The Drachma/Deutschemark Exchange Rate, 1980-1997: A Monetary Analysis

by

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Abstract: The validity of the monetary approach to the Drachma/Deutschemark exchange rate determination is investigated by utilising the Hendry (1993) approach. The statistical findings presented in this investigation of the determinants of the Drachma/DM nominal exchange rate are consistent with the monetary approach and support the MacDonald-Taylor proposition that when the approach is interpreted as a long-run equilibrium condition and is appropriately tested, then it appears to have a significant measure of validity.
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I. INTRODUCTION

Levich (1985), Taylor (1995) and Frankel and Rose (1995) in their surveys of the empirical literature on the determinants of the nominal exchange rate draw attention to the rather poor performance of the basic, flexible prices monetary approach to exchange rate determination. Early studies appeared to support the approach but their conclusions were undermined by a large volume of empirical studies since the early 1980s. However, more recently, MacDonald and Taylor (1991) and (1994) have presented statistical findings, which suggest that the monetary approach, when interpreted as a long run equilibrium condition, is not without empirical validity. Their basic proposition is that the proper testing of the approach requires the application of the Johansen multivariate cointegration technique.

This paper seeks to ascertain whether or not testing the monetary approach against the Drachma/Deutschemark exchange rate yields statistical findings consistent with the MacDonald-Taylor conclusion.

II. THE MONETARY EQUATION DATA

The basic monetary approach rests on two assumptions. First, it assumes that purchasing power parity holds continuously and, second, that the demand for money functions of the domestic and foreign economies are stable. Monetary equilibrium in the two economies is given by:

$$m = p + cy - di \qquad (1)$$

and

$$m^* = p^* + c^*y^* - d^*i^* \qquad (2)$$

where m is the logarithm of the money supply, p denotes the logarithm of the price level, y denotes the logarithm of real income and i is the interest rate with German variables being denoted by an asterisk.

Purchasing power parity is given by

$$\varepsilon = p - p^* \qquad (3)$$

where ε is the logarithm of the nominal exchange rate defined to be the price of the Mark in terms of Drachma.

Now if we substitute (1) and (2) into (3) and after rearranging we have the basic monetary equation:

$$\varepsilon = (m-m^*) - cy + c^*y^* - d^*i^* + di \qquad (4)$$

Dornbusch (1976) relaxed the assumption that purchasing power parity holds continuously by assuming that prices are sticky in the short-run while Frankel (1979) argued that the logic of the Dornbusch overshooting model implied that the real interest deferential should also be included as an explanatory variable. Significant though these contributions were, McDonald and Taylor (1992) observed that all variants of the monetary model collapse to a long run equilibrium condition of the form (4).

Our principal objective is to ascertain whether there exists a long run equilibrium relationship between the Drachma and the DM that is based on relationships implied by equation (4). Monthly not seasonally adjusted data for 1980(2) to 1997 (3), but adjusted for lags, were used to test the monetary model. Centered Seasonal Dummies were used to capture seasonallity and avoid possible exclusion of trend.

For the German money supply (M1), income (industrial production) and interest rate (interbank rate) variables the data used are from the IMF and OECD CD ROMs databases. Data for money supply (M1) and income (manufacturing production) relating to Greece are taken from the same sources while for the Greek interest rate the Bank of Greece Monthly Statistical Bulletin figures for the Treasury bill rate was used. Finally the Drachma/DM exchange rate data is taken from the Bank of Greece Monthly statistical Bulletin. Details for each variable are provided in the appendix.

The PcFiml module of GivWin (1997) as developed by Doornik and Hendry (1997) was used for the empirical work.

III. MULTIVARIATE COINTEGRATION ANALYSIS

We begin by testing the order of integration of the stochastic variables by employing the standard Dickey and Fuller (1979) tests and Augmented Dickey and Fuller tests, by following a procedure from the unrestricted model towards the most restricted model . The results presented in Table 1 indicate that the series ε, $(m\text{-}m^*)$, (y), (y^*), i^*, and i are all I(1) processes.

