Stock Prices and the Monetary Model of the Exchange Rate:

An Empirical Investigation

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October 2003

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Abstract

This paper develops an alternative version of the monetary model of exchange rate determination, which incorporates a stock price measure. This model is then tested using data from Canada and the USA, applying the cointegration and error correction methodology. In contrast to many previous tests of the monetary model, this version produces evidence of cointegration and stock prices have a highly significant effect on the exchange rate in both the short and long run. In addition the restricted version of the model outperforms a random walk in out of sample forecasting.

(JEL Classification: F 32)

1. Introduction

Although the asset market approach to exchange rate determination dominates theoretical exchange rate modelling, attempts to construct empirical models based on the asset approach have met with limited success. This is especially true of the flexible price monetary model, which was shown by Meese and Rogoff (1983) to provide inferior out-of-sample forecasts compared to a random walk. Furthermore attempts to produce the valid long-run equilibrium relationship implied by the monetary model have generally met with mixed success, particularly when the implicit restrictions of the model are applied. For example Meese (1986) and McNown and Wallace (1989), fail to find a valid long-run relationship for the conventional monetary model[1].

This paper develops and tests a version of the monetary model that incorporates stock prices. The analysis is motivated by earlier work by Friedman (1988) and Boyle (1990) that shows how the demand for money is determined in part by the level of the stock market. To date the only attempt to test the role of stock prices on the exchange rate is Smith (1992) who uses a Portfolio Balance approach[2]. We show that including the level of the stock market produces a valid long-run equilibrium relationship and correctly specified dynamic error correction model (ECM). The implicit restrictions of the model are then examined and it is shown that the ECM out-performs a random walk in out-of-sample forecasting.

The remainder of the paper is as follows. Section 2 outlines the theoretical case for including equities in the monetary model and discusses the econometric methodology used in the paper. Section 3 describes the data set and presents the time series results. Section 4 contains the conclusions and considers some implications for the integration of capital markets.

2. Stock prices and money demand

In the conventional monetary model the exchange rate adjusts to balance the international demand and supply of monetary assets. The demand for money is usually considered to be a function of the level of interest rates and income. However there is an increasingly good case for including equity prices as separate determinants of the demand for money. In particular Friedman (1988) and Boyle (1990)[3] provide empirical evidence describing the relationship between money demand and the level of the stock market, including a specific lag structure to the relationship, which due to a different methodology we do not attempt.

On the theoretical side, Friedman (1988) suggests four possible channels through which stock prices might directly effect money demands. Firstly as stock market fluctuations tend to outweigh fluctuations in income, stock market movements are generally associated changes in the wealth to income and hence money to income ratios. Secondly a rise in stock prices reflects an increase in the expected return from risky assets relative to safe assets. The implied increase in portfolio risk can be offset by an adjustment away from other risky assets such as long term bonds toward safer assets including money. Thirdly a rise in stock prices reflects an increased level of financial transactions and thus an increase in the demand for money. The above three 'wealth effects' all suggest a positive relationship between the level of the stock market and money demand. However as the real stock price rises equities become more attractive to investors causing a 'substitution effect' from equities for money. The relationship between equity prices, the demand for money and exchange rate is therefore an empirical question. As with Friedman (1988) we expect the wealth effect to dominate and thus we expect the demand for money and stock prices to be positively related. To capture these effects we incorporate a stock market variable into the standard money demand function,

$$m_t = p_t + \alpha y_t - \beta i_t + \chi s_t \tag{1}$$

Where m is the nominal demand for money, p is the price level, y is the real income level, i is the nominal rate of interest and s is the real level of the stock market (following Friedman (1988), a market index is used). All variables except the interest rate are in logarithms. Foreign money demands are given by,

$$m_t{}^* = p_t^* + \alpha y_t^* - \beta i_t^* + \chi s_t^* \tag{2}$$

Where * denotes a foreign variable. It is assumed that absolute PPP holds, so that,

$$p_t = p_t^* + e_t \tag{3}$$

Where e is the log of the exchange rate, defined as the domestic price of foreign currency. PPP is used only as a long-run equilibrium condition in this model, in the short run the error correction model allows deviations from PPP. The evidence on

