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Inflation-rate volatility and money demand: Evidence from less developed countries

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Abstract

Theory suggests that inflation-rate volatility should affect real money balances, although there is ambiguity about the sign of the effect. This paper uses techniques designed to accommodate nonstationary data, and the major results show that increases in the volatility of domestic inflation exert a significant negative effect on money demand in both the short run and the long run in each of the eight less developed countries (LDCs). Nonstationarity in the demand for money in most of our samples is resolved only when a proxy for inflation-rate volatility is included as a regressor. The interest rate variable, ignored in several LDC studies, exerts a statistically significant long-run effect on real money balances of all eight LDCs, except for Kenya.

JEL classification: E40; E41
Keywords: Money demand; Inflation; Cointegration; Structural instability

1. Introduction

Since the early 1980s, the inflation-rate volatility has increased substantially in most less developed countries (LDCs). A central question has been the effect of such high inflation-rate volatility on the demand for money behavior. Studies by Allen (1982), Choudhry (1999), Garner (1985), Klein (1977), Mizrach and Santomero (1990), Smirlock (1982), and Sweeney (1988), among others, have provided empirical evidence for the developed countries (DCs) (mainly the United States) on the impact of inflation-rate volatility on the money demand. Most of these studies have concluded that inflation-rate

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volatility has a statistically significant effect on DCs' money demand; except for Klein (1977), who finds a statistically significant positive effect of inflation-rate volatility on money demand, most previous studies have concluded that inflation-rate volatility reduces domestic money holdings. However, little is known about the extent to which this conclusion may be confirmed for LDCs.

Studies of LDCs' experience have been few due to the unavailability of data, and no clear pattern of results or conclusions has emerged from such studies. Blejer (1979) uses data from Argentina, Brazil, and Chile and finds statistically significant negative influence in each case. Carlos, Asiles, Honohan, and McNelis (1993) obtain a long-run negative effect and a short-run positive effect using the Bolivian data. Khan (1982) and Paul (1981) find a statistically significant positive effect of inflation-rate volatility for Pakistan and India, respectively. It should, however, be emphasized that unlike studies conducted for DCs, the above empirical inquiries of LDCs have derived their results from models that exclude the interest rate variable.[1] Clearly, standard money demand models that omit the interest rate variable are likely to prove poor predictors of money demand. In this paper, we argue that even if the deposit rates in LDCs are set by authorities, it is still a relevant alternative asset to money that the public can hold. Thus, empirical validity of the importance of the interest rate variable should be treated as an empirical matter. Furthermore, financial and economic reforms in several LDCs since the 1970s have brought about interest rate liberalization as well as varying interest rate data which make exclusion of domestic interest rate from the demand for money functions less logical. The analysis in this paper is focused on those LDCs for which interest rate data are available with reasonable variation and consistency.

Previous LDC studies as well as most studies for the DCs on this subject have relied on the conventional ordinary least squares (OLS) procedure with restrictive lag structures such as the partial adjustment model in estimating the money demand equation, ignoring the time series properties of the data. Because the types of data used in this analysis tend to be nonstationary, it is conceivable that the reported results are subject to the spurious regression problem.[2] Mindful of this, our analyses are conducted using recently developed cointegration and dynamic modelling techniques. The benefit of such a statistical approach is that it provides more efficient short-run and long-run coefficient estimates and is able to overcome many of the problems associated with time series models. For example, if cointegration is achieved between real money balance and its determinants, then the "spurious regression" problem is overcome and the correct statistical specification needed for the analysis has been identified.

The main purpose of this paper is to empirically investigate the impact of inflation-rate volatility on the money demand of eight LDCs over the quarterly period 1973:2 through 1999:4. The LDCs examined in the analysis include Indonesia, Kenya, Malaysia, Mexico, the Philippines, South Africa, Singapore, and Thailand. The evidence presented will add an extra dimension to this literature and provide a basis with which future studies can be compared.[3] Knowledge of the degree to which inflation-rate uncertainty

[1] For China, Arize and Malindretos (2000) find a statistically significant negative effect in both the long run and the short run. Their model includes the interest rate variable.

[2] According to Stock and Watson (2002, p. 461), it is a problem that arises in regression analysis when "stochastic trends can lead two or more series to appear related when they are not." Such a regression has estimators and test statistics that are misleading. According to Granger and Newbold (1974), an $R^2 > \mathrm{DW}$ is a good rule of thumb to suspect that the estimated regression is spurious. Also, Stock and Watson (2002, p. 462) point out that "one special case in which regression-based methods are reliable is when ... the series are said to be cointegrated."

[3] For other aspects of international comparisons, see Hirschberg, Maasoumi, Slottje, and Arize (2003).

(volatility) affects the demand for money is important for both the design of monetary policy and economic reform programs. For example, if inflation-rate volatility has adverse effects on domestic money holdings, stabilization programs in LDCs that have strongly emphasized the need to bring the macroeconomy into balance could be unsuccessful if inflation rates are very volatile. In addition, the intended effect of financial liberalization policy may be doomed by a variable inflation rate and could induce financial crisis to the detriment of real sector growth and savings. Also, a variable inflation rate is believed to hamper the value of money and could threaten to induce instability in the money demand function.

The objective of this paper is fourfold. The first objective is to determine whether there exists a stationary long-run relationship among real money balances, real gross domestic product, domestic interest rate, and inflation-rate volatility in eight LDCs. Tests are conducted using the broad definition of money. The second objective is to determine the sign, magnitude, and statistical significance of the effects of real income (real GDP), interest rates, and inflation-rate volatility on money demand. The third objective is to determine whether the estimated long-run cointegrating relationship between money demand and its determinants exhibits the desired property of structural stability using the testing framework in Hansen (1992). Such tests are important because parameter nonconstancy is indicative of model misspecification; see, for example, Ghysels and Hall (2000), who point out that "structural instability is another form of nonstationarity ... and neglected structural instability can bias inferences about other aspects of the model."

The fourth objective is to estimate the speed of adjustments and mean time lags of real balances in LDCs. In the literature of LDC economies, there is conflicting evidence about how quickly real balances respond to changes in income, interest rates and inflation-rate volatility. The traditional view (see, e.g., Aghevli, Khan, Narvekar, & Short, 1979) is that the speed of adjustment to equilibrium in most LDCs is close to unity because of higher risks and uncertainties attributable to economic and sociopolitical instability and the lack of a variety of financial assets available for the wealth holders to undertake portfolio switches.

On the other hand, Wong (1977) provides an opposite view by arguing that the speed of adjustment to equilibrium in LDCs may indeed be slow if a majority of wealth holders (households and entrepreneurs) in these economies are risk averters. The slowness or the existence of longer lags could arise because risk averters have a tendency to be conservative in adjusting their portfolios in the face of higher risks and uncertainties. He contends that the speed of adjustment should depend on the degree of risk aversion and, consequently, should be treated as an empirical matter.

The remainder of this paper is organized as follows. In Section 2, we examine the theoretical relationship between demand for money and inflation-rate uncertainty (volatility), followed by a discussion of the econometric methodology issues. In Section 3, we discuss the empirical results. Summary and concluding remarks are drawn in the last section.

2. Theoretical considerations and model specification

Economists have long suspected that inflation rates and inflation uncertainty are tightly linked, so that increased inflation entails higher inflationary uncertainty. A theoretical analysis explaining why inflation uncertainty increases with inflation has been conducted by Ball (1992). He develops a model of monetary policy in which a rise in inflation raises uncertainty about future inflation. Such uncertainty

arises because policymakers face a disinflation dilemma: They would like to disinflate, but are afraid of the recession that would occur as a result of their actions. On the other hand, the public does not know whether a policy of disinflation will occur because it depends on the tastes of future policymakers.

Blejer (1979) has stressed that if the rate of inflation becomes variable, forecasting of inflation becomes less precise, and more uncertainty about the future level of prices should be expected.[4] As uncertainty increases, it may affect the demand for money in opposite directions: On the one hand, it will increase the precautionary demand, and on the other hand, as uncertainty increases, the risk of holding real money balances rises relative to other assets, inducing changes in portfolio composition and substitution away from real balances. The net effect of increased variability in the rate of inflation cannot be determined a priori but is, rather, an empirical matter.[5]

Other theoretical studies (Boonekamp, 1978; Klein, 1977) have provided support for the positive effects of inflation uncertainty on the demand for money. Klein (1977) has argued that as uncertainty about the future value of real balances increases, individuals may hold more real balances to increase the probability that they can undertake all of their previously planned transactions. The action would be consistent with each individual's attempting to maintain a stable flow of monetary services. The action would also be positively related to a measure of unpredictability in the value of monetary services. Simply put, increased inflation uncertainty lowers the stream of monetary services from a given real balance, which thereby increases the demand for money.

Boonekamp (1978) has pointed out that the expected real rate of return to holding nominal balances may be decomposed into two parts: the negative of the expected rate of inflation and the variance in that rate. Since the expected return increases with the variance in his formulation, we might plausibly expect the variance measure to enter positively in a function describing the demand for money. Thus, he argues that it is reasonable to assume that market participants who are in the process of establishing a productive portfolio take into account the hedging characteristics of an asset in their decision making. Therefore, under inflation uncertainty, the hedging motive should lead to an increase in money holdings.