Next, starting from a VAR with 12 lags on all the stochastic variables through simplification tests the adequacy of 4 lags was established. This reduction was implemented by using likelihood ratio tests adjusted for degrees of freedom. As Table 2 reveals all residual correlations are low and the companion matrix of the dynamics had no eigenvalues outside the unit root circle. Summary statistics and diagnostics on both the individual equations and the VAR as a whole suggest that the residuals do not suffer from autocorelation while ARCH tests imply the absence of misspecification. The observed absence of normality, as Gonzalo (1994) points out, does not affect neither the number of cointegrating vectors nor the coefficients obtained from Johansen (1988).

Table 1: Unit root tests. Sample 1979:9 1997:3

Variable (x)	Unit Root in x ADF			Unit Root in Δx ADF		
	τ	τ_μ	τ_τ	τ	τ_μ	τ_τ
ε			1.73			-9.06
(m-m*)		3.40			-11.72	
y	-0.07			-7.56		
y*	1.17			-13.93		
i*	-1.18			-8.9		
i	-0.07			-12.93		

Significant at 1% level (**)
MacKinnon critical values for rejection of hypothesis of a unit root
3.46 for non-trended variables
4.00 for trended variables
τ no constant and trend in DGP
τ_μ with constant, without trend in DGP
τ_τ with constant, with trend in DGP

Table 2: Residual Correlations, Dynamic Analysis, Goodness of fit and evaluation of the system.

	ε	(m-m*)	y	y*	i*	i	
E	1.0000						
(m-m*)	0.1616	1.0000					
y	0.0833	-0.002	1.0000				
y*	0.0398	0.0049	0.0994	1.0000			
i*	-0.113	0.1055	0.0416	0.0949	1.0000		
I	0.0227	-0.043	0.0699	0.0229	-0.005	1.0000	

Eigenvalues of $\pi(1)$-I
$|\lambda_\alpha|$ 0.4759 0.1049 0.1049 0.06032 0.01311 0.00565

Eigenvalues of companion matrix
$|\lambda_b|$ 0.6624 0.6493 0.6493 0.5599 0.4327 0.5294 0.5294 0.3932 0.3932 0.6032 0.6032 0.4053 0.4053 0.4555 0.4555 0.5518 0.5121 0.8077 0.7937 0.8793 0.9505 0.9505 0.9479 0.9479

Statistic	ε	(m-m*)	y	y*	i*	i	VAR
$\hat{\sigma}$	0.0085	0.0160	0.0159	0.0073	0.0193	0.0093	
$F_{ar}(7,160)$	1.023	1.758	1.9766	0.6949	0.6842	0.6631	
$F_{arch}(7,153)$	0.5274	0.4867	0.2692	1.0453	1.9881	0.0308	
χ^2_{nd} (2)	126**	51**	31**	58**	126**	150**	
F^{v}_{ar} (252, 722)							0.9580
χ^2_{nd}							555**

As a further preliminary step to the cointegration analysis the break-point Chow test (1-step and N$\downarrow$ step) was used as an informal test for parameter constancy. Figure 1 indicates that for none of the individual equations do the test values for N$\downarrow$ step exceed the 1% significance level which is consistent with parameter constancy. On the other hand outliers are detected for the 1-step Chow test.

The next step in our empirical analysis is to test for cointegration by employing the multivariate cointegration technique proposed by Johansen (1988) and Johansen and Juselius (1990), with the estimation of a closed VAR model to the six dimensional vector X_t=[ε_t, (m-m*), y, y*, i*, i] The estimation results are reported in Table 3. The tests based on maximum eigenvalue, Tlog(1-λ), indicate no cointegration but the trace statistic, TΣlog(1-λ), indicates the existence of a unique cointegrating vector at 99% level of confidence, or two cointegrating vectors at 95% level of confidence. We chose the vector which indicates a cointegrating relationship at the 99% level of confidence.

