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Money Demand in Ireland 1972-1989: A Dynamic Equation*

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Abstract: The relationship between the demand for real money and its principal determinants is the vital link by which the actions of the monetary authorities affect the real economy. The estimation of money demand functions that are stable over a long period of time has traditionally, however, proved difficult. Here we attempt to overcome these difficulties by using econometric methods that analyse the time series properties of the data and that also rely on careful diagnostic testing. We achieve some success in estimating a relationship between real M3 and a small number of determinants that appears to be satisfactorily stable over an 18 year period.

I INTRODUCTION

For many countries the intermediate objective of monetary policy is the money supply, the level of interest rates or the exchange rate. Each of these variables interacts with the demand for money and so knowledge of the money demand function is necessary to assess the effects of monetary actions. Not surprisingly, given this fundamental role, much work has been done in estimating the demand for real balances equations. This empirical work has, in general, not been very successful. For instance, there have been large variations between different studies in the magnitudes of parameter estimates;

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particular variables have been found to be significant in some studies but insignificant in others; and parameter coefficients have been unstable over time causing the equations to forecast poorly. In fact, money demand research is arguably the most notorious area of econometric "failure".

In the 1970s when this breakdown became apparent the response was to try feverishly to find what Goldfeld (1976) called the "missing money". There are three main strands to the explanations of why previously estimated econometric relationships became unstable. The first is to state that the relationship between real balances, economic activity and interest rates postulated by theory simply does not exist over any reasonably long period of time. This is not a very useful response given that a money demand equation is the vital link between the actions of monetary authorities and the real economy. The second possible explanation is that the demand for money functions estimated in the early 1970s were too simple, and broke down because they failed to include some vital missing variables that, for instance, capture institutional changes, financial innovations and/or differences in inflation regimes.[1] A third explanation for the problems with these equations relates to the econometric methods used. In particular, it suggests that ignoring the dynamic properties and time series nature of the data will lead to equations that exhibit out of sample instability. This viewpoint was put forward by David Hendry, amongst others, in a large number of papers, e.g., Hendry (1983).

This paper follows the third of these approaches. We take a theoretically conventional money demand specification that suggests the demand for real balances is determined by some measure of the level of economic activity and interest rates. The innovative part of our work is in applying newly developed econometric techniques that look at the time series (including seasonal) properties of our data in order to uncover a stable relationship. We use the cointegration methodology, which focuses on the long-run properties of the data, to identify the combination of variables that will give an equilibrium relationship. The estimation of the dynamic equations borrows from the Hendry approach in our recognition of the importance of lags and of careful diagnostic testing, though our approach does not have all the characteristics of a Hendry-type equation.

The next section describes the econometric methodology in some detail. Section III shows how the equation we estimate relates to a conventional money demand specification. Section IV discusses our empirical results and the final section summarises and concludes.

1. Judd and Scadding (1982) survey this literature.

II ECONOMETRIC METHODOLOGY

The estimation of our equation is performed using a two stage approach. The first stage, which uses the cointegration methodology developed by Engle and Granger (1987), focuses on the long-run time series properties of the available data, enabling us to isolate a set of variables that give a stable equilibrium relationship between real balances and its determinants. Using cointegration to choose between variables is by now a much used technique,[2] so here we only briefly restate the method.

Variables which themselves are non-stationary in a particular way (i.e., are I(1) in the terminology of the cointegration literature) can produce stationary residuals when combined in a linear regression, if a long-run equilibrium relationship exists between the variables. Tests can be performed to determine the order of integration of the individual series and the presence or otherwise of cointegration between the sets of variables. We test for order of integration using Dickey-Fuller and Durbin-Watson statistics; however, since we know from previous work that the data are of a seasonal (monthly) nature we also use HEGY statistics developed in recent papers by Hylleberg, Granger, Engle and Yoo (1988) and Beaulieu and Miron (1989). These papers show that there is more than one form of integration of a series, in particular a series can have a unit root at one or more seasonal frequency. In this terminology, the Dickey-Fuller tests for first order of integration are testing for the existence of a unit root at the zero frequency. Our interest in this literature is that it shows how to test for zero frequency unit roots in the presence of different forms of seasonality.

Basically, the method involves putting a single series through a number of specifically chosen filters, regressing one of the filtered series on the others, and on lags and deterministic seasonals, and then testing the significance of parameters of this regression individually and in combination. Hylleberg *et al.* (1988) show these tests can be interpreted as demonstrating the existence or non-existence of unit roots at different frequencies.

Critical values for these tests have been produced by a Monte-Carlo study for monthly data by Beaulieu and Miron (1989). Once the order of integration and seasonal properties of a series of data are fully understood, the cointegrating regression can be estimated as before, the residuals of which should have a unit root at the zero frequency. However, the critical values for testing the seasonal properties of regressions of residuals are not yet available; we dealt with this problem by treating the residuals as an ordinary time series and using

2. Kearney and McDonald (1988) and Melnick (1988) are other examples of where it has been used in a money demand context.

the Beaulieu and Miron critical values. This is not strictly valid; but the Beaulieu and Miron critical values should form a lower bound for the correct values, and we interpret the rejection of the null hypothesis of a unit root at the zero frequency as not ruling out the existence of an equilibrium relationship between the variables. The second stage, which is dependent on the success of the first, uses the variables that have been found to be in long-run equilibrium to specify a dynamic equation, which is then estimated in first differences, with a number of lags. To ensure that this equation is consistent with the long-run equilibrium relationship estimated at the cointegration stage, the I(0) residuals from the cointegration regression are included in the dynamic regression as an error correction term. The statistical significance or otherwise of the coefficient on this term is a fuller test, albeit a weak one, of whether the cointegrating regression validly represents an equilibrium relationship. The final part of the analysis is to test the dynamic equation for stability, both within sample and out of sample.

III THE FORMAL MODEL

The model we estimate is a reparameterisation of a very simple money demand specification which postulates that the demand for real balances depends positively on some measure of economy wide activity and negatively on the opportunity cost of holding money. The question of what should be the measure of activity (the "scale variable") and the measure of opportunity cost has been much discussed in the literature (e.g., Mankiw and Summers (1986), Kearney and MacDonald (1988)). For the moment, we simply denote our scale variable "x" and our opportunity cost cariable "r".

