Can Currency Demand Be Stable Under a Financial Crisis? The Case of Mexico

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The paper finds strong evidence that real currency demand in Mexico remained stable throughout and after the financial crisis in Mexico. Cointegration analysis using the Johansen-Juselius technique indicates a strong cointegration relationship between real currency balances, real private consumption expenditures, and the interest rate. The dynamic model for real currency demand exhibits significant parameter constancy even after the financial crisis as indicated by a number of statistical tests. The paper concludes that the significant reduction in real currency demand under the financial crisis in Mexico could be appropriately explained by the change in the variables that historically explained the demand for real cash balances in Mexico. This result supports the Bank of Mexico's use of a reserve money program to implement monetary policy under the financial crisis. [JEL E41, C51, C52]

At the onset of the financial crisis in Mexico and the devaluation of the peso in December 1994, the Bank of Mexico (BOM) was prompted to adopt a floating exchange rate.[1] This had significant implications for the implementation

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[1]Prior to the devaluation of the peso, the exchange rate was allowed to fluctuate within a band.

of monetary policy where the exchange rate no longer could provide the nominal anchor for the economy. Consistent with its target for price stability, the BOM established a reserve money target (and, in particular, a limit on the annual growth of its credit) as a central element of its monetary program.[2] The annual target for reserve money was formulated by projecting the demand for reserve money (currency plus banks' current accounts at the central bank), taking into account the inflation target.[3] Since the established reserve money target is based on projections of the demand for reserve money using the historic relationship governing the demand for reserve money, underlying the adoption of a reserve money target is an assumption that the relationship governing the demand for reserve money (and thus currency) remained stable during the financial crisis.

This paper examines whether the demand for currency remained stable during and after the Mexican financial crisis. In other words, we study whether the process determining the demand for currency in Mexico remained unchanged, even after the change in the exchange rate system and the inception of the financial sector crisis. Under crisis conditions, the relationship establishing the demand for money (currency as well as broader monetary aggregates) could change for multiple reasons. For example, the interest elasticity of currency demand could decrease as a result of the larger risk associated with bank deposits. If the observed increase in bank deposit rates largely reflected the increase in risk associated with such deposits, currency demand would not be expected to change in response to changes in deposit interest rates since the *risk-adjusted* interest rates remained virtually constant. When estimating the relationship describing the demand for currency, this effect would be reflected as a change in the interest rate elasticity, since the interest rates used in the estimation are the reported rates and not the effective (risk-adjusted) rates. Furthermore, a *change* in the relationship could occur if agents, as a result of the crisis conditions, consider holding alternative financial assets. For example, in a situation where the interest rates on domestic financial assets are no longer attractive, when taking into consideration the risk associated with these assets (that is, the risk premiums offered on domestic deposit rates are not large enough), foreign assets (such as offshore U.S. dollar deposits) could become an attractive substitute for domestic currency. The opportunity cost of holding currency would, therefore, be better measured by the expected return on foreign financial assets, which could be summarized by the

[2]The Bank of Mexico uses reserve money as an intermediate target to fulfill its inflation objective since reserve money bears a reasonably stable relationship with the price level. To implement its monetary policy, central bank credit is managed through open market operations so that, in principle, the supply of reserve money would meet its projected demand. The reserve money target is periodically revised in response to changes in certain indicators including the evolution of the exchange rate, observed inflation vis-à-vis targeted inflation, the evolution of inflation expectations, and developments in the settlement of wages (Bank of Mexico, 1994, 1995).

[3]Since reserve requirements in Mexico were abolished in 1988, the demand for currency comprises much of the demand for reserve money. For example, the share of currency in reserve money amounted to about 73 percent in mid-1997. The remaining reserve balances maintained by banks are generally for payment purposes.

expected rate of exchange rate depreciation when the expected rate of depreciation is large.[4]

Despite its significance for the conduct of monetary policy, the stability of money demand is not generally addressed in the relevant literature on Mexico. A notable exception to this is Ramos-Francia (1993). In this paper, the author estimates the demand for M1 using the "general to specific" methodology developed by Hendry and Richard.[5] The stability of the estimated equation is also evaluated. The study, however, extends only to 1990 and does not cover the financial crisis. Other recent works include Rogers (1992), Arrau and De Gregorio (1993), Choudhry (1995), De Lemos Grandmont (1991), Kamin and Rogers (1996), Aboumrad (1996), Thornton (1996), and Desentis (1997). The latter two come closest to addressing this issue. Thornton examines the long-run stability of M1 and M2 but does not cover the financial crisis in 1994. Also, the stability of the dynamic model is not evaluated. Desentis's study covers the financial crisis period, but it falls short of evaluating the stability of the estimates.[6]

This paper addresses the above issues and examines whether the real demand for currency remained stable during and after the Mexican financial crisis. If currency demand was stable, real balances would have, in the long run, a proportional relationship with the volume of real transactions and the opportunity cost of holding currency; that is, these variables would be cointegrated. Utilizing the Johansen-Juselius (1990) cointegration techniques, this paper examines the long-run determinants of real currency demand during 1983:1–1997:6 using monthly data. In addition, the dynamics of real currency demand are estimated using an error correction representation of the data, and the stability of the dynamic model is examined.[7] The study period contains the inflationary debt crisis period, the stabilization period under the December 1987 stabilization plan (the Pacto), the ensuing financial crisis in December 1994, and the recovery period thereafter.

The results of this study suggest that real currency demand remained stable after the financial crisis in Mexico, despite the substantial reduction in the public's holdings of real currency balances after the devaluation in December 1994. Strong evidence is found of long-term stability of real currency demand indicated by the cointegration of real currency, private consumption expenditures, and inflation.

[4]The opportunity cost of holding currency vis-à-vis foreign financial assets is the sum of the interest rate offered on these assets and the expected rate of exchange rate depreciation. In crisis conditions, exchange rate depreciation is usually substantially higher than the actual returns offered on such assets, thus justifying the use of the expected exchange rate depreciation alone as a proxy for the opportunity cost of holding currency.

[5]See Hendry and Richard (1982, 1983).

[6]Desentis's study suffers from some weaknesses such as the use of the manufacturing production index as the scale variable and the treasury bill (CETES) rate for the opportunity cost of holding currency, which, as discussed below, are inadequate proxies for transaction demand and the short-term opportunity cost of holding currency. Also, the use of velocity as the cointegration relationship is ad hoc.

[7]To detect any misspecification of the dynamic model, its stability is evaluated not only under the financial crisis but also throughout the period studied. Stability is tested using various statistics, including several variants of the Chow test, the forecast χ^2 statistic, and the output of the recursive estimation of the error correction model.

Since inflation and interest rates move together in the long run (the results also indicate a cointegration relationship between these two variables), changes in the interest rate would encompass changes in inflation. This explains the insignificance of the interest rate as a determinant of long-term real currency demand. The stability test of the dynamic model indicates constancy of the estimated parameters throughout the period. The dynamic model's specification includes—in addition to the dependent variable's lags and the error correction term—lagged inflation and interest rate. Changes in real private consumption expenditures do not seem to have any significant effect on real currency demand in the short run.[8]

I. Data

The study uses (seasonally unadjusted) monthly observations for the period 1983:1–1997:6 for currency in circulation (M) deflated by the consumer price index (P).[9] Real private consumption expenditure (Y) was used as the scale variable to estimate the transaction demand for currency. Quarterly data was used and was repeated for each month of the same quarter.[10] The CPI inflation rate (Δp) and the interest rate on 60-day time deposits (R) were used as estimates for the opportunity cost of holding currency (as opposed to holding real and financial assets, respectively).[11] Whereas demand deposits are a closer substitute for cash than time deposits, interest-bearing checking accounts were only introduced in 1990 and, therefore, could not be used for the whole period under study (Figure 1).[12] The interest rate on one-month treasury bills (CETES) was also tried. The CETES rate was significant and negative in the cointegration vector but significant and positive in explaining the short-term dynamics, which indicates that, in the short run, the estimation captured the money supply reaction function of the Bank of Mexico rather than the demand function for money.

Figure 2 shows the short-term procyclical relationship between the CETES rate and currency (shown in the graph as a countercylical relationship between the

[8]This, however, could be related to the use of quarterly consumption data to construct the monthly series. The constructed series does not reflect intraquarter changes in actual real consumption where quarterly consumption data was repeated for each month of the same quarter.

