THE STABILITY OF MONEY DEMAND FUNCTIONS: MÉXICO, 1986 – 2002

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Abstract – Over the years 1989-1993, Mexico underwent financial liberalization, privatization of commercial banks, and opening to foreign markets. Using Mexican monthly data over 1986-2002, this paper studies several money demand specifications. The benchmark vectors have currency strongly cointegrated with income (from 0.75 to 1.15), interest rate (from –0.03 to -0.05), and the inflation rate (about –0.29), while specifications with M1 and M2 fail. A coefficient of 0.59 of the *currency-money ratio* (CUR/M1) would suggest that when the ratio falls (financial innovations rise), the demand for currency falls, but the evidence is weak. Several stability tests are performed around the 1991 episode of "M1 explosion" and around the 1994 currency crisis. These do not detect structural change in the latter for the traditional models but do so in the former for all models. The traditional models are thus not worse than the financial innovations one, perhaps because the financial innovations term is contaminated by factors related to the money supply process.

Keywords: Money demand, financial innovations, currency to money ratio, cointegration, error-correction model.

JEL Classification: E41, E44.

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INTRODUCTION

An important building block of many economic models is the existence of a *stable money demand function*. Following Friedman (1956) and Cagan (1956), empirical works have measured real money demand balances relative to income and the various opportunity costs of holding money, such as: interest rates, the inflation rate, and the expected depreciation rate of a currency. Given the conflicting evidence, the survey by Judd and Scadding (1982) mentions *financial innovations* as the possible cause of the instability in U.S. money demand after 1973. The approach through evolving financial innovations in the economy has been applied to developed economies by Siklos (1993) and, more recently, by Choudhry (2002) to the U.S. Applications to less developed countries include the studies by Arrau and De Gregorio (1993), Arrau et al. (1995), and Melnick (1995).

This vast empirical literature is not devoid of theoretical foundations. Arrau et al. (1995) develop a utility maximizing model with reductions in transaction costs, while Hamilton (1989) explores a change in the technology parameter that causes credit goods to become cheaper to produce relative to cash goods. This encourages substitution away from cash and towards new means of payments. Similarly, Glennon and Lane (1996) provide a theoretical framework to study a couple of developments in the U.S.: first, the introduction of checking account deposits that pay interest rates and, second, the elimination of interest rate restrictions on deposits.

This paper also pursues mispecifications in traditional money demand equations due to *financial innovations*. We verify the validity of such hypothesis for the Mexican economy along the period 1986:1-2002:9 under monthly data. The work seems to be more complex in an economy like Mexico, since there is no obvious time of including new checking accounts, such as the NOW and Super-NOW accounts that happened in the U.S. and are reviewed by Choudhry (2002) and Glennon and Lane (1996). Rather, financial reforms are less visible and more gradual in Mexico, as can be seen in the study by Jallath and Negrín (2001) on the incorporation of other means of payments than cash. Instead of an event like the introduction of new financial instruments, the Mexican economy has been, since the late 80s, subject to profound transformations in its financial system. Although elaborated under the De La Madrid administration (1982-1988), the financial reforms were adopted in succession after the inauguration of the Salinas term (1988-1994). The various stages of the reforms (liberalization, legal reforms, privatization of commercial banks, and integration into NAFTA) are documented in the local literature by, for example, Girón (1994).

We take into account such period of substantial reforms and estimate various money demand functions under traditional and alternative (augmented by financial innovations) specifications. Dealing with financial innovations, one has to face inevitably the institutional background. In the second half of 1991, for example, M1 surged in Mexico, mostly led by a sudden rise in demand deposits. This could reflect either the high capital inflows due to a stable economy or, more fundamentally, a shift into interest-bearing assets, as scrutinized in the 1991 Annual Report of the Bank of Mexico. Unfortunately, the econometric techniques with monthly data are not able to distinguish between these two hypotheses. We do test, however, whether the stable functional relationship in money demand holds under financial sophistication and fundamental changes in the economy. It turns out that it does but only in a very narrow sense of money and with only marginal statistic significance.

We use in this paper the ratio of currency to money (CUR/M1) as proxy of *financial innovations*. This happens to mirror the demand for money in Mexico since the ratio rises as demand for money falls and later fall as money demand rises (from 1990) with financial sophistication and price stabilization. As CUR/M1 goes down, less currency is used, and more demand deposits (checking accounts) are employed in transactions. Changes in the composition of M1 should reflect financial innovations since the ratio of currency to money "mirrors the spread of commercial banking" [Siklos (1993)].

The non-stationarity of the data leads to cointegration techniques, of which a partial list to the studies of money demand contains the works by Stock and Watson (1993), Hoffman et al. (1995), Haug and Lucas (1996), and Muscatelli and Spinelli (2000). We find partial support for the CUR/M1 ratio appearing statistically significant in the cointegrating vector when currency is the monetary aggregate used. A positive long-run coefficient of the currency-money ratio on money demand suggests that when financial innovations rise (CUR/M1 fall), demand for money falls. Conversely, when CUR/M1 rises (financial innovations fall) due to lower banking penetration, for example, the demand for money rises. Given the existence of the long-run cointegration results, error correction models (ECM) are applied next and imply a low adjustment to equilibrium. The degree of short-run adjustment varies between 6% and 13% of the deviation eliminated after 1 month in the model with currency as the relevant monetary variable.

We also employ several econometric techniques in order to verify the stability content of the three currency specifications: with only interest rates and income; with these augmented by the inflation rate; and with the currency to money ratio in addition to interest rates and income. Several stability tests are performed around the 1991 episode of "M1 explosion" and around the 1994 currency crisis. These do not detect structural change in the latter but do carry instability in the former. The tests also show that the traditional models are not worse than the financial innovations one, perhaps because the financial innovations term is contaminated by factors related to the money supply process as shown by Hamilton (1989).

Employing a ratio of monetary aggregates, this research complements studies on Mexican data that employ trends as proxies for financial sophistication. Studies such as Arrau and De Gregorio (1993) and Arrau et al. (1995), for example, either do not consider the early 1990s period of financial reforms or are plagued by small sample bias (38 observations). Similarly, Thornton (1996), Galindo (2000), Garcés Díaz (2002), and Khamis and Leone (2001) present cointegration analyses on Mexican data without financial innovations. The results in this paper are broadly consistent with Khamis and Leone (2001) in the sense of currency being the monetary aggregate found to be more stable, relative to broader aggregates. It differs from theirs in analyzing explicitly the hypothesis of financial innovations.

The paper is composed of five sections: section 2 discusses the three versions of demand for money, section 3 summarizes the data and features of the financial reforms in Mexico, while section 4 presents the estimations. Section 5 reviews the main results of the work and provides extensions for further research.

2. METHODOLOGY

The *standard money demand function* estimated in empirical studies is based on a log-linear version of Cagan's (1956) classical analysis of hyperinflation. It usually captures long-run movements in real income (y) and nominal interest rates (i):

$$m_t = c + \alpha_1 y_t + \alpha_2 i_t + \upsilon_t \qquad (1),$$

where: c is a constant, m are real money balances (money supply stock deflated by the price level), υ_t is a white-noise error normally distributed with variance σ^2, and the variables are in logarithms, except for the interest rate. The coefficients α_1 and α_2 are interpreted as the income elasticity of money and the interest semi-elasticity, respectively, whose signs are expected to be positive and close to one for α_1 and negative for α_2. For recent estimation of (1) for Canada under cointegration methods, see Haug and Lucas (1996).

The *augmented version of the money demand function* has the inflation rate (Δp) as an additional variable in (2) below, taking into account the different opportunity cost of holding money:

$$m_t = c + \beta_1 y_t + \beta_2 i_t + \beta_3 \Delta p_t + \omega_t \qquad (2),$$

where: β_1 and β_2 are as expected before, β_3 is expected to be < 0, and ω_t is the white noise error. Equation (2) has been estimated often in studies of high-inflation economies. Examples are Rossi (1989) for Brazil; Blanco and Garber (1986), Rogers (1992), and Goldberg (1994) for Mexico; Melnick (1995) for Israel; and Muscatelli and Spinelli (2000) for Italy. The inflation rate may appear in studies of industrial countries as well, as in: Cooley and LeRoy (1981) and Hendry and Ericsson (1991).

