# Uncovered Interest Rate Parity: The puzzle of the US dollar

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**Abstract**

*This paper shows that although the Uncovered Interest Rate Parity (UIP) condition fails to hold in general, it fails mostly in regressions involving the US dollar and holds generally well in regressions that do not involve the US dollar. We examine the US dollar-German Mark (USD-DM), the US dollar-Pound (USD-GBP), the US dollar-Yen (USD-YEN), the US dollar – Canadian dollar (USD-CAD), the US dollar- New Zealand dollar (USD-NZD) and the US dollar – Australian dollar (USD-AUD) and the cross rates NZD-AUD, CAD-NZD, CAD-AUD, GBP-CAD, GBP-NZD and GBP-AUD. We use both Fully Modified OLS and Two-Sided Dynamic Least Squares of Phillips-Hansen (1991) and Phillips-Loretan (1991) efficient estimators to test the UIP from 1989 to 2000. The results provide further evidence on the UIP anomaly involving U.S. dollar exchange rates, namely, interest rate differentials fail to predict the direction of exchange rate changes. We use evidence from survey expectations to shed more light to the empirical UIP failure …..…*

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## 1. Introduction

Uncovered interest rate parity is an important building element of many models in international economics, yet it finds very little empirical support in the literature. In fact, empirical studies often suggest that UIP fails dismally to predict not only the magnitude but also the direction of exchange rate changes.[2] Obstfeld and Rogoff (1996, p. 623) show a close graphical relationship between the trade-weighted dollar exchange rate and the real interest rate differential – uncovered interest rate parity, UIP – and ask why the visual impression does not come through in empirical analysis.

In their answer, Obstfeld and Rogoff (1996) appeal to Baxter's (1994) argument that in order to find a link between the exchange rate and the interest rate differentials, one should filter the data properly. At very high frequencies (i.e., cycles 2-5 quarters) one should expect to find nothing but noise related to the irregular component of the time series. Using a Band-Pass filter Baxter and King (1995) find a correlation between exchange rates and interest rate differentials at both trend and business cycles.

Empirically, there seems to be no consensus on the appropriate interest rate variable--short versus long-term rates--to be used in exchange rate determination models. For example, Frankel (1979) uses long-term interest rates as a proxy for anticipated inflation, Kim and Mo (1995) use both short and long term rates, Bartolini and Bodnar (1995), McNown and Wallace (1994, 1989), Johansen and Juselius (1992) and Baillie and Pecchenino (1991) use short-term interest rates. The argument against the use of short-term interest rates is that central banks intervene in the money market to smooth out short-term interest rate movements affecting thereby their information content. On these grounds, empirical analysis of exchange rates using short-term interest rates requires an additional equation representing the behavior of the central bank (McCallum, 1994). On the other hand, long-term interest rates are – we believe – the average of anticipated future short-term interest rates, which reflects anticipated returns to capital over the business cycle, risk and anticipated inflation (Brunner and Meltzer, 1989).[3] In this paper we use the 90-day interest rate. We experimented with different 90-day interest rates such as the T-bills rate, bankers acceptance rates etc but the results do not change significantly.

This paper shows that the UIP fails, but it seems to fail only when the US dollar is used and holds relatively well when the US dollar is not used. For example, the UIP does not hold in the USD-DM, USD-GBP, USD-YEN, USD-CAD, but it holds well in the

[2] Examples are Frankel (1979), McNown and Wallace (1989), Baillie and Pecchenino (1991), Johansen and Juselius (1992), Edison and Pauls (1993), McNown and Wallace (1994), McCallum (1994), Kim and Mo (1995), Bartolini and Bodnar (1995), Meredith and Chinn (1998) and Edison and Melick (1999). The theoretical debate about how interest rates affect the exchange rates is found in Dornbusch (1976), Frankel (1979), Mussa (1976, 1979), Frenkel (1976), Bilson (1978, 1979) and Christiano et al (1998). See Messe and Rogoff (1983) and Messe and Rogoff (1988) and MacDonald and Nagayasu (1999) for empirical research on *real* exchange rates and *real* interest rate differentials.