PPP as a long-run equilibrium condition is generally positive (Culver and Papell, 1999). Straightforward rearrangement of (1) - (3) yields,

$$e_t = \alpha_0 + \alpha_1(m_t - m_t^*) - \alpha_2(y_t - y_t^*) + \alpha_3(i_t - i_t^*) - \alpha_4(s_t - s_t^*) \tag{4}$$

The monetary approach assumes that domestic and foreign bonds are perfect substitutes so that Uncovered Interest Parity (UIP) holds,

$$i_t = i_t^* + [E(e_{t+1} \mid I_t) - e_t] \tag{5}$$

Where $E(e_{t+1} / I_t)$ is the rational expectation of the exchange rate one period into the future, conditional on the currently available information set I_t. Denoting the set of forcing variables as $X_t = [\alpha_0 + \alpha_1(m_t - m_t^*) - \alpha_2(y_t - y_t^*) + \alpha_4(s_t - s_t^*)]$, substituting (5) into (4) and solving for the exchange rate yields,

$$e_t = \frac{E(X_{t+j} | I_t)}{1+\alpha_3} + \frac{\alpha_3}{1+\alpha_3} E(e_{t+1} | I_t)$$

Solving this equation by forward iteration gives,

$$e_t = (1+\alpha_3)^{-1} \sum_{j=0}^{n} [\alpha_3 / (1+\alpha_3)]^j E(X_{t+j} \mid I_t) + \left(\frac{\alpha_3}{1+\alpha_3} \right)^n E(e_{t+n} | I_t)$$

Letting $j \to \infty$, or assuming that the solution is free from arbitrary speculative bubbles gives the forward-looking solution for the monetary exchange rate[4] (FLME),

$$e_t = (1+\alpha_3)^{-1}\sum_{j=0}^{\infty}[\alpha_3/(1+\alpha_3)]^j E(X_{t+j} \mid I_t) \tag{6}$$

As in Campbell and Shiller (1987) and Macdonald and Taylor (1993) the exchange rate should be cointegrated with the forcing variables X_t. This is illustrated by subtracting X_t from both sides of (6) to obtain,

$$e_t - X_t = -\frac{\alpha_3}{1+\alpha_3}X_t + \frac{\alpha_3}{(1+\alpha_3)^2}E(X_{t+1}|I_t) + \frac{\alpha_3^2}{(1+\alpha_3)^3}E(X_{t+2}|I_t) + \ldots\ldots.. +$$

Rearranging into first differences yields,

$$e_t - X_t = \frac{\alpha_3}{1+\alpha_3}\left(E(X_{t+1}|I_t) - X_t\right) - \frac{\alpha_3^2}{(1+\alpha_3)^2}E(X_{t+1}|I_t) + \frac{\alpha_3^2}{(1+\alpha_3)^3}E(X_{t+2}|I_t) + \ldots\ldots.. +$$

and,

$$e_t - X_t = \left(\frac{\alpha_3}{1+\alpha_3}\right)\Delta E(X_{t+1}|I_t) + \left(\frac{\alpha_3}{1+\alpha_3}\right)^2 \Delta E(X_{t+2}|I_t) - \frac{\alpha_3}{(1+\alpha_3)^3}E(X_{t+2}|I_t) + \ldots\ldots.. +$$

Which for all $j \to \infty$ gives,

$$e_t - X_t = \sum_{j=1}^{\infty}[\alpha_3/(1+\alpha_3)]^j E(\Delta X_{t+j} \mid I_t) \tag{7}$$

