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DEMAND FOR MONEY IN THAILAND

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Abstract:

After a brief review of recent literature, new estimates on a long run of annual observations of the Thai demand for all the standard measures of money are presented. The results demonstrate that the demand for real money balances is a stable function of a scale variable and a coherent measure of opportunity cost, with all the properties predicted by economic theory.

Introduction

The resurgence of interest in the demand function for money stimulated by the development of cointegration techniques has generated a substantial literature on emerging as well as advanced economies. Thailand, as the first victim of the 1997 crisis, is a particularly intriguing example; but existing (English-language) studies raise more questions than they resolve.

After a brief review of this literature, new estimates of the Thai demand for all the standard measures of money are presented. All the results demonstrate that the demand for real money balances is a stable function of a scale variable and a coherent measure of opportunity cost, with all the properties predicted by economic theory.

Literature Survey

All previous work adopts the Cagan (1956) specification of the demand function as a maintained hypothesis, utilises quarterly data in estimation, defines real GDP as the scale variable, and agrees that the standard measure of opportunity cost, a domestic interest rate, is not appropriate for Thailand. As quarterly data on GDP are of recent origin, use of interpolation or a proxy is necessitated. Alternative scale variables have not been investigated.

The treatment of opportunity cost raises more substantial difficulties. Chowdhury (1997) asserts that regulations on interest rates and limited availability of alternative financial assets raise doubts about the appropriateness of standard interest rate measures; but he does not explain why controls on rates should *ipso facto* limit their appropriateness, or investigate his doubts empirically. He suggests instead that the inflation rate 'reflects the opportunity cost of holding money much more adequately than interest rates', but he does not actually use it in estimation. Bahmani-Oskooee and Rehman (2005) make a similar untested assertion, despite some success in using a money-market rate reported by Bahmani-Oskooee and Techaratanachai (2001); but they do at least include inflation in their estimating equation. Own-rates on time and savings deposits are nowhere mentioned.

References to external influences on money demand appear in all three papers. Chowdhury includes a Eurodollar interest rate 'to measure the degree of capital mobility'. All three include a nominal exchange rate in *levels* in a search for evidence of currency substitution, but do not spell out the expectational assumptions underlying their specification. In a bizarre interpretation of currency substitution Chowdhury concludes that

> An increase in the real [previously defined as nominal] exchange rate, i.e. depreciation of the Thai baht, would increase the demand for broad money. A depreciation raises the value of foreign securities held by domestic residents and lowers the value of domestic securities held by foreigners, as valued in their own currency. This, in turn, increases the demand for domestic money.

Bahmani-Oskooee and Rehman report the same sign for the exchange rate on M1 but the opposite for M2. Their suggestion of a wealth effect as the explanation for the former, in an economy committed to making large net interest payments denominated in dollars, strains credulity.

Suspending the disbelief induced by these idiosyncrasies, all three papers claim evidence of cointegration. Chowdhury's preferred specification yields two cointegrating vectors for both M1 and M2, but only one of each is reported. Bahmani-Oskooee and Rehman produce internally inconsistent results. The F-tests in their table 1 reject the null of non-cointegration for Thai M1 for some lag-structures, but not at all for M2. The t-ratio on the error-correction term in their table 8 is clearly insignificant for M1, perhaps because they include lags up to order 9 on the dynamic terms, most of which are also insignificant. The claimed significance of the error-correction term for M2 is based on an inappropriate standard critical value.

In the aftermath of the 1997 crisis more recent investigations have focused on short-run interactions between monetary and real variables, using very short samples for estimation which preclude consideration of long-run properties. For example, Disyatat and Vongsinsirikul (2002) employ quarterly data for 1993(1)-2001(4), and Fung (2002) estimates a twelfth-order VAR using monthly data for 1989(1)-2001(6), subsequently split into pre- and post-crisis sub-samples.

In short, the existing literature sheds very little light on the demand for money in Thailand.