Apart from "spurious regressions" that affect some of the earlier studies, one possible explanation for mispecifications in either (1) or (2) is related to their lack of ability in verifying forces in the real economy towards sophistication of financial markets. A burgeoning literature postulates mispecifications in (1) and (2) due to financial innovations and incorporate the alternative money demand function as follows:

$$m_t = c + \alpha_1 y_t + \alpha_2 i_t + \alpha_3 f_t + \varepsilon_t \qquad (3),$$

where ε_t and ι_t are white-noise errors and ϕ_t captures financial innovations. Arrau et al. (1995) derive (3) from an utility maximizing model with reductions in transaction costs. Hamilton (1989) justifies α_3 after exploring a utility maximization model with 2 goods for an exogenous change in the payments technology. A fall in such parameter causes credit goods to become cheaper to produce relative to cash goods, which encourages substitution away from cash and towards new means of payments. The series f_t is statistically conceived as a *non-stationary process* that changes permanently the demand for money. In practice, however, what composes f_t is debatable and varies considerably across studies, such as: Arrau and De Gregorio (1995), Siklos (1993), and Melnick (1995).

As in Siklos (1993), this paper employs *monetary indicators* for f_t. We concentrate on factors that measure shifts in currency within demand deposits, which are a component of M1. As conjectured two decades ago, "the instability in money demand appeared to be in its demand deposit component" [Judd and Scadding, 1982, p. 998].

We focus on the *currency-money ratio* (CUR/M1) that measures the proportion of currency with respect to the M1 aggregate: as CUR/M1 goes down, less currency is used, and more demand deposits (checking accounts) are utilized for transactions. Changes in the composition of M1 should reflect financial innovations since the ratio of currency to money "mirrors the spread of commercial banking" [Siklos (1993)]. As banking expands, in an U-shaped curve, velocity of money falls at first and later rises with financial sophistication and less need for money. Changes in the velocity of money are thus positively related to changes in the currency-money ratio.

Another possible (monetary) indicator for f_t is the *M2/M1 ratio* that captures in the long run the growth of a systematic increase in broad monetary aggregates with respect to a narrower money stock. Ireland (1994) provides a model consistent with such trend for the U. S.: always up from around 2 in 1952 to around 4 in 1992, except for a fall in the early 90s after disinflation and regulation that is also observed in other industrial countries.[1] However, the drawback with the usage of M2/M1 in the money demand context is that it is not clear the expected value of the coefficients in the context of money demand. For instance, as M2/M1 moves up, the demand for money balances (with M2 as the money stock) has to move positively. But that contradicts the hypothesis that more sophistication implies less need for money. Similarly, when M1 is the money stock variable in the left hand side (LHS) instead, a fall in M2/M1 (due to a rising M1 relative to M2, for example) should make (M1/P) to grow according to the hypothesis but to fall algebraically.

Such problems do not distort our conjecture when either currency or M2 is the monetary aggregate. Under CUR/M1 as the indicator of financial innovations, the financial sophistication hypothesis implies that increases in the ratio CUR/M1 (less important is the banking sector in the overall economy) cause demand for money (measured by CUR or M2) to move up.

A final caveat deals with an econometric point. Hamilton (1989) shows that if there are significant disturbances to the money demand equation, a variable such as CUR/M1 might appear statistically significant in (3), even though it has nothing to do with money demand. To see this point, Hamilton (1989) explores when to test the hypothesis of $\alpha_3 = 0$ in (3) is a test of adequacy of (1) as a theory of money demand. Hamilton (1989) shows that the vector of parameters to be

[1] According to Ireland (1994), the M2/M1 ratio rose from 2 to 3.2 in Japan between 1953 and 1970; from 1.7 to 3.4 in Germany during the same period; from 1.3 to 2.9 in Korea between 1960 and 1970; and from 1.4 to 2.9 in Taiwan between 1952 and 1970.

estimated converges in probability to $Z_t = (\gamma\alpha_1, -\alpha_2\gamma, 1-\gamma)$, where γ lies between 0 and 1 and is a ratio of disturbances. In order for the regression estimate to give appropriate inference about the true money demand parameters, it must be that γapproaches 1 and then $Z_t = (\alpha_1, -\alpha_2, 0)$.

3. THE DATA

The data used in this study come from the *Bank of Mexico Information System*, covering the monthly period 1986:1-2002:9. They are available online at http://www.banxico.org.mx. The real money stock measures are: M1 (currency in circulation + demand deposits); and M2 (= M1 + short-term bank instruments), deflated by the price level (measured by consumer price index), not seasonally adjusted. Within the short-term bank instruments are those with maturity of up to 1 year and those called bank acceptances. The narrower definition of money we consider is part of M1, the currency in circulation, abbreviated in this paper by CUR.

The first year of observations, 1986, lies within the Miguel De La Madrid administration, which was marked by attempts to launch structural reform programs and negotiation of the external debt under stagflation. Stabilization at moderate inflation and real growth was achieved years later. The *process of financial reform* planned under the De La Madrid administration was put into practice from end-1988 onwards, already at the outset of the Salinas term. The stages were, according to Girón (1994): i) liberalization of financial markets at the end of 1988 and beginning of 1989 with free interest rates, competition, and authorization to pay interest on checking accounts; ii) legal reforms and deregulation towards the end of 1989; iii) privatization of commercial banks between June 1991 and August 1992; and iv) gradual openness to foreign capital with the signature of NAFTA by the end of 1993, coupled with more autonomy to the central bank. By 1995, the financial reform process was completed and the exchange rate regime had also changed into a (managed) float, following the currency crisis of December of 1994.

Figure 1 displays the demands for real balances depending on the (real) monetary aggregate: CUR, M1 and M2. It also contains the monthly inflation rate (dP), the monthly 1-month interest rate (i), and the index of industrial production (IP). The end of 1994 is a landmark in Mexican financial history with the currency crisis that erupted at the beginning of the Zedillo administration. This can be seen in figure 1 as the sudden rise in inflation and in nominal rates and the subsequent fall in output in 1995.

With respect to the monetary aggregates, it can be seen that M1 and M2 both fall by the end of 1994. In addition, both series rise after 1990 with price stabilization. Figure 1 also shows real currency (CUR), which displays a more seasonal pattern at the end of each year than the other two monetary indicators. This is due to the end of year bonus payments to Mexican workers (*aguinaldos*). The econometric work in the next section removes the seasonal pattern of each series by employing monthly dummies.

Our two measures of financial innovations are calculated from the monetary data and are reproduced in figure 2. While the ratio M2/M1 measures the extent of growth in the substitutes of money with respect to its narrowest aggregate, CUR/M1 (or simply f1) captures a decreasing trend in currency usage under innovations in the financial system. In the second half of 1991, for example, M1 surged in Mexico, mostly led by a sudden rise in demand deposits. This could reflect either the high capital inflows due to an increasingly stable economy or a shift into interest-bearing assets. See the discussion in Banco de México (1991). The monthly (not annualized) interest rate (i) is the (nominal) yield on one-month CETES instruments. The transactions variable is the index of industrial production (y), which has been extensively used instead of monthly personal income or monthly GNP data for developing economies by Rogers (1992), Goldberg (1994), and Arrau et al. (1995). The inflation rate is defined as the rate of change of the consumer price index (CPI). Except for the interest rate and for inflation rate, which is already a rate of growth, money and income variables are transformed into logarithms.

The fall in the currency-money ratio during the second half of 1991 is consistent with financial sophistication due to a higher penetration of commercial banking. The CUR/M2 ratio, not reported, is flat around the 0.10 level. The other proxy of financial innovations (M2/M1), reflects a similar pattern. As explained in the previous section, however, in the context of money demand, CUR/M1 should be a better indicator than M2/M1.