In sum, the effect of volatility of inflation on real money balances is an empirical issue because theory alone cannot determine the sign of the relation between real balance and inflation-rate volatility.

A simple real money demand model that includes a proxy for inflation-rate volatility may be represented as:

$$m_t^* = \alpha_0 + \alpha_1 y_t + \alpha_2 R_t + \alpha_3 \sigma_{it} + \varepsilon_t \tag{1}$$

where m_t^* is the logarithm of desired holdings of real money balances, which consists of currency outside the banks, demand deposits at the scheduled banks, and quasi-money divided by the consumer price index;[6] y_t is the logarithm of real GDP; R_t is a market rate of interest; σ_t is the logarithm of a measure of inflation-rate volatility and the subscript i refers to the two alternative

[4] Blejer (1979) points out that "although we assume here that greater variability of the inflation rate does imply more uncertainty, it does not necessarily have to be so. It is conceivable that variable inflation may be better predicted if, for example, the greater variability induces people to invest more resources in gathering information."

[5] Another effect of inflation uncertainty is that it reduces the usefulness of money as a unit of account, which also tends to reduce the demand for money.

[6] See Domowitz and Hakkio (1990, p. 30) for a detailed explanation of the importance of deflating by consumer price index and using real GDP as scale measure.

measures of inflation-rate volatility discussed below. The stochastic disturbance term is ε_t, and the intercept term is α_0.

As is customary, Eq. (1) has assumed that in the long run, any deviation of actual (observable) real money balances from desired (unobservable) should disappear so that it may be viewed as a cointegrating model (Hendry & Ericsson, 1991). The basic idea of cointegration is that two or more nonstationary time series may be regarded as defining a long-run equilibrium relationship if a linear combination of the variables in the model is stationary (converges to an equilibrium over time). Thus, if the money demand function describes a stationary long-run relationship among the variables in Eq. (1), this can be interpreted to mean that the stochastic trend in real money balances is related to the stochastic trends in the real income, interest rates, and inflation risk. In other words, even though deviations from the equilibrium should occur, they are mean reverting (Arize, 1996).

Real money balances in Eq. (1) are assumed to be an increasing function of real income (i.e., real GDP), as the usual budget conditions dictate; that is, α_1 is expected to be positive. In Baumol's (1952) and Tobin's (1956) transactions demand theory α_1 is .5, whereas in Friedman's (1956) version of the quantity theory it is unity. If the income elasticity of money demand is less than unity, economies of scale in cash management or holding are implied. On the other hand, an increase in the domestic interest rate (opportunity cost of holding money) increases the attractiveness of alternative assets and hence reduces the demand for money, so α_2 is expected to be negative. If the income elasticity of money demand is less than unity, economies of scale in cash management or holding are implied. On the other hand, an increase in the domestic interest rate (opportunity cost of holding money) increases the attractiveness of alternative assets and hence reduces the demand for money, so α_2 is expected to be negative.

In the Johansen (1995) procedure, maximum likelihood is applied to an autoregressive representation of the form given by Eq. (2):

$$\begin{bmatrix} \Delta m_t \\ \Delta y_t \\ \Delta R_t \\ \Delta \sigma_{it} \end{bmatrix} = \Gamma(L) \begin{bmatrix} \Delta m_{t-1} \\ \Delta y_{t-1} \\ \Delta R_{t-1} \\ \Delta \sigma_{it-1} \end{bmatrix} + \pi \begin{bmatrix} m_{t-1} \\ y_{t-1} \\ R_{t-1} \\ \sigma_{it-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_t^m \\ \varepsilon_t^y \\ \varepsilon_t^R \\ \varepsilon_{it}^\sigma \end{bmatrix} \tag{2}$$

where $\mathbf{\Gamma}(L)$ is a 4×4 matrix of polynomials in the lag operator, which shifts a series back in time, that is, $Ly_t = y_{t-1}$. Δ refers to the first-differenced operator. The Johansen tests are on the rank of the long-run impact matrix $\mathbf{\Pi}$. In the absence of cointegration, $\mathbf{\Pi}$ is a singular matrix (its rank, $r=0$). Hence, in our case, the rank of $\mathbf{\Pi}$ could be anywhere between zero, if no cointegrating vector exists, and four, the number of variables in the system. If $r=1$,there is a single cointegrating vector, whereas for $1 < r \leq 4$, there are multiple cointegrating vectors.

Finally, if there is a cointegration relationship, Johansen (1995) shows that an alternative way of representing the vector autoregression (VAR) is as a conditional model [see Eq. (3) below] so that the demand for money can be analyzed without reference to the remaining equations. The concept of

weak exogeneity implies that the cointegrating vector and the feedback coefficients enter only the real balance equation, so that inferences about those parameters can be conducted from a conditional model of the real balance alone without loss of information relative to a system analysis.[7] Thus, weak exogeneity permits a simpler modeling strategy, namely, a single-equation analysis rather than a system one. Even lacking weak exogeneity, single-equation modeling can proceed, treating the system-based cointegration coefficients as given (see Ericsson, 1998; Juselius, 1992). Work by Wolters, Terasvirta, and Lutkepohl (1998) points out that if m_t is weakly exogenous but one or more of the other variables are endogenous, "there may be a money demand relation in a conditional error correction equation, since conditioning on a variable that is not weakly exogenous may result in a cointegration relation if the error terms in the system are correlated."

To summarize, if cointegration is established, then the error-correction model (ECM) in Eq. (3) is appropriate for all nonstationary variables (Engle & Granger, 1987) in cointegration equation (1):

$$\Delta m_t = K_0 + \lambda\varepsilon_{t-1}^{m} + \sum_{j=1}^{6}(\delta_j\Delta y_{t-j} + \delta_{6+j}\Delta R_{t-j} + \delta_{12+j}\Delta\sigma t - j) + \sum_{j=0}^{6}\beta_j\Delta m_{t-j-1} + \varepsilon_{1t} \tag{3}$$

where all variables are defined above, and Eq. (3) is in $I(0)$ space so that statistical inference using standard t and F tests are valid. The disturbance term, the intercept term, and the error correction term are ε_{1t}, k_0, and ε_{t-1}^{m}, respectively.

Eq. (3) gives the short-run determinants of money demand and embodies both the short-run dynamics and the long-run relation of the series. At the more intuitive level, the presence of ε_{t-1}^{m} in Eq. (3) reflects that actual real money balances do not adjust instantaneously to their long-run determinants. Therefore, in the short run, adjustments are made to correct any disequilibrium in the long-run money demand. The λ is the error-correction coefficient and measures the response of the regressand in each period to departures from equilibrium conditions (Arize, 1996). The ECM therefore reflects how the system converges to the long-run equilibrium implied by Eq. (1), with convergence being assured when λ is between minus one and zero. Note that the error-correction parameter (λ) in the Eq. (3) has a negative sign. In addition, the value of λ depends on the normalization of the cointegrating vector. Before presentation of empirical results, the next section examines alternative measures of inflation-rate volatility.

2.1. Alternatives measures of inflation uncertainty

Various statistical measures of the second-order moment of economic time series have been suggested in the literature. The most commonly used in this literature is a moving sample standard deviation such as Eq. (4) below. Studies such as those by Blejer (1979), Paul (1981), and many others have included such time-varying measure of inflation-rate variability in their money demand models to account for

[7] Broadly speaking, weak exogeneity means that there is no loss of information with respect to the parameters of interest from failing to model the process determining the (weakly) exogenous variable.

periods of high and low inflation uncertainty. In general, the moving sample standard deviation may be written as:

$$\sigma_{0t} = \left[Z^{-1} \sum_{i=1}^{Z} \left(\frac{P_{t+i-1} - P_{t+i-2}}{P_{t+i-2}} \right)^2 \right]^{1/2} \tag{4}$$

where P is the consumer price index and z is the order of the moving average. As Pagan and Ullah (1988) point out, Eq. (4) could lead to an underestimation of the effect of inflation-rate volatility on decisions because it is not defined in relation to some information set.

To take into account that it is uncertainty—unpredictable component of a variable's movement—rather than variability per se, which matters most to economic agents as an alternative to the moving standard deviation procedure, two variance measures are used here. First, as Engle (1983, p. 287) notes, the conditional variance is "of more relevance to economic agents planning their behavior"; therefore, inflation-rate uncertainty is proxied by the Engle (1983) model (now well known as the autoregressive conditional heteroskedasticity [ARCH] model). This model specifies the variance of a variable as a linear function of the expected squares of the lagged value of the error term from an auxiliary regression determining the mean of the variable of interest.

