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

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Abstract

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.

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I. Introduction

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. This had significant implications for the implementation of monetary policy where the exchange rate no longer could provide the nominal anchor for the economy. Accordingly, the primary objective of the monetary program for 1995 was attaining price level stability through offsetting the inflationary effects of the devaluation. Consistent with this target, 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. As reserve requirements in Mexico were abolished in 1988, the demand for currency comprises much of the demand for reserve money.[2] Therefore 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 under the financial crisis.[3]

Despite its significance for the conduct of monetary policy, the stability of money demand is generally unaddressed 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 et. al.[4] 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 (1995), Aboumrad (1996), Thornton (1996), and Desentis (1997). The latter two come closest to addressing this issue. Thornton examines the

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[2]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.

[3]The Bank of Mexico's monetary management framework is based on formulating an annual target for reserve money, by projecting the demand for real reserve money (currency plus banks' current accounts at the central bank), and taking into account the inflation target. Therefore, in addition to an appropriate projection of the real demand for currency, determining the reserve target would also involve the projection of banks' settlement balances at the central bank. Central bank credit is managed so that, in principle, the supply of reserve money would meet its projected demand. The (nominal) reserve money target is periodically revised in response to changes in certain indicators including the evolution of the exchange rate, observed inflation vís-a-vís targeted inflation, the evolution of inflation expectations, and developments in the settlement of wages.

[4]See Hendry and Richard (1982 and 1983).

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.[5]

In this paper, we address the above issues and examine whether the real demand for currency remained stable during and after the financial crisis in Mexico. 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 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.[6] The study period contains the inflationary debt crisis period, the stabilization period under the December 1987 stabilization plan (the Pacto) and the ensuing financial crisis in December 1994, and ends with 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. This paper finds strong evidence of long-term stability of real currency demand indicated by the cointegration of real currency, private consumption expenditures and the deposit rate. Since inflation and interest rates move together in the long-run (the results also indicate a cointegration relationship between these two variables), the changes in the interest rate encompass changes in inflation. This explains the insignificance of inflation 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—the 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.[7]

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[5]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.

[6]To detect any mis-specification 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.

[7]This, however, could be related to the use of quarterly consumption data to construct the monthly series. The constructed series does not reflect intra-quarter changes in actual real

(continued...)

The paper is organized as follows. Section II describes the data used in the analysis. Section III presents the cointegration results. Section IV proceeds with the error correction model estimation. The stability of real currency demand during and after the financial crisis is presented in Section V. Section VI concludes.

II. 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).[8] 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.[9] The CPI inflation rate (π) and the interest rate on 60-day time deposits (i) were used as estimates for the opportunity cost of holding currency (as opposed to holding real and financial assets, respectively).[10] 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.[11] The interest rate on one-month CETES

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[7](...continued)
consumption where quarterly consumption data was repeated for each month of the same quarter.

[8]The data used in this study is from the International Financial Statistics (published by the International Monetary Fund.)

[9]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.

[10]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 (i and π) used in this paper denote monthly rates (unannualized). Except for i and π, all lower case variables denote the natural logarithm of the original variables.

[11]Figure 1 shows that these rates generally move together and therefore the estimation results
(continued...)

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 CETES rate and the inverse of real balances (i.e., (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 since bank deposit interest rates generally require a period of two to three months to adjust to changes in CETES rates (Figure 1).

Figure 3 presents the series for m-p, y, i and π for the period 1983:1-1997:6. All four series reflect the major macroeconomic episodes in this period. 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, 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, i and π start a downward trend. The graph indicates a clear positive relationship between m-p and y, and a negative relationship between m-p and both i and π, supporting the likelihood of a long-term cointegration relationship among these variables. Although the series reflect breaks during the study period, cointegration is still possible mainly because these breaks are present in all four 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:

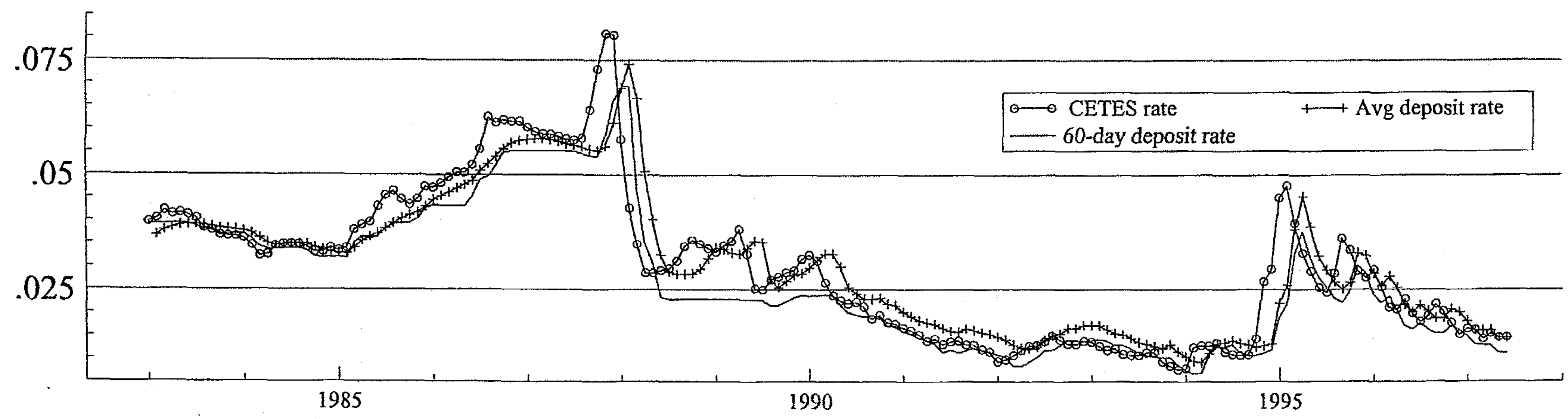
$$\frac{M^d}{P} = \gamma_0 Y^{\gamma_1} e^{(\gamma_3 i + \gamma_4 \pi)}$$