4. RESULTS AND DISCUSSION

4.1 Preliminaries

Consider first the unit root evidence in table 1 for the period 1986:01 to 2002:09 in Mexico. Two types of unit root tests are reported. The first is the usual Augmented Dickey-Fuller (ADF) test, in which we employ the sequential approach suggested by Ng and Perron (1995) for choosing the number of lags in the estimating equations that appear in parenthesis close to the ADF figure. We also employ the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) (1992) test, in which the null hypothesis is based on a stationary series.[2]

The ADF tests in table 1 show that, for all series, the unit roots null is not rejected in levels at the 5% level. In order to address the low power of unit root tests, ADF tests are reinforced by complementary KPSS tests. For the series in levels, the KPSS tests rejects the null of stationarity in all cases, which matches the results obtained by the ADF procedure. For example, in first differences, the only weak rejection of the null is for the currency to money ratio (f1), which is rejected at 10% according to the ADF test. However, the KPSS statistic does not reject the null in this case, which suggests the series is stationary in first differences. Documenting non-stationarity for all series in levels and stationarity in differences, the two unit root tests show consistency across tests. The series can be thus

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[2] For the KPSS test, table 1 contains values calculated from lag-length truncation parameter equal to 4. The Bartlett kernel is used for spectral estimation method. As a robustness check, we choose different choices of lag-length selection in the KPSS test using lag-lengths (L) of 0, 4 and 13 as computed from the formula lag = 0, or lag = integer $[L(T/100)^{.25}]$. These three choices match precisely the original choices of KPSS for lag-length, in which Monte Carlo simulations show that size distortions appear as L increases [KPSS (1992), p. 170]. The results using lag = 0 and the ones reported in table 1 are broadly consistent. If we did not correct for error autocorrelation (and had taken lag = 0), we would be assuming that iid errors are plausible under the null (of stationarity around a deterministic trend), which is not easily justified.

classified as integrated of order 1: I(1). This implies, if f1 is a legitimate variable in the right hand side (RHS) of the demand for money function, *permanent effects on the demand for money*.

4.2 The Long-Run

In order to proceed with the estimation by the Johansen cointegration procedure, the appropriate choice of the lag-length in the VAR is critical. Mechanical application of a single information criterion does not usually provide the best choice in terms of efficiency of the estimators. In fact, as shown by Stock and Watson (1993), results of cointegration in systems are highly sensitive to the choice of lags, the treatment of deterministic terms (constant, trends and seasonals), and stability of parameters.

Table 2 contains Johansen cointegration tests Mexico of equations (1) to (3) under the assumption of linear deterministic trend in the data, and that all series have stochastic trends. The VAR lag length is chosen by a search on a maximal order of $k = 18$ based on the highest degree of coincidence of all statistic criteria involved (likelihood ratio, final prediction error, Akaike information criterion, Schwarz information criterion, and Hannan-Quinn information criterion). If there is disagreement in the appropriate lag-length, we estimate two or more VARs with different lag lengths and apply the LM-test of serial correlation. The Breusch-Godfrey serial correlation LM test has under the null the hypothesis of no serial correlation up to lag order p. The column at the extreme right of the table contains this statistic, which does not reject the null in any case.

The VAR estimates upon which the Johansen methodology is based include seasonal dummies since the data originally are not seasonally adjusted. The coefficients of the seasonal dummies, however, are not reported in the tables. Table 2 shows that two of the specifications with currency as the monetary aggregate clearly reject the null of no cointegration. In particular, (CUR, i, y) and (CUR, i, y, ΔP) are the specifications in which both trace and maximum eigenvalue tests reject the null of no cointegration. For (CUR, i, y, f1), however, the evidence of cointegration is weaker since only in the maximal eigenvalue test one rejects the null at 5%. We report two vectors in each case: the unrestricted one and the restricted one with the interest rate set to zero.

The more standard specification, associated with equation (1), has a 0.75 income elasticity and –0.05 interest semi-elasticity. Under currency as the monetary aggregate within the specification (CUR, i, y, ΔP), the income elasticity goes beyond unit to 1.15, the interest semi-elasticity is -0.03, and the inflation rate has a negative effect on money demand of –0.29. All three parameters are statistically significant and are in agreement with economic theory. Restricting the estimation to a zero interest rate coefficient, the specification (CUR, i, y, ΔP) yields income elasticity of 1.54 and a negative response to the inflation rate of -0.46, and both are also statistically significant.

Gauged by the maximal eigenvalue test of the Johansen procedure, the currency to money ratio contributes to achieving long-run equilibrium within the specification (CUR, i, y, f1). In this case, the income elasticity is 0.49, the interest semi-elasticity is -0.11, and the effect of financial innovations turns out to be less than proportional at 0.59. The latter, however, is only marginally statistically significant. A possible rationalization is that, as currency falls with respect to M1, as is the case in Mexico in the early 90s, demand for money should fall. This is consistent with a more widespread banking sector in the Mexican economy. Another possibility is that f1 is capturing factors that are indeed related to the money supply process as shown by Hamilton (1989) and discussed in section 2. The interest rate and income coefficients are statistically significant in the estimation according to the Johansen procedure. Restricting the estimation to a zero interest rate coefficient, the specification (CUR, i, y, f1) yields a close to unit income elasticity and not significant response to the currency to money ratio.

The specifications with M1 and M2 display much weaker long-run results. For M1, in all three specifications there is only partial evidence of cointegration. For example, equation (1) yields a negative interest rate coefficient of –0.26 and a close to zero income elasticity at 0.08. These results are unchanged with the addition of the inflation rate in the right hand side. Note, however, for the specification with M1 that a fall in the f1 figure due to higher M1 must imply lower demand for money and thus the expected coefficient is negative. This is observed in both unrestricted and restricted specifications.

In M1 estimations, the interest semi-elasticity is relatively high in absolute value for two specifications at about –0.26. For the estimation with M2, the interest semi-elasticity appears statistically significant with the wrong sign (positive) under (M2, i, y, f1). Restricting the vector for a zero interest rate coefficient, income elasticity is estimated at 1.86 and the financial innovations term is not statistically significant. Therefore, under M1 and M2, evidence is much weaker and the interest semi-elasticity may even be estimated with the wrong sign.

Note that the income elasticity greater than unity, observed under equation (2) for currency and under equations (1) and (3) for M2, implies a downward trend in velocity that is commonly found in other studies.[3] Under currency as the monetary aggregate, the statistical evidence is strongest. The estimate of a coefficient on currency to money ratio of 0.59 matches the expected sign, although one can not rule out the identification point raised by Hamilton (1989). Without Δp and with f1, the income elasticity is below one (0.49), the interest semi-elasticity is low but negative (-0.11), and the financial innovations parameter is marginally positive as expected.

Our findings suggest that, under a narrow monetary aggregate (currency in circulation), the standard explanatory variables in money demand all match the theoretical sign. Besides, the *currency-money ratio* has a natural interpretation in the context of money demand. As currency falls relative to M1 (financial innovations rise), the long-run value of money demand falls, which is consistent with the financial sophistication hypothesis. Money demand equations, measured

[3] Thornton (1996) employs the Johansen cointegration method on quarterly Mexican data from 1980 to 1994 and obtains relatively large income elasticities (1.1 and 2.0 for M1/P and M2/P, respectively) and interest semi-elasticities of –0.7 in the case of M1/P and –0.3 in the case of M2/P. Using annual Mexican data from 1979 to 1997, Galindo (2000) finds 1.16 or 1.09 for the income elasticity and –0.007 or –0.008 for the interest semi-elasticity, depending on the long-run specification. Estimates are much lower for the income elasticity (0.44) when the partial adjustment model is used and lagged real balances enter as independent variables in the right-hand side. Such is also the case in Ortiz (1983) who estimates by OLS, corrected by Cochrane-Orcutt, a barely significant 0.10 income elasticity (t-statistics of 1.69) for the period 1960-1979 under a partial adjustment specification.

by currency, are partially cointegrated with income, interest rate and the financial innovations term. This result is not obtained under higher monetary aggregates.

4.3. The Short-Run and Stability Issues

After obtaining the factors that affect long-term real currency demand, deviations from the long-run relationship can be captured by an error correction model (ECM). In the short run, deviations from the long-term equilibrium reflect shocks to any of the variables. The theorem in Engle and Granger (1987) establishes that if a stationary relationship between nonstationary variables exists, an error correction representation of the data must exist. It is typical to incorporate short-run dynamics by estimating the ECM in first differences. As in Khamis and Leone (2001), we estimate the ECM for the period 1986:1 to 2002:9 by including originally up to 13 lags of all variables in each system, together with the lagged ECM term and monthly dummy variables:

$$(m\text{-}p)_t = c + \alpha\, EC_{t-1} + \Sigma\beta_i\, MD_t + \Sigma\gamma_i\, \Delta(m\text{-}p)_{t-i} + \Sigma\theta_i\, \Delta i_{t-i} + \Sigma\kappa_i\, \Delta p_{t-i} + \Sigma\phi_i\, \Delta f1_{t-i} + \upsilon_t \qquad (4),$$

where: c is the constant term, MD are monthly dummies from January through November, α is the error correction term that are the residuals obtained by running each of the long-run models (1) to (3), and all other variables are as defined earlier. Based on sequential reduction in lag length, we obtain the final model for each specification, which are reported in table 3.