Assume that the conditional mean and variance of inflation rate are generated as

$$\dot{P}_t = x_t\beta + u_t, \qquad u_t \sim N(0, \sigma_t^2) \tag{5}$$

$$\sigma_t^2 = f_t\alpha$$

where $\dot{P}_t$ is the rate of inflation between quarter $t-1$ and t, $\boldsymbol{x}_t$ is a vector of exogenous variables in the set Ω_t of information available at t and contributing to the conditional mean $x_t\beta$ of $\dot{P}_t$, and $\boldsymbol{f}_t$ is a vector of variables also in the information set at t and contributing to the conditional variance σ_t^2 of $\dot{P}_t$. Expectations of the mean and variance of inflation are, in effect, assumed to be rational with respect to the information sets $\boldsymbol{x}_t$ and $\boldsymbol{f}_t$, respectively.

Given a sample of n observations on $\{\dot{P}_t, \boldsymbol{x}_t, \boldsymbol{f}_t\}$, estimates $\{\hat{\alpha}_t, \hat{\beta}_t\}$ of the parameter vectors can be made by maximizing the log-likelihood function for the sample, namely:

$$\ln(L) = -\frac{n}{2}\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{n}\ln(\hat{\sigma}_t) - \sum_{t=1}^{n}(\hat{u}_t/\hat{\sigma}_t^2) \tag{6}$$

The $\hat{u}_t$ are the estimated residuals $\dot{P}_t - f_t, \hat{\beta}_t$, and the σ_t^2 are the estimated variances $f_t\ \hat{\alpha}$. Again assuming that the forecaster's utility function is of an appropriate form, these time-varying variances $\hat{\sigma}_t$ can be interpreted as measures of the uncertainty surrounding the one-quarter change in the inflation rate between $t-1$ and t.

To implement this model, specific assumptions must be made concerning the elements of the vectors $\boldsymbol{x}_t$ and $\boldsymbol{f}_t$ on which the mean of inflation rate is conditioned. The function for estimating the first moment, that is, the mean of inflation or expected inflation, is given as:

$$\dot{P}_t = \delta_0 + \delta_1\dot{P}_{t-1} + \delta_2\dot{P}_{t-2} + \delta_3\dot{P}_{t-3} + \delta_4\dot{P}_{t-4} + \mu_t^* \tag{7}$$

where $\dot{P}_t$ represents inflation, μ_t^* is the white noise residuals obtained from estimating Eq. (7) using OLS. It can be argued that Eq. (7) should also incorporate other factors such as terms of trade and stock prices, but they are not included because of unavailability of reliable data for these variables. Nevertheless, note that our model satisfies many of the requirements needed for the resulting expected inflation series to be rational approximations. These points should be borne in mind in assessing our empirical results.

The first measure is represented by an ARCH (autoregressive conditional heteroskedasticity) model introduced by Engle (1983), and the pth order ARCH model of the inflation process is formulated as follows:

$$\dot{P}_t = \dot{P}^* + \omega_t$$

$$\omega_t^2 = h_t + e_t$$

$$h_t = E(\omega_t^2 / I_{t-1}) = \alpha_0 + \alpha_1 \omega_{t-1}^2 + \cdots + \alpha_p \omega_{t-p}^2 \tag{8}$$

where h is the conditional variance and the information set, I_{t-1} includes information available through time $t-1$, ω_t is the white noise disturbance term, and p is the lag terms in the model.

The Lagrange multiplier (LM) test for the presence of ARCH effects was conducted for lags 1, 2, 3, and 4 (four ARCH models estimated). The LM test for the ARCH is TR^2, where T is the number of observations and R^2 is the coefficient of determination for the auxiliary regression. This test statistic is distributed as $\chi^2(p)$. The ARCH variable is measured as the logarithm of the conditional standard errors and is denoted as σ_{1t}. The computed LM test statistics are 7.99, 2.59, 0.83, 8.61, 17.65, 4.16, 0.13, and 21.21 for Indonesia, Kenya, Malaysia, Mexico, the Philippines, Singapore, South Africa, and Thailand, respectively. Five of these values are statistically significant at the 5% level. Therefore, based on these test results, the inflation series could be modeled as ARCH(1) for Indonesia, Mexico, the Philippines, Singapore, and Thailand.

The second measure uses the residuals, μ_t^*, to calculate a five-term moving average deviation around the predicted values of $\dot{P}_t$.[8] Specifically, if $\mu_t^* = \dot{P}_t - \dot{P}_t^*$, then σ_{2t} is calculated as

$$\sigma_{2t} = \ln\left[\left(\sum_{i=0}^{5} (\mu_{t-i}^*)^2 / 5\right)\right]^{1/2} \tag{9}$$

where ln is the natural logarithm. The consistency of Eq. (10) is enhanced by the omission of the current value of μ_t^* (see Pagan & Ullah, 1988, p. 97).

[8] There is no consensus as to what the optimal lag length in the moving average measure should be, especially with quarterly data. For example, Ghartey (1998) and Hendry and Ericsson (1991) used a four-quarter moving average; Dooley and Spinelli (1989) used a six-quarter, and Baba, Hendry, and Starr (1992) used an eight-quarter measure. This variable has also been measured by an eight-quarter moving standard deviation, and it yielded similar results to those reported in the text. The five-quarter moving standard deviation is preferred for degrees of freedom reasons.

3. Empirical results

3.1. The data and the unit root tests

The empirical work outlined here uses quarterly data from eight LDCs. The data are taken from the International Monetary Fund's International Financial Statistics (IFS) latest CD-ROM (2000), Supplement on money and various issues of IFS. Real money balances are defined as the broad monetary aggregate, divided by consumer prices (1990 = 100). Real GDP is used as a scale variable, and the level of interest rates is proxied by treasury bill rate (Kenya, Mexico, the Philippines, and South Africa) or money market rate (Indonesia, Malaysia, Singapore, and Thailand); a more detailed data description is given in Appendix A. The consumer price index of each country is used to measure inflation.

A prerequisite for testing for cointegration in a set of variables is to test for stochastic trends in the autoregressive representation of each individual time series. To this end, the results of applying three alternative approaches are reported in Appendix B: the augmented Dickey–Fuller (ADF), the Park and Choi (1988) $G(p,q)$ test statistic, and the Johansen (1995) test. Note that the last two procedures allow for formally testing the null hypothesis that the series is stationary against the alternative of nonstationary.

Without discussing each test statistic in detail, the ADF test results in Appendix B suggest that while it is reasonable to conclude for all variables in Eq. (1) that the null hypothesis of a unit root is accepted for the variables, some caution is necessary (unit root tests were also performed on the differenced variables). It is thus assumed that these series are integrated of order one. This is in line with the rejection of the null hypothesis of $I(0)$ by the Park–Choi and Johansen tests. These results are consistent with a large literature on the unit root properties of time series.

3.2. Cointegration tests[9]

For any set of $I(1)$ variables, Johansen (1992) has developed a system-based cointegration procedure to test the absence or presence of long-run equilibria among the variables in Eq. (1). The test utilizes two likelihood-ratio (LR) test statistics for the number of cointegrating vectors: namely, the trace and the maximal eigenvalue (λ-max) statistics; both have approximate chi-squared distributions. In the results reported here, we corrected these LR test statistics for finite sample bias by multiplying each test statistic by $[0.9(T - \text{number of estimated parameters})/\mathrm{T}] + 0.1$, as discussed by Cheung and Lai (1993).[10] The number of observations is represented by T, and the number of estimated parameters is the VAR order multiplied by the number of variables in the cointegration space.

[9] As suggested by a referee, we present results for all countries with each measure of inflation-rate volatility.

[10] We thank a referee for suggesting this adjustment factor. In an earlier version, we used the same formula but without taking into account 0.9 and 0.1.

Empirically, the lag order of the VAR is not known a priori, so some testing of the lag order may be fruitful to enhance the power of the Johansen procedure. For the system (m_t, y_t, r_t, σ_{1t}), that is, real money balances, real income, interest rate, and ARCH (1) conditional standard deviation, the order of the VAR applied in each cointegration test is three in Kenya, Malaysia, and Singapore, four in Indonesia and Thailand, five in Mexico, the Philippines, and South Africa. For the second system (m_t, y_t, r_t, σ_{2t}), that is, real money balance, real GDP, interest rate, and moving standard deviation, the order of the VAR applied in each cointegraton test is two in Kenya, three in Mexico and Singapore, five in Indonesia, Malaysia, and the Philippines and six in South Africa and Thailand. These lag lengths were determined by the information provided by the Sims' likelihood (LR) test, the Akaike information criterion, and the Ljung–Box statistic for serial correlation.[11]

After deciding the lag length in the VAR, we then followed the procedure in Johansen (1992) to test the joint hypothesis of both rank order and the deterministic components. This approach is important because the asymptotic distribution of the trace and the maximal test statistics is affected by the presence of a trend and/or constant term in the nonstationary part of the data-generation process. The model of each country is estimated under three alternative specifications, and the null of no cointegration is successively tested starting from the restricted version (Case 1). The first case under which we fail to reject the null will determine both the correct specification and the chosen rank. With one exception, an unrestricted constant model (Case 2) was chosen for all the countries in our sample. To perform the cointegration test, the conditioning vector includes centered seasonal and impulse or spike dummy variables (were applicable).[12] Similar dummies are included in the studies by Clements and Mizon (1991), Hendry and Doornik (1994), and many others.