[3] Recently, Meredith and Chinn (1998) used regressions in *differences* and showed evidence in favor of UIP(short term) when long-term interest rates are used instead of short-term rates.

GBP-CAD, and other cross rats not involving the USD such the NZD-AUD. The paper uses an asymptotically efficient and unbiased estimator (Phillips and Loretan, 1991) to estimate the coefficients. We use the Consensus Forecasts and survey data to measure the expected depreciation rates. It is assumed that this kind of treatment may overcome the "joint test" problem.[4] The findings of deviations from UIP are then explored by conducting econometric tests for the presence of risk premia in the foreign exchange market and the rationality of the exchange rate forecasts.

Section 2 describes the UIP regression that is used in this paper. Section 3 discusses the estimation method. The time series properties of the data are discussed in Section 4. Section 5 presents estimates of the UIP the regressions and interprets the results. Section 6 contains an analysis of the properties of exchange rate forecasts in relation to risk premia and rational expectations.

## 2. The UIP

We briefly discuss the uncovered interest rate parity condition. Using quarterly data, annualized 90-dayinterest rates and one quarter ahead expected exchange rate, UIP can be expressed by:

$$i_{t?s}/4 ? i^{*}_{t?s}/4 ? (E_t e_{t?s} ? e_t), \qquad (1)$$

where $i_t$ is time $t$ U.S. *annualized* nominal interest rate or yield to maturity on a pure discount bond that matures at time $s$. Similarly $i^{*}_t$ is the foreign magnitude, $e_t$ is the natural logarithm of the spot exchange rate, $E_t$ is the mathematical expectations operator conditional on the information set at time $t$. The term in parentheses is the expected depreciation rate. To test the UIP, this regression

$$i_{t?s}/4 ? ? ? [i^{*}_{t?s}/4 ? (E_t e_{t?s} ? e_t)] ? v_{t?s} \qquad (2)$$

is estimated and the hypotheses $? ? 0$, $? ? 1$ are tested, and the residuals $v_{t?s}$ is checked for whiteness. The expected depreciation rate $E_t e_{t?s} ? e_t$ assumes that Consensus forecasts or survey data are reasonable proxies for $E_t e_{t?s}$.

The order of integration of the variables in (1) and (2) matters for the method of estimation. In equation (2) if the domestic interest rates $i_t$ has a unit root the LHS is an I (1) process.

If the foreign interest rate $i^{*}_t$ is also an I(1) variable and the expected depreciation rate $(i^{*}_t ? ? E_t e_t)$ is an I(0) then the sum of the two variables is an I (1). Given such order of integration the estimation method of equation (2) must account for unit roots. If the $i_t$ and

[4] Assuming rational expectations, it is not clear whether the rejection of the UIP condition in OLS regressions implies the rejection of the rational expectations hypothesis or the UIP regression.

$(i_t^* ? ? E_t e_t)$ are cointegrated then an error correction model is appropriate to test for the UIP. If they are not cointegrated then a regression model in differences is more appropriate.

There are different ways to empirically estimate UIP and none of them escapes criticisms (see Edison, Melick, 1999).

## 3. The Phillips-Loretan (1991) method

We briefly explain the methodology used in this paper to estimate the coefficients that govern the long-run relationship between the exchange rate and the interest rate differentials. This method is based on the Triangular Representation Theorem.

Consider the following triangular system of equations:

$$y_{1t} ? ? ? ? ? y_{2t} ? u_{1t} \quad (3)$$

$$? y_{2t} ? u_{2t}, \quad (4)$$

where $u_t ? ? u_{1t} \; u_{2t} ? ?$ is a stationary vector.

Phillips and Loretan (1991) consider estimation and inference in the above system. Equation (3) can be estimated using single equation methods, and provided that the equation is appropriately augmented the asymptotic properties of the estimator can be readily determined. The asymptotic (distributional) properties hinge on the interrelationships that exist between $u_{1t}$ and $u_{2t}$, which are assumed to be stationary.