Data and Methodology

The long-run demand function is provisionally assumed to be of the Cagan form, in which the log of the real stock of money demanded is determined by the log of a real scale variable and a measure of opportunity cost (entered as a proportion). The assumption of zero-degree homogeneity in prices is subsequently tested, and alternative measures of the scale variable and of opportunity cost are examined. The long-run function is estimated within an unrestricted error-correction model which incorporates short-run dynamics. The null hypothesis of non-cointegration is tested by deletion of the 'level' variables, using non-standard critical F-values (Pesaran *et al*, 2001).

The data are annual, from the IMF *International Financial Statistics* Yearbooks (various issues), supplemented by reference to the Bank of Thailand's website and Nidhiprabh (1993). The advantage of ‘genuine’ data for the longest available period (from the early 1950s to 2002 in some cases) is believed to more than offset the larger number of observations in quarterly data for the purpose of examining long-run properties of demand for money.

Within this framework it is not necessary to test for the order of integration of the variables, but it is nevertheless informative. The ADF tests reported in table 1 show that all measures of the real money

Table 1 about here

stock, the two scale variables of primary interest, and the three measures of the cost of holding money (measured as proportions) are unambiguously I(1); any residual doubts about the order of demand deposits were resolved by inspection of the plots of the series and the autocorrelation function. Accordingly the upper bound of the non-standard F-distribution is used as the critical value in the variable-deletion tests of non-cointegration.

The serial correlation test included in the standard diagnostics reported for the estimates is for first-order; in all cases tests for up to third-order were conducted but revealed nothing to cause concern. All the CUSUM and CUSUMSQ plots for the reported equations were satisfactory. Predictive failure was tested with a break at end-1993; shifting the break to end-1996 to isolate the post-crisis period made no difference to the results.

A Digression on Measurement

The interest semi-elasticity (β) of the Cagan demand function plays a crucial role in the monetary model of exchange-rate determination, which expresses the spot exchange rate as a function of current and expected future fundamentals and the expected exchange rate at the horizon, with a per-

period discount factor of $\beta/(1+\beta)$. The student who struggles successfully through the hoops of purchasing power parity, uncovered interest parity, rational expectations and the associated algebraic manipulations to arrive at this result may be forgiven for questioning the purpose of the exercise on being informed by (for instance) Hallwood and MacDonald (2000) that $\beta \approx 0.02$, on the authority of Bilson (1978a); if so, the coefficient on expected future fundamentals asymptotes to zero so rapidly that they might as well be ignored. On the other hand Engel and West (2005) attribute an estimate of $\beta \approx 60$ to the same author in a different paper (Bilson, 1978b) in support of their contention that the quarterly discount factor is around 0.98, in which case expected fundamentals in the distant future matter a great deal. Is this an illustration of the rhetoric of economics?

The substantive issue here concerns the measurement of interest rates. Engel and West assert (p497) without discussion that

> Bilson uses quarterly interest rates that are annualized and multiplied by 100 in his empirical study. So his actual estimate [of β = 0.15] should be multiplied by 400 to construct a quarterly discount rate.

Three questions arise. Did Bilson use annual percentage rates of interest? Was his actual estimate β = 0.15? And should interest be expressed at a quarterly proportional rate? Bilson's (1978b, p61) own summary of the key result sheds light on the first two questions:

> The interest rate *elasticity,* which is equal to the product of the regression coefficient [-1.3853] and the interest rate, would be equal to -0.15 [presented as a stylised fact] if the rate of interest were 10.83 per cent.
>
> [Italics supplied].

Engel and West are clearly guilty of careless misreading in confusing elasticity and semi-elasticity; and Bilson's calculation demonstrates that interest is here entered as a proportion per annum.