Table 1 presents the cointegration test results. The estimated λ-max and trace test statistics (adjusted using the small sample correction noted above) and their attendant critical values to test for the presence or absence of long-run equilibria for the system (m_t, y_t, r_t, σ_{1t}) are reported in the upper tier of Table 1. For the λ-max and trace statistics, the null hypothesis is that there are, at most, r cointegrating vectors, where the alternative hypotheses are $r+1$ and at least $r+1$ for the λ-max and trace statistics, respectively.

Starting with the λ-max test results for the first system, an important result that emerges is that the null hypothesis of $r=0$ (no cointegration) is rejected at the 5% level in favor of $r=1$ in each country except South Africa. The calculated test statistics range from a low of 25.05 in South Africa to a high of 78.50 in Indonesia. The critical value at the 5% level are from Osterwald-Lenum (1992, p. 468) and are shown in Table 1. The results for South Africa are statistically

[11] Throughout this paper, the computations were done with EVIEWS, GAUSS, MICROFIT, and PcGive computer packages.

[12] Seasonal dummies are centered to ensure that they sum to zero over time (Johansen, 1995). A spike dummy variable is 1 for the ith observation and 0, otherwise, and it is used to adjust for the largest residuals. Following Hendry and Doornik (1994), the criterion used to detect an outlier is the studentized residuals (i.e., the ratio of the absolute value of the residuals to the standard deviation of the residuals) greater than three.

Table 1
Results from cointegration tests

Countries	Maximum Eigenvalue					Trace statistics			
	H_0	$r=0$	$r\leq 1$	$r\leq 2$	$r\leq 3$	$r=0$	$r\leq 1$	$r\leq 2$	$r\leq 3$
	H_1	$r=1$	$r=2$	$r=3$	$r=4$	$r\geq 1$	$r\geq 2$	$r\geq 3$	$r\geq 4$
A. Maximum eigenvalue and trace test for ARCH									
Indonesia		78.50	19.07	6.42	0.04	104.0	25.53	6.46	0.04
Kenya		32.27	12.14	6.75	0.52	51.67	19.40	7.26	0.52
Malaysia		31.51	15.34	13.15	0.43	60.42	28.92	13.58	0.43
Mexico		33.38	13.80	5.52	0.08	52.77	19.40	5.59	0.08
Philippines		28.40	12.13	3.98	1.63	46.15	17.75	5.62	1.63
Singapore		29.48	16.58	5.52	0.07	51.89	22.41	5.60	0.07
South Africa		25.05	16.88	0.69	0.37	43.00	17.95	1.06	0.37
Thailand		36.51	32.73	6.64	0.25	76.13	39.62	6.89	0.25
Critical value									
5%		27.07	20.97	14.07	3.76	47.21	29.68	15.41	3.76
10%		24.73	18.60	12.07	2.69	43.95	26.79	13.33	2.69
B. Maximum Eigenvalue and trace test for moving standard deviation									
Indonesia		46.09	19.85	3.97	0.70	70.61	24.52	4.67	0.70
Kenya		27.57	13.91	5.25	0.10	46.83	19.26	5.35	0.10
Malaysia		32.25	13.50	6.83	0.96	53.55	21.30	7.79	0.96
Mexico		28.37	11.44	4.36	0.43	44.60	16.23	4.79	0.43
Philippines		20.76	12.43	5.11	0.17	38.46	17.71	5.27	0.17
Singapore		39.10	13.34	12.79	0.01	65.24	26.14	12.79	0.01
South Africa*		35.64	13.07	9.16	2.56	60.43	24.79	11.72	2.56
Thailand		25.13	12.45	7.66	0.30	45.53	20.41	7.96	0.30
Critical value									
5%		27.07	20.97	14.07	3.76	47.21	29.68	15.41	3.76
5%*		31.46	25.54	18.96	12.25	32.99	42.44	25.32	12.25
10%		24.73	18.60	12.07	2.67	43.95	26.79	13.33	2.69

r denotes the number of cointegrating vectors. The critical values are from Table 1 of Osterwald-Lenum (1992). The 5%* shown above represents the critical values from Table 2* of Osterwald-Lenum (1992). It is used because South Africa's model allows for a trend and constant in the cointegration space; see Johansen (1992).

significant at the 10% level.[13] Furthermore, except for Thailand, the null hypotheses of $r\leq 1$, $r\leq 2$, and $r\leq 3$ cannot be rejected in favor of the alternative hypotheses of $r=2$, $r=3$, and $r=4$, respectively. These results indicate the presence of one cointegrating relationship for each country. In the case of Thailand, there are two cointegrating relationships.

For the trace test results, we obtain similar conclusions when the null hypothesis of $r=0$ is tested against the alternative hypothesis of $r\geq 1$ in each country except South Africa, which fails at the conventional levels. In addition, the null hypotheses $r\leq 1$, $r\leq 2$, $r\leq 3$ cannot be rejected in all countries except Thailand. In sum, we assume the presence of one cointegrating vector for the

[13] Johansen and Juselius (1990, p. 192) point out that "it seems reasonable . . . to follow a test procedure which rejects for a P value than the usual 5%."

countries in our sample. For Thailand, we assume the presence of two cointegrating vectors. Taken together, these findings suggest that there is a long-run equilibrium relationship among real balances, real income, interest rates, and inflation-rate volatility.[14]

The cointegration test results from estimating the second system (m_t, y_t, r_t, σ_{2t}) are shown in the lower tier of Table 1. The results indicate that the null hypothesis of $r=0$ (no cointegration) is rejected in favor of $r=1$ in all countries. The exception is the Philippines, where both the maximum eigenvalue and the trace test statistics are marginally above the 10% significant level so that the null hypothesis that there is not a unique long-run equilibrium relationship among the variables cannot be rejected at the 10% level. In sum, both the maximum eigenvalue and trace tests suggest the presence of one cointegrating relationship in seven of eight cases.

An economic interpretation of the results is facilitated by normalizing on real balances. The resulting values for real income and inflation-rate volatility are long-run elasticities, whereas the coefficient on the interest rate is a semielasticity. These estimated long-run elasticities and hypothesis test results are reported in Table 2.[15] Johansen's standard errors for the cointegrating estimates are from EVIEWS 3.0 and have been used to calculate the t ratios that are reported in Table 2.

Focusing first on the data for inflation-rate volatility variable for the first system (m_t, y_t, r_t, σ_{1t}) in the upper tier of Table 2, observe that the hypothesis that inflation-rate volatility is an important determinant of the demand for money (i.e., H_0: $\beta_3=0$) is supported in seven out of eight cases when a conditional standard deviation is used. For Indonesia, Kenya, Mexico, the Philippines, Singapore, and Thailand, the inflation-rate volatility is statistically significant at the 5% level; for South Africa, it is significant at the 10% level; and for Malaysia, it is nonsignificant at the conventional levels. In a similar vein, for the second system (m_t, y_t, r_t, σ_{2t}), we obtain the estimated coefficients on the inflation-rate variability that are statistically significant at the 5% level. Also, note that the signs on the inflation-rate volatility variable of each country are also consistent with the theory's a priori expectations. The negative sign in the case of Thailand for the first system was confirmed by following the approach in Juselius (1992).[16] Taken

[14] As suggested by a referee, we examined whether R^* (i.e., the rate of U.S. certificate of deposit plus the annualized quarterly change in the domestic per dollar exchange rate) enters the cointegration space. First, the ADF test suggests that R^* is $I(1)$, whereas the other unit root tests find that it is an $I(0)$. Nevertheless, we proceeded with Johansen's cointegration analysis by taking R^* to be $I(1)$, although it has no stochastic trend and may have nonconstant variance. When Eq. (1) is augmented with uncovered returns to investing in foreign assets, it is statistically nonsignificant. Also, when R^* is included, neither the foreign (U.S.) interest rate nor the change in the exchange rate when entered separately are statistically significant. Failure of this portfolio balance model can be attributed to R^* being dominated by exchange rate changes and having $I(0)$ properties. Another reason is that international investors are concerned with the (expected) returns from their investments; therefore, given the progress of the financial liberalization in the countries studied here, it is clear that if investments are uncovered, UIP then $R=R^*+[(E(e)-e)/e]$ or if covered CIP, then $R=R^*+[(f-e)/e]$ is inevitable. As Hamburger (1977, p. 31) concluded, the domestic interest rate and the foreign interest rate move together and when they do it is the domestic interest rate that determines the amount of money held. As a result, if [domestic interest rate] and [the foreign interest rate] are included in the same equation, only the coefficient of the former remains statistically significant.'' See also Ghosh and Arize (in press) for more on these issues.

[15] These elasticities are obtained by normalizing the estimates of the unconstrained cointegrating vectors on real balances, that is, by setting the estimated coefficients of real balances equal to -1 and dividing the cointegrating vector by the negative of the estimated m_t. The resulting values are the long-run elasticities.

[16] Also, looking at the dominant vector, observe that the estimated coefficient on the inflation-rate volatility variable is negative.