One commonly used functional form for modelling money demand is given by:

$$\frac{M}{P} = kx^{\alpha}e^{\beta r} \tag{3.1}$$

Where $\frac{M}{P}$ is real balances, k is a constant, α is the elasticity of money demand with respect to the scale variable and β is the semi-elasticity of money demand with respect to the opportunity cost variable. (The form reduces to the quantity theory of money if $\alpha = 1, \beta = 0$.) The log linear form is given by:

$$\left[\log \frac{M}{P}\right]_t = K + \sum_{i=1}^{\infty} \chi_i \left[\log \frac{M}{P}\right]_{t-i} + \sum_{i=0}^{\infty} \alpha_i \left[\log x\right]_{t-i} + \sum_{i=0}^{\infty} \beta_i r_{t-i} \tag{3.2}$$

However, this model would be difficult to estimate in the above form; the lags of a given variable would be highly collinear with each other. Following the Hendry approach we reparameterise the above model[3] by writing it in first difference form, with the level of each variable appearing only once:

$$\Delta\left[\log\frac{M}{P}\right]_t = \nu + \sum_{i=1}^{\infty} c_i\ \Delta\left[\log\frac{M}{P}\right]_{t-i} + \sum_{i=0}^{\infty} a_i\,\Delta\left[\log x\right]_{t-i}$$

$$+ \sum_{i=0}^{\infty} b_i\,\Delta r_{t-i} + \psi\left[\log\frac{M}{P} - \mu - \lambda\log x - \zeta r\right]_{t-1} \quad (3.3)$$

where the new parameters a_i, b_i, c_i, ψ, μ, λ, ζ, are particular linear combinations of the α_i, β_i, χ_i (see Hendry, 1989, p.13). The constant K in Equation 3.2 has been subdivided into ν and $\psi\mu$. The term in square brackets, which in Hendry's terminology is called the Error Correction Mechanism (ECM), gives the long-run equilibrium form of the model. By the Engle-Granger (1987) representation theorem, it comes from the cointegrating regression. The coefficient on this term ψ is a linear combination of the coefficients of the lags on real balances in 3.2 only, in particular:

$$\psi = \sum_{i=1}^{\infty} \chi_i - 1 \quad (3.4)$$

which, a priori, should be negative, if the equation is to be non-explosive.

To estimate this equation, we limit the numbers of lags to three (a testable limitation) and add a normal zero mean, error term, giving:

$$\Delta\left[\log\frac{M}{P}\right]_t = \nu + \sum_{i=1}^{3} c_i\,\Delta\left[\log\frac{M}{P}\right]_{t-i} + \sum_{i=0}^{3} a_i\,\Delta\left[\log x\right]_{t-i}$$

$$+ \sum_{i=0}^{3} b_i\Delta r_{t-i} + \psi\ ECM_{t-1} + E_t \quad (3.5)$$

The long-run equilibrium is estimated by regression separately, and the residuals of this regression are input as the ECM term in the dynamic equation.[4] If

3. See Cuthbertson (1985), p. 266 for an example of this type of reparameterisation and Gilbert (1989) for a discussion of the Hendry approach to econometrics.

4. As pointed out by a referee, an alternative approach would be to estimate Equation 3.5 in one stage using non-linear least squares.

the ECM term is a valid equilibrium representation, then the parameter should be significant, though as emphasised in Engle and Granger (1987), the t-test on the parameter is only a weak test for cointegration.

IV RESULTS

4.1 *Stage One: Choice of Variables*

Pragmatically, the appropriate choice of variables for money demand models is that which results in a stable function and which makes economic sense. Here we use the cointegration methodology, as outlined above, to guide us in our choice of variables.

From our previous work (Hurley and Guiomard, 1989) we decided to focus on M3 as our measure of money and Retail Sales Volume as the scale variable. We had a large number of possible opportunity cost variables: inflation, various domestic interest rates, and various foreign interest rates. Since Ireland is a small open economy, non-domestic variables could impinge on domestic money demand; the currency substitution theory of money demand suggests that foreign/domestic interest rates differentials will be an important determinant.[5] We allowed for this possibility by including a German and British three month interest rate in our data set.

The first step in any cointegration analysis is to determine the order of integration of the variables. We do this using the by now conventional Dickey-Fuller tests, in combination with the more unusual HEGY tests, which test for the presence of unit roots in series that may possibly be of a seasonal nature. The results of these tests are in Table 1. To summarise the table, all variables appear to be I(1); the null hypothesis of a unit root at the zero frequency cannot be rejected when the variables are in levels, but is rejected for the changes of variables at the 5 per cent level of significance. Some of the interest rates seem to be borderline I(0)/I(1) on the basis of the Dickey-Fuller tests; the seasonal order of integration tests indicate that they are more likely to be I(1). The other property of the data which is apparent in this table is the importance of deterministic seasonality: most of the variables exhibit significant F-statistics.

Having established that the variables are I(1), we test for cointegration. This involves regressing real M3 on retail sales volume and various opportunity cost variables. The initial results, reported in Table 2(a), of two and three variable cointegrating regressions are not supportive of the hypothesis of a long-run equilibrium relationship. We went on to estimate four variable co-integrating regressions, which include a second opportunity cost variable.

5. See Cuthbertson (1985), pp. 116-117.

Table 1: *Tests of Order of Integration 1971:12 to 1989:6*

Variable	*DF*	*ADF**	*# of Lags*	*HEGY Test** Zero Frequency*	*F-tests*** Deterministic Seasonals*
LM3RPI	-2.32	-2.7	12	-2.351	4.91
LRSVOL	-8.6277	-1.8713	15	-1.4516	15.418
RPINFL	-6.1641	-2.8765	15	-2.5328	5.7480
EBILL	-2.5976	-2.7165	1	-2.1997	1.497
RLONG	-1.8784	-2.1904	1	-2.3473	0.92576
RSPREAD	-4.6861	-3.3047	1	-3.1083	1.3228
RDUK	-2.9422	-2.2025	1	-0.91179	1.4361
RDGER	-2.9241	-2.9956	1	-2.1511	1.7472
ΔLM3RPI	-14.81	-3.3229	11	-3.637	3.231
ΔLRSVOL	-22.344	-3.277	14	-4.4543	9.5937
ΔRPINFL	-18.571	-6.943	13	-6.1124	5.0156
ΔEBILL	-13.999	-10.544	1	-4.7546	1.5723
ΔRLONG	-12.249	-8.2974	1	-4.2664	1.0029
ΔRSPREAD	-16.237	-11.947	1	-5.4528	1.38
ΔRDUK	-17.901	-13.313	1	-5.0527	1.691
ΔRDGER	-13.958	-10.596	1	-5.6262	1.8847
Critical 1% level	-3.51	-3.51		-3.85	2.28
Values 5% level	-2.89	-2.89		-3.31	1.81

* Ho: Existence of unit root; large values for the levels reject Ho, suggesting that the variable is not I(1).