Under rational expectations the forecasting errors are stationary, thus if the forcing variables in X_t are I(1), then the right hand side of (7) must also be stationary. Consequently if e_t is also I(1), then the exchange rate must be cointegrated with the variables $m_t, m_t^*, y_t, y_t^*, s_t$ and s_t^*. Thus a test for the FLME is to test for cointegration between the exchange rate and forcing variables[5] :

$$e_t = \beta_0 + \beta_1 m_t + \beta_2 m_t^* + \beta_3 yt + \beta_4 y_t^* + \beta_5 s_t + \beta_6 s_t^* + u_t \tag{8}$$

Where u_t is a random error term and,

$\beta_1 = -\beta_2$, $\beta_3 = -\beta_4$, $\beta_5 = -\beta_6$

$\beta_1, \beta_4 > 0, \beta_2, \beta_3 < 0$, $\beta_5, \beta_6 \lessgtr 0$

The sign on the stock market differential depends on the relative strengths of the income and substitution effect, although as with Friedman (1988), the wealth effect is assumed to dominate, producing a negative relationship. Bahmani-Oskooee and Sohrabian (1992), provide a further explanation of why exchange rates and domestic stock prices are negatively related. They suggest that an exogenous increase in domestic stock prices should result in a rise in domestic wealth. According to the portfolio approach, the rise in wealth ought to facilitate an increase in the demand for money and a rise in the interest rate. Higher interest rates should encourage a capital inflow, increased demand for the domestic currency, which results in an appreciation of the domestic currency. To represent dynamic market adjustments, we can rewrite the equilibrium model of (8) as an error correction model (ECM) to give;

$$\Delta e_t = b_0 + b_1 \Delta m_t + b_2 \Delta m_t^* + b_3 \Delta y_t + b_4 \Delta y_t^* + b_5 \Delta s_t + b_6 \Delta s_t^* \\ - \lambda [\mathrm{e_t} - \beta_1 m_t - \beta_2 m_t^* - \beta_3 y_t - \beta_4 y_t^* - \beta_5 s_t - \beta_6 s_t^*]_{t-1} + v_t \tag{9}$$

Where all terms must be stationary, that is integrated of order zero, denoted I(0), v_t is a random error term with a zero mean. Δ is the first difference operator and the speed of adjustment is given by λ. For values of λ close to unity, adjustment is very rapid, with the disequilibrium being totally eliminated within one period of time. For $0 < \lambda < 1$ the dynamic adjustment path will be monotonically convergent.

If there is evidence that the foreign and domestic coefficients satisfy the implicit restrictions of the monetary model, then the following restricted model is subsequently estimated:

$$e_t = \lambda_0 + \lambda_1 (m - m^*)_t + \lambda_2 (y - y^*)_t + \lambda_3 (s - s^*)_t + u_t \tag{10}$$

Where: $\lambda_1 > 0, \lambda_2 < 0, \lambda_3 \gtrless 0$

To represent dynamic market adjustments, we can again write the equilibrium model of (10) as an error correction model (ECM) to give;

$$\Delta e_t = a_0 + a_1 \Delta (m - m^*)_t + a_2 \Delta (y - y^*)_t + a_3 \Delta (s - s^*)_t \\ - \gamma [(e_t - \lambda_1 (m - m^*) - \lambda_2 (y - y^*) - \lambda_3 (s - s^*)]_{t-1} + u_t \tag{11}$$

3. Empirical Results

We initially estimate the equilibrium unrestricted model (8) and the dynamic unrestricted model (9) for the Canadian dollar against the US[6] dollar. The estimation is over the period January 1977 to December 1999, using monthly data extracted from *International Financial Statistics*, and the country's national accounts.. The income measure, as in other similar studies (Choudhry and Lawler, 1997) is real industrial production, the money supply is represented by M1 and the stock market is represented by the main market[7] index. The start of the sample period was chosen so as to avoid the period covered by the Bretton-Woods system of fixed exchange rates and the subsequent removal of capital controls in the USA then Canada.

All the variables were first tested for stationarity using the Augmented Dickey-Fuller (ADF) and Phillips-Perron tests. The results in Table 1 show that taking both tests into account all the variables tested are non-stationary. The number of lags in the ADF statistic were determined by the Akaike criteria. This requires all the variables in the ECMs to be first differenced and unless valid cointegrating vectors can be found the model is to be rejected, since the residuals from any regression of the exchange rate on the output, money supply and stock price variables will be non-stationary.