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. All variables

Figure 1. Interest Rates

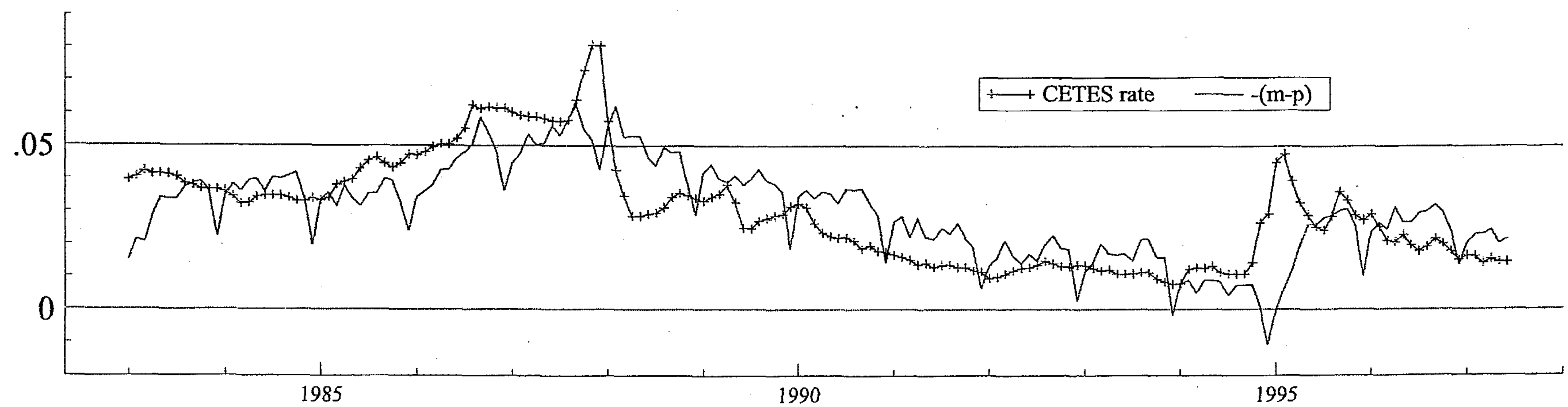
[11](...continued)
should not be affected by which rate is used in the specification.

Figure 1. Interest Rates



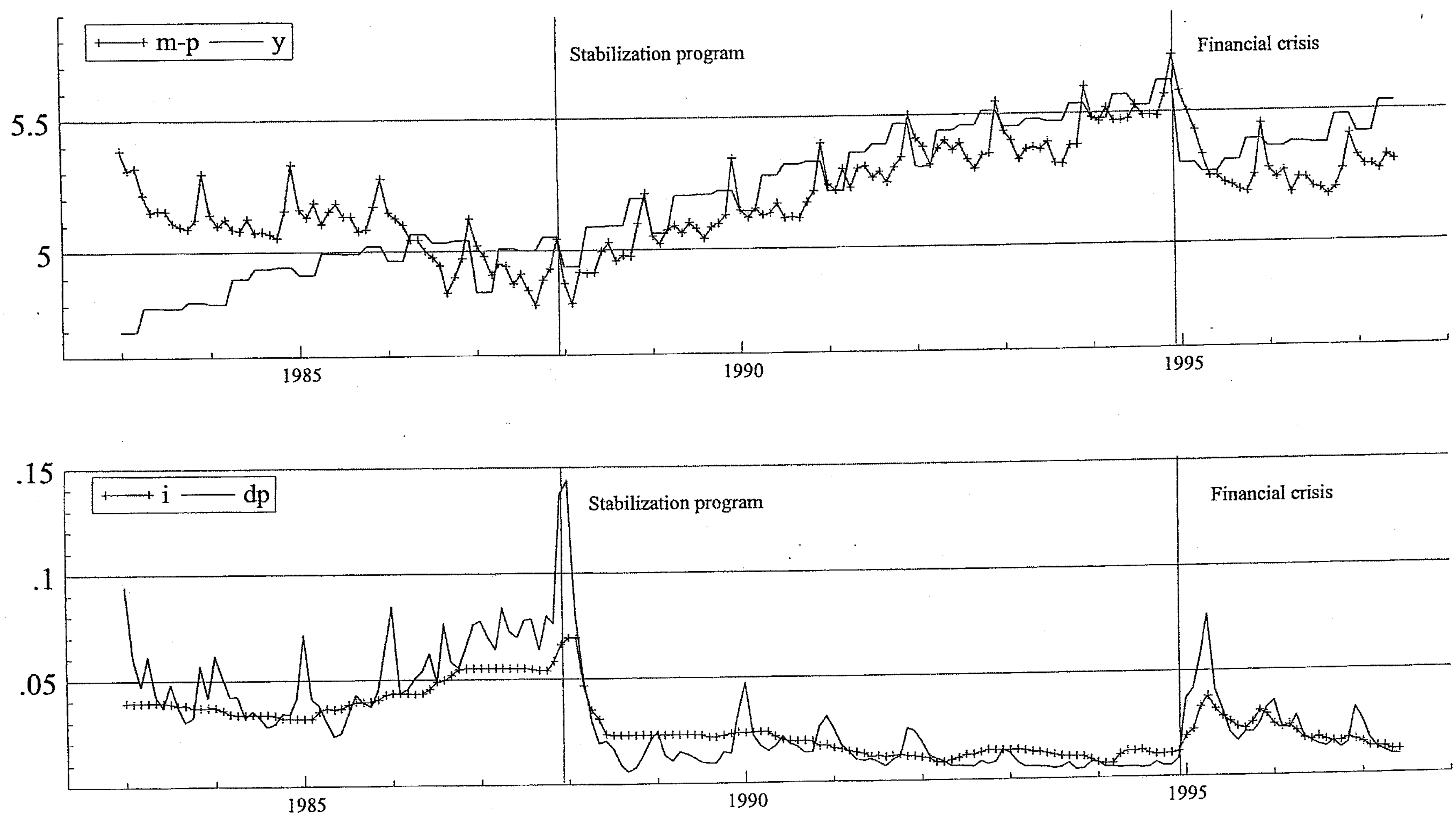
Source: Bank of Mexico and the International Financial Statistics, International Monetary Fund.