** Ho: A unit root at the zero frequency; large values reject Ho. Critical values are from Beaulieu and Miron (1989).

*** Ho: Deterministic seasonality is not present; large values reject Ho suggesting that the variable is deterministically seasonal.

This allows for the fact that changes in relative interest rates may affect demand for real balances by causing portfolio shifts; either between assets at different parts of the maturity spectrum, or between assets at home and abroad. Once again (see Table 2(b)) none of the regressions satisfy all the co-integration criteria; however, the three regressions with the difference between long and short interest rates have reasonably high HEGY statistics (over three in absolute value). These results could be interpreted as supporting the existence of a long-run equilibrium relationship between at least one of these sets of variables, with the low values of the CRADFs simply a result of the significantly seasonal nature of the data. Of these three regressions, the first seems the most likely to be useful; all the variables are significant, the R^2 is reasonably high and the CRDW is borderline between rejecting/accepting cointegration.

Table 2a: *Cointegration Regressions 1971:12 to 1989:6*

		Test Statistics						
Variables	*Cointegrating Regressions*	R^2	*CRDW*	*CRDF*	*CRADF*	*# of Lags*	*HEGY Test Zero Frequency*	*F-tests Seasonals*
LM3RPI, LRSVOL	LM3RPI = 5.1692 + 0.7168LRSVOL (0.5412) (0.1139)	15.9%	.179	−3.44	−1.16	14	−2.415	12.865
LM3RPI, LRSVOL, RPINFL	LM3RPI = 4.884 + 0.7991LRSVOL − 0.00834RPINFL (0.4588)(0.1019) (0.001128)	33.6%	.271	−4.3	−1.99	12	−2.35	7.859
LM3RPI, LRSVOL EBILL	LM3RPI = 5.762 + 0.555LRSVOL + 0.00119EBILL (0.5111)(0.1153) (0.0027)	23.1%	.131	−2.68	−0.65	12	−3.11	16.97
LM3RPI, LRSVOL RLONG	LM3RPI = 5.17 + 0.7164LRSVOL − 0.00056RLONG (0.522)(0.1178) (0.0036)	15.9%	.179	−3.44	−1.157	14	−2.42	12.86
	Critical Values,* 5% level		.367	−3.93	−3.67		−3.31	1.81

* The critical values are for three variable cointegrations. The appropriate 5% critical values for CRDW, CRDF, CRADF for two variables are, respectively 0.368, −3.37, −.3.17. The critical value for the HEGY-t quoted is approximate only.

Ho: Existence of unit root in the cointegration regression residuals; high (absolute) values of the test statistics reject the null hypothesis, suggesting that the residuals are stationary and the variables are cointegrated.

Table 2b: *Cointegration Regressions 1971:12 to 1989:6*

		Test Statistics						
Variables	*Cointegrating Regressions*	R^2	*CRDW*	*CRDF*	*CRADF*	*# of Lags*	*HEGY Test Zero Frequency*	*F-test Seasonals*
LM3RPI, LRSVOL RPINFL, RSPREAD	LM3RPI = 5.676 + 0.6301LRSVOL − 0.0068RPINFL − 0.0274RSPREAD (0.427) (0.0946) (0.0011) (0.0039)	46%	.309	−4.48	−2.01	11	−3.46	5.9*
LM3RPI, LRSVOL RPINFL, RDUK	LM3RPI = 4.911 + 0.7918LRSVOL − 0.0085RPINFL − 0.0077RDUK (0.455) (0.101) (0.0012) (0.0037)	35%	.28	−4.31	−1.84	12	−2.55	7.71
LM3RPI, LRSVOL RPINFL, RDGER	LM3RPI = 5.1501 + 0.7217LRSVOL − 0.0073RPINFL − 0.0147RDGER (0.4207) (0.0938) (0.001) (0.0023)	45%	.28	−4.06	−1.64	12	−2.34	7.63
LM3RPI, LRSVOL EBILL, RSPREAD	LM3RPI = 6.0467 + 0.5362LRSVOL − 0.0006EBILL − 0.0336RSPREAD (0.4703) (0.1056) (0.0032) (0.0053)	36%	.245	−3.92	−2.08	12	−4.3	10.369*
LM3RPI, LRSVOL EBILL, RDUK	LM3RPI = 6.0289 + 0.4825LRSVOL + 0.0185EBILL + 0.0139RDUK (0.5135) (0.1168) (0.0036) (0.0052)	26%	.107	−2.47	−0.43	19	−2.69	15.188
LM3RPI, LRSVOL EBILL, RDGER	LM3RPI = 5.496 + 0.06232LRSVOL + 0.0016EBILL − 0.0158RDGER (0.4884)(0.1104) (0.0033) (0.0032)	31%	.19	−3.14	−0.83	12	−2.5	17.6
LM3RPI, LRSVOL RLONG, RSPREAD	LM3RPI = 6.0467 + 0.5362LRSVOL − 0.0006RLONG − 0.0331RSPREAD (0.4703) (0.1057) (0.0031) (0.0041)	36%	.245	−3.92	−2.08	12	−4.3	10.37*
LM3RPI, LRSVOL RLONG, RDUK	LM3RPI = 5.1223 + 0.7351LRSVOL − 0.0032RLONG − 0.0662RDUK (0.5228) (0.1186) (0.0045) (0.0051)	17%	.195	−3.57	−1.09	12	−2.76	14.4
LM3RPI, LRSVOL RLONG, RDGER	LM3RPI = 5.2052 + 0.724LRSVOL − 0.0132RLONG − 0.0216RDGER (0.4596) (0.1037) (0.0036) (0.0028)	35%	.282	−4.08	−1.63	12	−2.74	16.94

Therefore, we tentatively conclude that the reported results indicate the existence of a long-run equilibrium relationship between real balances, retail sales, inflation and the term structure spread, and move on to the next stage of estimating a dynamic equation with these variables. The residuals from the first regression, Table 2(b), which we have just found to be I(0), will be used as an ECM in the dynamic equation. The significance of the parameter on the ECM is another test, albeit one of low power, of the validity of our conclusion above (see Engle and Granger, 1987).