The α estimates reported in table 3 indicate the speed with which deviations from the cointegrating equilibrium equations are corrected. Across models, the degrees of adjustment are as follows: 13% of the deviations are eliminated within a month in case of model (1), 12% in model (2) and 6% in model (3). In all cases, the α estimates are statistically significant at 5% or less, which reinforces the long-run equilibrium found by the Johansen procedure. These rates of adjustment are lower than the 13 to 27% range documented by Thornton (1996) for the quarterly period 1980:1 – 1994:4, and in line with the 9% reported by Khamis and Leone (2001) for the monthly 1984:3 – 1997:6 period. Employing residual-based tests, we do reject the null of no-serial correlation according to the Breusch-Godfrey test for model (3) at 5%. Residuals appear to be normal, except for model 2, and there is no evidence of ARCH serial correlation in all models.

We investigate whether the residuals from the ECM estimations are well behaved. This follows Hamilton (1989) in checking if there are significant disturbances to the money demand equation as pointed out in section 2. We find that there are no substantial variations across the ECM estimates in figure 3 as far as the occasional and few spikes coincide. If anything, the residuals from model 3 are higher than those from model 1 and from model 2. This goes against the idea that introducing a financial innovations term leads to better money demand estimations.

A critical issue is the extent to which these estimates are stable over time. In Mexico, the months around the currency crisis of December of 1994 are natural breakpoints since the exchange rate regime collapsed and output fell substantially in 1995. See figure 1, particularly with the money and output figures. It is also interesting to check the stability of the models around 1991:10 when M1 "exploded" in Mexico, as can be seen in figure 2. We perform in table 4 Chow breakpoint tests of stability of coefficients on the ECM. The null hypothesis of no structural break is not rejected in the currency crisis breakpoint of 1994 for models 1 and 2. It is rejected strongly, however, for model 3. The episode of M1 explosion seems, however, to be more serious. There is always a rejection of stability of coefficients at standard significant levels for all 3 models.

Figure 4 contains series of recursive estimates of the most interesting parameters of the ECM in equation (4). Consider first the error correction term that converges to –0.13 in model 1,to -0.12 in model 2, and to –0.06 in model 3. The recursive parameters show that these estimates become more accurate as time passes. The inflation and lagged interest rate coefficients in figure 4 are negative throughout and also converge to their final value, reported in table 3.

At the bottom of table 4, we report the results of running the CUSUM and CUSUMQ stability tests as developed by Brown et al. (1975) and recently applied to money demand by Bahmani-Oskooee and Shin (2002). In order to implement it, the sample period is broken and the CUSUM and CUSUMQ statistics are updated recursively. If the plot of the CUSUM and CUSUMQ stays within 5% significance level, the coefficient estimates are said to be stable. The plots for model 1 in figure 5 show minor deviations from 2002 for the CUSUM test. For model 2 in figure 6 with the inflation rate, the CUSUM test shows violation of stability in 1994-1995 while the CUSUMQ is well behaved. For model 3 with the currency to money ratio, the results in figure 7 also differ across tests. Now the CUSUM shows visible violation of stability after 2000 although the CUSUMQ is well behaved. These tests show that the traditional and augmented models are not worse (in the sense of stability) than the one that captures financial innovations.

FINAL REMARKS

The methodology for checking long-run relationships in money demand takes particular care in considering Mexican financial reforms around 1989-1993, which brings to our attention the *sudden fall in the currency to money ratio* in late 1991. M1 then surged in Mexico, mostly led by a sudden rise in demand deposits. This could reflect either the high capital inflows due to a stable economy or, more fundamentally, a shift into interest-bearing assets. While we are not able to distinguish between these two hypotheses, this paper shows that long-run vectors are strongly found with currency as the monetary aggregate. This suggests narrow money demand appears to be a function of certain variables under a period of substantial financial sophistication in Mexico in the late 80s, which is extremely useful for policymaking purposes.

Without the currency to money ratio, the income elasticity varies between 0.75 and 1.15 (with the inflation rate) and the interest semi-elasticity is negative from –0.03 to -0.05. The cointegration link is partial when the *currency-money ratio* is taken as the proxy of financial innovation. A positive long-run coefficient on the financial innovation measure (+0.59) means that, as the currency to money ratio falls, innovations rise and the long-run value of money demand falls, consistent with our hypothesis. On the short-run models, the degree of adjustment varies between 6 and 13% of the deviation eliminated after 1 month in the model with currency. Based on stability tests, we conclude that the traditional and augmented models are not less stable than the one that captures financial innovations. As noted by Hamilton (1989), this could reflect the fact that the identification issue is critical and the more general model is indeed capturing money supply factors instead of money demand. We leave a more elaborate exploration of this issue for further work. A harder task includes considering the critique in Barnett (1997) to the Mexican money demand literature as far as "Divisia" indexes mitigates measurement error problems.

This research complements Arrau et al. (1995) that suffers from small-sample bias and not much financial sophistication going on in the Mexican economy during 1980-1989. It also recasts, on financial grounds, estimates on Mexico by Thornton (1996), Galindo (2000), Khamis and Leone (2001) and Garcés-Díaz (2002) that cover cointegration *without financial innovations*.

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Table 1. Unit Root Tests (Monthly):
México: 1986.1-2002.9.

Unit Root Test and Series	Series in Levels for tests with constant and trend		Series in First Differences for tests with constant only	
	ADF Test	KPSS Test	ADF Test	KPSS Test
CUR	-2.03(15)	0.30***	-3.09(14)**	0.08
M1	-2.67(13)	0.36***	-3.31(13)**	0.08
M2	-1.98(13)	0.52***	-3.01(12)**	0.31
i	-2.66(16)	0.50***	-3.39(15)**	0.03
Y	-2.14(16)	0.29***	-3.85(15)***	0.04
ΔP	-2.32(12)	0.25***	-4.62(11)***	0.02
f1	-1.71(14)	0.65***	-2.83(13)*	0.17

Notes: The variables are defined as follows: CUR stands for the real value of currency in circulation; M1 stands for the real value of M1; M2 stands for the real value of M2; i stands for the one month (not annualized) Cetes rate; Y stands for the index of industrial production; ΔP stands for the monthly variation in prices as measured by the CPI (consumer price index); and f1 is the measure of financial innovation defined as currency to M1. In all cases, the Mexican CPI is the deflator. ADF (k) refers to the Augmented Dickey-Fuller t-tests for unit roots, k is the selected lag length, and KPSS refers to the test developed by Kwiatkowski, Phillips, Schmidt and Shin in Kwiatkowski et al. (1992). For the series in levels, the ADF (k) and KPSS of each entry are estimated with a constant and trend as suggested by visual plots in figure 1. For the unit root tests in first-differences the test has only a constant. In the ADF tests, k is determined by the Campbell-Perron's lag length selection procedure developed formally in Ng and Perron (1995). The method starts with an upper bound, k_{max}= 18, on k. If the last included lag is significant, choose k = k_{max}. If not, reduce k by one until the coefficient of the last lag becomes significant (we use the 5% value of the asymptotic normal distribution to assess significance of the last lag). If no lags are significant, set k = 0. In the KPSS test the spectral estimation method is the Bartlett kernel and the bandwidth is verified for different values of the lag truncation parameter. Reported in the table are the statistics for lag truncation = 4; see the text for explanation. The symbols *, **, and *** attached to the figure indicate rejection of the null of no-stationarity at the 10%, 5%, and 1% levels, respectively.

Table 2. Johansen Cointegration Tests of Money Demand: MEXICO (1986:1-2002:9).