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



Source: The International Financial Statistics, the International Monetary Fund.

Figure 3. Real currency, Real Private Consumption Expenditures, the Interest Rate and Inflation, 1983:1-1997:6



Source: The International Financial Statistics, International Monetary Fund.

(except for inflation) were found to be integrated of order one—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.[12]

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

Variable	ADF Statistic	Lags	Constant Included	Trend Included
m-p	-2.39	3	yes	yes
π[1]				
1983:1 - 1997:6	-3.02**	0	yes	no
1983:6 - 1997:6	-2.72	0	yes	no
i	-2.60	1	yes	yes
y	-2.10	0	yes	yes
Δ(m-p)	-6.42 ***	2	yes	no
$\Delta(\pi)$	-3.82 ***	11	no	no
Δi	-7.69 ***	1	no	no
Δy	-2.91 ***	11	no	no

1/The test for inflation is inconclusive; as indicated in the table, tests using sample sub-periods could not reject the null hypothesis of a unit root.
Note: The test period for all variables (except if indicated otherwise) is 1983:1-1997:6 minus the lags. ** and *** imply rejection of the null of a unit root at the 5 percent and 1 percent level respectively.

III. Long-Run Behavior and Cointegration

In this section, we examine the presence of cointegration between m-p, i, π, and y using the Johansen-Juselius procedure (Johansen (1988), and 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.

The Johansen-Juselius system-based procedure was applied to m-p, y , i, and π, with a constant and monthly dummies, and 13 lags. The lag length of the system was determined by estimating a regular VAR using the above variables; sequential lag reduction with Akaike

[12]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.

information criteria and likelihood ratio (LR) tests were used to determine the lag length (see Enders, 1995).[13] The constant was restricted to lie in the cointegration space. Accordingly, the hypothesized cointegration vector(s) is of the form: $\beta_1 m+\beta_2 y+\beta_3 i+\beta_4 \pi+a$.[14]

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 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 and π respectively). The top panel of Table 3 presents α and β' from the restricted cointegration estimation, where the number of cointegrating vectors is restricted to two. The coefficients on all variables remain the same as in the unrestricted estimation. 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 i (-0.45 and 9.73 respectively). Accordingly, the estimated income elasticity of real currency demand is smaller than one (0.45) and is close to the elasticity implied by the Tobin-Baumol model of 0.5.[15] The interest rate semi-elasticity is in line with other results in the literature; a 1 percentage point change in monthly interest rate

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[13]Additionally, lags were not omitted if their exclusion introduced serial correlation in the residuals.

[14]This restriction results in a general solution for the variables (in levels) that does not contain a time trend. That is, all changes in the variables are expected to be zero in equilibrium (when the cointegration vector $\beta_1 m+\beta_2 y+\beta_3 i+\beta_4 \pi+a$ is zero). Although the series for m and y seem to exhibit a trend , there have been, as mentioned above, major macroeconomic events that resulted in breaking this trend twice in the time period studied thus lessening the importance of the trend in this period. The presence of a trend in the series for π and i is even less clear. The chosen formulation is thus closely related to the data characteristics in the period studied. To confirm the above, a VAR including m-p, y , i, and π was estimated including a time trend; the time trend in the four equations was statistically insignificant at the 10 percent level.

[15]Ahumada (1994) estimates very similar income elasticity for currency demand in Argentina for the period 1977–88.

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	0.360	71.85**	28.10	107.00**	53.10
2	0.106	17.97	22.00	35.19*	34.90
3	0.081	13.52	15.70	17.22	20.00
4	0.023	3.70	9.20	3.70	9.20

II. Normalized α and β' Matrices

α (Weighing Matrix)

Variable				
m-p	-0.105	-0.362	-0.239	-0.007
π	0.007	-0.566	0.252	0.003
i	0.004	-0.102	0.040	-0.001
y	-0.002	-0.226	-1.767	0.003

β' (Cointegrating Vector)

Variable	m-p	π	i	y	Constant
	1.000	-3.136	13.573	-0.471	-4.027
	-0.027	1.000	-1.540	0.017	0.040
	0.021	-0.350	1.000	0.037	-0.357
	-1.417	-8.578	11.487	1.000	1.001

* and ** denote significance at the 5 percent and the 1 percent level, respectively.