The existence of long-run cointegrating vectors was tested for using Johansen's Maximum Likelihood Procedure (Johansen 1988; Johansen and Juselius 1990). The Johansen cointegration test is sensitive to the choice of lag length. To determine the most appropriate lag length, the Akaike criteria was used and in addition the residuals in the Johansen VAR were checked for misspecification. In the event of evidence of serial correlation extra lags were added until this was removed. According to Gonzalo (1994), the costs of over-parameterisation in terms of efficiency loss is marginal, but this is not the case in the event of under-parameterisation. When testing for cointegration, the question of whether a trend should be included in the long-run relationship arises. As with Hendry and Doornik (1994), the trend is restricted to the cointegrating space, to take account of long-run exogenous growth, not already included in the model.

The results for the cointegration test on the unrestricted model are contained in Table 2. The VAR included a lag length of 6, based on the methods mentioned earlier. The maximum eigenvalue test statistic reveals one significant cointegrating relationship, whereas the trace statistic suggests there are two cointegrating vectors. This indicates the presence of one cointegrating relationship based on the evidence of the stronger maximum eigenvalue test (Johansen and Juselius, 1990).

The normalised equation is reported in Table 3. All the variables are significant, except both the money supplies and US stock price variable. However the US money supply and both income variables are different to what we might expect but insgnificant. The restrictions implicit in the monetary model are presented in Table 4., and both individually and jointly indicate acceptance at the 5% level of significance. This suggests that the model can be investigated in its restricted form. The signs on the stock price variables supports the view of Friedman (1988), that the wealth effect dominates.

The results of the test for cointegration on the restricted model are also included in Table 2. Both the maximum eigenvalue and trace results provide evidence of a single cointegrating vector. The normalised equation is in Table 3. and again the coefficients are largely incorrectly signed. Only the stock price differential variable is significant, appearing to dominate the other two variables. Smith (1992) observes a similar result, although using a different model and methodology, as the influence of the stock prices completely dominates all other effects, particularly the effects of money and income.

The error correction models are included in Table 5. for the unrestricted model. As the main focus of the tests is on the exchange rate and stock prices these results alone are reported. The residuals from the cointegrating vector, lagged once, act as the error correction term. This term captures the disequilibrium adjustment of each variable towards its long-run value. The coefficient on the error correction terms in each individual equation represents the speed of adjustment of this variable back to its long-run value. A significant error correction term implies long-run causality from the explanatory variables to the dependent variables (Granger, 1988)[8]. In Table 5 the first statistic represents the sum of the coefficients on the lagged differences of the variables. The second statistic is a chi-square statistic indicating the significance levels of the sum of the coefficients. This can be interpreted as capturing the short-run dynamics in the model and indicates short-run causality between the variables.

In the exchange rate and stock price equations the error correction terms are insignificant, except for the US stock price equation. However for the exchange rate equation there is evidence of short-run causality from the Canadian and US stock market to the exchange rate, as well as short-run causality from Canadian income to the exchange rate. For both stock market equations there is less evidence of short-run causality, particularly running from the exchange rate to stock prices. This indicates causality predominantly runs from stock prices to exchange rates. A possible explanation for this is that there are more market participants in international stock markets than foreign exchange markets, so the former react more quickly to any new information. In the Canadian stock price equation, causality appears to run from output to stock prices. However there is no evidence of output affecting US stock prices, this may be because the US stock market is more dependant on international factors as a result of greater international participation in it.

The error correction results for the restricted model are included in Table 6. Once again the error correction term is only significant for the stock price equation. As with the unrestricted model, there is some evidence of short-run causality from stock prices to the exchange rate, but no evidence of causality in the other direction. The main feature of the stock price equation is the strong causality to the stock price differential from previous differentials. Both equations are well specified, although the explanatory power is low.