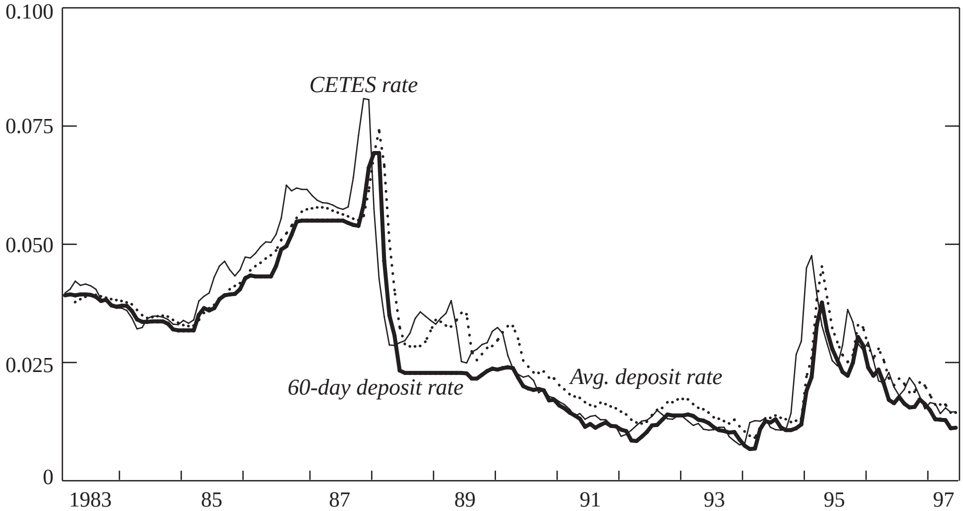
[9]The data used in this study are from the *International Financial Statistics*, published by the International Monetary Fund. All lower case variables denote the natural logarithm of the original variables.

[10]Other scale variables were also tested. GDP data (similar to the treatment of consumption data, quarterly data were repeated for the months of each quarter) as well as the monthly industrial production index were individually tested in the cointegration estimation. The results of the estimation were not robust; the sign and magnitude of the GDP coefficient and the industrial production index varied substantially with different sample sizes. The superiority of consumption expenditures as a scale variable for transaction demand is consistent with Friedman and Schwartz (1982) and Hall (1978) in that consumption is closely related to unobservable permanent income, which in turn is a better proxy for the volume of transactions.

[11]Laidler (1985) argues that because, for reasons not well understood, variations in nominal interest rates do not fully reflect variations in the expected inflation rate, this leaves room for the expected inflation rate to play a direct role in the demand-for-money function over and above that played by nominal interest rates. Both variables (R and Δp) used in this paper denote monthly rates (unannualized).

[12]Figure 1 shows that these rates generally move together and, therefore, the estimation results should not be affected by which rate is used in the specification.

Figure 1. Interest Rates



Sources: Bank of Mexico; *International Financial Statistics*, International Monetary Fund.

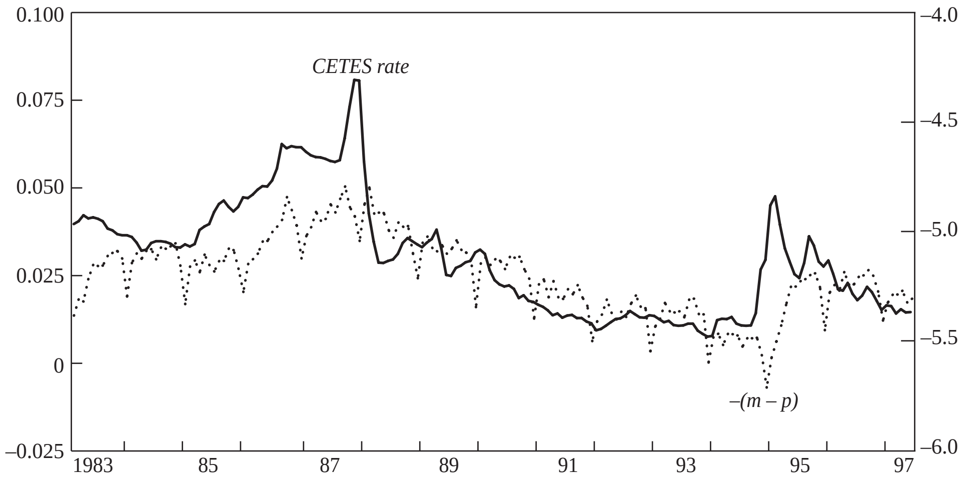
CETES rate and the inverse of real balances, $-(m-p)$, indicating the active use of government securities auctions for monetary management purposes. Whereas this does not generally affect the long-term relationship, it could affect the short-term estimation particularly when high frequency data is used; this is because bank deposit interest rates generally require a period of two to three months to adjust to changes in CETES rates (see Figure 1).

Figure 3 presents the series for $m-p$, y, Δp, the exchange rate (Δe), and velocity (v, defined as $y-[m-p]$) for the period 1983:1–1997:6. All series reflect the major macroeconomic episodes in this period. Panel (a) shows that $m-p$ declines following the debt crisis, reflecting the large demonetization of the economy in that period as a result of high inflation. It recovers substantially after the initiation of the Pacto in December 1987 and exhibits an abrupt decline at the onset of the crisis in December 1994, with some recovery evident in the early part of 1997. Real private consumption expenditures largely mirror $m-p$ behavior, although the effect of the debt crisis on y is less obvious.

In the same way, panel (b) shows that inflation and the interest rate move very closely together, with inflation significantly more variable. Both series increase steadily after the debt crisis and decline substantially in early 1988, reflecting the stabilization program. After a prolonged decreasing trend (with periodic increases in the interest rate in 1992 and 1994), both series increase substantially, reflecting the financial crisis in December 1994. After the initial few months of the crisis, R and Δp start a downward trend.

Panel (c) depicts a close relationship between the exchange rate change and inflation. Episodes of large devaluations (at the announcement of the Pacto and at the onset of the financial crisis in 1994) are accompanied by large spurts in infla-

Figure 2. The 28-day CETES Rate and Real Cash Balances



Sources: Bank of Mexico; *International Financial Statistics*, International Monetary Fund.
Note: Graph is rescaled by means and ranges.

tion, with subsequent stabilization in both variables. In general, the figures indicate a clear positive relationship between $m-p$ and y, and a negative relationship between $m-p$ and both R and Δp, supporting the likelihood of a long-term cointegration relationship among these variables.

The possibility of a cointegration vector among these variables is further supported by panel (d), which indicates a long-term comovement between v and Δp.[13] Furthermore, panel (b) indicates a strong long-term relationship between R and Δp, indicating the possibility of a second cointegrating vector between these two variables. Although the series reflect breaks during the study period, cointegration is still possible mainly because these breaks are present in all series, and the series continue to move together even in the presence of these breaks.