Table 2
Normalized cointegrating vectors and t ratios

Countries	Normalized cointegrating vector	H$_0$: $\beta_1=0$	H$_0$: $\beta_2=0$	H$_0$: $\beta_3=0$
A. Long-run elasticities and hypotheses tests				
Indonesia	$m_t=2.29y_t-0.090R_t-0.66\sigma_{1t}$	18.1	2.1	1.8
Kenya	$m_t=1.41y_t+0.043R_t-1.09\sigma_{1t}$	2.8	2.4	2.2
Malaysia	$m_t=1.47y_t-0.053R_t-12.31\sigma_{1t}$	6.0	0.7	0.8
Mexico	$m_t=1.92y_t-0.010R_t-0.45\sigma_{1t}$	5.4	6.1	4.5
Philippines	$m_t=2.52y_t-0.029R_t-0.55\sigma_{1t}$	8.0	2.7	2.3
Singapore	$m_t=1.10y_t-0.030R_t-0.47\sigma_{1t}$	10.6	2.5	2.8
South Africa	$m_t=1.05y_t-0.020R_t-4.20\sigma_{1t}$	2.8	1.7	1.4
Thailand	$m_t=1.46y_t-0.050R_t-0.780\sigma_{1t}$, $m_t=1.38y_t-0.080R_t+0.530\sigma_{1t}$	14.9	5.0	6.7
B. Long-run elasticities and hypotheses tests				
Indonesia	$m_t=2.33y_t-0.080R_t-0.32\sigma_{2t}$	28.9	2.8	2.0
Kenya	$m_t=0.91y_t+0.030R_t-0.34\sigma_{2t}$	3.6	3.0	2.3
Malaysia	$m_t=1.60y_t-0.029R_t-0.06\sigma_{2t}$	77.5	4.4	2.8
Mexico	$m_t=2.27y_t-0.010R_t-0.24\sigma_{2t}$	4.8	6.0	3.9
Philippines	$m_t=2.72y_t-0.030R_t-0.30\sigma_{2t}$	8.0	2.6	2.0
Singapore	$m_t=1.25y_t-0.018R_t-0.07\sigma_{2t}$	44.9	4.3	3.4
South Africa	$m_t=1.70y_t-0.040R_t-0.003\sigma_{2t}$	3.5	3.5	3.0
Thailand	$m_t=1.56y_t-0.010R_t-0.26\sigma_{2t}$	43.2	1.5	4.4

The values beneath, for example, H$_0$: $\beta_1=0$, are the t ratios. An * implies that a trend term was included (not shown here for simplicity).

together, our results imply a fairly large response of real money balances to changes in the inflation-rate volatility in LDCs.

A further, and perhaps more substantive, evidence for the importance of the inflation-rate volatility variable is that without such a variable we either fail to find cointegration in Kenya, Malaysia, the Philippines, Singapore, and Thailand or we find a weak evidence of a stationary money demand function in Indonesia, Mexico, and South Africa. In addition, as shown below, the constancy of the cointegration space for most of the countries in our sample is rejected without the inclusion of the inflation-rate volatility variable in the estimated cointegrating model.

Another appealing aspect of our results is that the interest rate variable, ignored in several LDCs studies, exerts a statistically significant long-run effect on real money balances. Irrespective of the measure of volatility used, the data in Table 2 indicate that the estimated coefficient on the interest rate variable is negative in sign and significantly different from zero at the conventional levels in seven out of eight countries. The exception is Kenya, where it is positive. The equilibrium elasticities (calculated by multiplying the impact elasticities by the mean of the variable) for the first system (m_t, y_t, r_t, σ_{1t}) are -1.49, $+0.59$, -0.30, -0.40, -0.45, -0.17, -0.23, and (-0.19) for Indonesia, Kenya, Malaysia, Mexico, the Philippines, Singapore, South Africa, and Thailand, respectively. For the second system (m_t, y_t, r_t, σ_{2t}) they are -1.33, $+0.41$, -0.17, -0.40, -0.47, -0.10, -0.47, and -0.11 for Indonesia, Kenya, Malaysia, Mexico, the Philippines, Singapore, South Africa, and Thailand, respectively.

The sign, magnitude and significance of the long-run elasticity of money demand with respect to real GDP reported also in Table 2 are consistent with previous studies (e.g., Aghevli et al., 1979). They range from 1.1 (Singapore) to 2.52 (the Philippines). These elasticities point to an important role for real GDP in determining the demand for money in LDCs. Each of the income elasticities is greater than unity, a verdict that is corroborated by the work of Aghevli et al. (1979) and several others. This evidence concerning the absence of economies of scale in money holdings in LDCs implies that as income rises, velocity tends to decline. A possible reason why the demand for money rises at a faster rate than income in LDCs may be the growing degree of monetization in these economies with more extensive branching of banking services to remote areas and the increasing assimilation of rural agents into the mainstream economic development efforts.

3.3. Constancy of the cointegration space

Having provided evidence concerning cointegration, it seems prudent to examine whether the estimated elasticities are constant over time (Judd and Scadding, 1982). Studies by Arize and Darrat (1994) and Ericsson (1998), among others, have documented that parameter constancy is a crucial issue in empirical money demand studies, and tests for parameter nonconstancy in cointegrated systems have been suggested by Hansen (1992). He suggests three tests for parameter instability in cointegrated systems. For the tests, the null hypothesis is that there is no structural change. Hansen's Lc test considers instability due to relatively constant parameter over the sample period. For this test, the null hypothesis is cointegration with no structural break at an unknown point in time, against the alternative of no cointegration among the variables. Hansen's Mean*F* test is designed to detect, instead, a slow shift of parameters and the tests for overall stability of the model. Hansen's Sup*F* test assumes as the alternative hypothesis that a sudden shift in regime occurs at an unknown point in time. Hansen's parameter stability testing framework is instructive in the sense that it allows testing for parameter constancy as well as cointegration.

Table 3 reports the test results from estimating Eq. (1) with and without the inflation-rate volatility variable. Two important conclusions emerge from these results. First, the estimated Lc test statistics and

Table 3
Hansen's parameter instability test

Country	Hansen's test								
	m,y,R			m,y,R,σ_1			m,y,R,σ_2		
	Lc	M*F*	Sup*F*	Lc	M*F*	Sup*F*	Lc	M*F*	Sup*F*
Indonesia	0.80 [.03]	8.84 [.01]	14.90 [.04]	0.48 [.20]	4.61 [.20]	9.61 [.20]	0.74 [.07]	6.25 [.13]	12.80 [.20]
Kenya	1.32 [.01]	9.38 [.01]	15.44 [0.04]	0.93 [.03]	5.49 [.20]	7.64 [.20]	0.92 [.09]	5.46 [.19]	7.89 [.20]
Malaysia	0.71 [.04]	7.02 [.07]	18.05 [.05]	0.80 [.06]	4.90 [.20]	12.08 [.20]	1.20 [.01]	12.89 [.01]	37.56 [.01]
Mexico	0.73 [.01]	14.22 [.01]	131.08 [.01]	0.41 [.20]	6.01 [.14]	12.75 [.20]	0.32 [.20]	6.52 [.20]	15.97 [.16]
Philippines	1.49 [.01]	13.50 [.01]	18.83 [.02]	0.54 [.19]	5.29 [.20]	13.02 [.20]	0.47 [.20]	4.00 [.20]	9.15 [.20]
South Africa	1.35 [.01]	12.87 [.01]	30.58 [.01]	0.43 [.20]	6.41 [.20]	15.00 [.20]	0.59 [.15]	6.81 [.09]	10.42 [.20]
Singapore	0.79 [.03]	8.99 [.03]	29.07 [.01]	0.27 [.20]	3.34 [.20]	7.17 [.20]	0.30 [.20]	3.29 [.20]	7.61 [.20]
Thailand	1.03 [.01]	8.78 [.01]	19.52 [.01]	0.47 [.20]	4.62 [.20]	8.30 [.20]	0.51 [.20]	5.70 [.20]	10.70 [.20]

M*F* represents Hansen's (1992) Mean *F*-test statistic. A *P* value (in square brackets) is reported for each test statistic. A low *P* value, say below .05 for a particular test statistic is interpreted as the stability of the parameters of the cointegrating vector.

their associated *P* values for the cases that exclude inflation-rate volatility show that the null hypothesis that real money demand is cointegrated with real income and interest rate can be rejected. This is so because all of the *P* values for L_c are below .05.

Second, from the models that include the inflation-rate volatility measure, we gather that most of the Lc test statistics are statistically nonsignificant at the 5% level. The exceptions are those for Kenya, for the case (m_t, y_t, r_t, σ_1), and for Malaysia, for the case (m_t, y_t, r_t, σ_2). In the former, Lc has a *P* value of .03, which is below the significance level of .05. Encouragingly, except the evidence for Malaysia, all the Mean*F* and Sup*F* test statistics are statistically nonsignificant at the 5% level when a measure of inflation-rate volatility is included. The evidence in Table 3 implies that the long-run demand for money function for these LDCs requires the inclusion of the inflation-rate volatility variable to exhibit the desired property of parameter constancy. Without such a variable, the equation appears seriously misspecified and structurally unstable, at least at the 5% level.