TABLE 3: Cointegration analysis 1980 (2) to 1997 (3) of ε

Ho:rank=p	-Tlog(1-μ) (eigenvalue test)	95%	-TΣlog(1- μ) (trace test)	95%	99%
p=0	37.17	39.4	111.4**	94.2	103.1
p≤1	29.26	33.5	74.26*	68.5	76.07
p≤2	22.76	27.1	44.99	47.2	54.46
p≤3	15.55	21.0	22.23	29.7	35.65
p≤4	6.474	14.1	6.683	15.4	20.04
p≤5	0.2083	3.8	0.2083	3.8	6.65

standardized β' eigenvectors

ε	m-m*	y	y*	i*	-i
1.0000	-1.502	-5.5295	2.0980	0.01760	-0.80114
-0.5805	1.0000	-1.8456	-0.3851	-0.06063	0.36203
0.6092	0.24484	1.0000	-4.3788	0.59725	-0.67094
-0.0064	-0.26193	-0.7016	1.0000	0.03426	-0.04302
74.842	-96.2300	14.7950	-219.700	1.00000	15.48900
0.84436	-0.79962	-0.16366	-1.6501	0.48225	1.0000

Standardized α coefficients

ε	-.02070	0.04765	-0.0160	0.02132	$-5.65e^{-005}$	-0.0007
m-m*	0.0031	-0.0048	0.01647	0.07856	0.00018	-0.0034
y	0.05050	0.06563	-0.0291	0.04485	0.00015	0.00103
y*	-0.0040	0.01927	-0.0001	-0.0523	0.00014	-0.0003
i*	-0.0083	-0.0773	-0.0834	-0.0055	0.00023	-0.0004
i	-0.0215	0.00576	0.01133	0.04359	0.00013	0.00145

*Critical values from Osterwald-Lenum (1992)

Table 4 presents the results of testing jointly the existence of a single cointegrating vector and long-run weak exogeneity of the variables m-m*, y, y* and i*, i for the parameters in the exchange rate equation, which are constrained for long run money neutrality and unitary money demand income elasticities. This implies a single row in the β' matrix and a single column in the α matrix of the form (*,0,0,0,0,0). The restrictions are data acceptable at the 5% level of significance. The normalised coefficients of the cointegrating vector have theory consistent signs while the loading coefficient is negative as expected. Thus the estimated cointegrating vector can be interpreted as a long run exchange equation.

TABLE 4: Restricted Cointegration analysis 1980 (2) to 1997 (3)

Standardized β' eigenvectors and $\alpha = A\theta$ coefficients

	ε	m-m*	y	y*	i*	-i
β'	1.0000	-1.0000	1.0000	-1.0000	0.2891	-0.9347
α	-0.0431	0	0	0	0	0

LR-test, rank=1: Chi^2(8) = 15.496 [0.0502]

Next the data was mapped to the I(0) space and the results are presented in Table 5. The model diagnostic tests indicate that the residuals in this parsimonious VAR are white noise and that there are no problems with parameter constancy.

The same table reports the coefficients on the ECMs in the six equations. Except for the exchange rate equation, these coefficients are insignificant. It follows then that the weak exogeneity conclusion is confirmed enabling, therefore, the inclusion of contemporaneous observations of the weakly exogenous variables m-m*, y, y*, i* and i in the estimation of a conditional dynamic exchange rate equation.

Table 5: Reduction to I(0). RLS estimates. Sample: 1980 (2) to 1997 (3)

Dε = 0.022*C+0.167* Dε_1-0.047*D(m-m*)_1+0.037* Dy_1 +0.037* Dy*_1-0.009* Di*_1
(0.006) (0.072) (0.037) (0.033) (0.076) (0.029)
+0.008* Di_1-0.031*CI_1
(0.029) (0.010)

D(m-m*) = 0.001*C+0.13* Dε_1-0.380*D(m-m*)_1+0.052* Dy_1 -0.087* Dy*_1+0.104* Di*_1
(0.010) (0.13) (0.068) (0.061) (0.139) (0.054)
-0.097* Di_1+0.004*CI_1
(0.128) (0.016)

Dy = 0.018*C-0.105* Dε_1-0.118*D(m-m*)_1-0.424* Dy_1 +0.155* Dy*_1-0.057* Di*_1
(0.011) (0.142) (0.070) (0.066) (0.151) (0.058)
-0.118* Di_1-0.030*CI_1
(0.138) (0.018)

Dy* = 0.055*C-0.09* Dε_1+0.018*D(m-m*)_1-0.004* Dy_1 –0.465* Dy*_1+0.011* Di*_1
(0.005) (0.061) (0.030) (0.028) (0.064) (0.025)
-0.043* Di_1-0.006*CI_1
(0.060) (0.008)