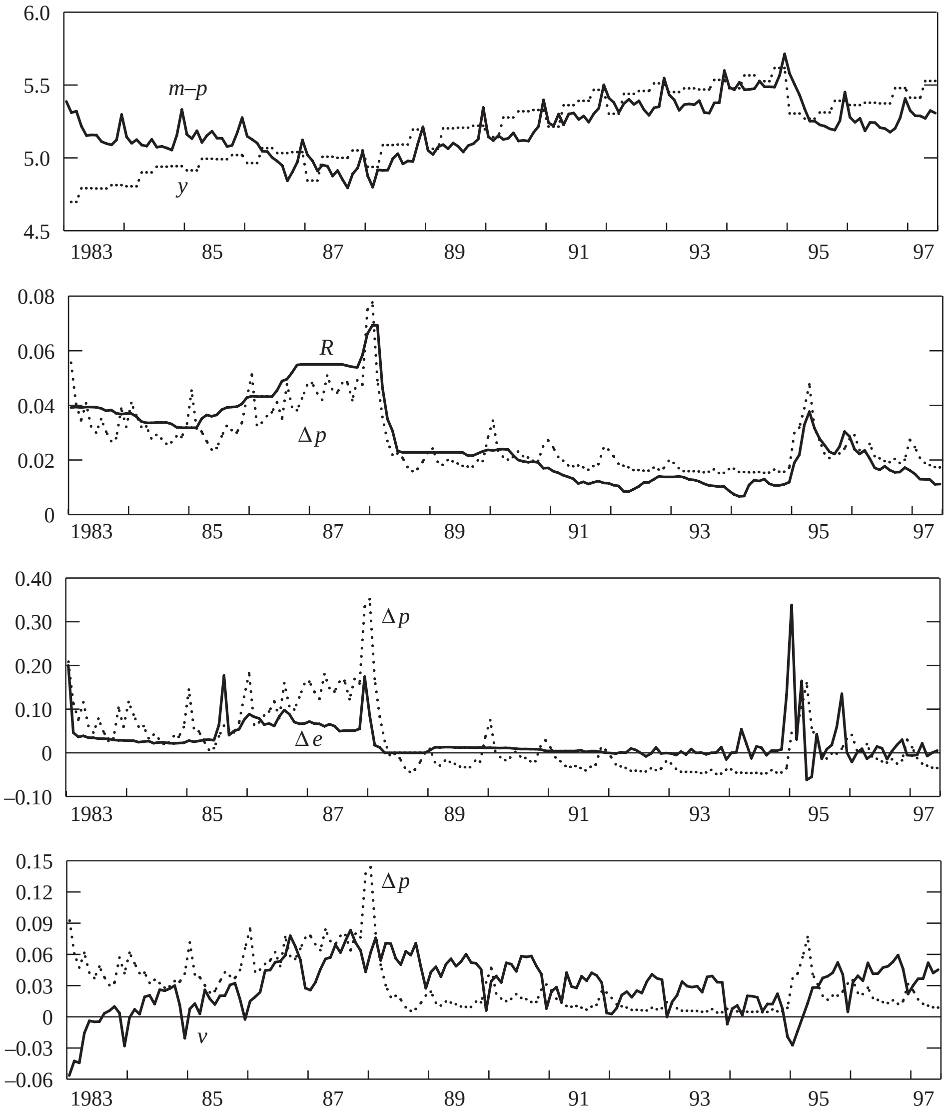
The logarithmic transformation of variables corresponds to the following specification of the long-run currency demand equation:

$$\frac{M^d}{P} = \gamma_0 Y^{\gamma_1} e^{\left(\gamma_2 R + \gamma_3 \Delta p\right)}.$$

The augmented Dickey-Fuller test (ADF) was used to test for the stationarity and order of integration of the series used in the estimation. Results are reported in Table 1. ADF tests indicate that both m and p are integrated of order two—$I(2)$.

[13]Since Δp moves closely with R and Δe, in principle, the cointegration vector could also include the latter two variables. As indicated later, Δe was found to be $I(0)$. It was included in the dynamic model but was found to be statistically insignificant.

Figure 3. Real Currency, Real Private Consumption, Inflation and Interest Rates, and Velocity, 1983:1–1997:6



Source: *International Financial Statistics*, International Monetary Fund.
Note: All graphs are rescaled by means and ranges.

Table 1. Augmented Dickey-Fuller (ADF) Tests for the Order of Integration of Individual Variables

Variable	ADF Statistic	Lags	Constant Included	Trend Included
m	–2.29	6	yes	no
p	–2.09	1	yes	no
$m-p$	–2.39	3	yes	yes
R	–2.60	1	yes	yes
y	–2.10	0	yes	yes
Δp				
1983:1–1997:6	–3.02*	0	yes	no
1983:6–1997:6	–2.72	0	yes	no
$\Delta(m-p)$	–6.42**	2	yes	no
ΔR	–7.69**	1	no	no
Δy	–2.91**	11	no	no
Δm	–1.93	5	yes	no
$\Delta\Delta p$	–3.82**	11	no	no
$\Delta\Delta m$	–5.42**	4	yes	no

Note: The test period for all variables (except if indicated otherwise) is 1983:1–1997:6 minus the lags. The symbols * and ** imply rejection of the null hypothesis of a unit root at the 5 percent and 1 percent level, respectively. The test for inflation is inconclusive; tests using sample subperiods could not reject the null hypothesis of a unit root. The symbol $\Delta\Delta$ indicates the second difference; that is, $\Delta\Delta x_t = \Delta x_t - \Delta x_{t-1}$.

Their cointegrating vector $(m-p)$ is of order one—$I(1)$.[14] Figure 4 presents the m and p series and their first differences. All other variables used in the estimation (that is, Δp, R, and y) were found to be $I(1)$. The test results for inflation are very sensitive to the time period chosen. Nevertheless, inflation was assumed to be $I(1)$, having the same order of integration as the interest rate.[15]

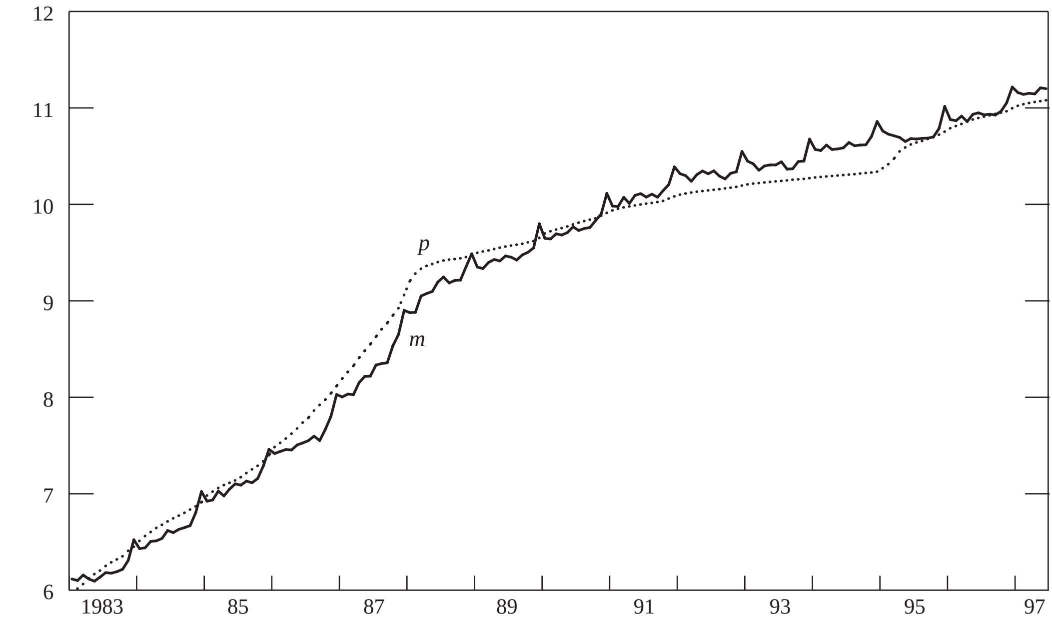
II. Long-Run Behavior and Cointegration

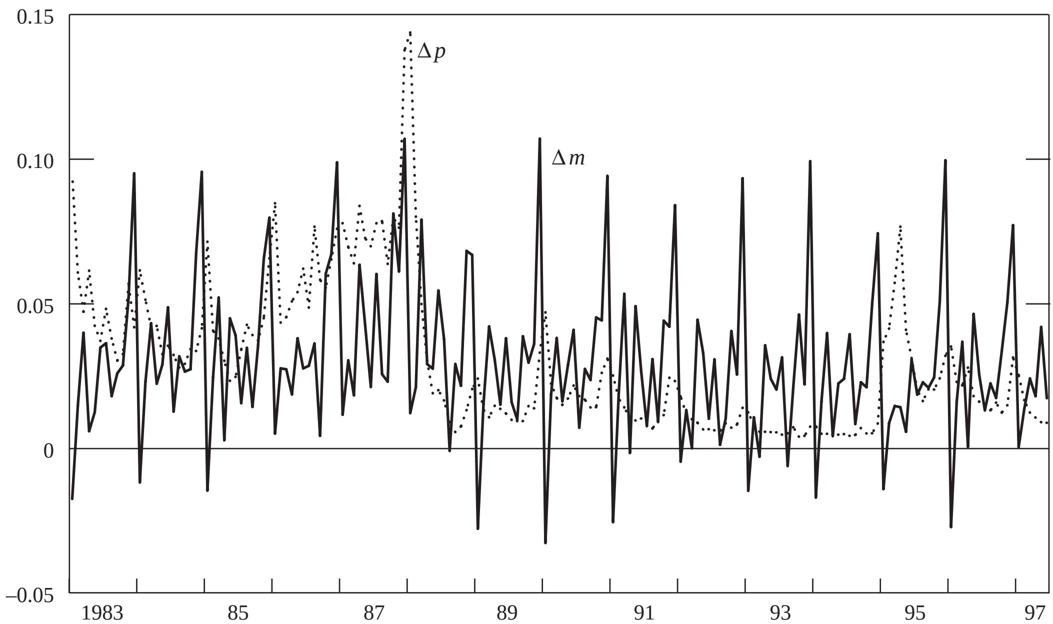
In this section, we examine the presence of cointegration between $m-p$, R, Δp, and y using the Johansen-Juselius procedure (Johansen, 1988; Johansen and Juselius, 1990). For cointegration to exist among variables, all the relevant variables must be integrated of the same order. This has already been established by the ADF tests above.