A further means of examining the speed with which the markets contained in this version of the monetary model return to their long-run equilibrium is to plot the persistence profiles following a system wide shock (Pesaran and Shin, 1996). As suggested by Pesaran and Shin (1996), the effects of a system wide shock on the cointegrating vector can be more informative than analysing variable specific shocks. This is due to the inherent ambiguities of impulse response analysis with regard to variable specific shocks in a cointegrating vector and because persistence profiles provide information about speeds of adjustment for the system as a whole, although the shock may have a lasting impact on the individual variables.

The persistence profile has a value of unity on impact, then tends to zero as the length of the time horizon increases, if the cointegrating vector is valid. Figure 1 contains the persistence profiles for both the restricted and unrestricted models. The unrestricted model appears to converge back to its equilibrium state much more quickly than the restricted model, with most of the adjustment occurring within a month. The restricted model on the other hand converges much more slowly, even appearing to overshoot to begin with.

A further test of the monetary model, is how well it forecasts out of sample. The exchange rate equation was estimated from January 1977 to December 1998 and 1999 was used for forecasting. As with other studies, the forecasting performance is compared to a random walk. In addition both the restricted and unrestricted models are compared to the forecasting performance of the Frankel Real Interest Differential[9] model. The root-mean-square error (RMSE) statistics from all four models are compared in Table 7. Ironically the worst performer is the unrestricted model, whilst the best is the restricted model. The Frankel model fails to beat the random walk over short time horizons, but over longer time horizons is the second best forecaster of the exchange rate. In addition the significance of each of the measures of forecast accuracy is tested using the Diebold-Mariano (1995) procedure, in which the squared forecast error differential (model forecast minus the benchmark random walk forecast) is regressed on a constant. Only the restricted model and Frankel model produce forecasts that are significantly different to the benchmark random walk model.

4. Conclusions

This paper has examined the relationship between the stock market and exchange rate applying the monetary model of exchange rate determination. The results indicate that in equilibrium , this version of the monetary model produces a cointegrating vector, in which stock prices are the most significant determinant. The dynamic results produce well specified error correction models, in which in the short-run stock prices are the most significant determinant of the exchange rate. However there is very little evidence that exchange rates have a significant effect on stock prices.

These results support those of other studies which indicate that in the short run equities are an important determinant of the exchange rate. These findings not only add to the increasing empirical evidence that foreign exchange markets and stock markets are closely related, but also suggests that in general, models of the equilibrium exchange rate must be extended to include equity markets in addition to bond markets. As with the portfolio balance model, the exclusion of equities from asset holders portfolios imposes excessively strong restrictions on the monetary model.

As with other studies of the Canadian-United States dollar exchange rate, the restrictions implicit in the monetary model of the exchange rate appear to hold over the post 1973 float as well as the 1950's float. This finding is supported by the forecasting performance of the models, in which the restricted model outperforms all the alternatives over short and long time horizons. These results add to other recent studies which portray the monetary models generally in a more favourable light, although more research on the monetary class of exchange rate models is still required.

End Notes

[1]Chrystal & Macdonald (1995) find evidence of a valid long-run relationship using divisia money. Choudhry & Lawler (1997) find evidence of a long-run relationship for the restricted monetary model using Canadian/US data for the 1950's float.

[2] Gavin (1989) provides a nice theoretical version of the sticky price monetary model of exchange rates in which stock prices have wealth effects on the demand for money and exchange rate.

[3] This contrasts with Friedman's (1956) paper that relates money demand to the rate of return on equities.

[4] An advantage of using the FLME, is that it produces a model in which stock prices are the explanatory variables along with income and money. If the conventional monetary model, with static expectations or Frankel real interest rate model had been used, both long and short interest rates would have been incorporated into the model, which could have produced problems of collinearity between the interest rates and stock price returns in the ECMs. In general the conventional FLME (without stock prices) has not been widely used as it generally fails to produce evidence of a valid long-run equilibrium relationship and is not a good predictor of the exchange rate.