Di* = 0.005*C-0.121* Dε_1+0.076*D(m-m*)_1+0.091* Dy_1 +0.153* Dy*_1+0.360* Di*_1
(0.013) (0.163) (0.085) (0.076) (0.174) (0.068)
-0.066* Di_1-0.011*CI_1
(0.162) (0.012)

Di = -0.112*C+0.08* Dε_1-0.055*D(m-m*)_1+0.054* Dy_1 -0.022* Dy*_1-0.0006* Di*_1
(0.031) (0.077) (0.039) (0.035) (0.081) (0.031)
+0.144* Di_1-0.021*CI_1
(0.074) (0.012)

Model diagnostic tests

Statistic	Dε	D(m-m*)	Dy	Dy*	Di*	Di	VAR
$\hat{\sigma}$	0.00865	0.01588	0.01721	0.00740	0.01991	0.00923	
$F_{ar}(7,177)$	1.321	1.0834	4.181**	0.84759	0.46375	0.50216	
$F_{arch}(7,170)$	0.5141	0.7232	2.21*	3.723**	2.61*	0.051	
χ^2_{nd} (2)	257**	54.71**	67.75**	83.36**	106.6**	127**	
$F_{ar}(252, 823)$							1.051
$F^{v}_{ar}(294, 1783)$							1.26**
χ^2_{nd} (12)							711**

The preferred specification of the error correction formulation of the long run exchange rate equation is presented in Table 6. All diagnostics tests are satisfied and the suggestion is that the estimated equation is a reasonably data coherent representation.

TABLE 6: FIML estimates. Sample: 1980 (3) to 1991 (1)

$$D\varepsilon = \underset{(0.006)}{0.021}*C + \underset{(0.07)}{0.155}*D\varepsilon\_1 + \underset{(0.039)}{0.085}*D(m-m^*) - \underset{(0.040)}{0.008}*D(m-m^*)\_1 + \underset{(0.036)}{0.013}*Dy + \underset{(0.036)}{0.043}*Dy\_1$$

$$+\underset{(0.085)}{0.089}*Dy^* + \underset{(0.085)}{0.096}*Dy^*\_1 - \underset{(0.031)}{0.073}*Di^* + \underset{(0.031)}{0.026}*Di^*\_1 + \underset{(0.068)}{0.020}*Di + \underset{(0.069)}{0.013}*Di\_1 - \underset{(0.009)}{0.03}*CI\_1$$

Model diagnostic tests

Statistic	$D\varepsilon$	statistic	VAR
$F_{ar}(7,172)$	1.3256	$F_{ar}(7,172)$	1.326
$F_{arch}(7,165)$	0.419	$F_{arch}(24,154)$	1.109
$F_{het}(90,88)$	1.2306	$F_{het}(90,98)$	1.2306
$\chi^2_{nd}(2)$	234.8**	$\chi^2_{nd}(2)$	234.8**

To check the constancy of the model over the sample, Figure 2, plots the actual and fitted values. The latter tracks the former reasonably well. Further, Figure 3, based on the recursive FIML estimates highlights the model's constancy in terms of residuals sums of squares, one step residuals with $\pm 2SE$, one step Chow forecasting failure test and N-step Chow stability test. The one step errors lie within their appropriate 95% confidence band while the values of the break-point Chow F-tests do not exceed the 1% significance level. In summary, then, the model appears to be reasonably constant.

IV. CONCLUSIONS

The statistical findings presented in this investigation of the determinants of the Drachma/DM nominal exchange rate are consistent with the monetary approach and support the MacDonald-Taylor proposition that when the approach is interpreted as a long-run equilibrium condition and is appropriately tested, then it appears to have a significant measure of validity.