4.2 *Stage Two: Dynamic Modelling*

Non-rejection of the stationarity of the cointegrating regression residuals means that it is valid to proceed to the next stage; the specification and estimation of the error correction money demand model, which contains both the long-run equilibrium money demand relationship and the short-run dynamics. Seasonal dummies will form part of the specification; our analysis of the time series properties of the data described in the last section has shown that deterministic seasonality is present in most of our individual data series.

We first estimated Equation 3.5 with the two opportunity cost variables suggested by the cointegration analysis; namely retail sales inflation and the spread between long and short interest rates. Our analysis was of the period up to June 1989, with six observations retained for tests of out of sample stability. The dependent variable in all of our models is the first difference of M3, deflated by the retail sales price index, in log terms. Three lags of each of the non-dummy variables were allowed for.[6] A trend was included to account for any possible increase in velocity of circulation over time: possibly capturing a financial innovation effect. It was found to be statistically significant. The parameter estimates and standard errors are reported in the first part of Table 3. At first glance, the model appears to fit reasonably well; the R^2 of 50 per cent is respectable for a model in first difference form, and all variables, except possibly the constant, appear significant. A closer look at the residuals, however, indicates that there are some problems with the model. From the way that the model was specified the residuals should be normally distributed with constant variance, and should not be serially correlated. Also, the model parameters should not change significantly over the forecast period.

We can test whether the model is well-behaved by looking at a series of diagnostics output at the end of Table 3.[7] One asterisk in the column marked

6. The GIVE package has a limit of 40 variables in any one model. Within this constraint up to four lags of each of the non-dummy variables could be included; however, the F-test of the reduction from a four lags model to a three lags model was not significant, implying that the fourth lag added little to the model fit.

7. For a discussion of these statistics see O'Reilly (1985), Hendry (1989) and Harvey (1981).

Table 3: *Dynamic Modelling*†

		Model 1 1972:3 to 1989:6 less 6 forecasts		*Model 2 1972:3 to 1989:6 less 6 forecasts*		*Model 3 1972:3 to 1989:6 no forecasts*	
Variable	*Lag*	*Coefficient*	*Std Error*	*Coefficient*	*Std Error*	*Coefficient*	*Std Error*
ΔLM3RPI	1	−.10482	.07053	−.12943	.06642	−.11880	.06609
ΔLM3RPI	2	.03906	.06967	−.00840	.06644	−.05607	.06632
ΔLM3RPI	3	.28177	.07117	−.24888	.06628	.22235	.06561
CONSTANT		.00731	.00708	.01035	.00656	.00934	.00668
TREND		.00005	.00002	.00005	.00002	.00005	.00002
ΔLRSVOL	0	.09713	.02744	.12356	.02554	.13555	.02580
ΔLRSVOL	1	.02471	.03200	.02126	.02944	.03118	.02989
ΔLRSVOL	2	.06077	.03120	.06302	.02888	.06882	.02922
ΔLRSVOL	3	.01134	.02854	.00683	.02649	.00623	.02685
ΔRPINFL	0	−.00054	.00027	−.00048	.00026	−.00045	.00026
ΔRPINFL	1	−.00042	.00026	−.00031	.00024	−.00027	.00024
ΔRPINFL	2	−.00064	.00025	−.00044	.00023	−.00041	.00024
ΔRPINFL	3	−.00020	.00026	−.00001	.00024	.00004	.00025
ΔRSPREAD	0	.00055	.00089	.00504	.00159	.00494	.00162
ΔRSPREAD	1	.000309	.00094	.00096	.00179	.00087	.00182
ΔRSPREAD	2	.00204	.00095	.00102	.00184	.00139	.00188
ΔRSPREAD	3	−.00042	.00080	.00229	.00115	−.00232	.00117
ECM	1	−.05088	.01539	−.05197	.01417	−.05406	.01442
ΔEBILL	0			.00590	.00613	.00595	.00165
ΔEBILL	1			−.00203	.00172	−.00206	.00174
ΔEBILL	2			−.00081	.00176	−.00043	.00178
ΔEBILL	3			−.00252	.00130	−.00254	.00133
EMS				.05977	.01258	−.06182	.01283

	Model 1		*Model 2*		*Model 3*	
Test Statistics	*Value*	*Sig*	*Value*	*Sig*	*Value*	*Sig*
R^2	$R^2 = .486$		$R^2 = .589$		$R^2 = .57$	
SE of Regression	$\sigma = .01276$		$\sigma = .01158$		$\sigma = .01183$	
Residual of Sum of Sqs	RSS = .0282		RSS = .0225		RSS = .0244	
Forecast χ^2	$\xi_1(6) = 2.45$	*	$\xi_1(6) = 2.60$	*	$\xi_1(0) =$	
AR Errors	$\xi_2(7) = 17.341$	*	$\xi_2(7) = 11.143$		$\xi_2(7) = 8.133$	
ARCH	$\xi_3(7) = 5.813$		$\xi_3(7) = 4.231$		$\xi_3(7) = 7.681$	
Normality	$\xi_4(2) = 19.922$	**	$\xi_4(2) = 2.555$		$\xi_4(2) = 1.937$	
Forecast CHOW	$\chi_1(6,173) = 2.16$	*	$\chi_1(6,168) = 2.27$	*	$\chi_1(0,170) =$	
Heteroskedasticity	$\chi_2(35,146) = .4888$		$\chi_2(35,146) = .4136$		$\chi_2(35,152) = .8846$	
Reset F	$\chi_3(1,172) = 2.998$		$\chi_3(1,167) = 2.86$		$\chi_3(1,173) = 2.027$	

†Monthly dummies were also estimated but are not included due to space considerations.

"significance" indicates that the null hypothesis is rejected at the 5 per cent level, but not at the 2.5 per cent level, two asterisks indicates rejection at the 2.5 per cent level.

Model 1 fits reasonably well in the prediction period – the null hypothesis of stability is only barely rejected but the errors are autocorrelated. The null hypothesis of normally distributed errors is strongly rejected by the data, as indicated by the Jarque-Bera statistic. On closer examination of the residuals, it appears that one observation has a particularly high residual; the model fails to fit at this one point. Looking at plots of the dependent variable in levels and first differences (Figure 1) shows that though the pattern in levels of M3 before and after EMS seems similar, the intercept term changes, giving one particularly large first difference. As has been noted by Patrick Honohan (1984), a once and for all jump in money demand is theoretically expected to occur when a currency union breaks up, as happened to the Irish/UK currency union at the beginning of the EMS. The inclusion of an EMS dummy which allows the constant term in our model to be different at this one observation should solve this problem.