[5] Testing for cointegration between the exchange rate and forcing variables is also a test for the presence of bubbles in the exchange rate. If cointegration is found and certain restrictions proved to hold, then the speculative bubble hypothesis is rejected. However this line of investigation is beyond the scope of this paper. Assuming UIP means the interest rate differential equals the expected rate of depreciation. In the absence of arbitrary bubbles, the rate of expected epreciation is some function of expected movements in fundamentals and so equation (8) must be true.

[6] Canada and the USA were used as both countries have financial systems based around financial markets, rather than the banking sector as in Germany or France. The UK was not used as in 1982 it changed the way in which it's main monetary aggregates were calculated.

[7] Stock market indexes are as follows: US; Standard and poor Composite index; Canada; Toronto stock market composite index.

[8] Given that the Johansen maximum Likelihood procedure is essentially a vector autoregression (VAR) based technique, it is more appropriate to produce the complete ECM rather than a parsimonious specification , in which the non-significant lags are omitted.

[9] The results for the Frankel real interest model are not included here, as this model has been tested on Canada and the USA over the 1950's float and the recent float in a number of other studies (Mcnown and Wallace, 1989, Choudhry and Lawler, 1997). The unrestricted Frankel real interest model did provide evidence of cointegration, however the restrictions on the domestic and foreign explanatory variables were rejected, so the restricted version of this model was not estimated.

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Table 1- The Augmented Dickey-Fuller (ADF) and Phillips-Perron Test for Unit roots

	ADF Test		Phillips-Perron Test	
Variables	Test for I(0)	Test for I(1)	Test for I(0)	Test for I(1)
E	-2.586	-3.007	-2.590	-28.894
CM1	0.688	-4.211	1.000	-25.294
UM1	-1.663	-2.213	-1.997	-24.645
CY	-2.485	-2.916	-2.656	-12.527
UY	-2.870	-2.767	-1.931	-5.640
CS	-0.824	-15.110	-0.942	-17.471
US	1.502	-15.272	2.089	-19.717
DM1	-0.470	-2.686	0.051	-20.287

DY	-2.944	-3.704	-1.922	-28.361
DS	0.191	-7.987	0.464	-18.292

Notes:. E is the exchange rate, CM1 and UM1 are Canadian and US M1 respectively, CY and UY are Canadian and US real income respectively, CS and US are Canadian and US real stock prices respectively, DM1, DY and DS are the differential between Canadian and US M1, real income and real stock prices respectively. For each variable the first column of statistics tests the null hypothesis that the series is I(1) against the alternative that it is I(0). The second column tests the null that the series is I(2) against the alternative that it is I(1). The critical values for both these tests at the 10% and 5% levels of significance are -2.56 and -2.89 respectively. The Phillips Perron test uses 40 Bartlett lags in each test. Using the same tests with a trend included does not materially change the results.

Table 2- Johansen Maximum Likelihood Test for Cointegration of the Unrestricted and Restricted models.

	Unrestricted Model		Restricted Model	
Vectors	Trace Test	Eigenvalue Test	Trace Test	Eigenvalue Test
$r = 0$	177.92*	50.52*	78.00*	46.01*
$r \leq= 1$	127.41*	43.31	31.98	16.92
$r \leq= 2$	84.09	30.47	15.07	9.75
$r \leq= 3$	53.63	22.51	5.32	5.32
$r \leq= 4$	31.11	16.36		
$r \leq= 5$	14.75	9.80		

$r \leq= 6$	4.96	4.96

Notes: Critical values of Johansen's Trace and Eigenvalue tests at the 95% level of significance are: $r = 0$; 147.27 and 49.32. $r \leq= 1$, 115.85 and 43.61. $r \leq= 2$, 87.17 and 37.86. $r \leq= 3$, 63.00 and 31.79. $r \leq= 4$, 42.34 and 25.42. $r \leq= 5$, 25.77 and 19.22. $r \leq= 6$, 12.39 and 12.39 respectively. A * indicates significance at the 5% level. For the Restricted Model: $r = 0$, 63.00 and 31.79. $r \leq= 1$, 42.34 and 25.42. $r \leq= 2$, 25.77 and 19.22. $r \leq= 3$, 12.39 and 12.39. Both tests included seasonal dummy variables.