3.4. Weak exogeneity

We also examined whether any variable in the four-variable system can be considered weakly exogenous (see Ericsson & Irons, 1994). The data in Table 4 show the results for testing each variable individually and for testing real GDP, interest rate and inflation-rate volatility, jointly.

Table 4
Results of weak exogeneity test

Countries		Real balances	Real income	Interest rate	Inflation-rate volatility	Joint test Wald $\chi^2(3)$
Indonesia	σ_1	− 0.03 (3.97)*	− 0.06 (1.26)	− 1.25 (0.96)	− 0.10 (1.07)	2.82
	σ_2	0.06 (1.86)	− 0.05 (2.01)	− 3.41 (0.78)	0.03 (0.43)	2.11
Kenya	σ_1	− 0.07 (2.55)*	0.03 (1.01)	0.12 (3.65)*	− 0.38 (3.21)*	7.86*
	σ_2	− 0.21 (2.94)*	0.08 (1.77)	0.97 (2.05)*	− 0.87 (1.38)	1.18
Malaysia	σ_1	− 0.09 (2.20)*	− 0.01 (1.36)	0.36 (1.12)	− 0.08 (5.83)*	3.98
	σ_2	− 0.26 (2.89)*	− 0.11 (1.15)	0.94 (2.42)*	− 0.85 (0.41)	1.84
Mexico	σ_1	− 0.17 (3.38)*	0.03 (3.01)*	− 28.13 (4.41)*	− 0.02 (0.08)	9.97*
	σ_2	− 0.11 (2.63)*	0.02 (1.90)	− 22.63 (4.15)*	− 0.05 (0.31)	1.81
Philippines	σ_1	− 0.10 (4.15)*	− 0.02 (1.00)	− 1.73 (1.17)	− 0.52 (2.92)*	3.16
	σ_2	− 0.09 (3.18)*	0.01 (0.60)	− 1.08 (2.30)*	− 0.25 (1.57)	2.89
Singapore	σ_1	− 0.11 (2.94)*	− 0.01 (1.73)	0.75 (0.66)	− 0.06 (1.18)	1.45
	σ_2	− 0.21 (5.40)*	− 0.03 (2.33)*	− 2.51 (1.06)	− 0.44 (0.44)	3.14
South Africa	σ_1	0.12 (1.70)	0.00 (0.28)	− 1.33 (1.61)	− 1.19 (4.88)*	5.73
	σ_2	− 0.20 (3.43)*	0.29 (1.16)	− 2.95 (3.38)*	− 1.28 (1.17)	2.18
Thailand	σ_1	− 0.05 (1.75)	− 0.06 (3.90)	0.05 (0.05)	− 0.68 (3.87)*	7.85*
	σ_2	− 0.01 (1.44)	− 0.11 (3.00)*	− 1.86 (0.57)	− 1.66 (3.57)*	6.73

The absolute *t* values are in parentheses, and σ_1 refers to the first system so uses the conditional standard deviation. The second system is represented by σ_2 so uses a moving average measure. For example, the feedback coefficient is − 0.03 with a *t* ratio of 3.97. Significance of the joint exogeneity of real income, interest rate, and inflation-rate volatility is presented by means of Wald statistics; thus, the degrees of freedom are three.

* *t* ratio or χ^2 value is significant at the 5% level.

The test results show that the null hypothesis of weak exogeneity cannot be rejected in 10 out of 16 cases for inflation-rate volatility variable and real income for 12 out of 16 cases, whereas the null hypothesis for the interest rate variable cannot be rejected in 9 out of 16 cases. The results also confirm that the real money demand variable should be considered endogenous. In addition, the null hypothesis of joint exogeneity of income interest rate and inflation-rate volatility is confirmed in 13 of the 16 cases at a significance level of 5%.

In the next section, we examine elements of the short-run dynamics. Recall that Engle's (1983) LM test suggests that ARCH(1) may be appropriate for the inflation series of Indonesia, Mexico, the Philippines, Singapore, and Thailand. For space considerations, the elements of the short run discussed in the next section are obtained from using σ_{1t} for the abovementioned five countries. The measure σ_{2t} is used for Kenya, South Africa, and Malaysia.

3.5. Error-correction model

The Granger representation theorem proves that if a cointegrating relationship exists among a set of $I(1)$ series, then a dynamic error-correction representation of the data also exists (Engle & Granger, 1987). The methodology used to find the error-correction representation is the general-to-specific paradigm (see Hendry, 1987). For this purpose, Eq. (3) was used. The one-lagged error-correction term ($\varepsilon_{t-1}^{\mathrm{m}}$) generated from the Johansen procedure was used. Then the dimension of the parameter space was reduced to a final parsimonious specification by sequentially insignificant restrictions or eliminating insignificant coefficients.

The results from the ECMs are summarized in Table 5. We report instrumental variable (IV) estimates, given the endogeneity of some of the regressors and the generated regressor problems associated with the use of the conditional standard deviation (Pagan, 1984, theorem 12). The IV estimators used here are outlined in Pagan and Ullah (1988, 1999). The list of instrumental variables consists of the constant term, the rank of σ_{1t}, the lagged error-correction term, five lags in the differences of all variables included in the long-run solution and the conditioning variables,

Table 5
Speed of adjustments and mean time lags for adjustments of real balances

Countries	Speed of adjustments ($\varepsilon_{t-1}^{\mathrm{m}}$)	Response of real balances to each regressor			
		Half-life adjustment	Mean time lag		
			Y	R	σ
Indonesia	− 0.041 (0.0107)	16.56	17.74	25.71	22.31
Kenya	− 0.170 (0.0447)	3.72	4.33	6.01	5.23
Malaysia	− 0.026 (0.0118)	26.31	28.66	38.71	32.95
Mexico	− 0.148 (0.0445)	4.33	1.97	6.94	6.27
Philippines	− 0.073 (0.0293)	9.14	8.98	17.09	12.53
Singapore	− 0.030 (0.0132)	22.76	14.96	34.43	34.93
South Africa	− 0.195 (0.0537)	3.20	1.24	5.22	3.34
Thailand	− 0.053 (0.0210)	12.73	12.52	18.87	13.99

The values in parentheses beside the speed of adjustments are the standard errors. Both the half-life and the mean time lag are in absolute terms and in quarters.

including the first-difference term of weakly exogenous variables. Note that the results from Sargan's test for legitimacy of the instrument set (not reported here) support the validity of the instrument set.

Considering that each regressand in Table 5 is cast in first difference, the empirical results suggest that the statistical fit of each model to the data is satisfactory, as indicated by the values of adjusted R^2, the standard error of estimate (SEE), and F value for testing the null hypothesis that all the right-hand side variables as a group except the constant term have a zero coefficient. As is clear from the data in the table, all of the diagnostic tests support the statistical appropriateness of the equations.

Table 6
Results from ECMs (instrumental variable procedure)