The other problem, that of autocorrelation, indicates that there is perhaps a missing variable; we decided to try using ΔEBILL[8] as a separate variable in addition to having it in the model as part of ΔRSPREAD. There is no particular reason why the equal but opposite signs constraint on EBILL and RLONG that appears in the long-run equilibrium equation should also hold in the short-run dynamic formulation of the model.

Model 2 includes these two extra variables; with three lags of the ΔEBILL variable. The coefficient on the EMS dummy is significant, with a t-statistic of over three, as can be seen from ζ_4, the Jarque-Bera statistic, the non-normality problem has disappeared. The lags of the ΔEBILL terms are jointly significant, though not all of the coefficients are individually significant. As would be expected when extra variables are added, the R^2 has increased and the Standard Error (SE) of the regression is smaller. Heteroskedasticity does not appear to be a problem.

However, there is some evidence of parameter instability over the prediction period; the significance value on the forecast CHOW F-test goes from 0.049 (just rejected at 5 per cent level) in Model 1 to 0.039 in Model 2. The null hypothesis that the model fits as well out-of-sample as it does in-sample is rejected at the 5 per cent level, though not at the 2.5 per cent level. The fit of the model over the prediction period is displayed in Figure 2. We tentatively accepted the model as adequate, though not as stable out of sample as we would have liked.

8. The acronyms used for each variable are set out in an appendix.

Figure 1

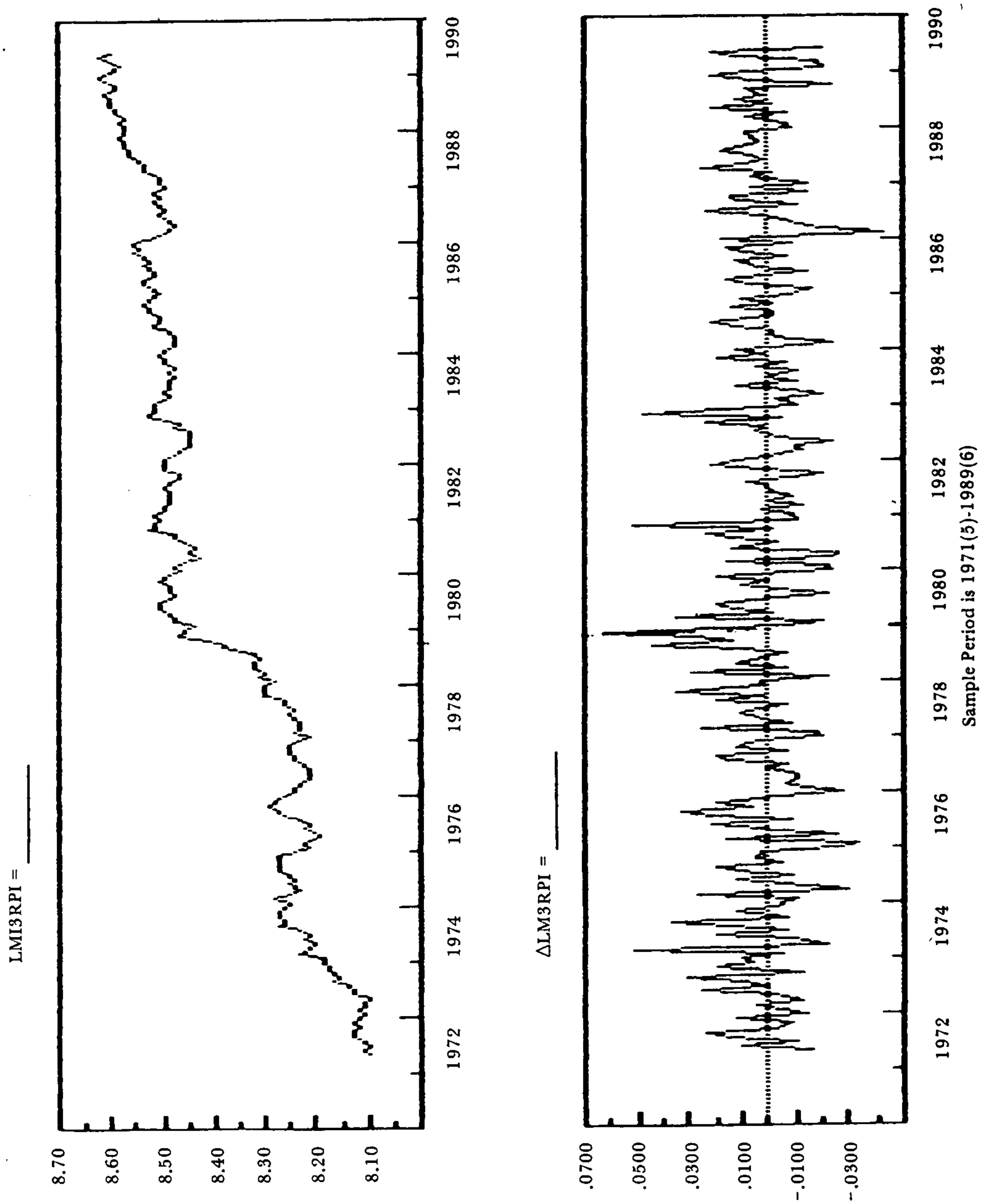
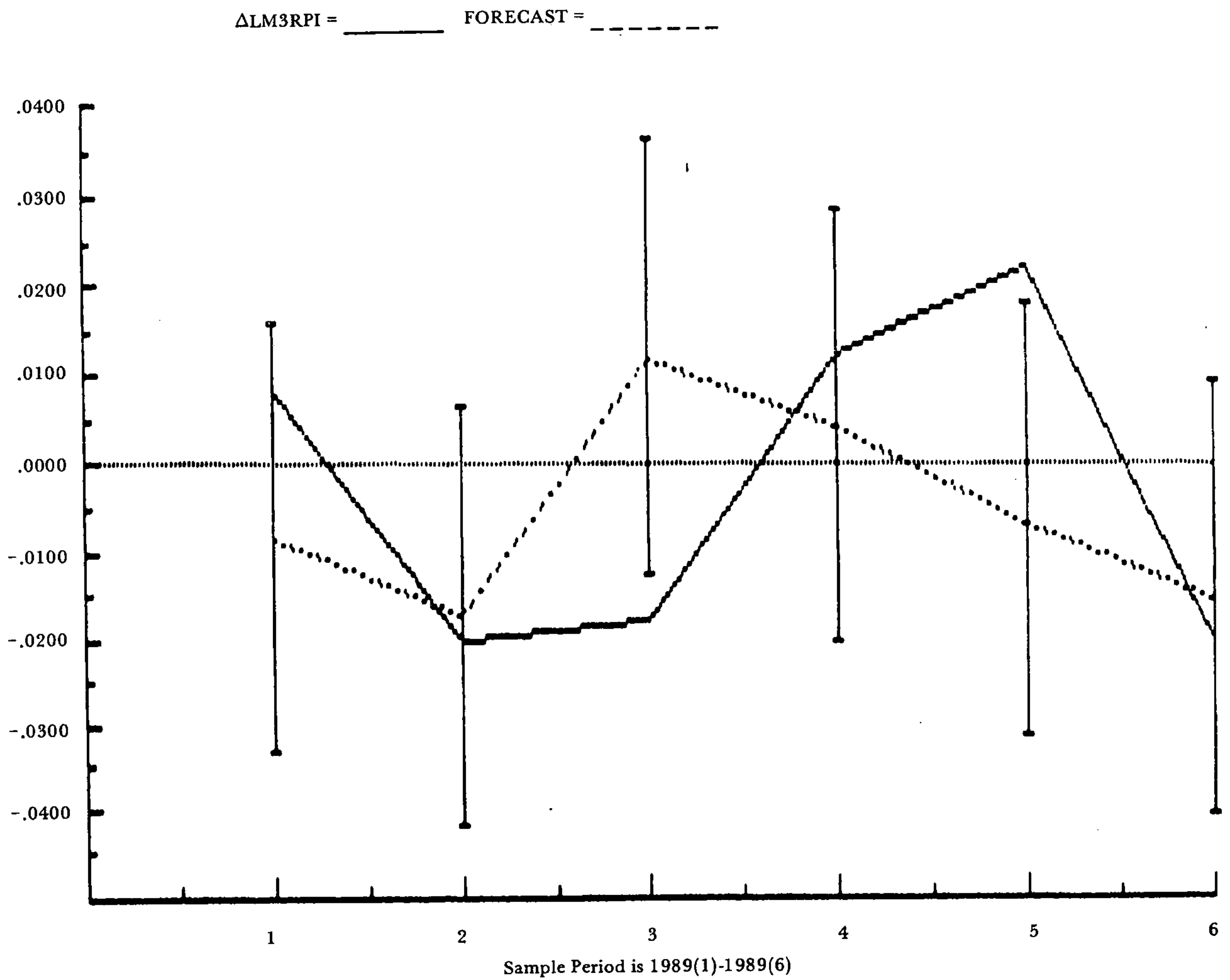