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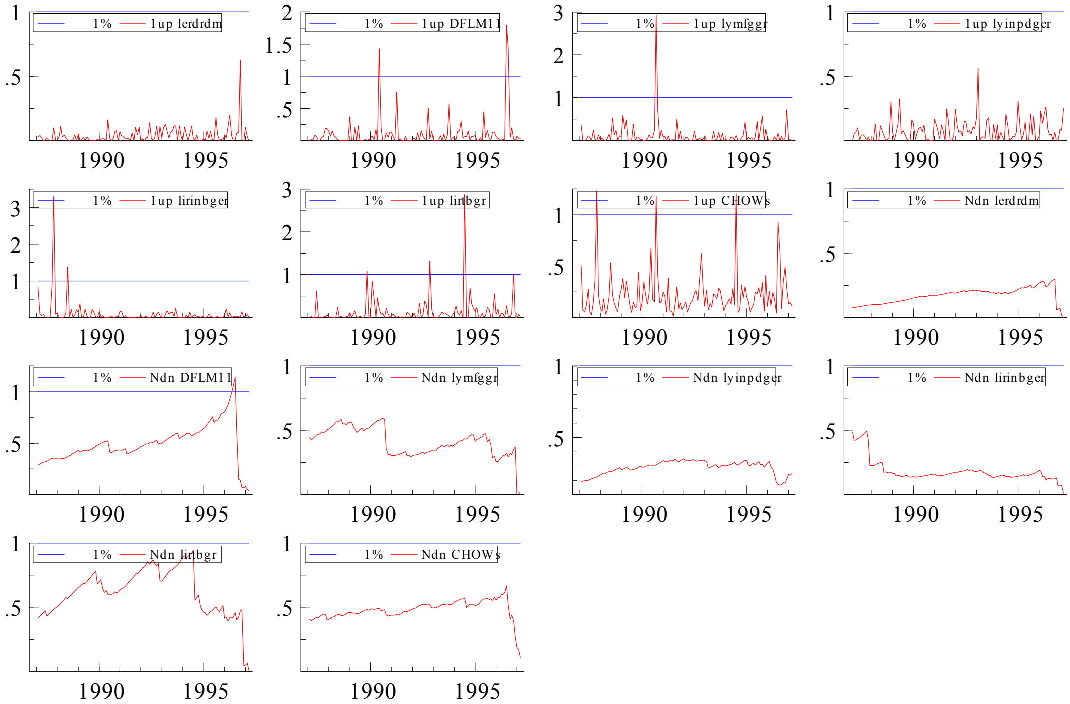

