989) is used to perform the cointegration rank test for each country separately and provides the basis for panel test on cointegration rank proposed by Larsson, *et al.* (2001). Ahmad and Iqbal (2002) show that, Hurlin and Venet (2001) step by step testing procedure may be helpful to reduce the dynamic panel data bias when time dimension is 30 years. So we use, Hurlin and Venet (2001) approach for causality testing in dynamic panel data. We consider a balanced panel of 15 developing countries over the 1971-2001 period. IPS test results show, money (both narrow and broad) and prices (CPI) clearly I(1), contain a unit root. Income (Real GDP) and interest rate are clearly reject the hypothesis of unit root in some cases (without time trend). Panel cointegration test shows series are cointegrated. Country-by-country results based on Johansen multivariate likelihood-based inference also support the findings of panel test that money and income are cointegrated. Panel causality test results indicate, irrespective of the choice of lag order, we reject the null hypothesis of homogeneous causality. The results show a heterogeneous bi-directional causality for some countries.

The reminder of this paper is set out as follows: Section 2 describes the panel tests for unit root, cointegration and causality. We discuss the data in Section 3 and present the main results in Section 4. Conclusions of the paper are presented in Section 5.

2. UNIT ROOT, COINTEGRATION, AND CAUSALITY IN PANEL DATA

We start with panel unit root test.

2.1. Panel Unit Root Test

Levin and Lin (1992), consider the following model:

$$y_{i,t} = \rho_i y_{i,t-1} + z'_{i,t} \gamma + u_{i,t} \quad \dots \quad \dots \quad \dots \quad \dots \quad \dots \quad (1)$$

($i=1, \dots, N; t=1, \dots, T$)

Where, $z_{i,t}$ is the deterministic component and $u_{i,t}$ is a stationary process. $z_{i,t}$ could be zero, one, the fixed effects, μ_i , or fixed effect as well as a time trend. The Levin and Lin (LL) test assume that $u_{i,t}$ are *iid* $(0, \sigma_u^2)$ and $\rho_i = \rho$ for all i . The LL test is restrictive in the sense that it requires ρ to be homogeneous across i . Im, Pesaran

³For more detail see, *Journal of Econometrics*. Vol. 68, Special Issue on Panel Data.

and Shin (2002) (IPS) allow for a heterogeneous coefficient of $y_{i,t-1}$ and propose an alternative testing procedure based on averaging individual unit root test statistics. IPS suggested an average of the augmented Dickey-Fuller (ADF) tests when $u_{i,t}$ is serially correlated with different serial correlation properties across cross-sectional units, i.e.; $u_{i,t} = \sum_{j=1}^{p_i} \alpha_{ij} u_{i,t-j} + \varepsilon_{it}$. Substituting this $u_{i,t}$ in (1) we get:

$$y_{i,t} = \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + z'_{it} \gamma + \varepsilon_{it} \quad \dots \quad \dots \quad \dots \quad (2)$$

where,

$$p_i = 0, 1, 2^4$$

The null hypothesis is:

$$H_0 : \rho_i = 1$$

for all i and the alternative hypothesis is:

$$H_a : \rho_i < 1$$

For at least one i . The IPS t -bar statistic is defined as the average of the individual ADF statistic as:

$$\bar{t} = \frac{1}{N} \sum_{i=1}^N t_{pi} \quad \dots \quad \dots \quad \dots \quad \dots \quad \dots \quad \dots \quad (3)$$

where t_{pi} is the individual t -statistic of testing $H_0 : \rho_i = 1$ in (2). It is known for a fixed N as $T \rightarrow \infty$

$$t_{pi} \Rightarrow \frac{\int_0^1 W_{iz} dW_{iz}}{[\int_0^1 W_{iz}^2]^{1/2}} = t_{iT} \quad \dots \quad \dots \quad \dots \quad \dots \quad \dots \quad (4)$$

IPS assume that t_{iT} are iid and have finite mean variance. Then

$$\frac{\sqrt{N} \left(\frac{1}{N} \sum_{i=1}^N t_{iT} - E[t_{iT} / \rho_i = 1] \right)}{\sqrt{Var[t_{iT} / \rho_i = 1]}} \Rightarrow N(0,1) \quad \dots \quad \dots \quad \dots \quad (5)$$

as $N \rightarrow \infty$ by the Lindeberg-Levy central limit theorem. Hence

$$t_{IPS} = \frac{\sqrt{N} (\bar{t} - E[t_{iT} / \rho_i = 1])}{\sqrt{Var[t_{iT} / \rho_i = 1]}} \Rightarrow N(0,1) \quad \dots \quad \dots \quad \dots \quad (6)$$

⁴As proposed by Madala and Wu (1999).

as $T \rightarrow \infty$ followed by $N \rightarrow \infty$ sequentially. The values of $E[t_{iT}/\rho_i=1]$ and $\text{Var}[t_{iT}/\rho_i=1]$ have been computed by IPS vis simulations for different values of T and p_i' s.