If the variance-covariance matrix of $u_t$ is block diagonal (so that the partitioned elements of the vector do not co-vary) and $u_t$ is also (*i.i.d.)* then equation (3) can be estimated using least squares. The estimates will be normally distributed, at least asymptotically, and will be equivalent to maximum likelihood estimates of the parameters of the system. If the variance covariance matrix of $u_t$ is block diagonal but the residuals are autocorrelated (but also block diagonal) the regression can be augmented with lagged error correction terms, counteracting the autocorrelation of $u_{1t}$. This solution works because the regressors – $y_{2t}$ – are super-exogenous. They are independent of the past and future history of $u_{1t}$, i.e. they are orthogonal to $? u_{1t} ? ???$ .

If $u_{1t}$ and $u_{2t}$ are *correlated* one can alleviate the problems that this causes by augmenting the regression with leads and lags of $? y_{2t}$. One can think of this as projecting $u_{1t}$ against leads and lags of $u_{2t}$ and then applying the Frisch –Waugh - Lovell theorem (see Davidson and MacKinnon, 1993). The error from the projection is uncorrelated with $u_{2t}$, and the simultaneity problem is thus dealt with. The inclusion of leads eliminates feedback from

$u_{1t}$ back to $u_{2t}$ and it is important for valid conditioning. The basic problem is that, *a priori,* we do not know whether $u_{1t}$ correlates with leads of $u_{2t}$ or vice versa.

The inclusion of lag differences eliminates the simultaneity problem, while the autocorrelation problem may need to be remedied by including lags of the error correction term too. Phillips and Loretan then estimate the following equation using least squares:

$$y_{1t} = \alpha + \beta y_{2t} + \sum_{k=1}^{K} d_{1k} \Delta y_{1t-k} + \alpha \beta y_{2t-k} + \sum_{i=p}^{p} d_{2i} \Delta y_{2t-i} + \varepsilon_{1t} \qquad (5)$$

Phillips and Loretan demonstrate that the parameter estimates from estimating this single equation are equivalent to maximum likelihood estimation of the system and hence are efficient. Furthermore, the parameters are asymptotically normally distributed.

The distributions of the parameters of the cointegrating vector – and corresponding statistics – depend on the null hypotheses concerning the interrelationships. If the series are integrated but not cointegrated the regression is spurious, and the asymptotic theory is found in Phillips (1986).

Barnhart, McNown and Wallace (1999) used this technique to test the forward rate unbiasedness hypothesis and demonstrated that the estimator is most unbiased among many different methods used in testing. The assumptions of unit roots and cointegration need to be tested first, which is the next step.

## 4. The Data

Quarterly data covering the period March 1989 to March 2000 are used in this paper. The exchange rate is defined as the foreign currency price of the US dollar. The exchange rates are the USD-DM, USD-GBP, USD-YEN, and the USD-CAD period averages. We also use the USD-NZD and the USD-AUD rates. These exchange rates are used because New Zealand has followed a clean float since it floated its currency in 1985, and Australia's foreign exchange rate intervention is very small. We also use these two currencies to show that the two capital markets are actually fully integrated and contrast that with the USD-CAD to show that while the US and Canadian capital markets are integrated, the UIP condition does not pass the test.

For interest rates we use short-term, 90-day interest rates. The reason is that we want to show that although the UIP fails to hold when the 90-day rates are used, it does hold in cases that do not involve the US dollar. Thus the explanations put forward regarding the inadequacy of the data is not as solid as we thought.[5] Before estimating the UIP condition, we briefly discuss the time series properties of the data because this is quite crucial for the estimation of the regressions.

[5] The source of the data is the Reserve Bank of New Zealand. We used different 90-day interest rates such the T-bills, bankers acceptance rate for the US, Interbank rates for the UK, Japan and Australia, and the Canadian 90-day deposits rate. For New Zealand the average bank bill rate is used.

There seems to be an agreement among economists that exchange rates, during the post Bretton-Woods era, may contain stochastic trends. The literature on unit roots and cointegration is vast and it will not be reviewed here. Suffice to say that there is a valid concern among economists about the appropriateness of the tests for unit roots and their power against stationary alternatives. The choice of a particular testing methodology is not straightforward. Ultimately, one may not be able to determine whether a particular time series contains a unit root or not. It seems inevitable, however, that one must make a choice.