[14]Johansen (1994) has a similar result for U.K. data.

[15]The rate of exchange rate depreciation was also tested and was found to be $I(0)$. It was statistically insignificant when included in the dynamic model as an explanatory variable.

Figure 4. A Graphical Analysis of the Order of Integration of Currency and Prices





Source: *International Financial Statistics*, International Monetary Fund.
Note: All graphs are rescaled by means and ranges.

Table 2. Unrestricted Cointegration Results

I. Eigenvalues and Related Test Statistics

		Maximal Eigenvalues		Eigenvalue Trace	
	Eigenvalue	Statistic	95% Critical value	Statistic	95% Critical value
1	0.224	40.46**	31.5	85.90**	63.0
2	0.148	25.59*	25.5	45.43*	42.4
3	0.077	12.89	19.0	19.85	25.3
4	0.043	6.96	12.3	6.96	12.3

II. Normalized α and β' Matrices

α (Weighing Matrix)

Variable				
$m-p$	–0.064	–1.653	0.023	–0.039
R	–0.001	0.105	–0.004	0.017
y	0.002	–0.612	–0.121	–0.041
Δp	–0.010	0.875	–0.019	–0.002

β' (Cointegrating Vectors)

Variable	$m-p$	R	y	Δp	Trend
	1.000	–17.956	–2.013	20.505	0.0040
	0.029	1.000	–0.008	–0.429	0.0004
	0.085	2.423	1.000	0.514	–0.0014
	0.182	–1.660	–0.062	1.000	–0.0005

Notes: Estimation period is 1984:3–1997:6. The symbols * and ** denote significance at the 5 percent and the 1 percent level, respectively.

The Johansen-Juselius system-based procedure was applied to $m-p$, y, R, and Δp, with a constant and monthly dummies, trend, and three lags.[16] The trend was restricted to lie in the cointegration space.[17] Accordingly, the hypothesized cointegration vector is of the form $\beta_1 m + \beta_2 y + \beta_3 R + \beta_4 \Delta p + \beta_5 t$. Table 2 presents the cointegration estimation results. The top panel presents the eigenvalues and the maximal and trace test statistics. The eigenvalue trace statistics indicate the existence of two cointegrating vectors at the 5 percent significance level. The maximal eigenvalue statistics indicate the existence of only one cointegration vector (significant at the 1 percent level). Based on the above statistics, and to ensure that our estimate does not ignore any potential cointegration vector, we proceed on the basis that two integration vectors are indicated by the data. The second panel of

[16]The lag length was determined by estimating a regular VAR using the above variables as follows: starting with 13 lags, Akaike information criteria and likelihood ratio (LR) tests were used for sequential lag reduction (see Enders, 1995).

[17]This restriction allows for linearly (but not quadratically) trending variables and ensures invariance to the trend coefficient in the cointegrating vector.

Table 3. Restricted Cointegration Estimation

I. Just-Identified System with Cointegration Vectors Restricted to Two

	α (Weighing Matrix)		β' (Cointegrating Vectors)				
Variable			$m-p$	R	y	Δp	Trend
$m-p$	–0.105 (0.025)	–0.362 (0.774)	1.000 (—)	0.000 (—)	–1.416 (0.594)	8.409 (1.670)	0.003 (0.001)
R	0.007 (0.002)	–0.566 (0.057)	0.000 (—)	1.000 (—)	0.033 (0.020)	–0.674 (0.056)	–0.000 (0.000)
y	0.004 (0.014)	-0.102 (0.447)					
Δp	–0.002 (0.007)	-0.226 (0.223)					

II. Final Restrictions on Elements of α and β'[1]

	α (Weighing Matrix)		β' (Cointegrating Vectors)				
Variable			$m-p$	R	y	Δp	Trend
$m-p$	–0.105 (0.018)	0.000 (—)	1.000 (—)	0.000 (—)	–1.000 (—)	8.650 (1.287)	0.002 (0.001)
R	0.000 (—)	0.077 (0.039)	0.000 (—)	1.000 (—)	0.000 (—)	–0.709 (0.045)	–0.000 (—)
y	0.000 (—)	0.000 (—)					
Δp	0.000 (—)	0.715 (0.155)					

Notes: Estimation period is 1984:3–1977:6. Numbers in parenthesis indicate standard errors.
[1]The χ^2 statistic corresponding to the final restriction is $\chi^2(8) = 7.8301$ [0.4502]. The number

Table 2 presents the normalized weighing matrix α and the matrix of potential cointegrating vectors β'. The two normalized cointegration vectors are indicated by the first and second rows of β' (normalized on $m-p$ and R, respectively). The top panel of Table 3 presents α and β' from the (just-identified) restricted cointegration estimation, where the number of cointegrating vectors is restricted to two, and the coefficient on R in the first vector and on $m-p$ in the second vector are restricted to zero. The bottom panel in Table 3 presents the final restricted cointegration estimation with zero restrictions on certain elements of the α and β' matrices that could not be rejected at the 10 percent level; the restrictions were tested individually and combined using χ^2 statistics.

The first cointegration vector appears to reflect deviations from long-term currency demand; it has the expected coefficient signs on y and Δp. The unit restriction on the (long-term) income elasticity of real currency demand could not be rejected at the 5 percent significance level.[18] The semi-elasticity of inflation is in line with other results in the literature; a 1 percentage point change in the monthly inflation rate results in 8.7 percent change in real currency demand.[19] The nominal interest rate does not affect long-term currency demand: The zero restriction on the interest rate coefficient could not be rejected at the 5 percent level. The trend is also significant and implies a gradually decreasing demand for real currency over time.[20] Based on the restricted model, the estimated long-run real currency demand is:

$$m - p = -0.002{*}t + y - 8.650\,\Delta p. \tag{1}$$

The second cointegration vector describes a stationary relationship between R and Δp alone, thus confirming a stationary real interest rate.[21] A stationary real interest rate further supports the previous result that the nominal interest rate is not significant in the determination of the long-run demand for currency: Given a stationary real interest rate, R and Δp move closely together (see Figure 3), thus revealing similar information in the long run; that is, long-run changes in R (given a stationary real exchange rate fluctuating around a constant mean) merely reveal information on the changes in inflation expectations. It is reasonable, therefore, that the long-run estimation of currency demand would include either R or Δp, but not both variables together.

Zero restrictions on the first column of the α matrix indicate that weak exogeneity cannot be rejected for R, y, and Δp. In other words, any deviation from the long-run equilibrium for real currency demand feeds back only into real currency demand.[22] Zero restrictions on the second column of the α matrix indicate that deviations from the long-run real interest rate feeds back only into inflation and the nominal interest rate.

III. The Error Correction Model

The estimated cointegration relationship reveals factors affecting long-term real currency demand. In the short run, deviations from this relationship could occur, reflecting shocks to any of the relevant variables. Furthermore, the dynamics

[18]Ahumada (1994) estimates a Tobin-Baumol income elasticity (0.5) for currency demand in Argentina for the period 1977–88.

[19]See Laidler (1985). For Argentina, Ahumada (1994) finds a similar result where inflation dominates the interest rate effect in the long run. Ahumada's estimate of the long-run semi-elasticity for inflation is 2.3.

[20]One possible interpretation is that the trend is proxying for the maximum historical inflation rate; that is, a ratchet effect that increases over time (see Piterman, 1988; Kamin and Ericsson, 1993).

[21]Zero restrictions on the coefficient of m and y in the second vector cannot be rejected at the 5 percent level. The restriction of equal coefficients on R and Δp was rejected at the 10 percent level. The estimated real interest rate is therefore $R - 0.71\,\Delta p$. The term comprising current inflation could be viewed as a measure of expected inflation.