2.2. Panel Cointegration Test

Larsson, Lyhagen, and Lothgren (2001) presented a likelihood-based (LR) panel test of cointegration rank in heterogeneous panel models based on the average of the individual rank trace statistics developed by Johansen (1995). The likelihood ratio test statistic (also called the trace statistic) of the reduced rank hypothesis for country i $H_i(r) : \text{rank}(\Pi_i) \leq r$, against the full rank alternative for the bivariate model $H_i(2) : \text{rank}(\Pi_i)=2$, is given by,

$$LR_{iT}(H_i(r)/H_i(2)) = -T \sum_{j=r+1}^2 \ln(1 - \hat{\lambda}_j), \quad \dots \quad \dots \quad \dots \quad (7)$$

Where $\hat{\lambda}_j$ is the j th ordered eigen value obtained from a certain eigen value problem that is specific to the chosen model, the deterministic components (intercepts and trends) in the model. Johansen (1995, Chapters 6 and 11) presents the various alternatives in great detail along with the asymptotic distribution of trace statistic is a complex function of vector-valued Brownian motions. In what follows we let Z_k , $k=p-r$, denote the asymptotic distribution of the trace test $LR_{iT}(H_i(r)/H_i(2))$.

Larsson, *et al.* (2001) proposed a panel test of the hypothesis that all of the N countries in the panel have the same (maximum) number of cointegrating relationships among the P variables in a general p -variate VECM.

$$H_0: \text{rank}(\Pi_i) = r_i < r \text{ for all } i=1, \dots, N$$

against the full rank alternative for all countries.

$$H_1: \text{rank}(\Pi_i) = p \text{ for all } i=1, \dots, N.$$

and

$$Z_{\overline{LR}}(H(r)/H(p)) = \frac{\sqrt{N}(\overline{LR}_{NT}(H(r)/H(p)) - E(Z_k))}{\sqrt{\text{Var}(Z_k)}} \quad \dots \quad \dots \quad (8)$$

where the $\overline{LR}$ -bar statistic $\overline{LR}_{NT}(H(r)/H(p))$ is defined as the average of the N individual trace statistics $LR_{iT}(H(r)/H(p))$ statistic as $\overline{LR}_{NT}(H(r)/H(p)) = \frac{1}{N} \sum_{i=1}^N LR_{iT}(H(r)/H(p))$, and $E(Z_k)$ and $\text{Var}(Z_k)$ is the mean and variance of the asymptotic trace statistic. The proposed testing procedure is the sequential procedure suggested by Johansen (1988). First, the hypothesis that $r = 0$ is tested. If this

hypothesis is rejected, the hypothesis that $r=1$ is tested. This sequential procedure is continued until the null is not rejected or the hypothesis $r=p-1$ is rejected. This procedure gives the rank estimate r . Johansen (1995) has shown that this procedure asymptotically yields the correct size of the trace statistic. As the trace statistic diverges to infinity with T when the true rank is larger than the hypothesised rank this is also true for panel rank statistic $\overline{LR}_{NT}$ and the standardised static $Z_{\overline{LR}}$. To perform the panel rank test the expected value $E(Z_k)$ and $\text{Var}(Z_k)$ of the asymptotic trace statistic are needed for the calculations of the standardised panel rank statistic $Z_{\overline{LR}}(H(r)/H(2))$. These moments can be obtained from stochastic simulations as described in Johansen (1995, Chapter 15).

2.3. Fixed Coefficients Approach

Hurlin and Venet (2001), proposed an extension of the Granger (1969) causality definition to panel data models with fixed coefficients. Consider the following model:

$$y_{i,t} = \sum_{K=1}^P \gamma^{(K)} y_{i,t-K} + \sum_{K=0}^P \beta_i^{(K)} x_{i,t-K} + v_{i,t} \quad \dots \quad \dots \quad \dots \quad (9)$$

With $P \in N^*$ and $v_{i,t} = \alpha_i + \varepsilon_{i,t}$, where $\varepsilon_{i,t}$ are i.i.d. $(0, \sigma_\varepsilon^2)$. Contrary to Nair-Reichert and Weinhold (2001), they assumed that the autoregressive coefficients $\gamma^{(k)}$ and the regression coefficients slopes $\beta_i^{(K)}$ are constant $\forall k \in [1, P]$. Also assumed that parameters $\gamma^{(k)}$ are identical for all individuals, whereas the regression coefficients slopes $\beta_i^{(K)}$ could have an individual dimension. In model (9), Hurlin and Venet (2001), consider four principal cases.