Digressing further, this is not their only careless slip: in the same footnote they cite MacDonald and Taylor (1993) as adopting Bilson's (1978a) estimate of $\beta = 0.15$, rather than 60; but the number they adopt is actually 0.015. They may take some consolation from the fact that other contributors to this arcane literature have also displayed carelessness: Bilson (1978a, p85), whose calculations reveal that interest rates are here entered as annual percentage rates, offers a baffling comment:

> US studies of the demand for money have found interest rate *elasticities* of approximately *0.015* if the nominal rate of interest is 10 per cent. [Italics supplied].

In a context where precise numbers are crucial, experienced professionals might reasonably have been expected to exercise the degree of care customarily required of undergraduate students.

The remaining question is substantive. Expressing interest at quarterly rates in a quarterly model seems intuitively plausible, but the larger issue of measurement units is less easily resolved. In a static model it appears that the investigator is completely free to choose between proportions and percentages, or for that matter basis points, with dramatically different implications for the value of the interest semi-elasticity and for the properties of the monetary model. The unrestricted error-correction model provides an answer. With the dependent variable expressed as a proportionate rate of change, the conformable measure of the interest rate is as a proportion. Engel and West got the numbers wrong and in consequence overstated their case that the discount factor is almost one, but they selected the appropriate unit of measurement.

Demand for Cash

Cash in the hands of the public plays a major role in the Thai monetary system, accounting on average over the sample for 70% of M1 and 11% of M2. Attempts to model the determinants of demand for cash produced results similar to those reported by Reilly and Sumner (2005) for Sri Lanka. The only interest rate available for the full sample is the central bank's discount rate, of rather limited direct relevance to the personal sector but a highly visible influence on other rates. With either this or the inflation rate (DLP, defined on the GDP deflator) to represent the cost of holding cash, the diagnostics were unsatisfactory and the CUSUM test failed; in addition, with the former measure the null of non-cointegration was not rejected. Including both together did not improve matters; but a coherent measure of opportunity cost as the highest available rate of return on alternative assets, defined as

$$OPC_{RD} = MAX\ (RD, DLP)$$

produced the more promising results detailed in equation 1 of table 2.

Table 2 about here

The fit is appreciably better than with either of the standard cost measures; all the dynamic terms (indicated by prefix D) are well-determined; apart from a marginal problem with the functional form all the diagnostics are satisfactory; and the F-test in the last row is consistent with a stable long-run relationship among the level variables. In contrast to Sri Lanka there were only four years (1955, 1957 and 1973-74) in which DLP > RD. Entering the two components of OPC_{RD} separately yielded

-1.43 OPC(RD) -1.30 OPC(DLP)

[2.69] [3.79]

with a p-value for the hypothesis of equal coefficients of 0.67. The same property did not hold, however, for the first-differences of the opportunity cost components; instead these dynamic terms enter as separate continuous variables. Entering OPC in logs or the first-difference of the level rather than log of the interest rate caused a deterioration in the fit, as in the UK (Chadha et al, 1998) and Sri Lanka.

Homogeneity was tested by respecifying equation 1 in nominal terms. The appropriate coefficient restrictions were easily satisfied, with p-values of 0.37 for the level terms, 0.43 for the differences, and 0.51 for all constraints. All subsequently reported regressions were re-estimated with the addition of LP(-1) and DLP as a check for non-homogeneity, but these extra terms were always insignificant.

The unrestricted error-correction model facilitates tests of weak exogeneity. Respecifying equation 1 with DLY or DOPC as dependent variable and appropriate modification of the dynamics yielded F-values for the variable deletion test of 2.17 and 3.86 respectively, far below the critical values. Hence the interpretation of equation 1 as a demand function for money is confirmed.

Close substitutes for GDP (domestic absorption, GNP) as scale variable made no material difference to the results, but a later start of the estimation period enables comparison with consumers' expenditure, in equations 2 and 3. Equation 3 (also 5) uses the consumption deflator in the construction of OPC, the real stock of cash and hence the dependent variable, so comparison of fit relies on the standard error of the regressions, not the (strictly) incomparable coefficients of determination. The improvement in fit with consumers' expenditure is noticeable though hardly dramatic, and the residuals now exhibit heteroscedasticity.