Estimates→ Specification↓	H_0 r=0 r≤1 r≤2 r≤3	Estimated Cointegrating Vectors	Lags	Serial Corr. LM-st.
With CUR:				
CUR, i, y	Trace 31.57* 7.08 1.35	CUR= -0.05 i + 0.75 y (0.01) (0.14)	6 [FPE, AIC]	10.93 [0.28]
	Max. Eigenvalue 24.49* 5.73 1.35	CUR= +1.05 y (0.21)		
CUR, i, y, Δp	Trace 59.05** 20.14 9.16 1.10	CUR= -0.03 i + 1.15y - 0.29 Δp (0.01) (0.19) (0.06)	3 [HQ]	12.18 [0.73]
	Max. Eigenvalue 38.91** 10.98 8.06 1.10	CUR= +1.54 y - 0.46 Δp (0.17) (0.08)		
CUR, i, y, f1	Trace 43.83 15.86 4.29 0.54	CUR= -0.11 i + 0.49 y + 0.59 f1 (0.02) (0.24) (0.34)	3 [AIC, HQ, FPE]	14.43 [0.57]
	Max. Eigenvalue 27.97* 11.56 3.76 0.54	CUR= +1.06 y - 0.37 f1 (0.27) (0.39)		
With M1:				
M1, i, y	Trace 27.94 5.62 0.62	M1= -0.26 i + 0.08 y (0.05) (0.57)	3 [LR, FPE, AIC, HQ]	9.36 [0.41]
	Max. Eigenvalue 22.32* 5.00 0.62	M1= +2.56 y (0.65)		
M1, i, y, Δp	Trace 48.76* 22.23 5.92 1.09	M1= -0.26 i + 0.08 y – 0.96 Δp (0.05) (0.57) (0.29)	3 [LR, FPE, AIC]	8.59 [0.93]
	Max. Eigenvalue 26.53 16.31 4.83 1.09	M1= +5.20 y – 3.27 Δp (1.37) (0.69)		
M1, i, y, f1	Trace 42.35 14.04 4.09 0.51	M1= -0.11 i + 0.40 y – 1.99 f1 (0.02) (0.22) (0.32)	3 [HQ, FPE]	16.98 [0.39]
	Max. Eigenvalue 28.30* 9.95 3.58 0.51	M1= +1.05 y – 2.84 f1 (0.29) (0.43)		

Notes: The money and income variables are in logarithms. The variables are defined in table 1. The null hypothesis, when r = 0, is that there are no cointegrating vectors against the alternative of at most one vector. Seasonal monthly dummies are included in the VAR estimation. The test statistic (for n = 3,4 variables in the system) is $-2\ln(Q) = -T\Sigma_{i=r+1}(1-\lambda_i)$, where T is the sample size, r is the number of cointegrating vectors, and λ_i are the n-r smallest eigenvalues from the correlation matrix between ΔX_t and X_{t-k} after regressing these 2n variables on a constant, lags of ΔX_t and monthly dummies. The lag-length adopted is obtained by inspection of the modified likelihood ratio (LR) test, the final prediction error (FPE), the Akaike (AIC), Schwarz (SC), and Hannan-Quinn (HQ) information criteria. Below the reported lag-length for each specification are the criteria adopted, decided on the basis of maximum coincidence of statistical criteria and by inspection of the serial correlation LM-test, whose null hypothesis is no serial correlation at lag order h. The mark ** indicates significance at the 1% level and * at the 5% level.

Table 2
(cont.) Johansen Cointegration Tests of Money Demand: MEXICO (1986:1-2002:9).

Estimates→ Specification↓	H_0 r=0 r≤1 r≤2 r≤3	Estimated Cointegrating Vectors	Lags	Serial Corr. LM-st.
With M2:				
M2, i, y	Trace 19.25 8.15 0.15	M2= -0.02 i + 1.61 y (0.02) (0.21)	4 [FPE, AIC]	5.42 [0.80]
	Max. Eigenvalue 11.10 8.01 0.15	M2= +1.79 y (0.14)		
M2, i, y, Δp	Trace 37.11 16.05 6.75 2.01	M2= +0.01 i + 2.20 y – 0.35 Δp (0.02) (0.29) (0.11)	5 [FPE, AIC]	10.94 [0.81]
	Max. Eigenvalue 21.05 9.30 4.74 2.01	M2= +2.03 y – 0.28 Δp (0.16) (0.09)		
M2, i, y, f1	Trace 37.27 10.95 2.20 0.01	M2= +1.21 i + 6.23 y – 12.23 f1 (0.24) (2.97) (4.39)	2 [SC, HQ]	15.13 [0.52]
	Max. Eigenvalue 26.32 8.75 2.19 0.01	M2= +1.86 y + 0.008 f1 (0.26) (0.37)		

Notes: The money and income variables are in logarithms. The variables are defined in table 1. The null hypothesis, when r = 0, is that there are no cointegrating vectors against the alternative of at most one vector. Seasonal monthly dummies are included in the VAR estimation. The test statistic (for n = 3,4 variables in the system) is $-2\ln(Q) = -T\Sigma_{i=r+1}(1-\lambda_i)$, where T is the sample size, r is the number of cointegrating vectors, and λ_i are the n-r smallest eigenvalues from the correlation matrix between ΔX_t and X_{t-k} after regressing these 2n variables on a constant, lags of ΔX_t and monthly dummies. The lag-length adopted is obtained by inspection of the modified likelihood ratio (LR) test, the final prediction error (FPE), the Akaike (AIC), Schwarz (SC), and Hannan-Quinn (HQ) information criteria. Below the reported lag-length for each specification are the criteria adopted, decided on the basis of maximum coincidence of statistical criteria and by inspection of the serial correlation LM-test, whose null hypothesis is no serial correlation at lag order h. The mark ** indicates significance at the 1% level and * at the 5% level.

Table 3. Error Correction Models of Money Demand: MEXICO (1986:1-2002:9).

Model and Estimates→ Variables ↓	Model CUR = f (i, y)	Model CUR = g (i, y, Δp)	Model CUR = h (i, y, f1)
C	1.60*** (0.38)	1.31*** (0.18)	0.96*** (0.29)
ECM_{t-1}	-0.13*** (0.04)	-0.12*** (0.02)	-0.06*** (0.02)
ΔCUR_{t-1}	-0.52*** (0.07)	-0.54*** (0.05)	-0.55*** (0.08)
ΔCUR_{t-2}	-0.31*** (0.08)	-0.34*** (0.07)	-0.31*** (0.09)
ΔCUR_{t-5}	0.10* (0.06)	0.13** (0.05)	0.10** (0.06)
ΔCUR_{t-6}	0.14* (0.07)	0.11* (0.06)	0.14** (0.07)
ΔCUR_{t-10}		-0.15*** (0.05)	
Δp_{t-1}	-0.03*** (0.008)		-0.02** (0.01)
Δi_{t-1}	-0.014*** (0.003)	-0.02*** (0.002)	-0.02*** (0.003)
Δi_{t-3}		-0.01*** (0.003)	
Δi_{t-13}	-0.01*** (0.004)	-0.01*** (0.003)	-0.01*** (0.004)
Sample Size	187	187	187
Adj. R^2	0.84	0.86	0.82
D.W. Stat.	1.82	1.96	1.83
JB-Normality	4.17 [0.12]	9.05* [0.01]	8.83 [0.01]
Breusch-Godfrey Serial Cor. LM	2.53* [0.06]	0.06 [0.98]	3.10** [0.02]
Serial Cor. ARCH	0.89 [0.56]	0.95 [0.51]	0.80 [0.66]

Notes: As in table 2, the money and income variables are in logarithms and the variables are as defined in previous tables. The lagged ECM term represents lagged residuals of the Johansen cointegration vector of currency as function of interest rate, income, and either inflation or the currency to money ratio, depending on the model specification. Up to 13 lags of each variable is included in the more general regression. We then remove all the insignificant terms and reach the ECM specification reported in the table. Monthly dummy coefficients, omitted in the table, from January to November are included in each equation and their coefficients are always statistically significant. The Newey-West HAC standard errors and covariance matrix is employed with lag truncation = 4. The mark *** indicates rejection of the null at the 1% level, ** indicates rejection at the 5% level, and * at the 10% level.

Table 4. Stability Tests of Money Demand: MEXICO (1986:1-2002:9).