Table 3. Restricted Cointegration Estimation

I. Cointegration Vectors Restricted to Two

α (Weighing Matrix)			β' (Cointegrating Vector)				
Variable			m-p	π	i	y	Constant
m-p	-0.105	-0.362	1.000	-3.136	13.573	-0.471	-4.027
π	0.007	-0.566	-0.027	1.000	-1.540	0.017	0.040
i	0.004	-0.102					
y	-0.002	-0.226					

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

α (Weighing Matrix)			β' (Cointegrating Vector)				
Variable			m-p	π	i	y	Constant
m-p	-0.094	0.000	1.000	0.000	9.733	-0.445	-4.360
π	0.024	0.802	0.000	-0.769	1.000	0.000	0.040
i	0.007	0.160					
y	0.000	0.000					

1/The χ^2 statistic corresponding to the restriction is $\chi^2(4) = 0.377$ [0.9843]. The number in brackets indicates the significance level.

results in 9.73 percent change in real currency demand.[16] Inflation does not affect long-term currency demand: the zero restriction on the inflation coefficient could not be rejected at the 5 percent level. These results indicate that despite the high instability of the Mexican economy in the period studied including the financial crisis in 1994–95, the demand for currency exhibits a stable long-term relationship. Based on the restricted model, the estimated long-run real currency demand is:

$$m - p = 4.36 + 0.45\,y - 9.73\,i$$

[16]See Laidler (1985). For Argentina, Ahumada (1994) finds that inflation dominates the interest rate effect in the long run. The estimated long-run semi-elasticity for inflation is 2.3.

The second cointegration vector describes a stationary relationship between i and π alone thus confirming a stationary real interest rate.[17] A stationary real interest rate further confirms the insignificance of inflation in the determination of the long-run demand for currency. Given a stationary real interest rate, i and π move closely together (Figure 3), thus revealing similar information in the long run; that is, long-run changes in i (given a stationary real exchange rate fluctuating around a constant mean) merely reveal information on the changes in inflation expectations. Therefore, it is reasonable that the long-run estimation of currency demand would include either i or π, but not both variables together.

Zero restrictions on the first column of the α matrix indicate that any deviation from the long-run equilibrium for real currency demand feeds back into real currency demand, the interest rate and inflation but not into real consumption. 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.[18]

IV. 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 governing the short-run behavior of currency demand (i.e., the short-run elasticities of currency demand) are different from those in the long-run. Engle and Granger (1987) showed that if there exists a cointegrating relationship between nonstationary variables, there must exist an error correction representation of the data. In this section, based on the estimation of the cointegration relationship between m-p, i, 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

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[17]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 i and π was rejected. The estimated real interest rate is therefore $i-0.77\pi+0.04$. The term comprising current inflation and the constant term could be viewed as a measure of expected inflation.

[18]Weak exogeneity is therefore rejected for π and i. A single equation approach for the determination of the cointegration relationship is therefore not appropriate. 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). Consistent with this, we use the cointegration relationship estimated from the system in modeling the (single equation) error-correction representation.

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 1983:1-1997:6 minus the included lags. The model was initially estimated by including 13 lags for all variables (Δ(m-p), i, y, and π) in addition to the lagged error correction term. The final lag structure was determined based on the significance of each lag as well as the significance of combined lags for each variable. The final model is:

$$\Delta(m-p)_t = \Sigma\, \alpha_i MD_i + \sum_{i=1}^{11} \beta_i\, \Delta(m-p)_{t-i} + \gamma\, EC_{t-1} + \sum_{i=1}^{7} \delta_i\, \Delta\pi_{t-i} + \theta\, \Delta i_{t-12}$$

Where MD_i denotes monthly dummies and EC denotes the error correction term.[19] Table 4 presents the results of the estimation.[20] 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 significant level. The residuals appear to be white noise (AR F), homoscedastic (ARCH F), and normally distributed (NORM χ^2). Figure 4 shows the residuals of the estimated equation. Except for three observations through the whole period, the residuals are within two standard errors from their zero mean.[21]

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[19]In addition to a constant term, dummies for January through November were included. In the final specification, only the dummy variables were significant; the constant term was not significant at the 10 percent level and thus was omitted from the specification.

[20]The detailed specification of individual lags is presented in the Appendix. The possibility of an additional effect of the 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.

[21]The model underestimates currency demand in most of 1994. This could be related to some shifting of checking accounts holdings into cash due to the imposition of transaction fees on checking account 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).

(continued...)