2.3.1. Homogeneous Non-causality Hypothesis (HNC)

The first case corresponds to the homogeneous non-causality (HNC) hypothesis. Conditionally to the specific error components of the model, this hypothesis implies that there does not exist any individual causality relationships:

$$\forall i \in [1, N] \quad E(y_{i,t} / \bar{y}_{i,t}, \alpha_i) = E(y_{i,t} / \bar{y}_{i,t}, \bar{x}_{i,t}, \alpha_i) \quad \dots \quad \dots \quad (10)$$

In model (9), the corresponding test⁵ is defined by:

$$\begin{aligned} H_o : \beta_i^{(K)} &= 0 \quad \forall i \in [1, N], \forall k \in (1, p) \quad \dots \quad \dots \quad \dots \quad \dots \quad (11) \\ H_a : \exists(i, k) / \beta_i^{(K)} &\neq 0 \end{aligned}$$

⁵Here, we do not consider instantaneous non-causality hypothesis.

In order to test these Np linear restrictions, we compute the following Wald Statistic:

$$F_{hmc} = \frac{(RSS_2 - RSS_1)/(Np)}{RSS_1/[NT - N(1+p) - p]} \quad \dots \quad \dots \quad \dots \quad \dots \quad (12)$$

Where, RSS_2 denotes the restricted sum of squared residual obtained under H_0 and RSS_1 corresponds to the residual sum of squares of model (9).

If the realization of this statistic is not significant, the homogeneous non-causality hypothesis is fail to rejected. This result implies that the variable x is not causing y in all the N countries of the samples. The non-causality result is then totally homogeneous and the testing procedure with goes no further.

2.3.2. Homogeneous Causality Hypothesis (HC)

The second case corresponds to the homogeneous causality (HC) hypothesis, in which there exist N causality relationships:

$$\forall i \in [1, N] \quad E(y_{i,t} / \bar{y}_{i,t}, \alpha_i) \neq E(y_{i,t} / \bar{y}_{i,t}, \bar{x}_{i,t}, \alpha_i) \quad \dots \quad \dots \quad (13)$$

In this case, they assumed that the N individual predictors, obtained conditionally to $\bar{y}_{i,t}, \bar{x}_{i,t}$ and α_i , are identical:

$$\forall (i, j) \in [1, N] \quad E(y_{i,t} / \bar{y}_{i,t}, \bar{x}_{i,t}, \alpha_i) = E(y_{i,t} / \bar{y}_{j,t}, \bar{x}_{j,t}, \alpha_j) \quad \dots \quad \dots \quad (14)$$

If we reject the null hypothesis of non-homogeneous causality (HNC), two configurations could appear. The first one corresponds to the overall causality hypothesis (homogeneous causality (HC) hypothesis) and occurs if all the coefficients β_i^K are identical for all k and non-null. The second on, which is the more plausible, is that some coefficients β_i^K are different for each individual. Thus, after the rejection of the null hypothesis of HNC, the second step of the procedure consists in testing if the regression slope coefficients associated to $x_{i,t-k}$ are identical. This test corresponds to a standard homogeneity test. Formally, the HC test is the following:

$$\begin{aligned} H_o : & \forall k \in [1, p] \quad / \beta_i^k = \beta^k \quad \forall i \in [1, N] \\ H_a : & \exists k \in [1, p], \exists (i, j) \in [1, N] / \beta_i^k \neq \beta_j^k \quad \dots \quad \dots \quad \dots \end{aligned} \quad (15)$$

The HC hypothesis implies that the coefficients of the lagged explanatory variable $x_{i,t-k}$ are identical for each lag k and different from Zero. Indeed, if we have rejected, in the previous step, the HNC hypothesis $\beta_i^K = 0 \quad \forall (i, k)$, this standard specification test allows testing the homogeneous causality hypothesis.

In order to test the HC hypothesis, we have to compute the following 'F' statistics:

$$F_{hc} = \frac{(RSS_3 - RSS_1)/[p(N-1)]}{RSS_1/[NT - N(1+p) - p]} \quad \dots \quad \dots \quad \dots \quad \dots \quad (16)$$

Where RSS_3 corresponds to the realisation of the residual sum of squares obtained in model (9) when one imposes the homogeneity for each lag k of the coefficients associated to the variable $x_{i,t-k}$.

If the F_{hc} statistics with $P(N-1)$ and $NT-N(1+P)-P$ degrees of freedom is not significant, the homogeneous causality hypothesis is fail to rejected. This result implies that the variable x is causing y in the N countries of the samples, and that the autoregressive processes are completely homogeneous.