[22]It is useful to note that in the presence of weak exogeneity, a single equation approach for the determination of the cointegration relationship would also be appropriate. However, in the absence of weak exogeneity, once the relevant hypothesis on β is tested using a full system, one can move to a single equation estimation, in which one can interpret and make the usual inference on the remaining parameters, keeping β fixed (Johansen, 1994).

governing the short-run behavior of currency demand (that is, the short-run elasticities of currency demand) are different from those in the long run. Engle and Granger (1987) showed that if a cointegrating relationship between nonstationary variables exists, an error correction representation of the data must exist. In this section, based on the estimation of the cointegration relationship between $m-p$, Δp, and y, we proceed with the estimation of the error correction representation, taking into account both the deviations from the long-run relationship and the short-run dynamics of real currency demand. In this representation, short-term dynamics are modeled by estimating in first differences. Adjustments in response to the deviation of real currency demand from the long-run trend are taken into account by including the error correction term estimated in the previous section. The vector describing deviations from the long-run real interest rate (the second cointegration vector) is not included in the currency demand equation, as the feedback coefficient for that vector in the currency equation (α) was not significantly different from zero. The stability of the estimated error correction model (in the whole estimation period and also specifically under the financial crisis) is discussed in the next section.

The error correction model was estimated for the period 1983:4–1997:6. The model was initially estimated by including 13 lags for all variables ($\Delta(m-p)$, ΔR, Δy, and $\Delta\Delta p$) in addition to the lagged error correction term and monthly dummies. Sequential reduction in the lag length of each variable was then carried out based on the significance of each lag, as well as the significance of the combined lags for each variable. The sequential reduction and reparametrization resulted in the following conditional model:

$$\Delta(m-p)_t = C + \Sigma\alpha_i MD_i + \sum_{i=1}^{11}\beta_i\Delta(m-p)_{t-i} + \gamma\, EC_{t-1} + \sum_{i=0}^{2}\delta_i\Delta\Delta_6 p_{t-i} + \sum_{i=12}^{13}\theta\,\Delta R_{t-i} \quad (2)$$

where C denotes a constant, MD_i denotes monthly dummies for January through November, and EC denotes the error correction term. The term $\Delta\Delta_6 p_{t-i}$ denotes $\Delta\Delta p_{t-i} - \Delta\Delta p_{t-i-6}$. Table 4 presents the results of the estimation.[23] All coefficients have the expected signs. The F-statistics indicate that all coefficients are significant at the 1 percent level. The diagnostic statistics listed in the table indicate that the equation is well specified; none of the statistics are significant at the 5 percent significance level. The residuals appear to be white noise (*AR F*), homoskedastic (*ARCH F*), and normally distributed (*NORM* χ^2). Figure 5 shows the residuals of the estimated equation. Except for six observations throughout the whole period, the residuals are within two standard errors from their zero mean.[24]

[23]The detailed specification of individual lags is presented in the Appendix. The possibility of an additional effect of exchange rate changes on real currency demand was tested using the *LM* statistic for omitted variables. The results indicate that exchange rate changes do not have a significant effect.

[24]The model underestimates currency demand in most of 1994. This could be related to some shifting of checking account holdings into cash due to the imposition of transaction fees on checking accounts in 1994, in addition to a decline in the use of credit cards due to the adoption of stricter regulations on overdue balances by banks (Bank of Mexico, 1995).

Table 4. The Error Correction Model

Dependent variable: $\Delta(m-p)$
Sample period: 1984:3–1997:6

	Coefficient Estimate	F-Statistics	Lags Included
$\Sigma\Delta(m-p)$	–0.770 (0.254)	F(11,131) = 13.60**	1–11
$\Sigma\Delta\Delta_6 p$	–2.2 (0.356)	F(3,131) = 13.73**	0–2
$\Sigma\Delta R$	–4.790 (1.000)	F(2,131) = 11.44**	12 and 13
EC_{t-1}	–0.091 (0.017)	F(1,131) = 28.77**	N/A
$R^2 = 0.904$	$\sigma = 2.63\%$		
AR $F(7,124)$ = 1.719 [0.111]	*DW* = 1.99	*ARCH* $F(7,117)$ = 0.697 [0.674]	
X_i^2 F(45,85) = 1.283 [0.161]	*RESET* $F(1,130)$ = 0.001 [0.972]	*NORM* χ^2 (2) = 1.691 [0.429]	

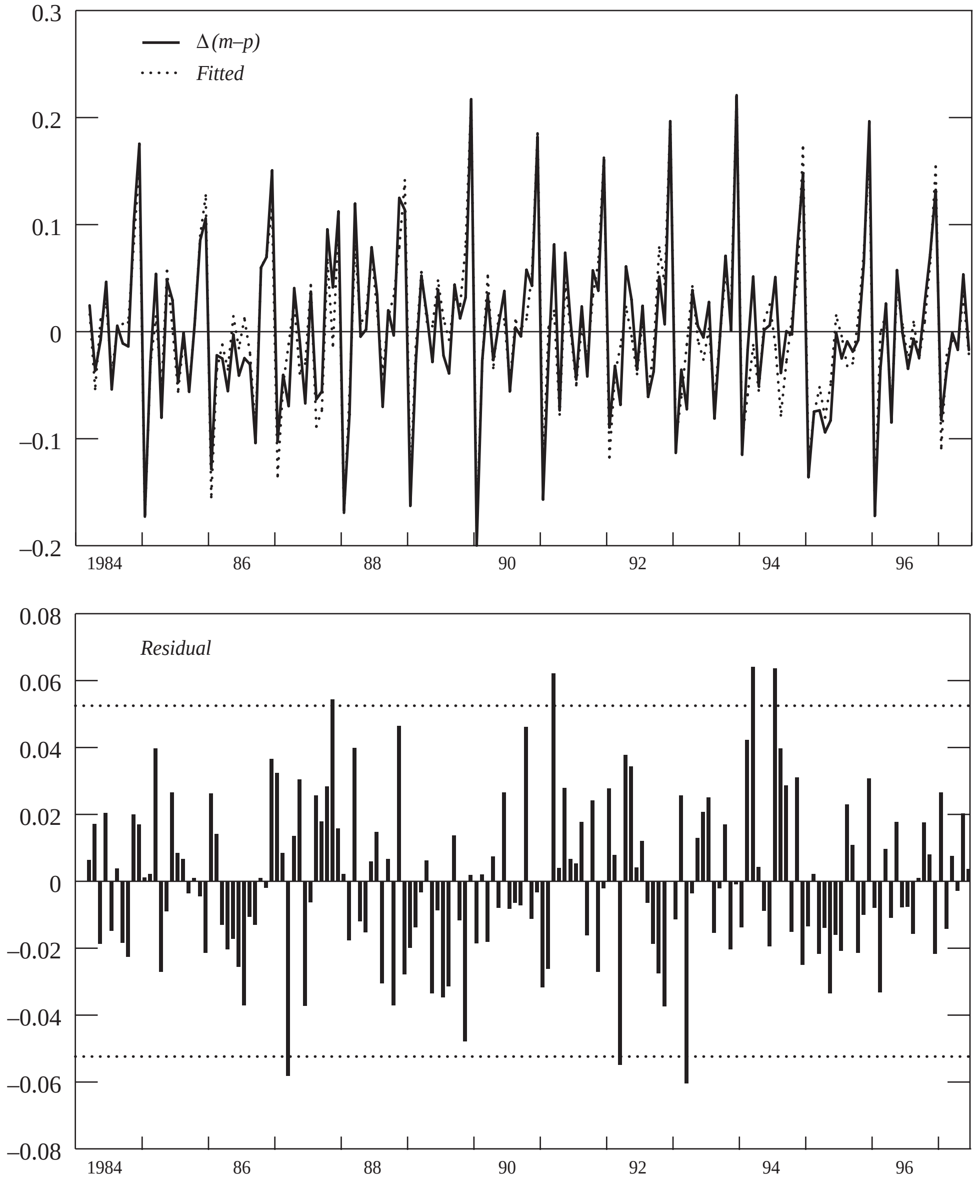
Notes: Numbers in parentheses indicate standard errors. N/A indicates not applicable. The symbols * and ** indicate significance at the 5 percent and the 1 percent level, respectively. Numbers in brackets indicate the significance level of the corresponding statistic. $\Delta\Delta_6 x = \Delta\Delta x_t - \Delta\Delta x_{t-6}$. *AR* $F(q,T-K-q)$ is the LaGrange multiplier (*LM*) statistic for the *q*-th order autocorrelation (Harvey, 1981); *ARCH* $F(q,T-K-q)$ is the *LM* statistic for the *q*-th order autoregressive conditional heteroskedasticity (Engle, 1982); *NORM* χ^2 is the Jarque and Bera (1980) statistic for the normality of the residual; X_i^2 $F(q,T-K-q)$ is White's (1980) statistic for heteroskedasticity quadratic in regressors; *RESET* $F(q,T-K-q)$ is Ramsey's (1969) statistic for nonlinearity (functional form misspecification); and *DW* is the Durbin-Watson statistic for serial correlation.