Figure 4. Error Correction Model: Regression Residuals

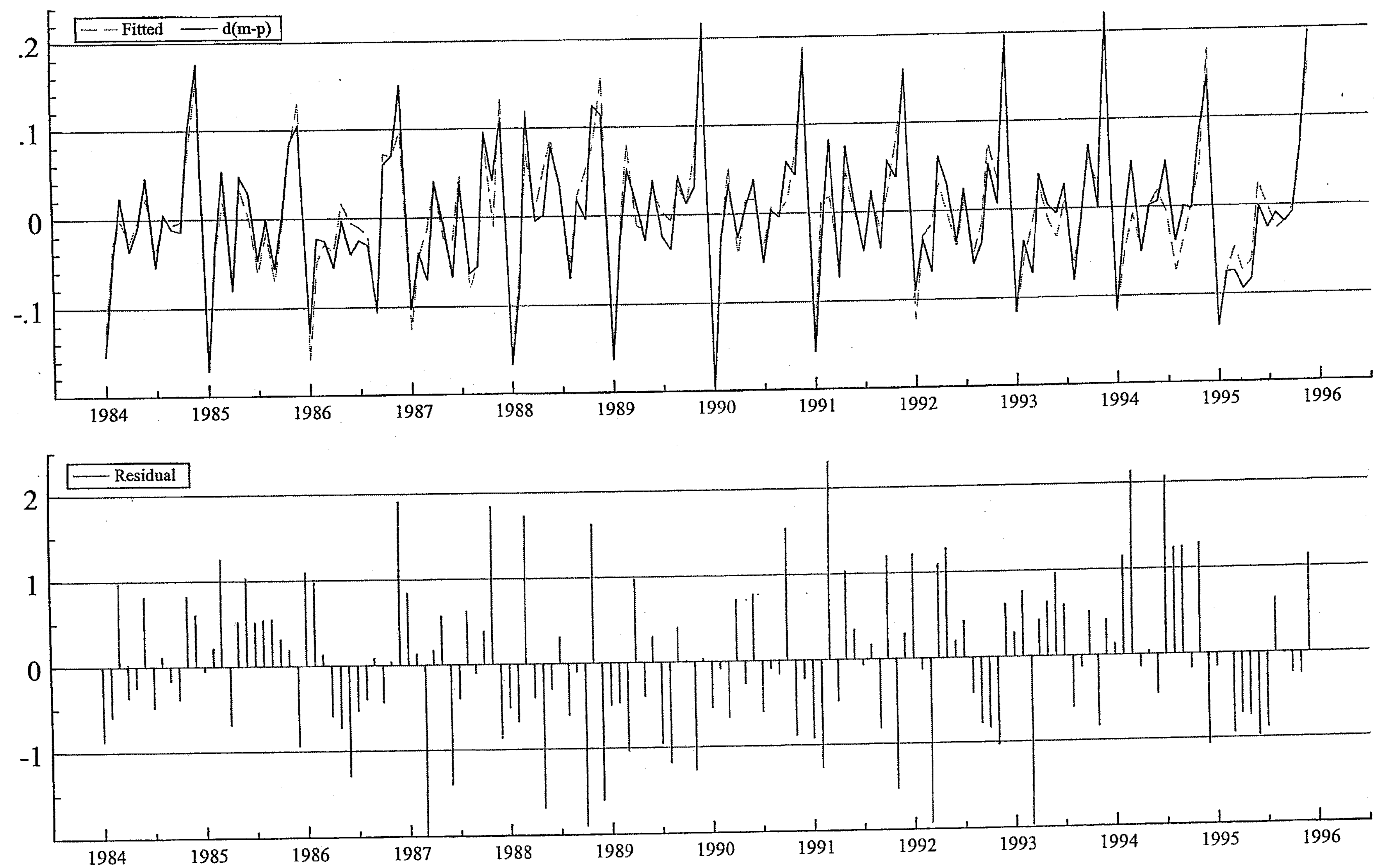

Table 4. The Error Correction Model, 1983:1-1997:6

Dependent variable: $\Delta(m\text{-}p)$

	Coefficient Estimate	F-Statistics	Lags Included
$\sum\Delta$(m-p)	-0.418	F(11,131)=10.36***	11
$\sum\Delta\pi$	-0.854	F(7,131)=8.67***	7
Δi_{t-12}	-3.400	F(1,131)=15.06***	na
EC_{t-1}	-0.114	F(1,131)=190.25***	na
$R^2 = 0.90$	$\sigma = 2.72\%$		
AR F(7,124) = 2.06 [0.0528]	DW=2.12	ARCH F(7,117) = 0.54 [0.801]	
X_i^2 F(51,79) = 1.32 [0.130]	RESET F(1,130) = 0.17 [0.681]	*NORM* $\chi^2 = 0.520$ [0.771]	

*** significant at the 1 percent level. Numbers in brackets indicate the significance level of the corresponding statistic.

Notes: 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 heteroscedasticity (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 heteroscedasticity quadratic in regressors, RESET F(q,T-K-q) is Ramsey's (1969) statistic for non-linearity (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 interest rate). In the short-run, they adjust their holdings by 11 percent of the past month's deviation from the 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.

V. 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 were the effect of the traditional determinants of currency demand could change and other variables could become significant (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.[22] First we evaluate the stability of the estimated relationship over the entire estimation period, including the financial crisis; i.e., 1983:1-1997:10 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 (i.e., 1994:12-1997:10).[23]

Figures 5 and 6 show the series of recursive estimates of parameters (and the 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 7 shows the sequence of break point Chow statistics for the forecast sequence {1987:8-1997:6, 1987:9-1997:6, ...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 7 shows the sequence of one-step Chow tests. Again, only five points are above the 5 percent level (none of which occurs after the inception of the crisis in 1994). The bottom panel shows the one-step residuals (±2 standard errors). Except for one observation in 1994 which slightly lies outside the range of the standard errors, all forecasted errors lie within the range. All three tests confirm the constancy of the estimated parameters within the estimation period inclusive of the financial crisis.