2.3.3. Heterogeneous Causality Hypothesis (HEC)

The third case corresponds to the heterogeneous causality hypothesis (HEC). Under HEC hypothesis, they assumed first that there exists at least one individual causality relationships (and at the most N), and second that individual predictors, obtained conditionally to $\bar{y}_{i,t}, \bar{x}_{i,t}, \bar{\lambda}_t$ and α_i , are heterogeneous.

$$\exists i \in [1, N] \quad E(y_{i,t} / \bar{y}_{i,t}, \alpha_i) \neq E(y_{i,t} / \bar{y}_{i,t}, \bar{x}_{i,t}, \alpha_i) \quad \dots \quad \dots \quad \dots \quad (17)$$

$$\exists (i, j) \in [1, N] \quad E(y_{i,t} / \bar{y}_{i,t}, \bar{x}_{i,t}, \alpha_i) \neq E(y_{j,t} / \bar{y}_{j,t}, \bar{x}_{j,t}, \alpha_j) \quad \dots \quad \dots \quad (18)$$

2.3.4. Heterogeneous Non-causality Hypothesis (HENC)

The last case corresponds to the HENC. In this case, they assumed that there exists at least one and at the most $N-1$ equalities of the form

$$\exists i \in [1, N] \quad E(y_{i,t} / \bar{y}_{i,t}, \alpha_i) = E(y_{i,t} / \bar{y}_{i,t}, \bar{x}_{i,t}, \alpha_i) \quad \dots \quad \dots \quad (19)$$

The third step of the procedure consists in testing the heterogeneous non-causality hypothesis (HENC). For that, we consider the following test:

$$\begin{aligned} H_o : \exists i \in [1, N] / \forall k \in [1, p] \beta_i^K &= 0 \\ H_a : \forall i \in [1, N], \exists k \in [1, p] / \beta_i^K &\neq 0 \quad \dots \quad \dots \quad \dots \quad \dots \quad (20) \end{aligned}$$

They proposed here to test this last hypothesis with two nested tests. The first test is an individual test realised for each individual. For each individual $i = 1, \dots, N$, test the nullity of all the coefficients of the lagged explanatory variable $x_{i,t-k}$. Then, for each i , test the hypothesis $\beta_i^K = 0, \forall k \in [1, p]$. For that, we compute N statistics:

$$F_{henc}^i = \frac{(RSS_{2,i} - RSS_1)/p}{RSS_1/[NT - N(1+2p) + p]} \quad \dots \quad \dots \quad \dots \quad \dots \quad (21)$$

Where $RSS_{2,i}$ corresponds to the realisation of the residual sum of squares obtained in model (9), when one imposes the nullity of the k coefficients associated to the variable $x_{i,t-k}$ only for the individual i .

A second test of the procedure consists in testing the joint hypothesis that there are no causality relationships for a sub-group of individuals. Let us respectively denote I_c and I_{nc} the index sets corresponding to sub-groups for which there exists a causal relationships and there does not exist a causal relationship. In other words, we consider the following model $\forall t \in [1, T]$:

$$y_{i,t} = \sum_{k=1}^p \gamma_i^k y_{i,t-k} + \sum_{K=0}^p \beta_i^K x_{i,t-k} + v_{i,t} \quad \dots \quad \dots \quad \dots \quad \dots \quad (22)$$

with $\beta_i^K \neq 0$ for $i \in I_c$
 $\beta_i^K = 0$ for $i \in I_{nc}$

Let $n_c = \dim(I_c)$ and $n_{nc} = \dim(I_{nc})$. Suppose that $n_c/n_{nc} \rightarrow \theta < \infty$ as n_c and n_{nc} tend to infinity. One solution to test the HENC hypothesis is to compute the Wald statistic.

$$F_{henc} = \frac{(RSS_4 - RSS_1)/(n_{nc}p)}{RSS_1/[NT - N(1+p) - n_c p]} \quad \dots \quad \dots \quad \dots \quad \dots \quad (23)$$

Where RSS_4 corresponds to realisation of the residual sum of squares obtained in model (9) when one imposes the nullity of the k coefficients associated to the variable $x_{i,t-k}$ for the n_{nc} individuals of the I_{nc} sub-group.

If the HENC hypothesis is fail to rejected, it implies that there exists a sub-group of individual for which the variable x does not cause the variable y . The dimension of this sub-group is then equal to n_{nc} . On the contrary, if the HENC hypothesis is rejected, it implies that there exists causality relationships between x and y for all individual of the panel.

3. THE DATA

Section 3 describes the data. In this paper we use a balanced panel of 15 developing countries for the period 1971–2001 to analyse the dynamic relationship between money and income. We examine the contemporaneous correlation of money and income, and check for evidence of Granger causality between money and income. The panel unit root, panel cointegration, and panel causality tests (causality tests with a lag order from one to five) are implemented. In a first step, a two-variable model is estimated containing only money and income. In a second step, a

three-variable system with the price level added is analysed. Finally, in a third step the interest rate is included in the model. The data used in this study, are annual observations of money, measured as a narrow (M_1) and broad (M_2) aggregate; income, measured as real GDP (at 1995 prices); the price level, measured as the consumer price index (CPI) (1995=100) denoted by (P); and a short-term interest rate (R) (see Appendix for more detail). We used CPI to increase the sample of countries: using the GDP deflator results in a reduction in country coverage. All the variables, except interest rate, in our data set are transformed into natural logarithms for the usual statistical reasons. The data source for all the series is IMF publication "International Financial Statistics" [CD-ROM (2001)]. Our criteria for including a country in our data set are follows. The country must not be highly developed in 1970s (according to World Bank definition). It must have 31 continuous annual observations (because we are using balanced panel data techniques) on the variables of interest and its population exceed 1 million in 1998. Fifteen countries were found to meet this criteria as follows: Costa Rica, Guatemala, India, Indonesia, Korea, Malaysia, Mexico, Pakistan, Paraguay, Philippine, Singapore, South Africa, Sri Lanka, Thailand, and Turkey.