Figure 1: Break point Chow tests

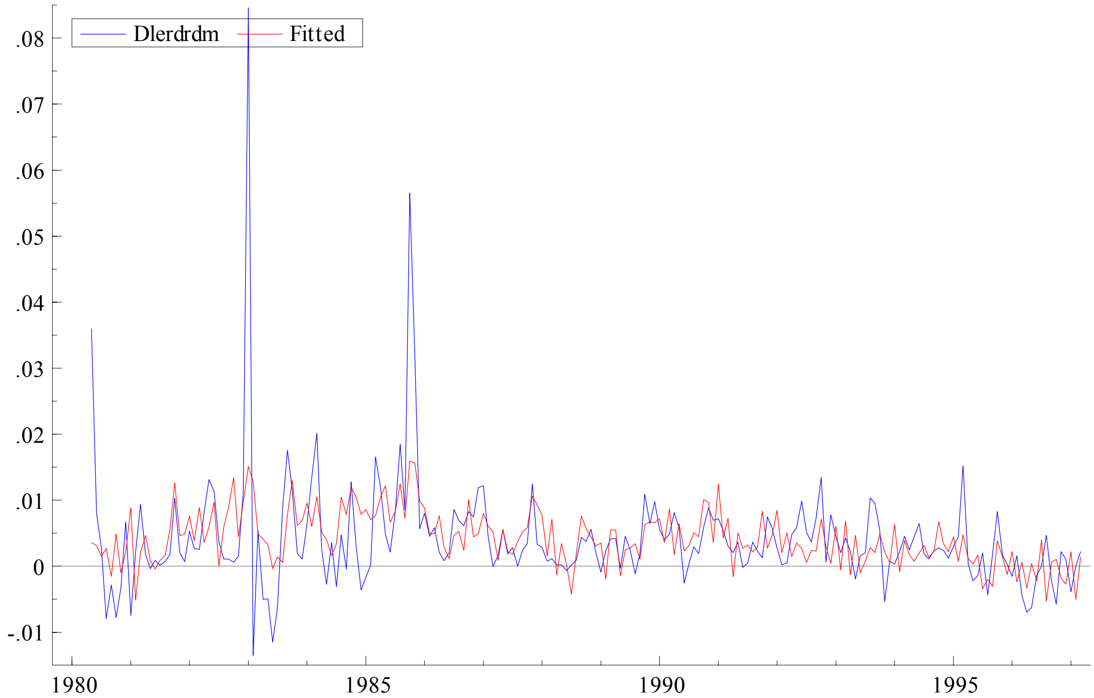

Figure2: Fitted and Actual values from the conditional dynamic model

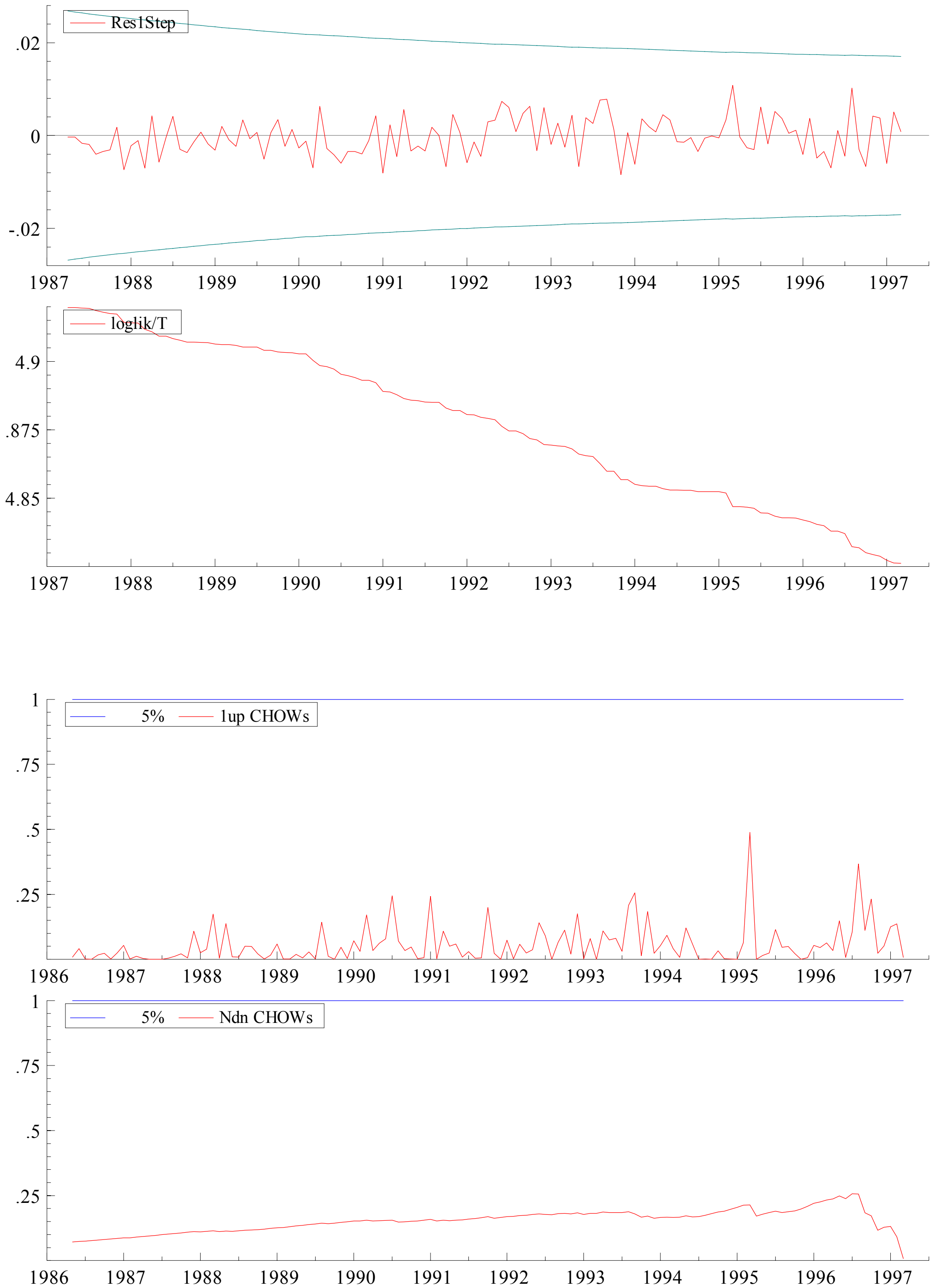

Figure 3: Model stability and constancy test