The estimated equation has a clear economic interpretation. Agents determine their real holdings of currency in the long run based on transaction needs and the opportunity cost of holding currency (the inflation rate). In the short run, they adjust their holdings by 9.1 percent of the past month's deviation from equilibrium. In addition to this disequilibrium effect, agents respond with a lag to interest rate changes and also to changes in inflation. It is notable that the transaction level does not seem to affect short-run demand for real currency. This, however, could be due to the approximation of monthly real private consumption expenditures using quarterly data. Finally, the coefficient estimates on the monthly dummies (see Appendix) indicate a notable increase in demand for real currency in December.

IV. Parameter Stability Under the Financial Crisis

Parameter constancy is an additional, crucial property to ensure a well-specified equation. The potential for parameter instability increases significantly during (and possibly after) a financial crisis, where the effect of the traditional determinants of currency demand could change and other variables could become significant

Figure 5. Error Correction Model: Regression Residuals



(such as the general confidence level in the economy or the rate of currency devaluation). In this section, we evaluate the constancy of the parameters during and after the financial crisis using a number of statistics.[25] First, we evaluate the stability of the estimated relationship over the entire estimation period, including the financial crisis using the one-step up and the break point Chow tests. Parameter constancy is also confirmed by the sequence of parameter estimates using OLS recursive estimation. Second, using the forecast Chow and χ^2 statistics, we evaluate the constancy of the parameters for two periods: the first 13 months of the crisis covering the period 1994:12–1995:12, and the entire period since the onset of the financial crisis in December 1994 (that is, 1994:12–1997:6).[26]

Figures 6 and 7 show the series of recursive estimates of parameters (and an interval of $\pm 2\sigma$ around the estimates) attached to the main regressors. These estimates are well inside the standard errors and become more accurate with time as more information is accumulated; the standard errors decrease and parameter estimates are more stable. Some parameters exhibit a small shift in the first quarter of 1995. The forecast Chow and χ^2 statistics that are presented below indicate that these shifts are not large enough to cause any significant parameter instability. To further confirm this result, individual parameter instability statistics based on Hansen (1992) were computed (see Appendix). All individual parameter statistics indicate parameter constancy. The top panel in Figure 8 shows the sequence of break point Chow statistics for the forecast sequence {1987:6–1997:6, 1987:7–1997:6, . . .1997:5–1997:6}. None is significant at the 5 percent level, indicating that constancy of the estimated parameters cannot be rejected for the whole sequence of forecasts. The middle panel in Figure 8 shows the sequence of one-step Chow tests. Again, only four points are above the 5 percent level (none of which occurs after the inception of the crisis in 1994). Both sequences of Chow tests confirm the constancy of the estimated parameters within the estimation period inclusive of the financial crisis. Furthermore, as indicated by the one-step residuals in the bottom panel, except for two observations (in 1991 and 1994) that lie slightly outside the range of $\pm 2\sigma$, all forecasted errors lie within the range.

For further confirmation of the stability of the parameter estimates during and after the financial crisis, we calculate the forecast Chow and χ^2 statistics for the periods 1994:12–1995:12 and 1994:12–1997:6. These statistics are reported in Table 5. The two statistics indicate that parameter constancy cannot be rejected at the 5 percent significance level for the two periods. Figure 9 shows the forecast $\Delta(m - p)$ versus the actual observations for the period starting with the financial crisis in December 1994 and their estimated standard errors ($\pm 2\sigma$). The second quarter of 1995 appears to display the largest forecast error in the whole period. Nevertheless, the actual observations still lie within the range of two standard errors of the forecasts. In conclusion, the above tests provide strong evidence of the stability of real currency demand in Mexico despite the substantial effects of the financial crisis on the Mexican economy.

[25]All tests presented below employ the null hypothesis of parameter constancy. The rejection of the null hypothesis implies the rejection of parameter constancy over the period tested.

[26]The latter is identical to the break point Chow statistic evaluated at 1994:12.