The performance of the central bank's discount rate and the commercial banks' deposit rate can be compared over the same period. The presumption that the latter is a more appropriate measure of opportunity cost is fully confirmed by the contrast between equations 2 and 4. The measure of opportunity cost in equation 4,

$$OPC_{RDEP} = \text{Max}\ (RDEP,\ DLP),$$

for which DLP > RDEP in 1979-80 as well as 1973-74, produces better-determined dynamic effects and a better fit than in equations 2 and 3, but does not rectify the heteroscedasticity problem. That disappears when both modifications are combined, in equation 5, which appears to be completely satisfactory.

All specifications strongly reject the null of non-cointegration. Breaking the estimation period in 1993 produces no evidence of predictive failure. The speed of adjustment parameter (on the lagged stock) and the long-run semi-elasticity with respect to the different measures of opportunity cost (OPC coefficient divided by the speed of adjustment) are largely unaffected by the specification changes in equations 2-5; equation 1 implies slower adjustment and a larger semi-elasticity in the full sample, but estimates of the long-run scale elasticity are remarkably robust across all specifications.

Despite these positive results there is one apparently embarrassing problem. It will not have escaped the reader's attention that the end-date of estimation precedes the end of the data series by 4 years. The reason is that inclusion of the remaining data points would produce a serious breakdown of all the demand functions reported in table 2: specifically, the prediction error for equation 1 in 1999 is 24%! The enormous increase in real cash balances in that year is the crucial piece of evidence adduced by Burnside *et al* (2001) in support of their argument that the Asian currency crisis of 1997 is explicable in terms of the 'first-generation' model of speculative attack. Despite the manifest fiscal rectitude of the Thai and other affected economies in the period preceding the crisis, large prospective deficits were caused by the implicit bailout guarantees to failing financial systems, and the anticipation that these would be at least partially financed by seignorage led to the collapse of the fixed exchange rate regimes when the expectation was formed, rather than when large-scale money creation began.

There is, however, an alternative explanation which focuses on money demand rather than supply as the source of disturbance. Burnside *et al* use semi-annual observations, and the IMF data on money stocks used here are dated at end-year: 31/12/99 is indistinguishable from 1/1/2000. Anticipation of millennial disruption, rather than monetary irresponsibility, may have been responsible for the bulge in real cash balances. Examination of higher-frequency data is certainly suggestive. Monthly observations show that 84% of the year-on-year increase occurred in December, by which time readers of *The Bangkok Post* were well aware that demand for cash was running at unprecedented levels. *The Nation* reported (2/1/00) that the only Bangkok bank open on New Year's Day had enjoyed brisk business as new and old customers disposed of cash hoards now seen to be unnecessary. A similar phenomenon was observed in the other crisis economies, but also elsewhere. Even a casual scrutiny of *International Financial Statistics* suggests that this form of the millennium bug affected New Zealand, Japan, Fiji, Samoa, the Solomon Islands, Poland, Qatar, Saudi Arabia and the United Arab Emirates. In economies where cash in the hands of the public did not increase, commercial banks' cash reserves did: *The Economist* (18/12/99, p121) noted a 60% growth in US bank reserves in September-October under the heading 'monetary pest control'. Resort to Google will rapidly confirm that these press references are illustrative, not exhaustive.

A regression of the forecast errors in 1994-2002 from (re-estimated) equation 1 yields an insignificant negative intercept, coefficients which sum to zero on dummies for 1999 and the following year, an R^2 of 0.85 and satisfactory diagnostics. Fortunately millennia are sufficiently rare events as to make further examination unnecessary.

Demand Deposits

Modelling the determinants of demand deposits proved much simpler. The main differences are that inflation makes no contribution to the measurement of opportunity cost, and consumers' expenditure is clearly inferior to GDP as the scale variable. Both findings are consistent with the predominant use of demand deposits for business purposes rather than personal expenditure. No role was found for dynamics. No millennial problem emerged, so estimation is over the full sample period; the error in 1999 is (unsurprisingly) negative, but small. The results are summarised in table 3.