Long-run Model is: $CUR = f(i, y) + \varepsilon_t$ (1)

$CUR = g(i, y, \Delta p) + \varepsilon_t$ (2)

$CUR = h(i, y, f1) + \varepsilon_t$ (3)

F-Statistics → Episodes and Dates ↓	Estimates of Chow Breakpoint Model CUR = f (i, y)	Model CUR = g (i, y, Δp)	Model CUR = h (i, y, f1)
Around M1 Explosion			
1991:8	2.85*** [0.00]	1.86*** [0.02]	2.97*** [0.00]
1991:9	2.81*** [0.00]	1.86*** [0.02]	2.97*** [0.00]
1991:10	2.87*** [0.00]	1.91** [0.01]	3.18*** [0.00]
1991:11	2.83*** [0.00]	1.91** [0.01]	3.17*** [0.00]
1991:12	2.81*** [0.00]	1.81** [0.02]	3.15*** [0.00]
Around Currency Crisis			
1994:10	1.31 [0.18]	0.78 [0.74]	1.95** [0.01]
1994:11	1.34 [0.17]	0.77 [0.76]	1.95** [0.01]
1994:12	1.34 [0.17]	0.78 [0.74]	1.96** [0.01]
1995:1	1.33 [0.17]	0.75 [0.78]	1.94** [0.01]
1995:2	1.50* [0.09]	0.78 [0.74]	2.13*** [0.006]
CUSUM	Deviations from 2002	Well-behaved	Deviations from 2002
CUSUMQ	Well-behaved	Small deviations in 1994-1995	Well-behaved

Notes: The Chow Test of Stability of Coefficients tests the null hypothesis of no structural change against structural change in the parameters of the ECM model. The mark *** indicates rejections of the null at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level. For the CUSUM and CUSUMQ statistics, the sample period is broken and the statistics are updated recursively and are plotted against the breakpoints. Figures 5, 6 and 7 contain the graphs of these tests. If their plots stay within 5% significance level, the coefficients are said to be stable.

Figure 1. Plots of real currency (CUR), real M1 (M1), and real M2 (M2). Plots of the rate of inflation measured by CPI (dP), of the (monthly) opportunity cost of money: the one month Cetes interest rate (i) and of the index of industrial production (IP).

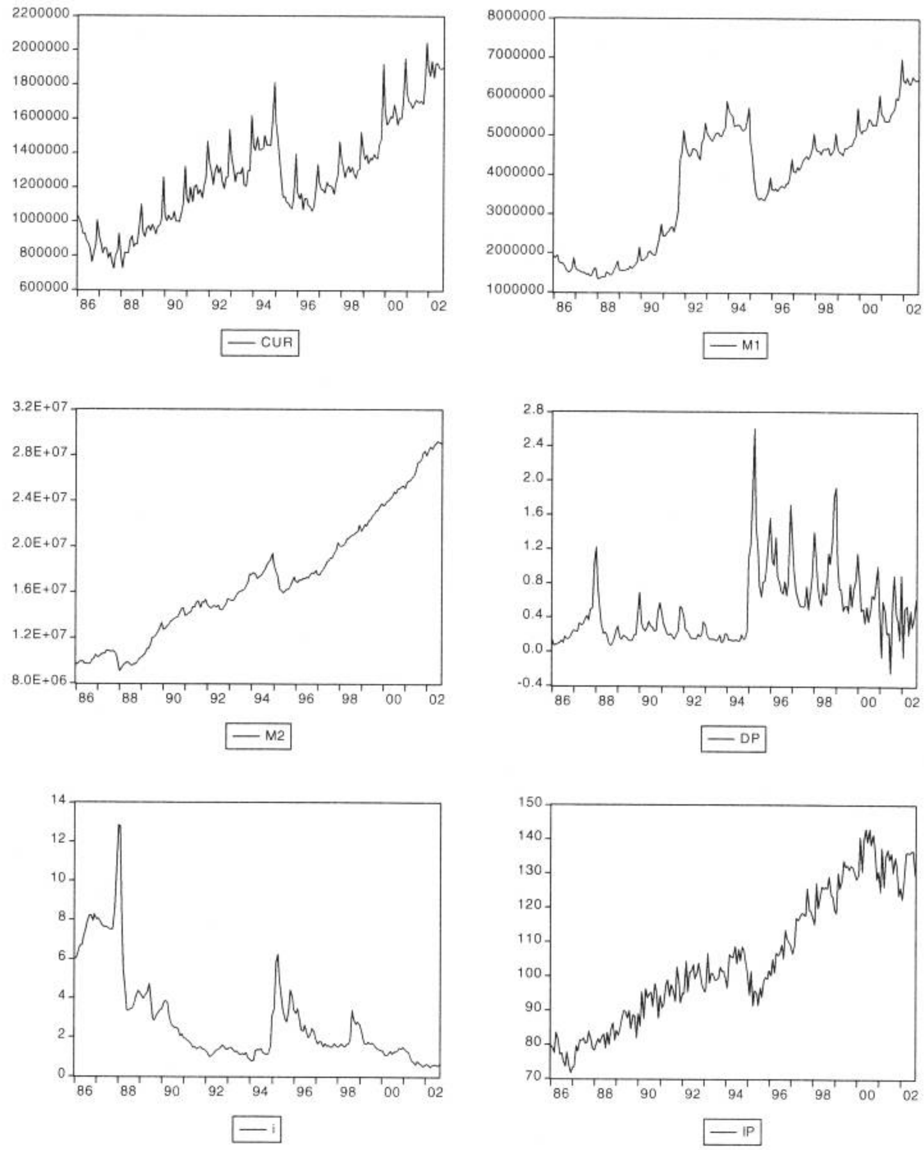
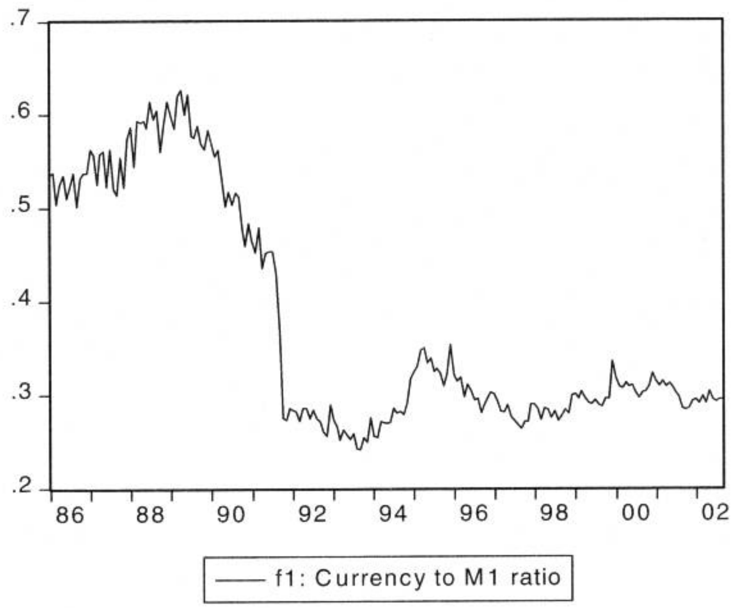


Figure 2. Plots of the two measures of financial innovations. f1 is the currency to money ratio and f2 is M2/M1, the ratio of aggregates M2 to M1.