Table 3- Normalised Equations of the cointegrating vectors.

	Unrestricted Model		Restricted model		
Variable	Coefficient	Significance	Variable	Coefficient	Significance
E	-1.000	0.651	CE	-1.000	0.237
CM1	1.318	0.513	DM	-1.015	1.117
UM1	0.139	0.024	DY	0.858	0.036
CY	4.394	4.724*	DS	-3.138	11.129*
UY	-6.360	5.904*			
CS	-1.942	5.963*			
US	1.594	1.866			

Notes: The significance of the coefficients were tested using the LM statistic which tests the restriction that the coefficient is equal to zero.(($\chi^2_{0.5}(1) = 3.841$) . A * indicates significance at the 5% level.

Table 4- Restriction Tests on the coefficients of the following variables

Null Hypothesis	Chi-square statistic
H1: CM1=1,UM1=-1	0.372
H3: CY=-UY	1.412
H4: CS=-US	0.144
H5: CM1=-UM1; CY=-UY; CS=-US	4.312

Notes: Critical Values are 3.84 and 7.815 (5%)

Table 5- Error Correction Model Results for the Unrestricted Model

	ΔE	ΔCS	ΔUS
Constant	0.017 [0.305]	-0.126 [0.607]	0.481 [2.529]*
res_{t-1}	-0.004 [0.328]	0.035 [0.736]	-0.107 [2.452]*
$\sum \Delta E$	0.096 (0.619)	-0.090 (1.900)	-0.031 (0.938)
$\sum \Delta CM1$	0.084 (0.343)	1.022 (2.774)	1.581 (8.594)*
$\sum \Delta UM1$	0.187 (0.645)	-0.478 (0.030)	1.504 (0.191)
$\sum \Delta CY$	-0.318 (3.994)*	1.161 (4.283)*	-0.001 (0.073)
$\sum \Delta UY$	0.324 (1.274)	-1.606 (4.839)*	-1.082 (1.236)
$\sum \Delta CS$	0.147 (8.931)*	-0.068 (0.745)	-0.376 (3.840)**

$\sum \Delta US$	-0.103 (4.924)*	0.147 (0.131)	0.047 (0.026)
R^2	0.187	0.206	0.213
SC(12)	1.658	2.022	0.827
SC(6)	1.417	1.019	1.021
Reset	0.077	0.232	1.573
Heteroskedasticity	0.522	0.204	0.122
ARCH(12)	0.482	0.155	0.989

Notes: res denotes the error correction term; R^2 is the coefficient of determination; DW is the Durbin-Watson statistic; SC(i) are the ith order tests for serial correlation; ARCH(i) is Engle's (1982) test for the i'th autoregressive conditional heteroskedasticity. These test statistics all follow the F-distribution, critical values are: F(6,222)=2.14, F(12,216)=1.80, F(1,227)=3.89. The values in square brackets represent t-statistics for the constant and ect. The values in ordinary brackets represent Wald statistics, which follow a chi-square distribution, critical value 3.842. All equations include seasonal dummies. A * indicates significance at the 5% level, ** 10% level.

Table 6- Error Correction Model for the Restricted Model

	ΔE	ΔDS
Constant	0.003 (0.705)	0.034 (4.112)*
res_{t-1}	-0.001 (0.678)	0.147 (4.921)*
$\sum \Delta E$	-0.075 (0.418)	-0.691 (1.208)
$\sum \Delta DM$	0.073 (0.545)	0.035 (1.311)
$\sum \Delta DY$	0.061 (0.064)	1.266 (0.253)
$\sum \Delta DS$	0.101 (3.733)**	-0.129 (13.055)*
R^2	0.08	0.189
SC(12)	1.592	0.746