Table 3 about here

In all cases the null of non-cointegration is firmly rejected by the F-test which, in the absence of dynamics, coincides with the F-value for the regression. The deposit rate produces a marginally better fit than the central bank discount rate, but it also raises concerns about predictive failure. The more rapid adjustment suggested by equation 8 does not change the estimates of the long-run income elasticity and opportunity cost semi-elasticity. The former, at around 0.65, is conspicuously lower than the estimate of about 0.95 for cash. This difference, the superior performance of consumption as scale variable for cash, and the absence of dynamics in the deposit demand function all raise doubts about the usefulness in Thailand of combining the two categories of money in conventional M1.

A major advantage of unrestricted error-correction modelling is that the long- and short-run parameters are estimated simultaneously. The unusual absence of dynamics in this instance provides a rare opportunity to compare this approach with alternatives on level terms. The static equation estimated over the full sample in the first step of the Engel-Granger procedure produced a similar value to equation 6 for the (long-run) income elasticity but a much lower (absolute) value for the interest semi-elasticity, 1.7 against 5.5. Both the standard tests on the residuals, the ADF of -3.51 (-3.92) and the CRDW of 0.80 (0.99), failed to reject the null of non-cointegration at the 5% level (whose critical values are parenthesised). Proceeding nevertheless to the second stage, the error-correction model confirmed the absence of dynamics but narrowly rejected non-cointegration: the t-ratio for the error-correction term was -3.96 (-3.92). This rose to -0.03 when the implied error-correction term from equation 6 was added to implement Davidson and MacKinnon's (1981) test of specification; the t-ratio on the latter was -2.44, falling to -4.87 when the Engel-Granger term was deleted.

The Johansen procedure fared little better. With 1 lag in the VAR as indicated by the choice criteria, estimation with unrestricted intercept and restricted trend showed the latter to be highly insignificant. With restricted intercept and no trend both tests indicate 2 cointegrating vectors, in the first of which opportunity cost appears with the wrong sign, as does the disequilibrium term in the error-correction model. The second was highly correlated (0.99) with that implied by equation 6, with an identical long-run income elasticity, a slightly lower interest semi-elasticity, but a much higher speed of adjustment (0.79). It attracted a t-ratio of -4.81 (-3.92) in the error-correction model; but with a coefficient of only (-)0.57, close to the ratio of the estimated adjustment parameters and far below the expected value of unity.

The obvious conclusions to draw from this methodological digression are that the Johansen procedure is extremely sensitive and both the alternative procedures entail a serious risk of bias. The comparisons cannot establish that the equivocal results of both these methods are ‘wrong’ and the unambiguously positive ones from the unrestricted error-correction model are ‘right’, but they do demonstrate that the latter has substantially greater explanatory power.

Interest-Bearing Deposits

The question of primary interest in relation to time and savings deposits (SAV) is their relative responsiveness to opportunity cost and their own-rate of return. Equation 9 in table 4 shows that the magnitudes of response, in both levels and (log-) differences, are very similar. The hypothesis of coefficients which sum to zero cannot be rejected for levels (p=0.33), differences (p=0.17) or both (p=0.29), and these parameters all appear to be well-determined. Imposing the constraints in equation 10 entails little cost in terms of fit.

Table 4 about here

As SAV accounts for much the greater part of broad money, these results carry over to the M2 regressions in equations 11 and 12. Indeed the p-values for the interest constraints are considerably higher (at 0.81, 0.98 and 0.97 respectively), so their imposition carries no cost. The only additional feature is some weak evidence of a dynamic role for GDP. The coefficients on the GDP level are not well-determined, but the higher point-estimate of the long-run income elasticity, around 1.4, than was found for other aggregates is plausible. The alternative scale variable and alternative measures of opportunity cost performed less well. The speed of adjustment is comparatively leisurely in all cases. The forecast error for 1999 is negative and substantial for SAV (almost 50% greater than its standard deviation), negative but tiny for M2. The hypothesis of non-cointegration is decisively rejected in all the equations, and none of the diagnostics causes any concern.