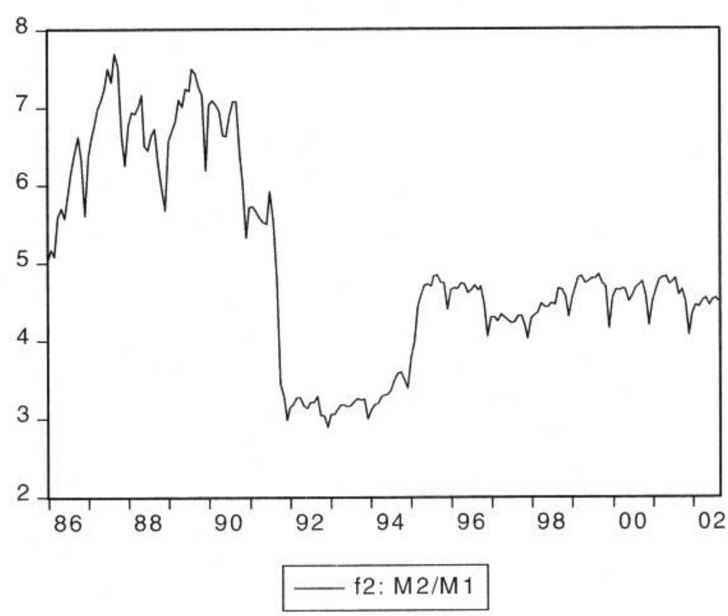
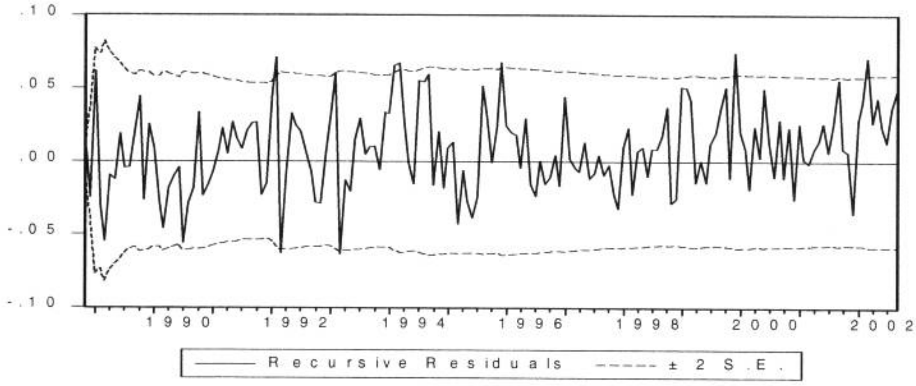
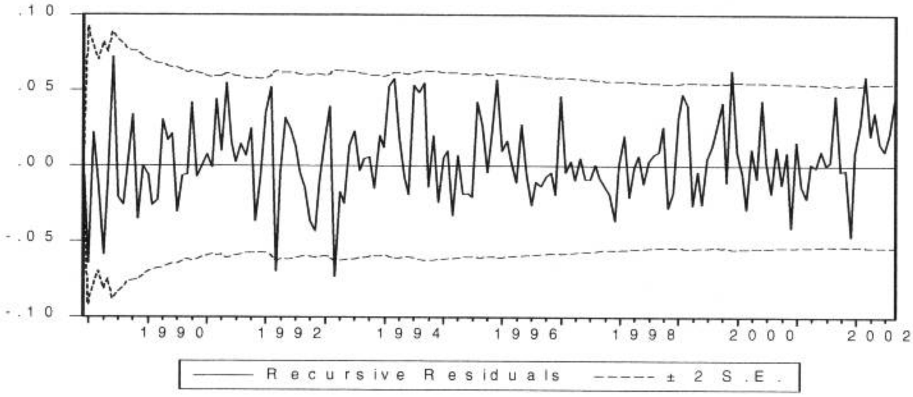


Figure 3. Recursive Residuals across Models.

Model 1:



Model 2:



Model 3:

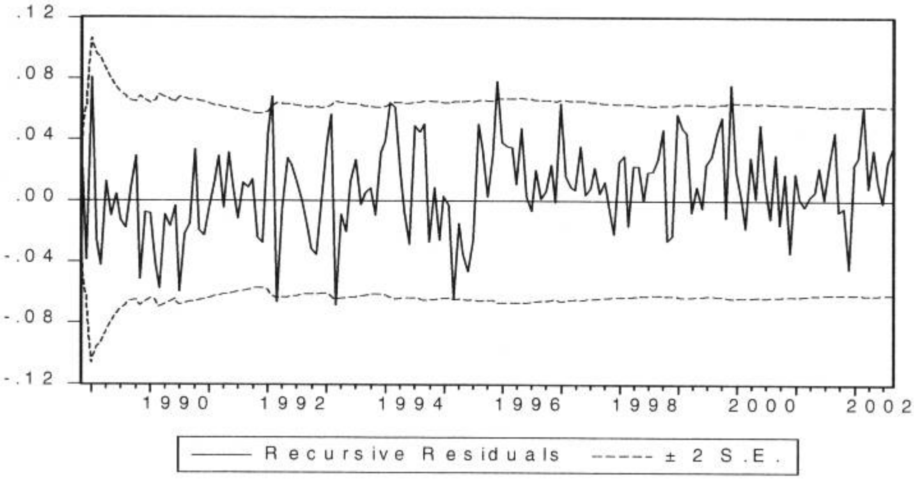
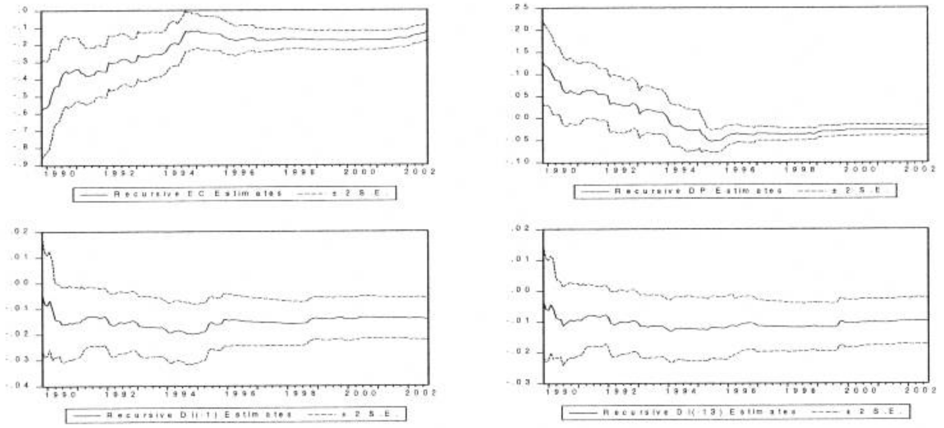
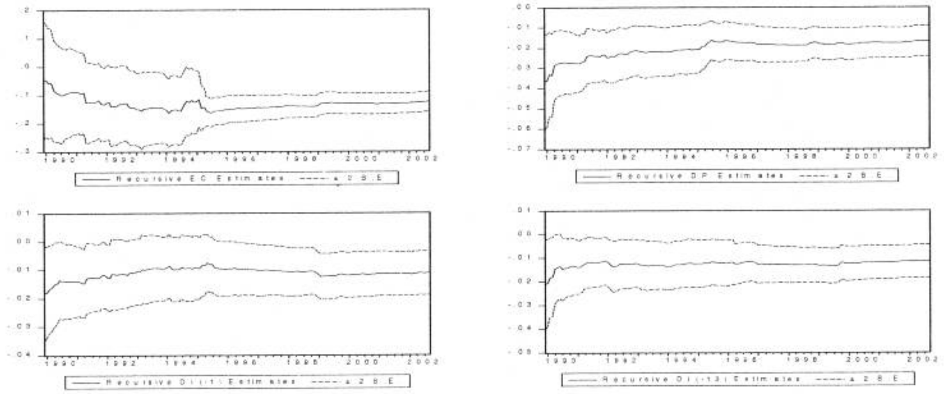


Figure 4. Recursive Error-Correction, Inflation and Interest Rate Parameter Estimates across Models.

Model 1:



Model 2:



Model 3:

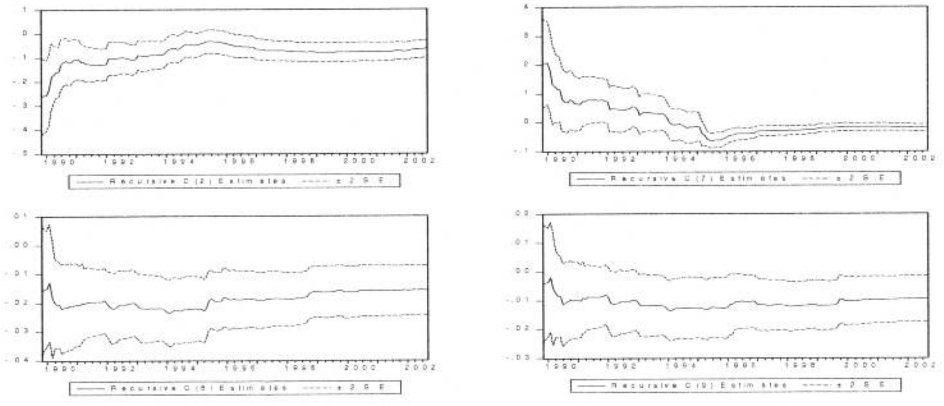
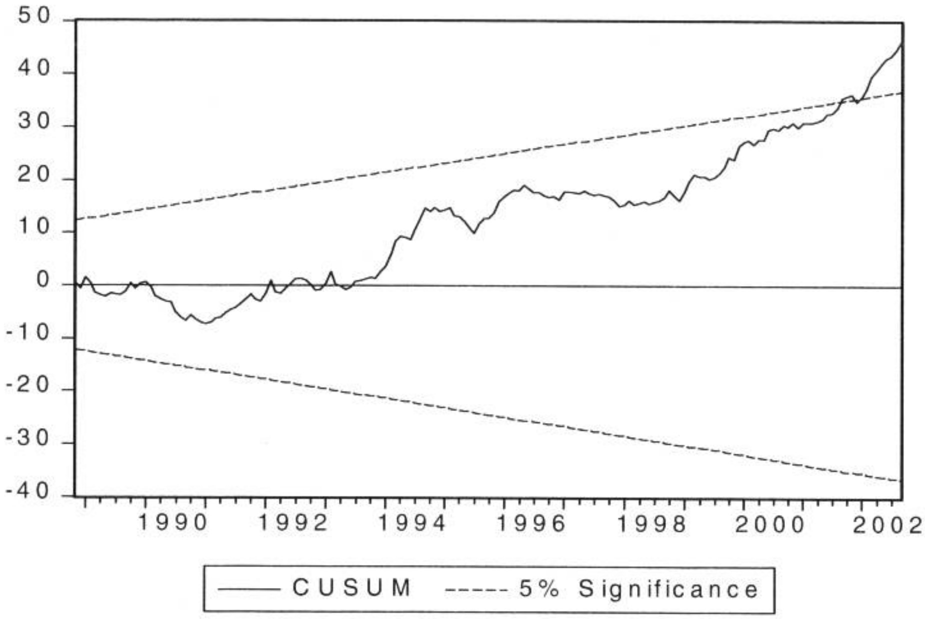


Figure 5. CUSUM and CUSUMQ Tests of Stability of Coefficients: ECM Associated with Equation (1).



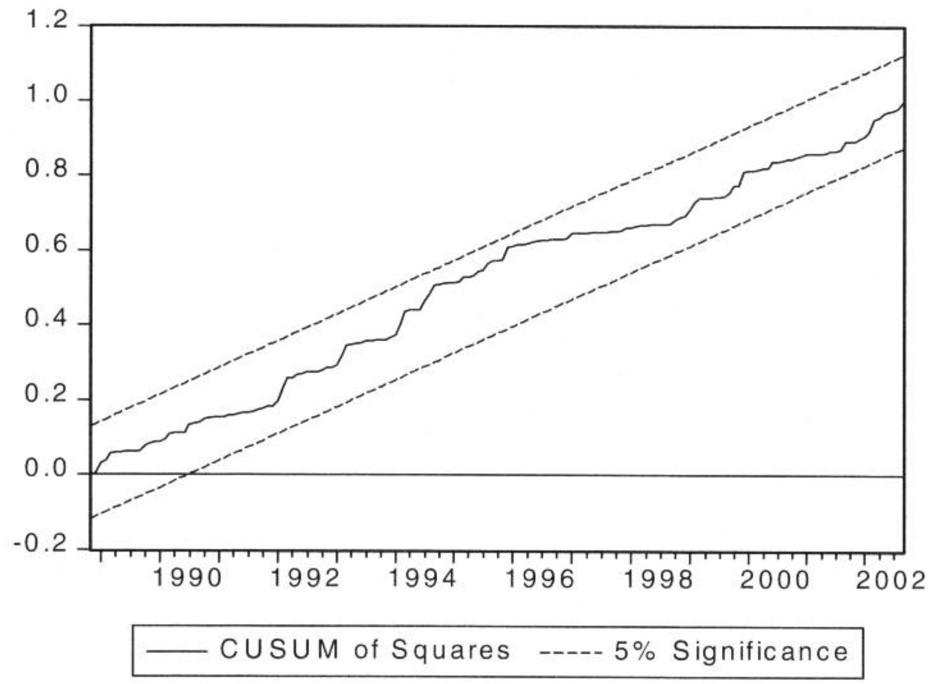
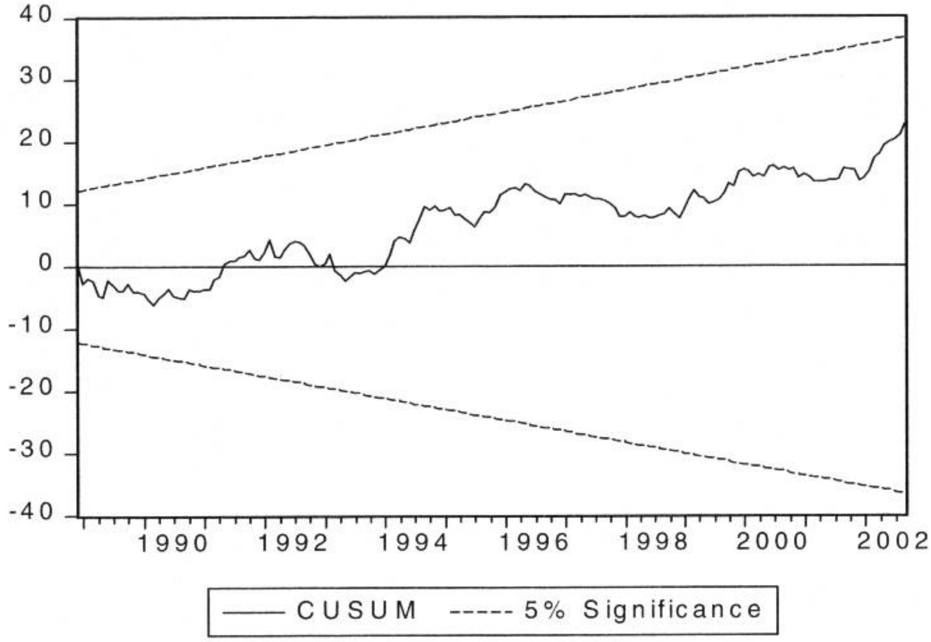


Figure 6. CUSUM and CUSUMQ Tests of Stability of Coefficients: ECM Associated with Equation (2).



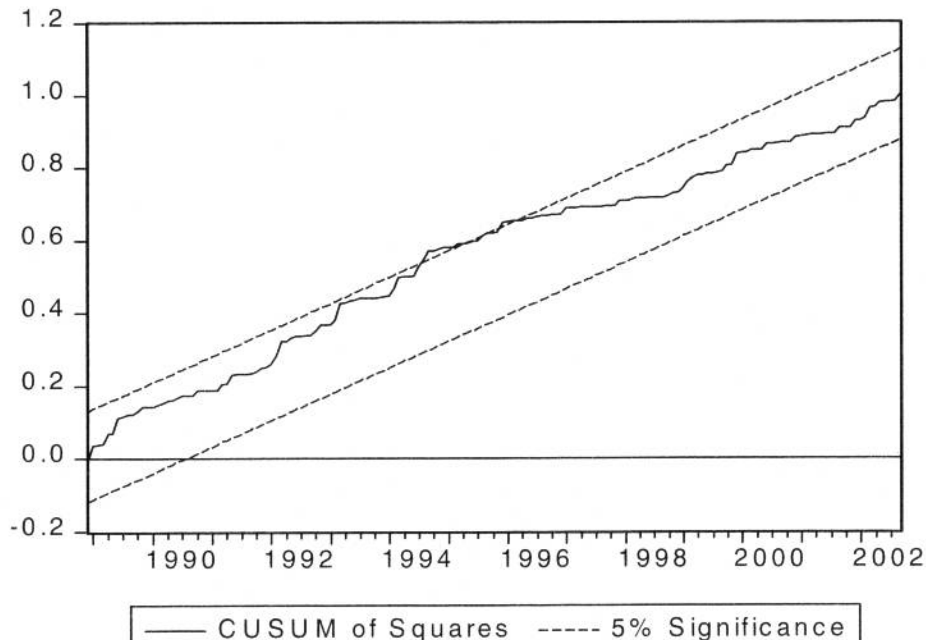


Figure 7. CUSUM and CUSUMQ Tests of Stability of Coefficients: ECM Associated with Equation (3).