4. EMPIRICAL RESULTS

In this section we summarise the results based on panel unit root test, panel cointegration test and panel causality test.

4.1. Panel Unit Root (IPS) Test Results⁶

The preliminary step in our analysis is concerned with establishing the degree of integration of each variable. For this purpose we test for existence of unit root at level and differences of each series in our sample of 15 developing countries. If all the series are stationary, than traditional estimation methods can be used to estimate the relationship among the variables, in this case, money (both M_1 and M_2), income (Y), CPI (P) and interest rate (R). If, however, at least one of the series is non-stationary then more care is required. In the first case we assume that none of the individual series in our model contains a time trend. Thus, it is assumed for each series Y_{it} that $E(\Delta Y^*_{it})=0$. This means that each series could contain a non-zero intercept but not a time trend. The results based on IPS t -bar statistic are reported in Table 1. The null hypothesis of the test is reported that the variable contains a unit root (non-stationary) and the alternative hypothesis is stationery.

As it is a one-sided test, a statistic less than $-2.05(-1.9)$ would cause rejection at 1 percent (5 percent) of the null hypothesis of non-stationary. Money (both M_1 and M_2) and CPI (P) clearly fail to reject the null hypothesis of non-stationary (unit

⁶Individual country results of unit root test based on augmented Dickey-Fuller (ADF) test are available on request. Here we discuss only IPS test results.

Table 1

Panel Unit Root (IPS) Test

Series	IPS Statistics	Inference
P=0		
Constant		
M1	-1.02	Fail to reject H_0
M2	-1.41	Fail to reject H_0
Y	-2.09	Reject H_0
P	2.95	Fail to reject H_0
R	-1.97	Reject H_0
Constant and Trend		
M1	1.08	Fail to reject H_0
M2	1.78	Fail to reject H_0
Y	5.04	Fail to reject H_0
P	2.65	Fail to reject H_0
R	-0.65	Fail to reject H_0
P=1		
Constant		
M1	-1.63	Fail to reject H_0
M2	-0.95	Fail to reject H_0
Y	-2.67	Reject H_0
P	1.59	Fail to reject H_0
R	-2.16	Reject H_0
Constant and Trend		
M1	-0.65	Fail to reject H_0
M2	-0.58	Fail to reject H_0
Y	-1.27	Fail to reject H_0
P	-0.30	Fail to reject H_0
R	-1.56	Fail to reject H_0
P=2		
Constant		
M1	-1.75	Fail to reject H_0
M2	-1.45	Fail to reject H_0
Y	-3.87	Reject H_0
P	-1.31	Fail to reject H_0
R	-1.27	Fail to reject H_0
Constant and Trend		
M1	-0.24	Fail to reject H_0
M2	-0.81	Fail to reject H_0
Y	-1.22	Fail to reject H_0
P	-1.0	Fail to reject H_0
R	-1.49	Fail to reject H_0

Without Time Trend (Constant)

Critical Value at 1percent = -2.05

Critical Value at 5percent = -1.90

Critical Value at 10 percent = -1.82

With Time Trend (Constant and Trend)

Critical Value at 1percent = -2.68

Critical Value at 5percent = -2.53

Critical Value at 10 percent = -2.45

root) as shown in Table 1. On the other hand, income (Y) clearly reject the null hypothesis of unit root in all cases and rate of interest (R) is also reject the null hypothesis except at $P=2$. However, our assumption that there is no time trend, especially in the case of money (both M_1 and M_2), income (Y), and CPI (P) may not be very appropriate (note: visual inspection of the data show a time trend⁷). Therefore, we test stationarity allowing for a time trend. Table 1 also reports, the results of the panel unit root test (IPS) with time trend. All the series are found non-stationary (we fail to reject H_0 of non-stationary). Given the presence of non-stationary variables in both specifications (with and without time trend), we now proceed to test for cointegration.

4.2. Panel Cointegration Test Results

Given the results of the IPS test, it is possible to apply cointegration methodology in order to test for the existence of a stable long run relationship between the money and income. In Table 2 and Table 4 we report results of cointegration tests based on the likelihood ratio test statistic (also called the trace statistic) by Johansen (1995) and Larsson, *et al.* (2001) likelihood-based (LR) panel test of cointegration, respectively. Tests have the same null hypothesis that is, there is no cointegration and the alternative hypothesis there is cointegration.