LMI3RPI =
8.70
8.60
8.50
8.40
8.30
8.20
8.10
1972
1974
1976
1978
1980
1982
1984
1986
1988
1990
ΔLM3RPI =
.0700
.0500
.0300
.0100
-.0100
-.0300
1972
1974
1976
1978
1980
1982
1984
1986
1988
1990
Sample Period is 1971(5)-1989(6)

Figure 2


ΔLM3RPI =
FORECAST =
.0400
.0300
.0200
.0100
.0000
-.0100
-.0200
-.0300
-.0400
1
2
3
4
5
6
Sample Period is 1989(1)-1989(6)

Our final step is to re-estimate the reduced model for the complete time period (up to mid-1989), with no observations put aside for forecasts, to facilitate the calculation of elasticities. This is reported in Table 3 in the third column.

4.3 *Properties of the Dynamic Model*

We now look more thoroughly at the stability of our final model (Model 3). We have already found that there is a break at end 1978, which we allowed for by introducing an EMS dummy. However, it is possible that the change of regime engendered by Ireland's entry into the EMS and the breaking of the currency union with the UK could have caused a more fundamental change in the pattern of money demand; any or all of the parameters in the model could have changed. To test whether the same model fitted adequately in both the pre- and post-EMS period, separate models were estimated for the two periods, and their combined errors can be compared with the error from the model which constrains all the parameters to be equal in the two regimes. This is a simple CHOW test. The statistic, which has an F-distributional form, is:

$$F(28,163) = 0.74316$$

i.e., though there is some gain in explanatory power (a smaller residual sum of squares) from allowing parameters to differ pre/post-EMS, the difference is small, and is insignificant in the statistical sense.

A graphical test for model stability, where the possible point of structural change is not known, is the CUSUMSQ test.[9] This is shown in Figure 3. If there is no structural break during the complete time period the statistic should climb steadily from zero to one, without wandering around too much. The test is only an approximate one, but the upper and lower bounds are 1 per cent bounds, which, in this instance, are never breached. This indicates that there has not been a structural change in pattern of real money demand in Ireland in the last 18 years; a very strong result, given the discussion of money demand "breakdown" in the international literature.

Two types of elasticities can be inferred directly from the model. Since the model is in log form the parameter on retail sales volume in Model 3 can be interpreted directly as an impact elasticity, i.e., a 1 per cent rise in retail sales volume will cause 0.136 per cent of an increase in real M3. The parameters on the opportunity cost terms are semi-elasticities. In the long run, when all changes are zero, the model reduces to the cointegration regressions; the cointegration regression parameters are therefore interpretable as long-run elasticities and semi-elasticities.