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

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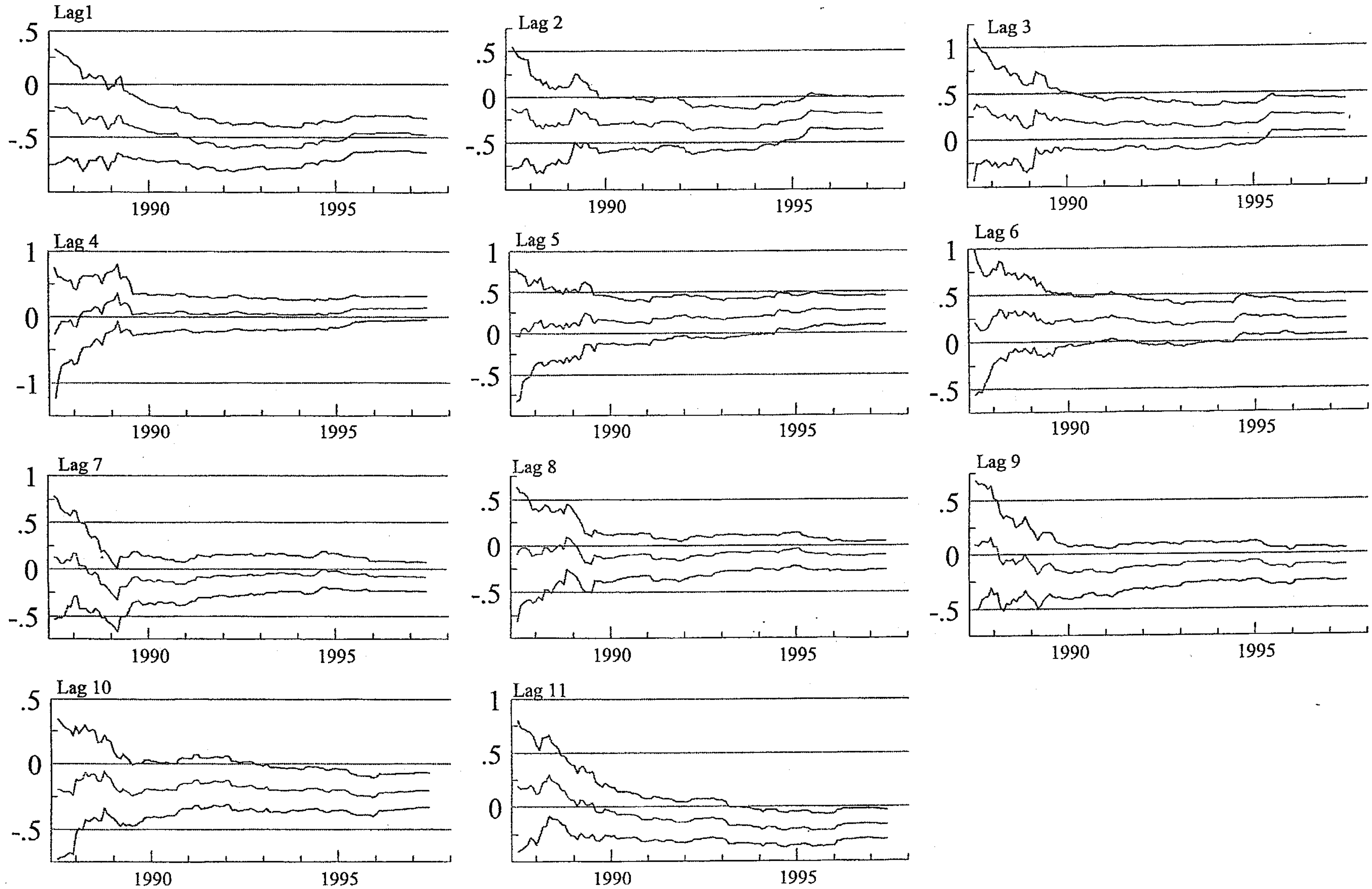
[22]All tests presented below employ the null hypothesis of parameter consistency. The rejection of the null hypothesis implies the rejection of parameter constancy over the period tested.

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

Figure 5. Recursive Parameter Estimates: Δ(m-p)