Conclusions

Economic theory is alive and well in Thailand. The demand for money in all its forms is a stable function of a few key variables. There is no evidence of money illusion. Agents adjust their (personal) money holdings in line with inflation when, but only when, real interest rates are negative. The differential between opportunity cost and the own-rate of return is a key determinant of demand for quasi-money. Precautionary cash balances are increased (dramatically in 1999) in response to uncertainty.

Models estimated on annual data clearly cannot shed light on currency substitution or other extremely short-term influences, but it is doubtful whether quarterly models (even if appropriately specified) would be much more informative. Until the 1997 crisis, the Thai baht had moved closely with the US dollar except during the latter's appreciation in the early 1980s, even after a currency basket was substituted for the formal dollar peg, and anticipations of the crisis developed at a very late stage. Burnside *et al* (2001) observe that forward premia did not begin to rise significantly until mid-May, only 6 weeks before the initial devaluation. There was a conspicuous jump of 47% in foreign currency deposits with Thai commercial banks and Thai branches of foreign banks in April, but this increased the stock to only 0.6% of total deposits; and half of this increase was eliminated by end-June. Any currency substitution did not show up in end-year data: the M2 residual from equation 12 was positive and 30% larger than the standard error of the regression in 1997; the increase in time deposits dwarfed the rise in foreign currency deposits.

What models using long runs of annual data can do is to determine whether a stable low-frequency relationship exists among the variables of interest, and to estimate its parameters with more confidence than would be possible with more but higher-frequency observations over a shorter period.

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Table 1: ADF Tests

		Null Hypothesis	
		I(1)	I(2)
Cash	LM0	-1.70	-4.97
Demand deposits	LDD	-3.41	-7.56
Savings deposits	LSAV	-0.42	-4.88
Broad money	LM2	-1.30	-4.97
GDP	LY	-2.20	-4.03
Consumers' expenditure	LC	-1.84	-4.49
Central bank discount rate	RD	-0.23	-6.69
Opportunity cost	OPC	-1.05	-7.16
Deposit rate	RDEP	-1.09	-5.54

Notes to table 1:

Prefix L denotes natural logarithm.

Augmentation as needed to ensure white-noise residuals.

The variants of OPC (described in the text) all have similar ADF values to that tabulated.

Approximate (because the series differ in length) critical value (at 5% level) for first 6 entries in column 1, which include a deterministic trend, is -3.6, and -3.0 for the remainder and for column 2.

Table 2: Demand for Cash

Equation	1	2	3	4	5
Sample	1954-1998	1967-1998			
Intercept	-0.16	-0.30	-0.50	-0.43	-0.64
	[2.09]	[1.72]	[2.73]	[2.57]	[3.90]
LMO(-1)	-0.15	-0.23	-0.24	-0.24	-0.28
	[1.95]	[1.99]	[2.91]	[2.29]	[3.50]
LY(-1)	0.14	0.21		0.23	
	[2.33]	[2.18]		[2.58]	
LC(-1)			0.26		0.29
			[3.25]		[3.99]
OPC_{RD}(-1)	-1.25	-1.31	-1.48		
	[3.92]	[3.55]	[4.55]		
OPC_{RDEP}(-1)				-1.20	-1.36
				[3.70]	[4.95]
DLRD	-0.11	-0.09	-0.06		
	[2.54]	[1.67]	[1.29]		
D2LP	-0.41	-0.38	-0.62	-0.46	-0.59
	[3.11]	[1.93]	[3.02]	[2.41]	[3.12]
DLRDEP				-0.14	-0.13
				[2.71]	[2.71]
DLY	0.43	0.50		0.65	
	[2.11]	[2.13]		[3.29]	
DLC			0.70		0.79
			[3.93]		[5.20]