As seen from Table 2, in case of money (M_1) and income (Y), the most common selected rank is $r=1$ (12 of the 15 countries have $r=1$), that indicate a cointegrating relation between M_1 and Y for these countries. Guatemala, India, and Thailand, have $r=0$, which state no cointegration between M_1 and Y for these countries. The case for M_1 , Y , and P shows the selected ranks are $r=1$, $r=2$, for 12 countries (6 countries in each category). Three countries (India, Korea, and Turkey) have the selected rank is $r=0$. But in case of M_1 , Y , P , and R , 14 countries have cointegration (with selected ranks from $r=1$ to $r=3$). Only India has rank=0. Eleven countries have $r=1$ in M_2 , Y case. Four countries (Korea, Paraguay, South Africa, and Thailand) have rank=0. In case of M_2 , Y , and P , 10 countries have cointegrating relationship with rank order from $r=1$ to $r=2$. Rest of the five countries have rank=0. The case for M_2 , Y , P , and R , we found cointegration in 14 countries with rank order from $r=1$ to $r=3$. Only India has rank $r=0$.

As we know, that the panel rank test is one sided and a level α test of the hypothesis $H_0: \text{rank}(\Pi_i)=r_i \leq r$ for all i , is rejected if $Z_{LR}(H(r)/H(P)) > Z_{1-\alpha}$ is the standard normal $(1-\alpha)$ quantile. Therefore, in these three cases (M_1 , Y), (M_1 , Y , P), and (M_1 , Y , P , R), we reject the null hypothesis that the largest rank in the panel is $r=0$ (as seen from Table 4). The cases for (M_2 , Y), (M_2 , Y , P), and (M_2 , Y , P , R) shows the same results (rejection of null hypothesis that the largest rank in the panel is $r=0$) in all three cases. So we reject the null hypothesis of no cointegration in all

⁷These results are available on request.

Table 2

Table 2

Table 3

*Simulated Moments of Z_k for the Model $H^*_{(r)}$*

$K = p-r$	$E(Z_k)$	$Var(Z_k)$
1	2.34	8.34
2	11.45	25.47
3	29.37	38.89
4	48.31	60.34

The moments are obtained from the procedure described in Johansen (1995; Chapter 15) using 10,000 replicates.

Table 4

Panel Cointegration (Larsson, et al.) Test

	M_1, Y			M_1, Y, P				M_1, Y, P, R				
	$r=0$	$r=1$	Rank (r_i)	$r=0$	$r=1$	$r=2$	Rank (r_i)	$r=0$	$r=1$	$r=2$	$r=3$	Rank (r_i)
$Z_{LR}^{-1}(H(r)/H(4))$	7.71*	2.35	1	10.66*	3.86*	1.20	2	14.7*	5.27*	2.33*	1.101	3
	M_2, Y			M_2, Y, P				M_2, Y, P, R				
	$r=0$	$r=1$	Rank (r_i)	$r=0$	$r=1$	$r=2$	Rank (r_i)	$r=0$	$r=1$	$r=2$	$r=3$	Rank (r_i)
$Z_{LR}^{-1}(H(r)/H(4))$	7.99*	1.51	1	8.12*	2.99	1.02	1	14.49*	2.69	2.23*	1.55	2

Panel Rank Test has critical values:

$r = 0 \Rightarrow 7.14$

$r = 1 \Rightarrow 3.32$

$r = 2 \Rightarrow 2.13$

$r = 3 \Rightarrow 1.64$

six cases. If a cointegrating rank of one or two (in this case may be up to four) is found the individual series are both non-stationary and cointegrated or stationary, respectively. If, however, the estimated of the system is zero, one of the series can be non-stationary and the other stationary or both series can be non-stationary but not cointegrated. Based on panel cointegration rank test we can say, money (both M_1 and M_2), income (Y), prices (P), and interest rate (R) are cointegrated. Therefore, there is a stable long-run relationship between money and income.

4.3. Panel Causality Test Results⁸

First we test the causality from money (both M_1 and M_2) to income (Y).

4.3.1. Causality from Money to Income

The results based on panel causality test are presented in Table 5. These results show that the homogeneous non-causality hypothesis (HNC) is strongly

⁸Since all series (M_1 , M_2 , Y , P , R) are non-stationary, so first we transformed all series into stationary, then applied panel causality tests. All procedures are available on request.

Table 5

rejected for both cases (narrow money (M_1), income (Y), inflation (P), and interest rate (R)) and (broad money (M_2), Y , P , and R). This is irrespective of the choice of lag order. That means causality from money (both M_1 and M_2) and income (Y) cannot be rejected for our sample of 15 countries. After the rejection of the HNC hypothesis, we test homogeneous causality (HC) hypothesis. This hypothesis imposes the strict homogeneity (identical slope coefficient), of the relationship between money (both M_1 , and M_2) and income (Y). HC hypothesis is also rejected for all cases ($M_1 \rightarrow Y$; $M_1 \rightarrow Y, P$, and $M_1 \rightarrow Y, P, R$) and ($M_2 \rightarrow Y$; $M_2 \rightarrow Y, P$, and $M_2 \rightarrow Y, P, R$). These results (rejection of HC hypothesis) are also true for all lag orders. Results also confirm the relative heterogeneity of the 15 developing countries sample. Indeed, it is not surprising that these countries do not follow identical policies and same economic structure. The results exhibit different relationship between money (both M_1 and M_2) and income (Y) in different countries.