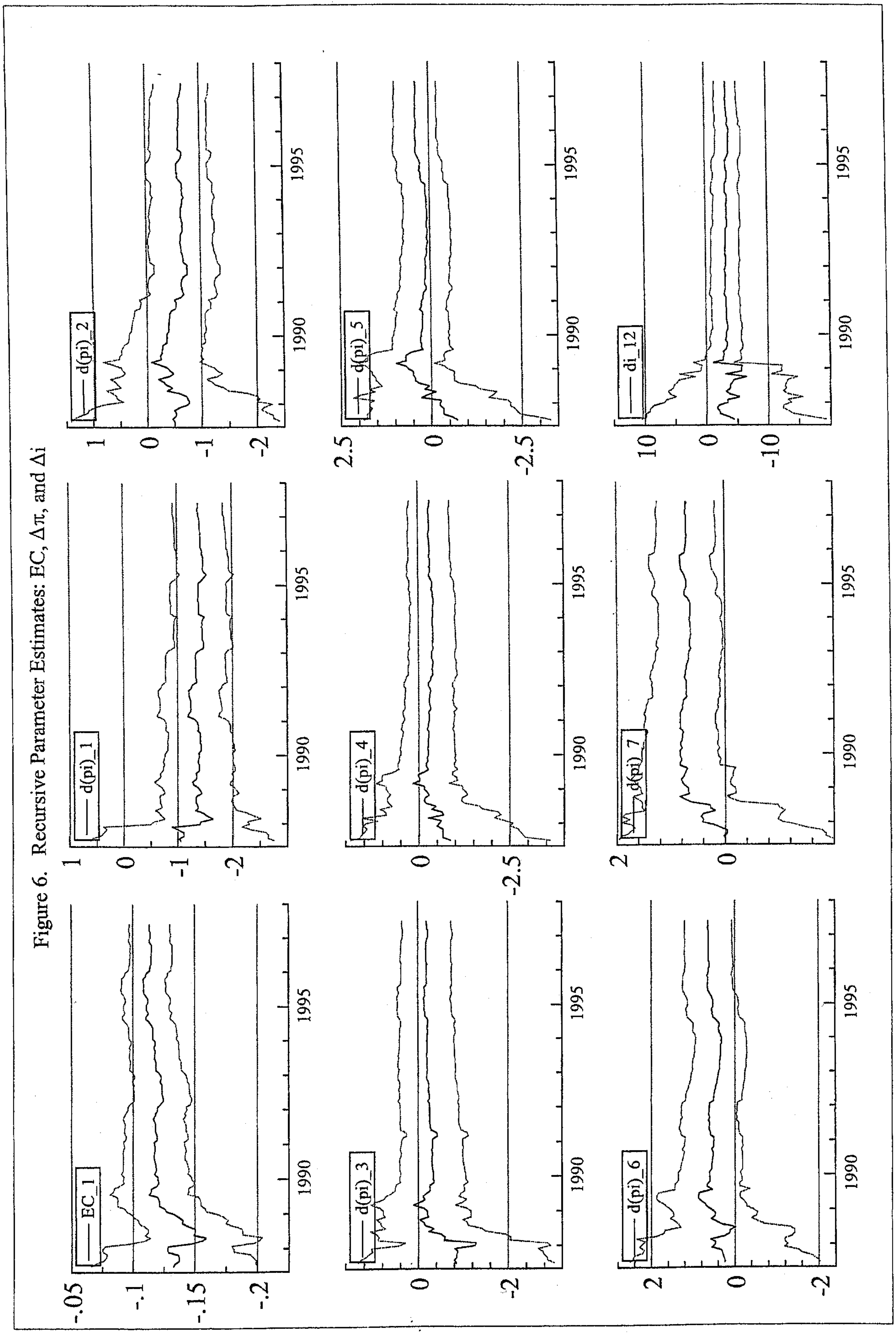
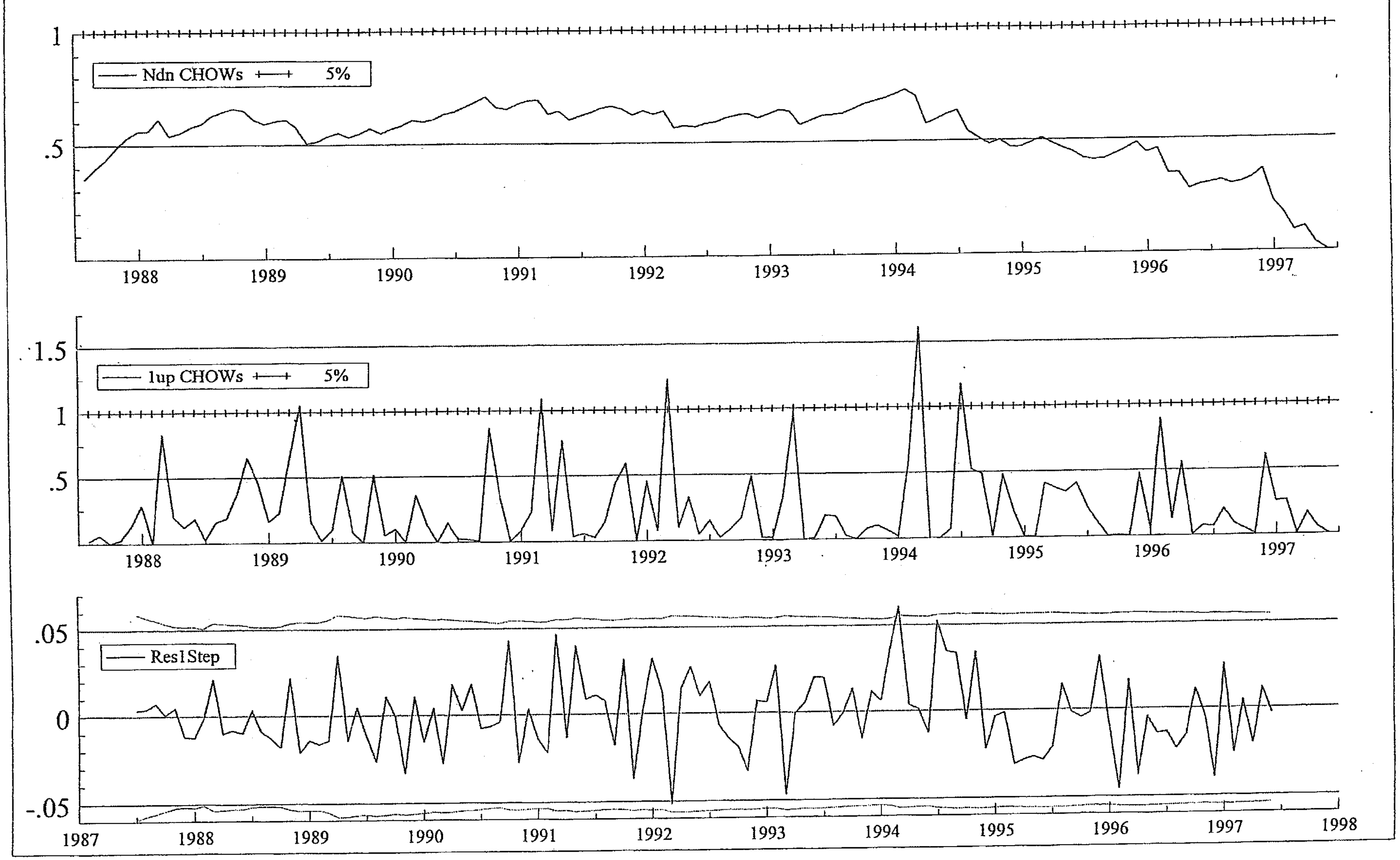


Figure 6. Recursive Parameter Estimates: EC, $\Delta\pi$, and Δi

Figure 7. One-Step and Break Point Chow tests and One-Step Residuals



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 8 shows the forecast Δ(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.