Figure 6. Recursive Parameter Estimates: $\Delta(m\text{–}p)$

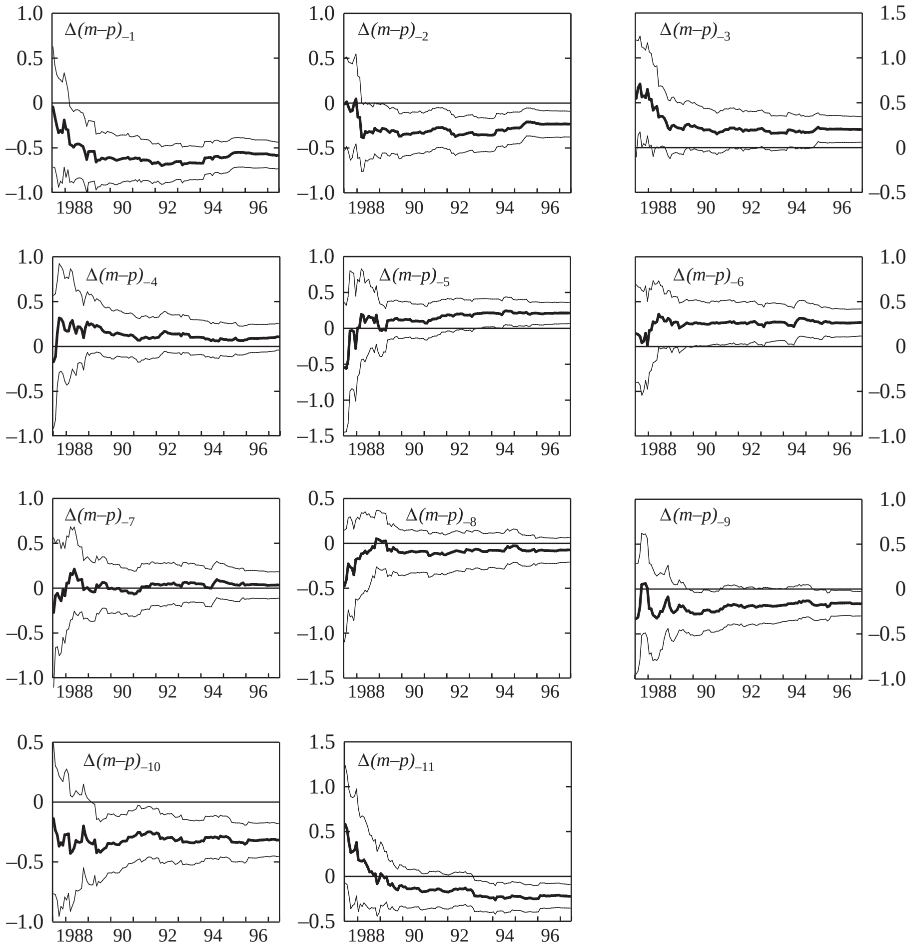


Figure 7. Recursive Parameter Estimates: ΔR, $\Delta\Delta p$, and *EC*

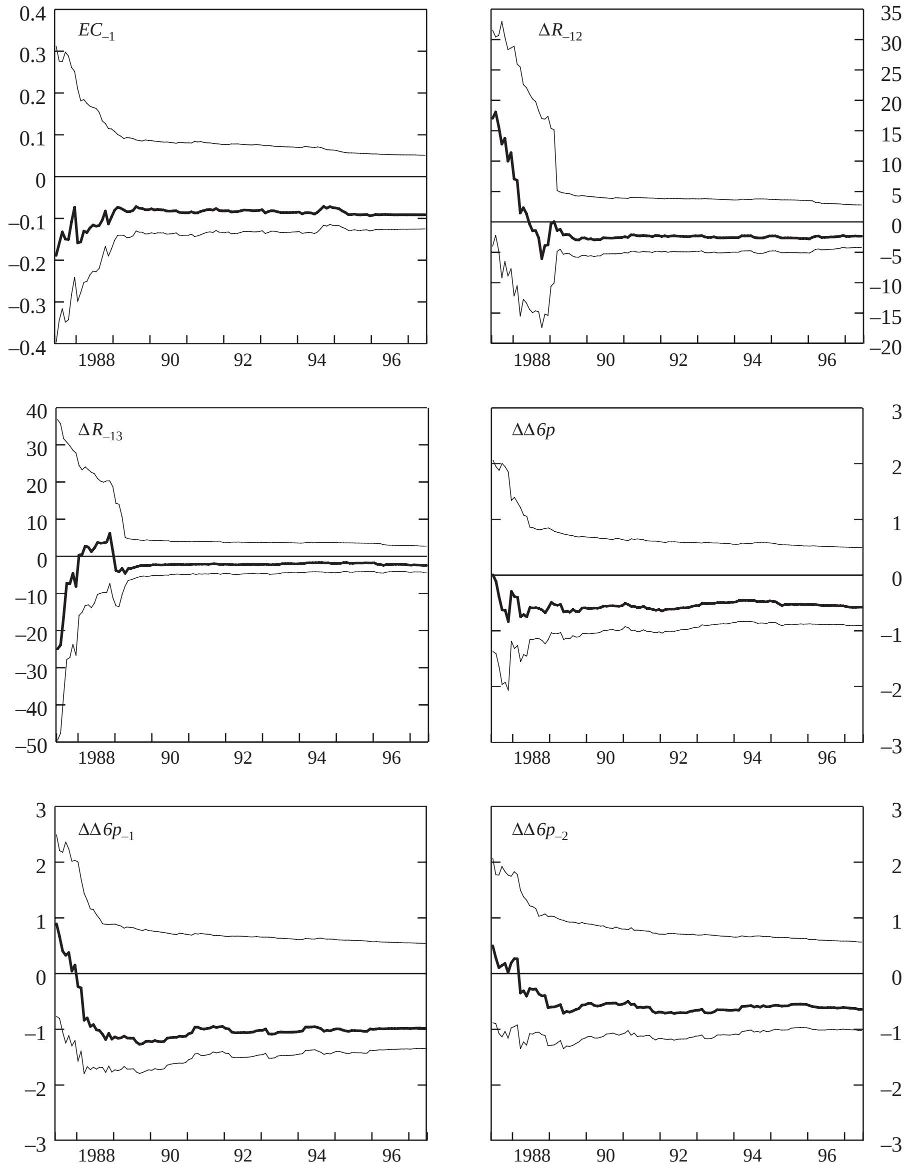


Figure 8. One-Step and Break Point Chow Tests and One-Step Residuals

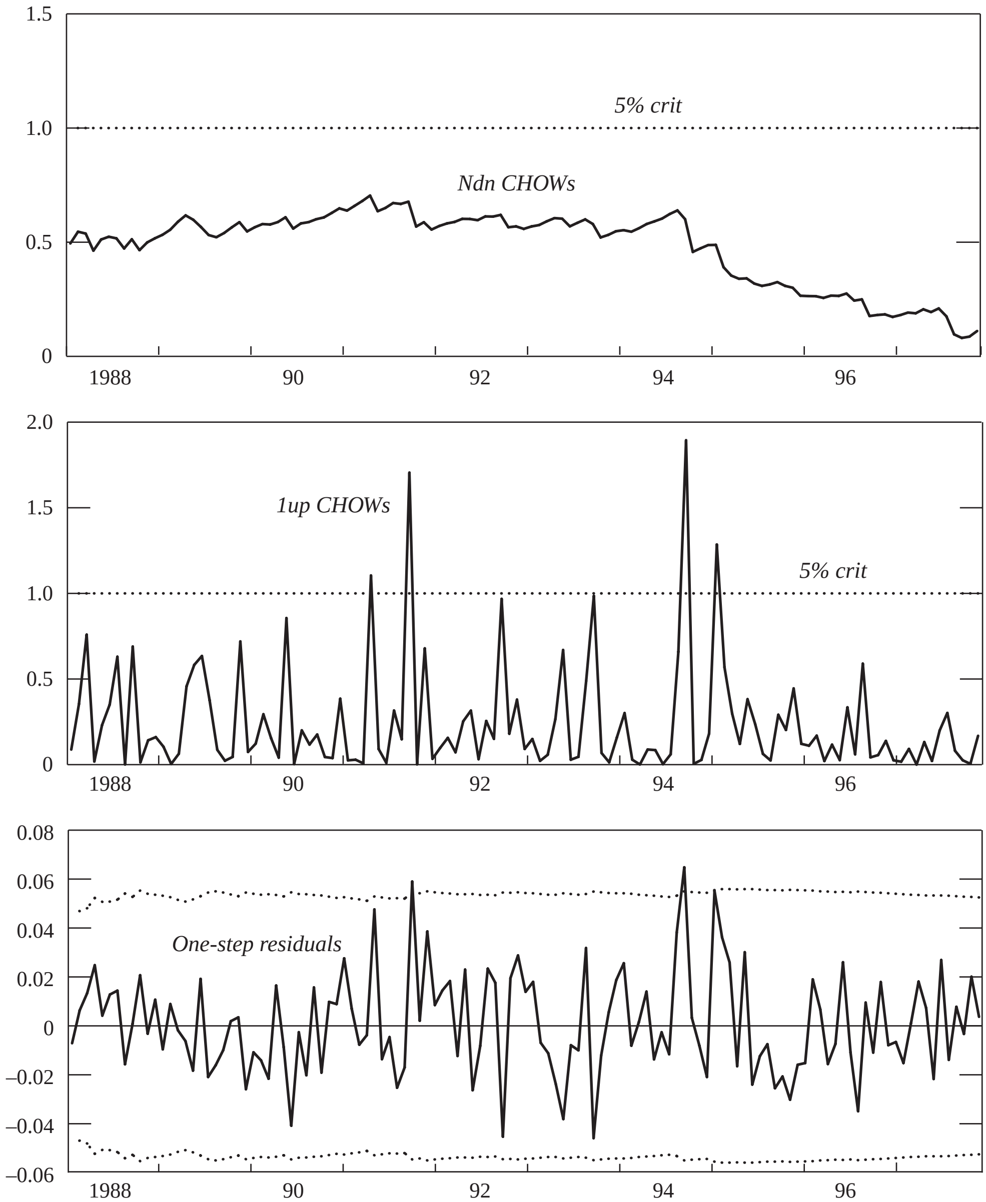
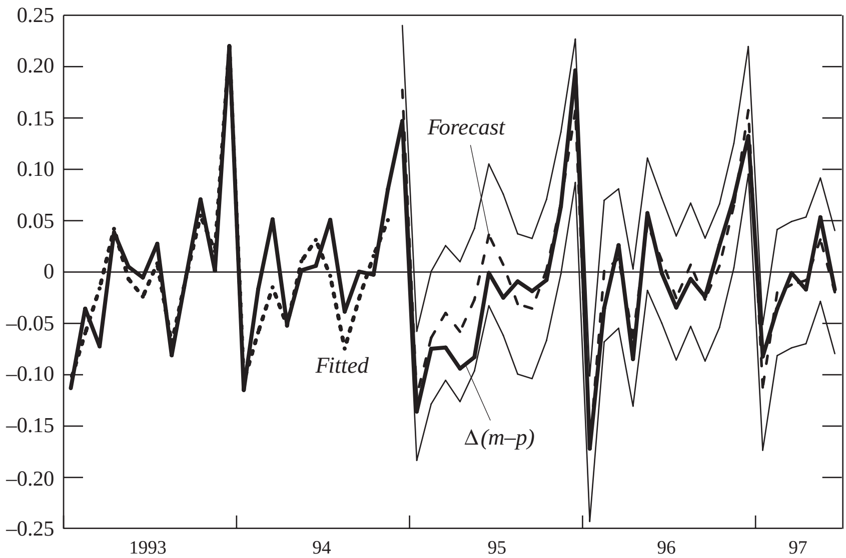


Table 5. Forecast Chow and χ^2 Statistics

	1994:12–1995:12	1994:12–1997:6
Chow F(. , .)	0.64 [0.31]	0.50 [0.99]
Forecast χ^2	14.91 [0.81]	21.79 [0.89]

Note: Estimation period starts at 1984:3. The null hypothesis for both tests is that parameters in the original and forecast periods are equal. Numbers in brackets indicate the significance level of the corresponding statistic.