SC(6)	0.320	0.524
Reset	1.510	3.913
Heteroskedasticity	1.025	0.007
ARCH(12)	0.795	0.920

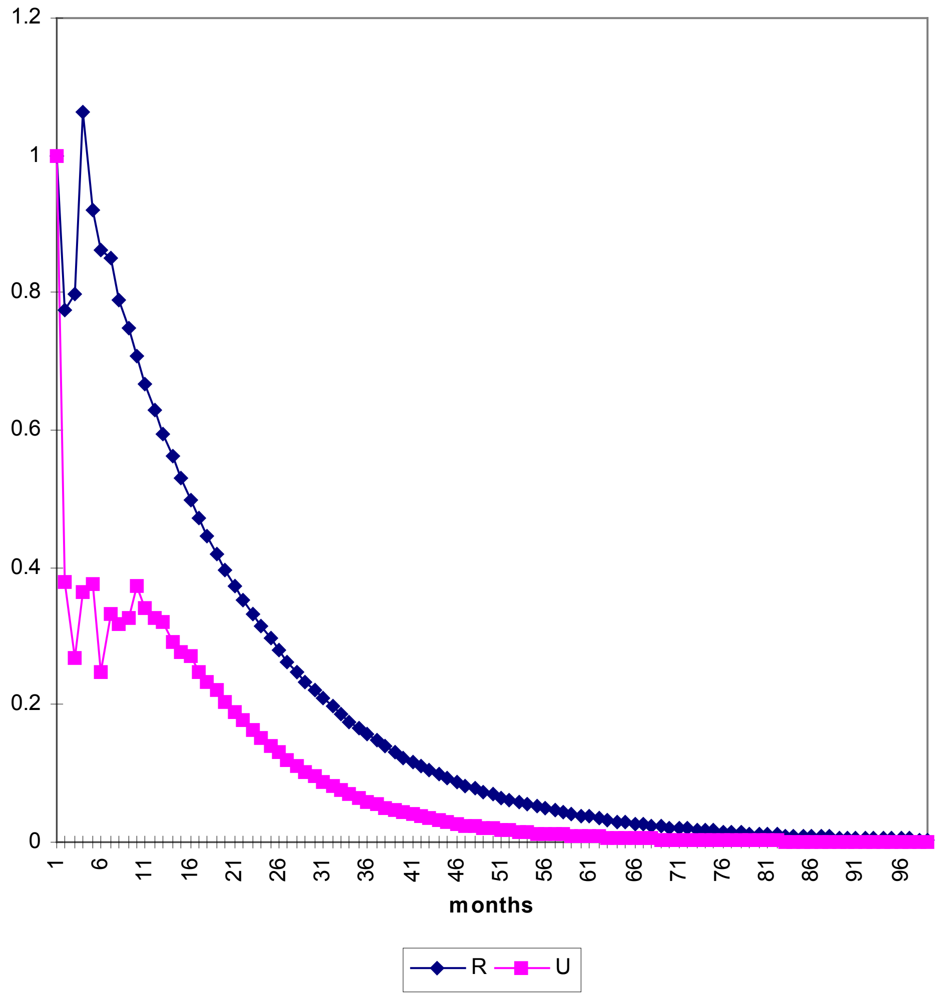
Notes: See Table 4

Table 7- RMSE Statistics for Forecasts using the Competing models

Models	3 Months	6 Months	9 Months	12 Months
Random Walk	0.010	0.017	0.017	0.016
Unrestricted Model	0.013	0.017	0.018	0.017
Restricted Model	0.009	0.016*	0.016*	0.015*
Frankel Model	0.011	0.016*	0.016*	0.015*

Notes: A * indicates a significant Diebold-Mariano test statistic at the 5% level. The test uses the standard Newey-West adjustment, with Bartlett weights and a lag window of 2.

Figure 1- Persistence Profiles of the Effect of a System Wide Shock on the Cointegrating Vector.



Notes: R is the persistence profile for the restricted model and U is the persistence profile for the unrestricted model.