Countries	Estimated equations
Indonesia	$\Delta m_t = \underset{(0.01)}{0.024} - \underset{(0.01)}{0.041}\varepsilon^m_{t-1} + \underset{(0.06)}{0.121}\Delta m_{t-1} + \underset{(0.001)}{0.002}\Delta y_{t-4} - \underset{(0.01)}{0.017}\Delta R_{t-1} - \underset{(0.01)}{0.016}\Delta\sigma_{1t-1} - \underset{(0.01)}{0.010}\Delta\sigma_{1t-2}$ $\bar{R}^2$=.5, SEE = 0.02, F = 10.9, DW = 1.8, $Z_1(4) = 4.3$, $Z_2(1) = 3.7$, $Z_3(2) = 9.7$, $Z_4(1) = 0.01$, $Z_5(1) = 2.3$
Kenya	$\Delta m_t = \underset{(0.09)}{0.352} - \underset{(0.04)}{0.17}\varepsilon^m_{t-1} + \underset{(0.10)}{0.132}\Delta m_{t-4} + \underset{(0.0004)}{0.001}\Delta y_{t-2} - \underset{(0.001)}{0.002}\Delta^2 2R_t + \underset{(0.01)}{0.023}\Delta\sigma_{2t}$ $\bar{R}^2 = 0.4$, SEE = 0.05, F = 7.9, DW = 2.0, $Z_1(4) = 4.6$, $Z_2(1) = 0.1$, $Z_3(2) = 3.5$, $Z_4(1) = 0.04$, $Z_5(2) = 0.2$
Malaysia	$\Delta m_t = \underset{(0.37)}{0.818} - \underset{(0.01)}{0.026}\varepsilon^m_{t-1} + \underset{(0.10)}{0.451}\Delta m_{t-1} - \underset{(0.10)}{0.298}\Delta m_{t-2} + \underset{(0.0005)}{0.001}\Delta y_{t-1} + \underset{(0.01)}{0.002}\Delta R_t - \underset{(0.004)}{0.007}\Delta\sigma_{2t-2}$ $\bar{R}^2 = 0.3$, SEE = 0.02, F = 8.4, DW = 1.9, $Z_1(4) = 3.4$, $Z_2(1) = 0.04$, $Z_3(2) = 2.1$, $Z_4(1) = 1.4$
Mexico	$\Delta m_t = \underset{(0.05)}{0.169} - \underset{(0.04)}{0.148}\varepsilon^m_{t-1} + \underset{(0.10)}{0.397}\Delta m_{t-1} - \underset{(0.10)}{0.299}\Delta m_{t-2} + \underset{(0.002)}{0.005}\Delta y_t + \underset{(0.002)}{0.006}\Delta y_{t-2}$ $+ \underset{(0.002)}{0.006}\Delta y_{t-3} - \underset{(0.0005)}{0.001}\Delta^2 R_t - \underset{(0.02)}{0.051}\Delta\sigma_{1t} - \underset{(0.02)}{0.023}\Delta\sigma_{1t-4}$ $\bar{R}^2 = 0.5$, SEE = 0.07, F = 7.9, DW = 2.0, $Z_1(4) = 5.1$, $Z_2(1) = 0.2$, $Z_3(2) = 4.2$, $Z_4(1) = 24.9$
Philippines	$\Delta m_t = \underset{(0.01)}{0.012} - \underset{(0.03)}{0.073}\varepsilon^m_{t-1} + \underset{(0.11)}{0.33}\Delta m_{t-4} + \underset{(0.0008)}{0.003}\Delta y_{t-4} - \underset{(0.0015)}{0.003}\Delta R_{t-1} - \underset{(0.01)}{0.022}\Delta\sigma_{1t-1} - \underset{(0.01)}{0.039}\Delta\sigma_{1t-2}$ $\bar{R}^2 = 0.4$, SEE = 0.05, F = 9.9, DW = 2.3, $Z_1(4) = 6.6$, $Z_2(1) = 0.4$, $Z_3(2) = 68.1$, $Z_4(1) = 2.3$ $Z_5(1) = 0.2$
Singapore	$\Delta m_t = \underset{(0.11)}{-0.176} - \underset{(0.01)}{0.030}\varepsilon^m_{t-1} + \underset{(0.08)}{0.180}\Delta m_{t-2} + \underset{(0.05)}{0.193}\Delta^2 y_t + \underset{(0.04)}{0.269}\Delta^2 y_{t-1}$ $- \underset{(0.02)}{0.003}\Delta R_{t-3} - \underset{(0.003)}{0.006}\Delta^2\sigma_{1t-2} - \underset{(0.003)}{0.009}\Delta^2\sigma_{1t-3}$ $\bar{R}^2 = 0.0$, SEE = 0.02, F = 17.9, DW = 1.7, $Z_1(4) = 5.6$, $Z_2(1) = 0.2$, $Z_3(2) = 1.5$, $Z_4(1) = 0.4$
South Africa	$\Delta m_t = \underset{(0.18)}{0.001} - \underset{(0.05)}{0.195}\varepsilon^m_{t-1} + \underset{(0.11)}{0.179}\Delta m_{t-2} + \underset{(0.11)}{0.185}\Delta m_{t-3} + \underset{(0.29)}{0.395}\Delta y_{t-1} - \underset{(0.001)}{0.002}\Delta^2 R_t - \underset{(0.01)}{0.016}\Delta\sigma_{2t-2}$ $\bar{R}^2 = 0.2$, SEE = 0.03, F = 3.3, DW = 2.0, $Z_1(4) = 1.7$, $Z_2(1) = 0.3$, $Z_3(2) = 0.7$, $Z_4(1) = 2.5$
Thailand	$\Delta m_t = \underset{(0.004)}{0.007} - \underset{(0.02)}{0.053}\varepsilon^m_{t-1} + \underset{(0.08)}{0.162}\Delta m_{t-1} + \underset{(0.08)}{0.421}\Delta m_{t-4} + \underset{(0.05)}{-0.073}\Delta^3 y_{t-1} - \underset{(0.0009)}{0.001}\Delta R_{t-2} - \underset{(0.002)}{0.005}\Delta^2\sigma_{1t}$ $\bar{R}^2 = 0.6$, SEE = 0.02, F = 13.0, DW = 1.9, $Z_1(4) = 2.2$, $Z_2(1) = 1.3$, $Z_3(2) = 0.7$, $Z_4(1) = 0.6$

The number in parentheses beneath the coefficient estimates are the standard errors. $\bar{R}^2$ is the adjusted coefficient of the multiple determination, SEE is equation standard error, DW is Durbin–Watson statistic, and Z_1 is Breusch–Godfrey chi – squared statistic for residual correlation of up to fourth order. Z_2 is RESET statistic with $\chi^2(1)$. Z_3 is Bera and Jarque normality test with $\chi^2(2)$. Z_4 is the test for heteroskedasticity with $\chi^2(1)$.

Several features of the results in Table 5 deserve mention. First, the coefficient of the error-correction term is statistically significant in each of the eight cases and is always negative. These findings support the validity of an equilibrium relationship among the variables in the cointegrating equation. This implies that overlooking the cointegratedness of the variables would have introduced misspecification in the underlying dynamic structure.

Second, the change in real money balance per quarter that is attributed to the disequilibrium between the actual and equilibrium levels is measured by the absolute values of the error-correction term of each equation. There is considerable intercountry variation in the adjustment speed to the last period's disequilibrium, with South Africa having the largest and Malaysia, the smallest. This implies that the adjustment of real balance to changes in the regressors may take about 5 quarters in South Africa to slightly more than 38 quarters in Malaysia. The results point to the existence of market forces in the money market that operate to restore long-run equilibrium after a short-run disturbance.

Finally, the dynamics of the equations show that changes in real income, lagged changes in real balance, interest rates, and inflation-rate volatility have significant short-run effects on real balances. These results are summarized in Table 6.

In almost all cases, the mean-time lags of real income and inflation-rate volatility are smaller than those of the interest rate. This suggests, for example, that there is a faster response of real balance to real income changes than to interest rate changes. In sum, the empirical results indicates that inflation-rate volatility, ignored in several previous LDCs money demand studies, has a short-run effect on real balance, in addition to its long-run effect. Thus, neglect of such a variable can produce biased results.

4. Summary and concluding remarks

As set out in the introduction, the primary aim of this paper was to investigate empirically the impact of inflation-rate volatility on money demand of eight LDCs by using new time series techniques. Also examined is the empirical importance of including the interest rate variable as the opportunity cost of holding money in LDCs. Previous time series studies that have been concerned with the effects of inflation-rate volatility on demand for money in LDCs have used Koyck-type models and have, in contrast to those focusing on DCs, omitted the interest rate variable from their money demand models.

In this paper, the basis of our analysis is a money demand function, estimated on quarterly data (1973:2–1999:4) for each of the eight countries. In the specific function considered, real money balances depend on real income, interest rate, and inflation-rate volatility. Tests were first conducted without the inflation-rate volatility variable in the money demand function to evaluate whether such a variable belonged in the money demand function of LDCs. The estimation and evaluation were done within the framework of cointegration and error-correction modelling. Such a framework allows attention to be paid to the "spurious regression problem" and is econometrically superior to the Koyck scheme. Since it was found that inflation-rate volatility plays a statistically significant role in the demand for money in LDCs, Hansen's parameter constancy tests were used to assess the constancy of a long-run, equilibrium relationship linking real money balances, real income, interest rate, and inflation-

rate volatility. This is essential for appropriate policy conclusions to be inferred from the estimated results. Lastly, the short-run dynamics of the model are evaluated by using various measures, namely, the speed of adjustment and the mean time lag for adjustment of real balances to changes in the explanatory variables. A number of key results have emerged from our econometric work.

First, the cointegration analyses reveal that real money balances, real income, the interest rate, and inflation-rate volatility are, more often than not, both necessary and sufficient to pin down the equilibrium money demand of these LDCs. Our results concerning the effects of inflation-rate volatility on real money balances suggest that there is a negative and statistically significant long-run relationship between real money balances and inflation-rate volatility in each of the eight LDCs.

Furthermore, and perhaps more substantive, evidence for the contribution of the inflation-rate volatility is that its inclusion appears necessary in most of the countries for the estimated money demand equation to exhibit the desired property of a constant parameter cointegration. Thus, neglect of such a variable can produce biased results. It can be deduced that traditional money demand studies for other LDC economies that do not include a variable representing the influence of inflation risks are potentially misspecified. In addition, the results suggest that monetary policy actions aimed at stabilizing the domestic economy can generate results that are, at best, uncertain, if policymakers ignore the stability, as well as the level, of inflation.

Monetary policy in these economies should be conducted through a formal process of inflation targeting. Targeting consumer price inflation can benefit these economies in many ways by providing a coordination device for inflation expectations as well as a yardstick of accountability for central banks. An important element of such a policy would be to have an explicit numerical target range for inflation that allows consumer price inflation not to be targeted closely. That is, the central bank's strategy should be to eliminate deviations between inflation and its target rate gradually and not to keep inflation at target period by period. As long as inflation remains within the stated range, the central bank can pursue other policy objectives (e.g., the rate of economic growth, unemployment, etc.). However, the central bank must make the inflation target its overriding objective whenever inflation attempts to break out of its permissible range.