9. See Harvey (1981), p. 152.

Figure 3

CUSUMSQ = ------ lower = ______ upper = ______

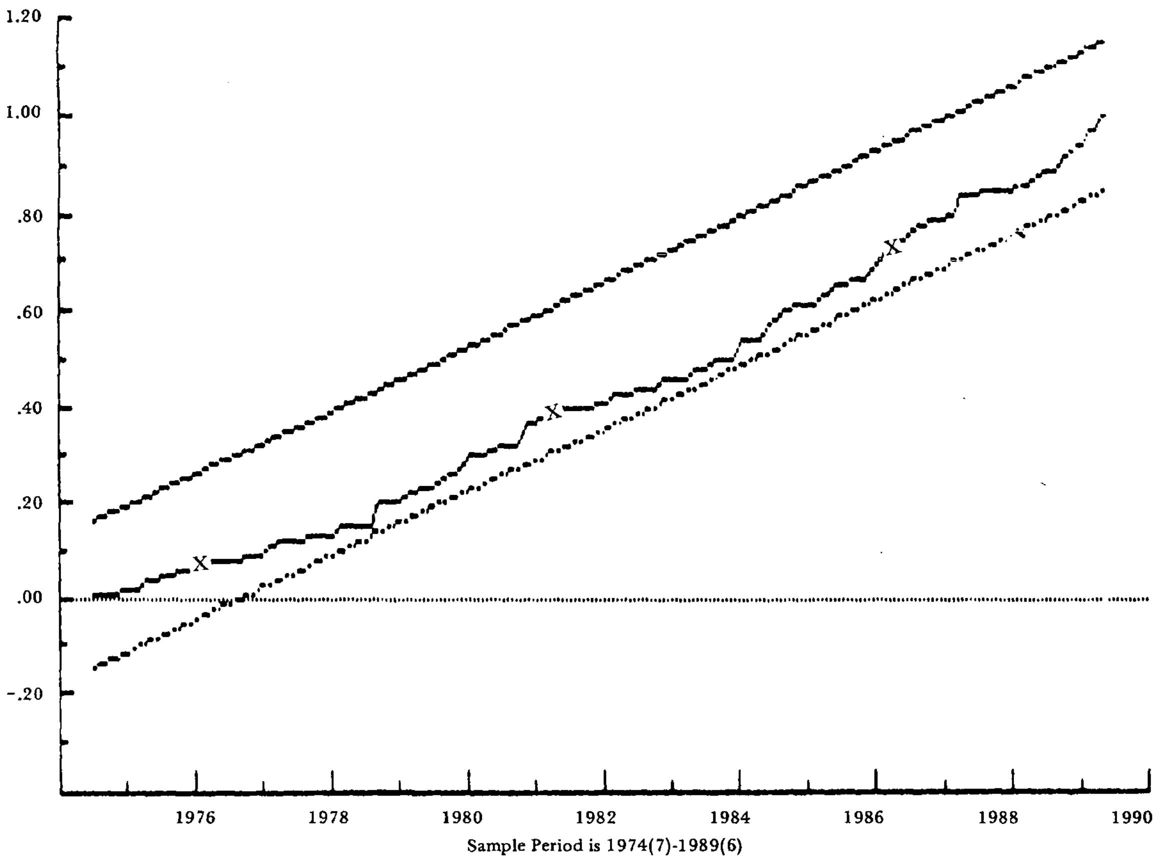
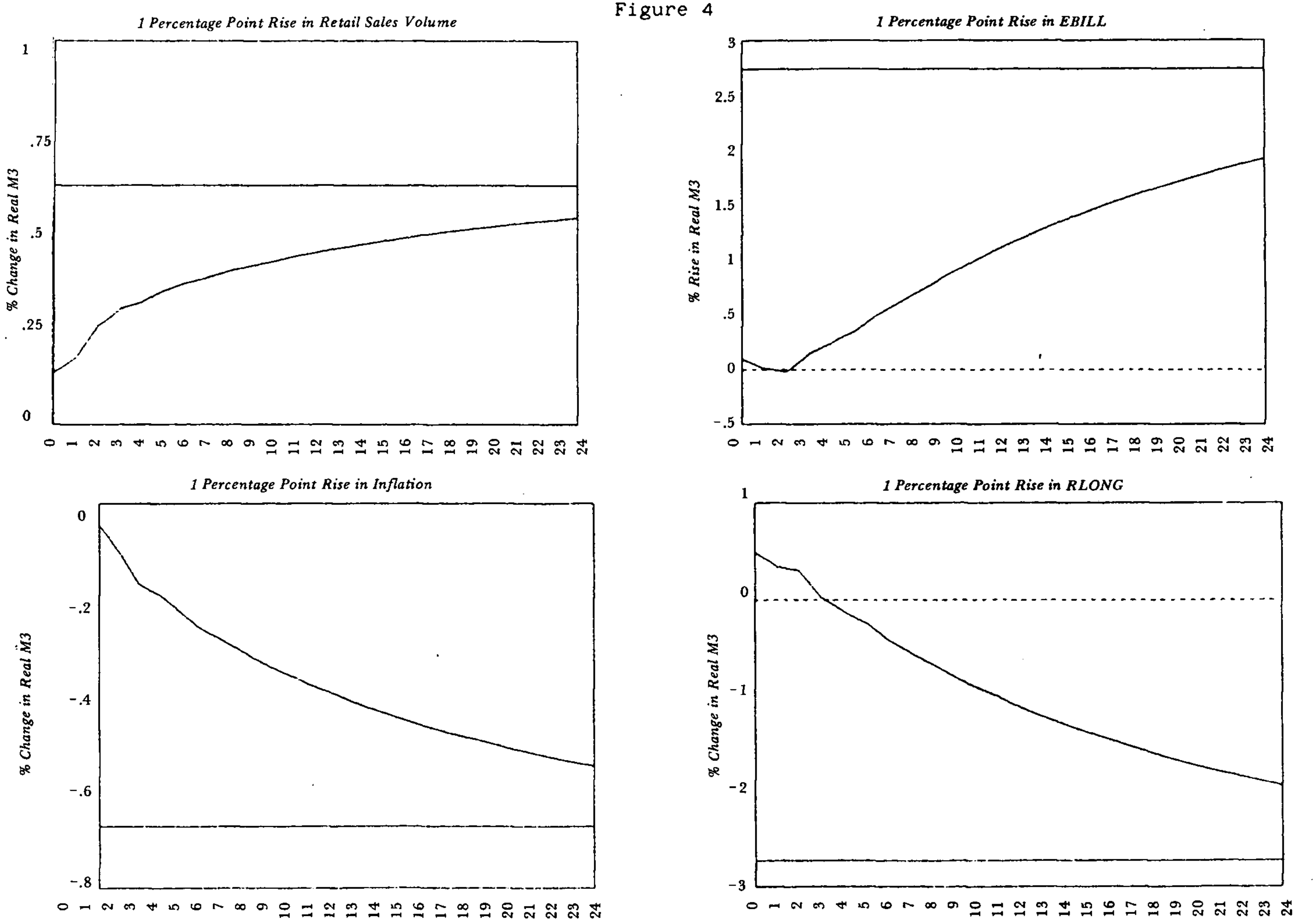