R^2	0.58	0.65	0.75	0.69	0.80
SER%	4.31	4.22	4.05	3.96	3.65
Serial Correlation	0.82	0.51	0.92	0.29	0.81
Functional Form	0.04	0.13	0.22	0.44	0.12
Normality	0.25	0.65	0.28	0.58	0.88
Heteroscedasticity	0.18	0.32	0.02	0.06	0.37
Predictive Failure	0.27	0.44	0.80	0.22	0.30
F (Levels exclusion)	6.13	5.40	9.76	5.83	11.43

Notes to table 2:

The dependent variable (DLMO), the corresponding lagged stock and the inflation rate included in OPC (and its difference) are constructed using the GDP deflator in equations 1, 2 & 4, and the CPI in equations 3 & 5. SERs (but not coefficients of determination) are comparable across these groups of equations.

Absolute t-ratios in parentheses below coefficients.

Diagnostic tests for serial correlation, functional form, normality, heteroscedasticity, and predictive failure, in F-form where available, reported as p-values.

Prediction period is 1994-1998.

Upper-bound critical values for the variable exclusion test in the last row are 4.13 (90%), 4.86 (95%) and 6.31 (99%).

Table 3: Demand Deposits

Equation	6	7	8
Sample	1953-2002	1967-2002	
Intercept	-0.24	0.08	-0.16
	[1.57]	[0.37]	[0.79]
LDD(-1)	-0.42	-0.46	-0.61
	[4.01]	[3.44]	[4.26]
LY(-1)	0.29	0.27	0.39
	[3.99]	[3.22]	[4.18]
RD(-1)	-2.31	-2.63	
	[3.27]	[3.22]	
RDEP(-1)			-3.11
			[3.87]
R^2	0.33	0.37	0.41
SER%	11.99	11.41	11.05
Serial Correlation	0.99	0.16	0.89
Functional Form	0.99	0.82	0.73
Normality	0.19	0.36	0.54
Heteroscedasticity	0.72	0.96	0.94
Predictive Failure	0.30	0.41	0.06
F (Levels exclusion)	7.58	6.23	7.35

Additional notes to table 3:

Dependent variable is DLDD.

Prediction period is 1994-2002.

Table 4: Interest-Bearing Deposits

	9	10	11	12
Equation				
Dependent Variable	DLSAV		DLM2	
Intercept	-0.26	-0.42	-0.58	-0.62
	[0.53]	[0.91]	[1.29]	[1.89]
LEVEL(-1)	-0.11	-0.14	-0.18	-0.19
	[1.60]	[2.19]	[1.91]	[2.95]
LY(-1)	0.14	0.20	0.25	0.27
	[1.57]	[1.66]	[1.72]	[2.63]
OPC_{RD}(-1)	-1.32		-1.29	
	[3.89]		[4.15]	
RDEP(-1)	1.66		1.36	
	[4.88]		[4.84]	
DIF(-1)		-1.50		-1.34
		[4.78]		[5.30]
$DLOPC_{RD}$	-0.12		-0.11	
	[3.36]		[3.88]	
DLRDEP	0.17		0.11	
	[3.44]		[2.51]	
DLDIF		-0.14		-0.11
		[4.23]		[4.42]
DLY			0.30	0.30
			[1.91]	[1.98]

R^2	0.66	0.62	0.67	0.67
SER%	4.07	4.17	3.32	3.21
Serial Correlation	0.25	0.10	0.34	0.32
Functional Form	0.68	0.10	0.74	0.74
Normality	0.48	0.42	0.57	0.55
Heteroscedasticity	0.76	0.93	0.14	0.15
Predictive Failure	0.69	0.37	0.82	0.86
F (Levels exclusion)	9.86	14.93	7.75	11.81

Additional notes to table 4:

Estimation period is 1967-2002.

LEVEL denoted the stock whose first difference is the dependent variable.

The upper-bound critical value of the F test for levels deletion in equations 9 and 11 is 5.62 (99%).