The empirical results, derived as they are from the LDCs' data, are consistent with the theoretical considerations discussed in Section 2 and also confirm previous research done for DCs (e.g., Choudhry, 1999, p. 636), which suggests that inflation-rate volatility has a significant negative impact on real money balances. Our findings also corroborate the evidence in Arize and Malindretos (2000) that inflation-rate volatility has both long- and short-run negative effects on the demand for real money balances. In addition, the findings of this paper agree, at least in spirit, with those of Blejer (1979) for Argentina, Brazil, and Chile.

Second, it is found that for most of the cases, the interest rate variable, ignored in several previous LDCs money demand studies, plays a statistically significant role as the relevant opportunity cost variable in explaining the long-run demand for real money balances. Our results demonstrate that variation in interest rates in LDCs has clearly been sufficient to justify use as the opportunity cost variable in the money demand function. Furthermore, our results imply that any monetary policy that affects interest rates may affect the demand for real balances.

Finally, our results suggest that real income has a positive and statistically significant effect on money demand behavior in LDCs. In almost all cases, the income elasticity is greater than unity, implying the absence of economies of scale in money holdings in LDCs. Our income elasticities appear to be consistent with previous work; see, for example, Aghevli et al. (1979), who report 1.85, 1.65, 1.54, 1.33, and 1.49 for Indonesia, Malaysia, the Philippines, Singapore, and Thailand, respectively. Information about the size of real income elasticity is important for appropriate monetary expansion. Our results suggest that monetary policies should be directed towards increasing real income to enhance monetary accumulation. They suggest further that the high income elasticity of around two in Indonesia, Mexico, and the Philippines could be evidence that the authorities should do more to improve public confidence in the available domestic financial assets.

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Appendix A. Data, definitions and sources

This appendix describes the raw data, sources and construction of variables used for empirical tests. All data were obtained from the IMF's IFS CD-ROM (2000), the Supplement on Money and the IFS published monthly by IMF.

m: Real money balances: This variable is defined as the natural logarithm of the ratio of (M2/*P*), where M2 is the sum of IFS, line numbers 34 and 35. The raw money stocks are end-of-period monthly observations. The quarterly figures are computed as weighted average from monthly data; for example, the first quarter is 1/6 December + 1/3 Januray + 1/3 February + 1/6 March. *P* is the consumer price index obtained from IFS, line number 64 (1990 = 100).

R: Domestic interest rate: For Kenya, Mexico, the Philippines, and South Africa, the interest rate is proxied by the treasury bill rate, that is, IFS, line number 60c. The money market rate (IFS, line 60b) is used to proxy interest rate for Indonesia, Malaysia, Singapore, and Thailand.

y: Real GDP: This variable is taken from IFS, line 99b.p or 99b.r and 99bvp. Annual data are available for each of the eight countries, whereas quarterly data are available for six countries: Malaysia (1988:1–1999:4), Mexico (1980:1–1999:4), Thailand (1996:1–1999:4), the Philippines (1981:1 1999:4), Singapore (1984:3–1999:4), and South Africa (1973:2–1998:1). There are no quarterly data for Indonesia and Kenya. Against this background, it was necessary to interpolate the annual series to obtain quarterly series for Indonesia, Kenya and the six countries mentioned above which have quarterly observations for some of sample periods. For example, the data for Malaysia for the sample period 1973:2 through 1987:4, were obtained from annual data by applying a quadratic interpolation method outlined and used in Bergstrom (1990). The procedure is presented below, and it has been used by other researchers (e.g., Arize, Osang, & Slottje, 2000).

Let x_{t-1}, x_t, and x_{t+1} be three successive annual observations of variable x. If the quadratic function that passes through the three points is such that

$$\int_0^1 (as^2 + bs + c)\mathrm{d}s = x_{t-1} \tag{A1}$$

$$\int_1^2 (as^2 + bs + c)\mathrm{d}s = x_t \tag{A2}$$

$$\int_2^3 (as^2 + bs + c)\mathrm{d}s = x_{t+1} \tag{A3}$$

then, integrating and solving for a, b, and c gives

$$a = 0.5x_{t-1} - x_t + 0.5x_{t-1} \tag{A4}$$

$$b = -2x_{t-1} + 3x_t - x_{t-1} \tag{A5}$$

$$c = 1.8333x_{t-1} - 1.1666x_t + 0.333x_{t-1} \tag{A6}$$

The four quarterly figures within a given year can now be interpolated, respectively, using

$$\int_1^{1.25} (as^2 + bs + c)\mathrm{d}s = 0.05468x_{t-1} + 0.23438X_t - 0.039067x_{t+1} \tag{A7}$$

$$\int_{1.25}^{1.50} (as^2 + bs + c)\mathrm{d}s = 0.00781x_{t-1} + 0.26563X_t - 0.02344x_{t+1} \tag{A8}$$

$$\int_{1.50}^{1.75} (as^2 + bs + c)\mathrm{d}s = -0.02344x_{t-1} + 0.26562X_t + 0.00781x_{t+1} \tag{A9}$$

$$\int_{1.75}^{2} (as^2 + bs + c)\mathrm{d}s = -0.0391x_{t-1} + 0.23437X_t + 0.05469x_{t+1} \tag{A10}$$

Multiplication by four expresses the series at annual rates.

Appendix B. Unit root test results

Countries	Method	m	y	r	σ_1	σ_2	Countries	m	y	r	σ_1	σ_2
Indonesia	ADF H_0: $I(1)$	−2.29	−1.46	−0.49	−0.49	−2.11	Philippines	−1.78	−3.38	−2.44	−2.79	−3.16
	$G(p,q)$ H_0: $I(0)$	15.48[0.03]	13.40[0.06]	29.54[0.00]	18.08[0.01]	19.43 [0.01]		12.46[0.02]	15.14[0.03]	8.06[0.05]	13.59[0.06]	13.77[0.06]
	Johansen (σ_1)	94.06[0.00]	93.45[0.00]	10.49[0.01]	58.28[0.00]	–		32.26[0.00]	30.50[0.00]	22.82[0.00]	15.50[0.00]	–
	Johansen (σ_2)	44.81[0.00]	44.30[0.00]	4.67[0.19]	–	30.42 [0.00]		25.18[0.00]	22.63[0.00]	13.23[0.00]	–	12.16[0.00]
Kenya	ADF H_0: $I(1)$	−2.28	−2.15	−2.42	−2.13	−2.33	South Africa	−2.80	−0.09	−2.27	−2.99	−2.99
	$G(p,q)$ H_0: $I(0)$	4.96 [0.03]	3.50[0.06]	3.55[0.06]	4.96[0.03]	2.74 [0.07]		6.53[0.04]	17.78[0.01]	18.63[0.01]	17.83[0.01]	21.21[0.00]
	Johansen (σ_1)	35.05[0.00]	28.82[0.00]	27.21[0.00]	24.22[0.00]	–		36.44[0.00]	37.15[0.00]	31.36[0.00]	10.18[0.01]	–
	Johansen (σ_2)	28.99[0.00]	21.49[0.00]	14.92[0.00]	–	26.14 [0.00]		29.63[0.00]	43.83[0.00]	30.67[0.00]	–	35.11[0.00]
Malaysia	ADF H_0: $I(1)$	−1.01	−2.69	−2.89	−3.67	−3.41	Singapore	−2.79	−3.01	−3.22	−1.96	−1.05
	$G(p,q)$ H_0: $I(0)$	18.08[0.01]	9.97[0.01]	24.45[0.00]	3.90[0.42]	13.83 [0.05]		23.59 [0.00]	15.26[0.03]	16.43[0.02]	6.58[0.04]	11.37[0.00]
	Johansen (σ_1)	34.71[0.00]	34.66[0.00]	22.44[0.00]	2.49[0.47]	–		30.35[0.00]	30.82[0.00]	13.72[0.00]	15.18[0.00]	–
	Johansen (σ_2)	32.11[0.00]	31.94[0.00]	16.28[0.00]	–	20.29 [0.00]		39.85[0.00]	39.40[0.00]	34.97[0.00]	–	34.29[0.00]
Mexico	ADF H_0: $I(1)$	−2.33	−2.61	−2.29	−2.85	−3.12	Thailand	−2.08	−1.98	−2.93	−2.62	−1.80
	$G(p,q)$ H_0: $I(0)$	10.09[0.04]	13.57[0.01]	6.07[0.02]	9.46[0.05]	12.57 [0.01]		8.20[0.04]	7.27[0.07]	9.42[0.02]	8.08[0.04]	7.81[0.05]
	Johansen (σ_1)	42.03[0.00]	42.26[0.00]	30.89[0.00]	32.04[0.00]	–		42.40[0.00]	41.92[0.00]	24.39[0.00]	2.15[0.54]	–
	Johansen (σ_2)	31.07[0.00]	31.27[0.00]	21.01[0.00]	–	28.67 [0.00]		32.62[0.00]	32.19[0.00]	27.25[0.00]	–	13.35[0.00]

Critical value for ADF test at 5% level is −3.46, and Johansen test has $\chi^2(3) = 7.82$ at 5% level.

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