Figure 4


1 Percentage Point Rise in Retail Sales Volume
% Change in Real M3
1 Percentage Point Rise in EBILL
% Rise in Real M3
1 Percentage Point Rise in Inflation
% Change in Real M3
1 Percentage Point Rise in RLONG
% Change in Real M3

Graphs of the effect over time of a *ceteris paribus* increase in the dependent variable of 1 per cent (1 percentage *point* increase for the variables already expressed as percentages) on the independent variable are given in Figure 4. The straight line in each of the graphs is the long-run elasticity from the co-integration regression. A number of comments can be made about these graphs:

1. Changes in the dependent variables affect real M3 over a relatively long time span; after 24 months only two-thirds to three-quarters of the total effect has occurred.
2. The long-run elasticity of retail sales volume is significantly less than unity.
3. Interest rate changes have quite sizeable effects on real M3, in the long run a 1 per centage point increase in the long interest rate leads to a 2 per cent fall in money demand, *ceteris paribus*. However, it should be noted that, since term structure theory suggests that interest rates are interlinked and move broadly in line with each other, a 1 per cent point in the term structure spread in one period is an unusually large change. Normally, it could be expected that a 1 per cent point change in one of the interest rates would affect other rates, so that the change in the interest rate spread (at least in the longer run) would be less than 1 per cent point. More precise estimates of complete effects of changes in individual interest rates (i.e., a relaxation of the *ceteris paribus* assumption) are beyond the scope of a single equation model.
4. Changes in inflation have smaller but still statistically significant effects on real M3. A 1 percentage point sustained rise in inflation leads to a long-run decrease in money demand of .68 per cent of 1 per cent; roughly £70 million, given that the current stock of nominal M3 is about £10 billion.

V CONCLUSION

In this paper we have estimated a new Irish money demand equation for the years 1972 to 1989. Why, one may ask, another money demand study, given that there have been numerous previous studies? (The longer, technical paper, version of this article contains a survey of these.) There are two main answers.

First, we wished to see whether modern econometric methods would successfully produce a stable and conventional relationship over a lengthy period. Secondly, we wanted to utilise the high frequency (monthly) data which were available, in order to have an equation that might be of practical use in a Central Banking context.

These efforts appear to have been reasonably successful. Intensive testing of the equation, reported above, suggests that the parameters are indeed stable over the 18 year period. This matters for monetary policy because, as Browne and O'Connell put it in an earlier study:

> It is important to know whether or not the demand for money function is stable. In a closed economy, a stable function enables the authorities to know the effects of their open-market operations on the level of interest rates and, secondly, a stable function is a necessary condition for assessing the effects of monetary actions on income. In an open economy with virtually perfect capital mobility, a stable demand for money function enables the authorities to adapt domestic credit policy to attain an external reserves target (Browne and O'Connell, 1977).

REFERENCES

BEAULIEU, JOSEPH, and JEFFREY A. MIRON, 1989. "Evidence on Unit Roots at the Seasonal and Zero Frequencies", Unpublished manuscript, M.I.T.

BROWNE, F.X., and T. O'CONNELL, 1977. "The Demand for Money Function in Ireland: Estimation and Stability", Central Bank of Ireland Research Paper 16/76.

CUTHBERTSON, KEITH, 1985. *The Supply and Demand for Money*, London: Basil Blackwell.

ENGLE, ROBERT F., and CLIVE W.J. GRANGER, 1987. "Cointegration and Error Correction: Representation, Estimation and Testing", *Econometrica*, Vol. 55, No. 3, pp. 251-276.

GILBERT, CHRISTOPHER L., 1989. "Econometric Theory and Econometric Models", *The Economic and Social Review*, Vol. 21, No. 1, pp. 1-26.

GOLDFELD, STEPHEN M., 1976. "The Case of the Missing Money", *Brookings Papers on Economic Activity*, 3: 1976, pp. 683-730.

HARVEY, A.C., 1981. *The Econometric Analysis of Time Series*, London: Philip Allen Ltd.

HENDRY, DAVID F., 1989. *PC-GIVE: An Interactive Modelling System*, University of Oxford.

HENDRY, DAVID F., 1983. "Econometric Modelling: The Consumption Function in Retrospect", *Scottish Journal of Political Economy*, Vol. 30, No. 3, pp. 193-220.

HONOHAN, PATRICK, 1984. "The Break-up of a Currency Union Increases the Demand for Money", *European Economic Review*, Vol. 25, pp. 235-238.

HURLEY, MARGARET, and CATHAL GUIOMARD, 1989. "Determinants of Money Demand in Ireland: Rounding-Up the Usual Suspects", *The Economic and Social Review*, Vol. 21, No. 1, pp. 139-148.

HYLLEBERG, S., R.F. ENGLE, C.W.J. GRANGER, and B.S. YOO, 1988. "Seasonal Integration and Cointegration", Discussion Paper 88-32, University of California, San Diego.

JUDD, JOHN P., and JOHN L. SCADDING, 1982. "The Search for a Stable Money Demand Function: A Survey of the Post-1973 Literature", *Journal of Economic Literature*, Vol. XX, No. 3, pp. 993-1023.

KEARNEY, COLM, and RONALD MACDONALD, 1988. "The Stability of Monetary Velocities in the UK and USA 1971-1985: A Cointegration Analysis", Central Bank of Ireland, Technical Paper 6/RT/88.

MANKIW, N. GREGORY, and LAURENCE H. SUMMERS, 1986. "Money Demand and the Effects of Fiscal Policies", *Journal of Money Credit and Banking*, Vol. 18, No. 4, pp. 415-429.

MELNICK, RAFI, 1988. "The Demand for Money in Argentina 1978-1987: Before and after the Austral Program", Department of Economics, Boston University, mimeo.

O'REILLY, LIAM, 1985. "The Hendry *et al.* Model Building Philosophy and the GIVE Package", Central Bank of Ireland Technical Paper 3/RT/85.

APPENDIX

Data Acronyms

LM3RPI:	The log of M3 deflated by the retail sales price index.
LRSVOL:	The log of the retail sales volume index.
RPINFL:	Retail sale price index percentage inflation, annualised.
EBILL:	The exchequer bill rate, as a percentage.
RLONG:	The yield on 15 year bonds, in percentage form.
RSPREAD:	The difference between RLONG and EBILL.
RDUK:	The UK T-bill rate minus Irish EBILL.
RDGER:	The German T-bill rate minus Irish EBILL.