Figure 9. Actual Observations and Forecasts, 1993–97



V. Conclusions

The paper finds strong evidence that real currency demand in Mexico remained stable throughout and after the financial crisis in Mexico. Cointegration analysis using the Johansen-Juselius technique indicates a strong cointegration relationship between real currency balances, real private consumption expenditures, and inflation. The dynamic model for real currency demand exhibits significant parameter constancy even after the financial crisis as indicated by a number of statistical tests. We therefore conclude that the significant reduction in real currency demand related to the financial crisis in Mexico could be appropriately explained by the change in the variables that historically explain the demand for real currency balances in Mexico. This result confirms that the BOM's use of a reserve money program to implement monetary policy under the financial crisis was appropriate.

APPENDIX

The Error Correction Model, 1984:3–1997:6
Explanatory variable: $\Delta(m-p)$

	Coefficient Estimate	Standard Error	Hansen's Instability Statistic
$\Delta(m-p)_{t-1}$	–0.584**	0.074	0.14
$\Delta(m-p)_{t-2}$	–0.234**	0.072	0.15
$\Delta(m-p)_{t-3}$	0.202**	0.071	0.09
$\Delta(m-p)_{t-4}$	0.108	0.074	0.03
$\Delta(m-p)_{t-5}$	0.214**	0.073	0.27
$\Delta(m-p)_{t-6}$	0.267**	0.075	0.14
$\Delta(m-p)_{t-7}$	0.035	0.074	0.12
$\Delta(m-p)_{t-8}$	–0.072	0.068	0.08
$\Delta(m-p)_{t-9}$	–0.163*	0.068	0.10
$\Delta(m-p)_{t-10}$	–0.318**	0.069	0.20
$\Delta(m-p)_{t-11}$	–0.222**	0.066	0.10
$\Delta\Delta_6 p_t$	–0.576**	0.164	0.07
$\Delta\Delta_6 p_{t-1}$	–0.982**	0.180	0.05
$\Delta\Delta_6 p_{t-2}$	–0.642**	0.189	0.23
ΔR_{t-12}	–2.348*	0.926	0.08
ΔR_{t-13}	–2.448**	0.912	0.28
EC_{t-1}	–0.091**	0.017	0.15
Constant	0.139**	0.016	0.05
MD_1	–0.171**	0.020	0.16
MD_2	–0.244**	0.023	0.10
MD_3	–0.282**	0.026	0.04
MD_4	–0.236**	0.026	0.11
MD_5	–0.238**	0.021	0.18
MD_6	–0.235**	0.021	0.09
MD_7	–0.181**	0.022	0.24
MD_8	–0.193**	0.021	0.11
MD_9	–0.182**	0.025	0.17
MD_{10}	–0.126**	0.025	0.06
MD_{11}	–0.126**	0.022	0.18

+Significant at the 10 percent level; * significant at the 5 percent level; ** significant at the 1 percent level. MD_i = 1–11 denotes monthly dummies for January through November.

REFERENCES

Aboumrad, A. J. Guillermo, 1996, “Instrumentación de la Política Monetaria con Objetivo de Estabilidad de Precios: El Caso de México,” *Monetaria,* Vol. 1 (January–March), pp. 69–114.

Ahumada, Hildegart, 1994, “A Dynamic Model of the Demand for Currency: Argentina 1977–1988,” in *Testing Exogeneity: Advanced Texts in Econometrics* (Oxford: Oxford University Press), pp. 191–218.

Arrau, Patricio, and José De Gregorio, 1993, “Financial Innovation and Money Demand: Application to Chile and Mexico,” *The Review of Economics and Statistics*, Vol. 75 (August), pp. 524–30.

Bank of Mexico, 1994, *The Mexican Economy 1994—Economic and Financial Developments in 1993, Policies for 1994* (Mexico City: Bank of Mexico).

———, 1995, *The Mexican Economy 1995—Economic and Financial Developments in 1994, Policies for 1995* (Mexico City: Bank of Mexico).

Choudhry, Taufiq, 1995, “High Inflation Rates and the Long-Run Money Demand Function: Evidence from Cointegration Tests,” *Journal of Macroeconomics,* Vol. 17 (Winter), pp. 77–92.

De Lemos Grandmont, Renato, 1991, “Multivariate Cointegration in the Presence of Structural Breaks: The Case of Money Demand in Mexico,” Research Paper No. 4895 (Montreal; University of Montreal, Center for Research and Economic Development).

Desentis, Samuel Alfaro, 1997, “La Demanda Oportuna de Billetes y Monedas en México,” *Gaceta de Economía,* Instituto Tecnológico Autónomo de México, Año 3, No. 5, pp. 265–81 (Mexico City: Otoño).

Enders, Walter, 1995, *Applied Econometric Time Series* (New York: John Wiley and Sons).

Engle, R. F., 1982, “Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation,” *Econometrica*, Vol. 50 (July), pp. 987–1007.

———, and C.W. J. Granger, 1987, “Cointegration and Error Correction: Representation, Estimation, and Testing,” *Econometrica,* Vol. 55 (March), pp. 251–77.

Friedman, M., and A. J. Schwartz, 1982, *Monetary Trends in the United States and the United Kingdom: Their Relation to Income, Prices, and Interest Rates, 1867–1975* (Chicago: University of Chicago Press).

Hall, R. E., 1978, “Stochastic Implications of the Life Cycle Permanent Income Hypothesis: Theory and Evidence,” *Journal of Political Economy,* Vol. 86, No. 6, pp. 971–87.

Hansen, B. E., 1992, “Testing for Parameter Instability in Linear Models,” *Journal of Policy Modeling,* Vol. 14 (August), pp. 517–33.

Harvey, A. C., 1981, *The Econometric Analysis of Time Series* (Oxford: Phillip Allan).

Hendry, D., and J. F. Richard, 1982, “On the Formulation of Empirical Models in Dynamic Econometrics,” *Journal of Econometrics*, Vol. 20, No. 1, pp. 3–33.

———, 1983, “The Econometric Analysis of Time Series,” *International Statistical Review,* Vol. 51 (August), pp. 111–63.

Jarque, C. M., and A. K. Bera, 1980, “Efficient Tests for Normality, Homoscedasticity, and Serial Independence of Regression Residuals,” *Econometric Letters,* Vol. 6, No. 3, pp. 255–59.

Johansen, S., 1988, “Statistical Analysis of Cointegration Vectors,” *Journal of Economic Dynamics and Control,* Vol. 12, No. 2/3, pp. 241–49.

———, 1994, “Testing Weak Exogeneity and the Order of Cointegration in U.K. Money Demand Data,” in *Testing Exogeneity: Advanced Texts in Econometrics* (Oxford: Oxford University Press), pp.121–43.

———, and K. Juselius, 1990, “Maximum Likelihood Estimation and Inference on Cointegration—With Application to the Demand for Money,” *Oxford Bulletin of Economics and Statistics*, Vol. 52, No. 2, pp. 169–210.

Kamin, S., and N. Ericsson, 1993, “Dollarization in Argentina,” International Finance Discussion Paper No. 460 (Washington: Board of Governors of the Federal Reserve System).

———, and J. Rogers, 1996, “Monetary Policy in the End-Game to Exchange-Rate Based Stabilizations: The Case of Mexico,” *Journal of International Economics*, Vol. 41 (November), pp. 285–307.

Laidler, David, 1985, *The Demand for Money: Theories, Evidence, and Problems* (New York: Harper and Row).

Piterman, S., 1988, “The Irreversibility of the Relationship Between Inflation and Real Balances,” *Bank of Israel Economic Review*, Vol. 60 (January), pp. 72–83.

Ramos-Francia, Manuel, 1993, “The Demand for Money in an Unstable Economy: A Cointegration Approach for the Case of Mexico,” Discussion Paper Series No. 9306 (Mexico City: Centro de Investigación Económica).

Ramsey, J. B., 1969, “Tests for Specification Errors in Classical Linear Least-Squares Regression Analysis,” *Journal of the Royal Statistical Society, Series B*, Vol. 31, No. 2, pp. 350–71.

Rogers, H. John, 1992, “The Currency Substitution Hypothesis and Relative Money Demand in Mexico and Canada,” *Journal of Money, Credit, and Banking*, Vol. 24, No. 3, pp. 300–18.

Thornton, John, 1996, “Cointegration, Error Correction and the Demand for Money in Mexico,” *Weltwirtschaftliches Archiv.*, Vol. 132, No. 4, pp. 690–99.

White, H., 1980, “A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity,” *Econometrica*, Vol. 48, No. 4, pp. 817–38.