Given the results (rejection of HNC and HC hypothesis), one must test heterogeneous causality relationships (HENC hypothesis). In Table 7, the realisations of the individual F^i_{HENC} are reported. These results indicate that money (M_1) Granger causes income (Y) only in 5 (Costa Rica, Guatemala, Indonesia, Pakistan, and South Africa) countries of our panel of 15 developing countries. However, the causal relationship ($M_1 \rightarrow Y$) is very much sensitive to choice of lag orders. Rest of the countries (10 countries) has no causal relationship in case of M_1 and Y . In second step we included consumer price index (denoted by P) in the model and test the F^i_{HENC} hypothesis among money, income and prices ($M_1 \rightarrow Y, P$). We found (From Table 7) causality relationship for 8 countries (Guatemala, India, Indonesia, Mexico, Pakistan, Philippines, South Africa, and Sri Lanka). Only Sri Lanka has the causal relationship ($M_1 \rightarrow Y, P$) independent of the lag orders choice. Then we included short-term interest rate (R) in the model and again test F^i_{HENC} hypothesis. From Table 7, we can see, only 6 countries (Costa Rica, Korea, Paraguay, Thailand, and Turkey) fail to reject the null hypothesis of non-causality. Rest of the 9 countries has a causal relationship among $M_1 \rightarrow Y, P$, and R . Except Sri Lanka (reject the Null hypothesis for all lags), results are sensitive to choice of lag orders.

Then we used broad money (M_2) as definition of money and test causality between money (M_2) and income (Y). As seen from Table 8, only 5 countries (Costa Rica, Guatemala, Korea, Mexico, and Pakistan) reject the null hypothesis of non-causality. Rests of the 10 countries have no causal relationship between money (M_2) and income (Y). Only Costa Rica has rejected the null hypothesis for all lag orders. Then we test causality (F^i_{HENC}) hypothesis among $M_2 \rightarrow Y, P$, and found 7 countries (Costa Rica, Guatemala, India, Malaysia, Pakistan, Sri Lanka, and Turkey) have causal relationship among M_2, Y , and P , but these results are not independent from the lag orders choice. After that, we included short-term interest rate (R) in the model and found (again from Table 8) causal relationship in 9 countries (Costa Rica,

Table 6

Table 7

Table 8

Indonesia, Mexico, Pakistan, Paraguay, Philippines, South Africa, Sri Lanka, and Thailand) of the panel. Except Pakistan, results depend on lag orders.

4.3.2. Causality from Income to Money

The results of the inverse causality tests, from income (Y) to money (M_1 , and M_2) based on F_{HNC} and F_{HC} are reported in Table 6. The homogeneous non-causality hypothesis (F_{HNC}) and homogeneous causality hypothesis (F_{HC}) are strongly rejected for all three cases, “ $Y \rightarrow M_1$ ”, “ $Y \rightarrow M_1, P$ ”, and “ $Y \rightarrow M_1, P, R$ ”. These results are independent from the lag orders. As seen from Table 6, homogeneous causality test (F_{HNC}, F_{HC}) results are same for “ $Y \rightarrow M_2$ ”, “ $Y \rightarrow M_2, P$ ”, and “ $Y \rightarrow M_2, P, R$ ” cases (but we found heterogeneous causal relationship in all cases). Our results seem to confirm the rejection of the hypothesis of same “relationship” for our 15 developing countries panel. That means a homogeneous statistical model cannot represent the effect of income (Y) on money (both M_1 and M_2) in our sample of 15 developing countries. Then we go for testing heterogeneous relationship among the said variables (M_1, M_2, Y, P, R). This is evident from Table 9, that reports a causal relationship between income and money ($Y \rightarrow M_1$) for 6 countries (India, Korea, Paraguay, Singapore, South Africa, and Sri Lanka), but these results are not independent from the lag orders (except Singapore). We found (Table 9) causality in 10 countries of the panel, when we included prices (P) in the model ($Y \rightarrow M_1, P$). We found only 5 countries (Malaysia, Mexico, Paraguay, Philippines, and Thailand) have no causal relationship in case of $Y \rightarrow M_1, P$. Seven countries (Costa Rica, Indonesia, Pakistan, Paraguay, Singapore, South Africa, and Sri Lanka) have a causal relationship among ($Y \rightarrow M_1, P, R$). Only South Africa and Sri Lanka have the relationship irrespective of the choice of lag orders. 4 countries (India, Korea, Singapore and Sri Lanka) have causality between $Y \rightarrow M_2$ (results are reported in Table 10). As seen from Table 10, five countries (India, Pakistan, Paraguay, Singapore and Sri Lanka) have causal relationship in case of $Y \rightarrow M_2, P$. The causal relationship relatively more strong in case of $Y \rightarrow M, P, R$. We found causality in 9 countries of the panel. These results are sensitive to the choice of lag orders. Guatemala, Korea, Malaysia, Paraguay, Thailand and Turkey fail to the reject the null of non-causality in case of $Y \rightarrow M_2, P, R$.