Table 5. Forecast Chow and χ^2 Statistics

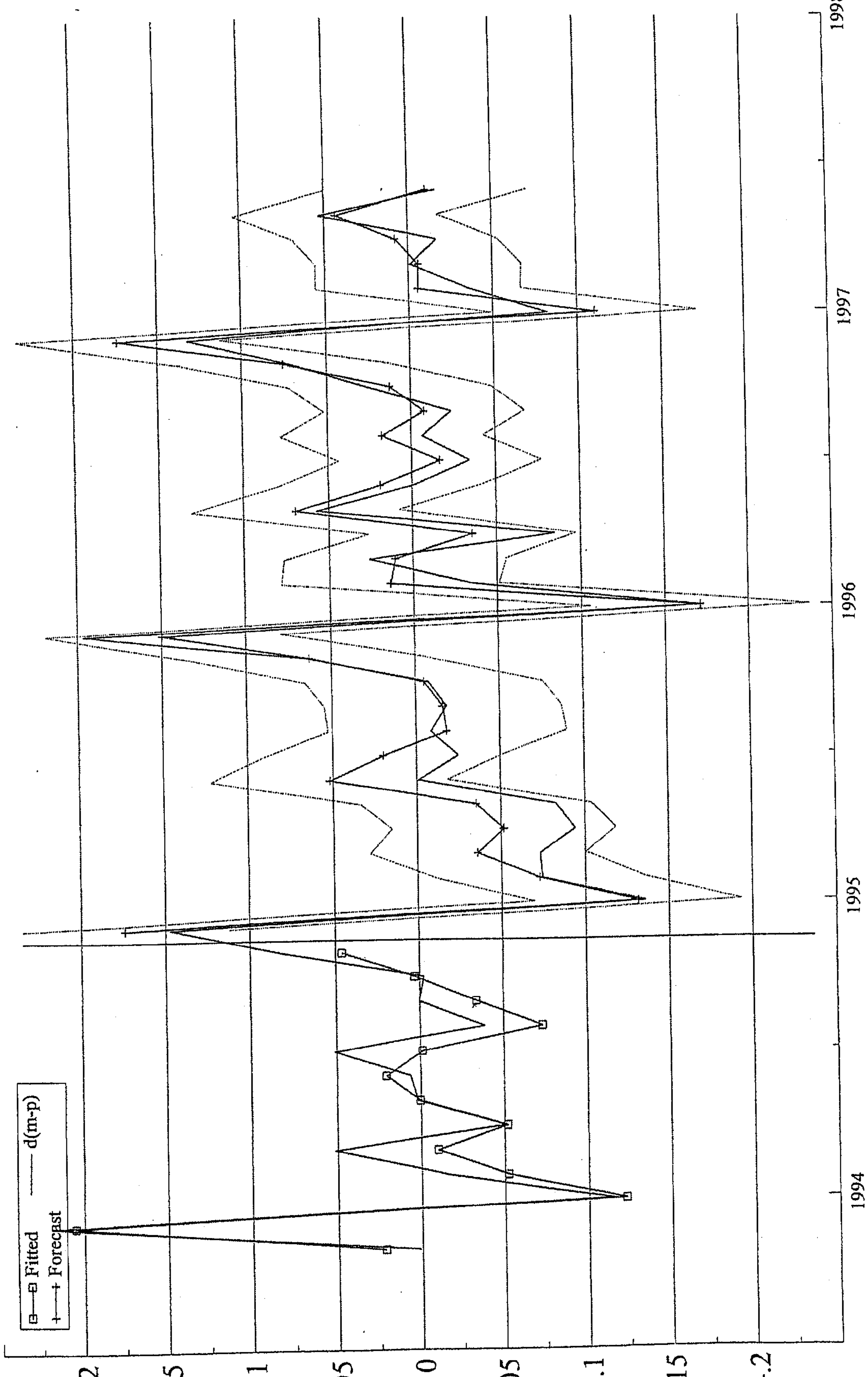
	1994:12-1995:12	1994:12-1997-6
Chow F(. , .)	0.76 [0.70]	0.74 [0.83]
Forecast χ^2	16.86 [0.21]	30.93 [0.28]

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.

VI. 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 indicate 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. 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.

Figure 8: Actual Observations and Forecasts, 1994:12-1997:6



The Error Correction Model, 1983:1-1997:6

Explanatory variable: Δ(m-p)

	Coefficient Estimate	Standard Error	Hansen's Instability Statistic
$\Delta(m\text{-}p)_{t-1}$	-0.478***	0.079	0.13
$\Delta(m\text{-}p)_{t-2}$	-0.183**	0.088	0.19
$\Delta(m\text{-}p)_{t-3}$	0.250***	0.089	0.17
$\Delta(m\text{-}p)_{t-4}$	0.139	0.088	0.09
$\Delta(m\text{-}p)_{t-5}$	0.274***	0.086	0.13
$\Delta(m\text{-}p)_{t-6}$	0.240***	0.083	0.14
$\Delta(m\text{-}p)_{t-7}$	-0.083	0.077	0.21
$\Delta(m\text{-}p)_{t-8}$	-0.107	0.077	0.07
$\Delta(m\text{-}p)_{t-9}$	-0.101	0.074	0.12
$\Delta(m\text{-}p)_{t-10}$	-0.204***	0.067	0.28
$\Delta(m\text{-}p)_{t-11}$	-0.164**	0.066	0.15
$\Delta\pi_{t-1}$	-1.383***	0.224	0.04
$\Delta\pi_{t-2}$	-0.672***	0.252	0.26
$\Delta\pi_{t-3}$	-0.192	0.271	0.21
$\Delta\pi_{t-4}$	-0.320	0.277	0.16
$\Delta\pi_{t-5}$	0.371	0.288	0.05
$\Delta\pi_{t-6}$	0.630**	0.278	0.10
$\Delta\pi_{t-7}$	0.712***	0.268	0.03
Δi_{t-12}	-3.400***	0.876	0.02
EC_{t-1}	-0.114***	0.008	0.16
MD_1	-0.197***	0.019	0.10
MD_2	-0.242***	0.024	0.18
MD_3	-0.292***	0.028	0.07
MD_4	-0.240***	0.029	0.29
MD_5	-0.237***	0.025	0.12
MD_6	-0.214***	0.024	0.13

The Error Correction Model, 1983:1-1997:6

Explanatory variable: $\Delta(m-p)$

	Coefficient Estimate	Standard Error	Hansen's Instability Statistic
MD_7	-0.143***	0.023	0.13
MD_8	-0.200***	0.023	0.09
MD_9	-0.203***	0.026	0.09
MD_{10}	-0.147***	0.027	0.08
MD_{11}	-0.121***	0.024	0.29

*Significant at the 10 percent level; ** significant at the 5 percent level; *** significant at the 1 percent level.
MD_i, =1 to 12 denotes monthly dummies for January through November.

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