In all, these results obtained from the individual country approach are broadly consistent with those obtained from the panel test based results. The only differences, which are observed, relate to whether the relationship is bi-directional or uni-directional. Thus, a coherent picture seems to emerge from a whole range of causality tests. In order to summarise the results which are reported in Table 2. The results from individual country cointegration (Johansen approach) we found cointegration in 12 to 14 countries in case of “ $M_1 \rightarrow Y, P, R$ ” and 10 to 14 countries have cointegration in case of “ $M_2 \rightarrow Y, P, R$ ”. Panel cointegration test (Larsson, *et al.*

Table 9

Table 10

approach) results confirmed the stable relationship among money (both M_1 and M_2), income (Y), inflation (P) and short-term interest rate (R). The summary tests based on Hurlin and Venet approach provide rejection of the homogeneity view that money-income nexus is same in all 15 countries. Panel causality test results show a complete heterogeneity. Only 5 countries reject the null hypothesis of non-causality in case of $M_1 \rightarrow Y$. But when we included inflation (P) and interest rate (R) in the model number of countries (with causal relationship) increase from 5 to 8 and 9, respectively. We found same results (more or less) when we replaced M_1 with M_2 . 5 countries ($M_2 \rightarrow Y$), 7 countries ($M_2 \rightarrow Y, P$), and 9 countries ($M_2 \rightarrow Y, P, R$) have the causal relationship. We found evidence of reverse causation in 6 countries ($Y \rightarrow M_1$), in 10 countries ($Y \rightarrow M_1, P$) and in 7 countries ($Y \rightarrow M_1, P, R$). We also found clear evidence of reverse causal relationship in 4 countries ($Y \rightarrow M_2$), in 5 countries ($Y \rightarrow M_2, P$) and in 9 countries ($Y \rightarrow M_2, P, R$). There is however, evidence from six countries (Costa Rica, India, Indonesia, Pakistan, South Africa and Sri Lanka), which suggests that, the relationship between money (both M_1 and M_2) and income is bi-directional (inflation (P) and interest rate (R) is also included in the model, so the model is, $M_1 (M_2), Y, P, R$). Guatemala and Singapore also has bi-directional causal relationship between money (only in case of M_1) and income. We also found evidence of bi-directional causality in 4 countries (Korea, Mexico, Paraguay and Philippines) when we used M_2 as definition of money. Malaysia, Thailand and Turkey have a weak evidence of uni-directional causal relationship between money and income.

5. CONCLUSIONS

The main purpose of the present endeavor has been to re-investigate the issue of causality among key aggregate macro-variables across a sample of diverse countries through employing relatively recent and more advance econometric techniques to answer the main empirical proposition whether clustering of diversified economies could conform the importance of money as significant informative tool in setting monetary policy. In general, our results, with *a priori* expectations, do not support an out-and-out rejection of money as an informative economic variable when it comes to setting or evaluating monetary policy, particularly in the case of developing economies. However, in accordance with the earlier empirical evidence, the causal relationship between money and the two variables viz; income and prices appeared to be fairly heterogeneous across diverse sample of 15 developing countries. Many of the countries in the grouping conform *a priori* expectation, while other do not display obvious similarities. Whereas most of the evidence seems to favour the view that, the relationship between nominal money and real output is bi-directional.

It is also evident from our causality tests is that the results are very much country specific. This highlights the dangers from lumping together in cross-section equations countries with very different economic experiences. Which may reflect different institutional characteristics, different policies and differences in their implementation. Thus, it could be ascertained that, economic policies would be country-specific and their success depends on the effectiveness of the institutions which implement them. Therefore, there would be no 'wholesale' acceptance of the view that 'money leads income' and there would be no 'wholesale' acceptance of the view that, money follows income, as well.

APPENDIX

All annual data have been taken from the IMF's International Financial Statistics (CD-ROM) for the period 1971-2001. The following further describes the data for each country.

Prices: Consumer Price Index (CPI), 1995 = 100, line 64.

Money: Narrow Money (M1) line 34 and $M_2 = M_1 + \text{quasi money}$, line 5.

Income: Real GDP (1995 prices), line 90.

Short-Term Interest Rate: Money Market Rate, line 60B, and for some countries, Discount Rate, Line 60.

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