Because there is no consensus on the issues of unit root in long-term interest rate differentials and cointegration between the exchange rate and the interest rate differentials, it is perhaps useful to subject the data to various different tests. This strategy reduces the risk of being on the wrong side. A decision can then be based on whether the results of various tests converge or not. For example, when different tests for unit roots (presumably all of them have the same null hypothesis that the variable has a unit root) move in one direction, e.g., indicating a unit root, one can be a little more confident in the results than when the tests diverge.

In this paper, four methods to test for unit roots are used. All of them have the same null hypothesis of a unit root and are available in most statistical packages. These are the DF test (Dickey and Fuller, 1979, 1981), the ADF test (Said and Dickey, 1984), the $Z$ test (Phillips, 1987 and Perron, 1988), and the DF-GLSu test (Elliott, 1999).[6]

We report some of the results of the unit root tests in table 1. The tests that are considered in this paper seem to agree with conventional wisdom that the log of the exchange rates $e_t$ and the log of the one quarter ahead forecast $E_t e_{t+1}$ have unit roots. The difference $E_t e_{t+1} - e_t$ is I(0). We do not report these results because they consistent with the general findings in the literature. However, both the domestic (The US in this case) and the foreign 90-day interest rates (Germany, UK, Japan, Canada, New Zealand and Australia) are I (1) and the sum $i_t^* - E_t e_{t+1}$ is an I(1). We report these results.

We do not report the results of the tests for the cross rates such as GBP-CAD or GBP-NZD etc because the null hypothesis of unit roots could not be rejected either.

Testing the null hypothesis of no cointegration between $i_t$ and $i_t^* - E_t e_{t+1}$ is not straightforward. The main reason is that our sample is rather a short span, which is not a good representative of the long-run relationship underlying the tests for no cointegration. Also, the Engle-Granger ADF test is weak against stationary alternatives. We will assume that these two variables are cointegrated.[7]

[6] Perron (1997) could also be used, but we have not used it in this paper. It allows for breaks in the data. This is important because the ADF test may incorrectly indicate that the data have unit roots when in fact they have breaks.

[7] The empirical evidence is weak using this test.

## 5. Estimation of the Long-run Coefficients

As we explained earlier, this single-equation estimation technique of Phillips and Loretan is asymptotically equivalent to a maximum likelihood on a full system of equations under Gaussian conditions. This technique provides estimators that are statistically efficient, and whose t-ratios can be used for inference in the usual way. Most importantly, this method takes into account both the serial correlation of the errors and the endogeneity of the regressors that are present when there is a *cointegration* relationship. Recently, Barnhart, McNown and Wallace (1999, pp. 288-289) provided evidence that this procedure produces no bias in the estimates of the parameters in forward premium equation. We estimate the regression given in equation (5) using non-linear least squares.[8]

We are mostly interested in the magnitudes $\hat{a}$, $\hat{b}$ because they are the parameter estimates that determine the *long-run* relationship between the interest rate $i_t$ and $i_t^* ? ?E_t e_{t?1}$ conditional on the dynamics.

Results of the nonlinear least square estimation are reported in table 2. We report the t-ratios and the p-values. These results are consistent with the Johansen-Juselius test results both in magnitudes and in quality (not reported). The intercept is not significant except in the case of the USD-DM. The slope coefficient is negative in all the cases involving the USD, except in the case of Canada, New Zealand and Australia. However, the coefficient is significantly different from one in the case USD-CAD and USD-AUD. It statistically equals to one in the case of the USD-NZD. UIP holds perfectly.

The second block of results in table 2 involves cross rates other than the USD. The results are remarkably different. The intercept ? is zero, except in the cases of the GBP-CAD, GBP-NZD and the GBP-AUD. The slope coefficient is positive in all cases, and statistically not different from one in the cases of the AUD-NZD, CAD-NZD and GBP-CAD.

The various diagnostics of the residuals indicate that the residuals are white noise, serially uncorrelated, and homoskedastic.

We conclude that the exchange rates contain unit roots. The 90-day interest rate is an I (1) variable. Tests for cointegration are rather difficult to justify in short spanned data so we assumed that the interest rate at home and $i_t^* ? ?E_t e_{t?1}$ are cointegrated. The UIP condition

[8] The Phillips Information Criteria (Phillips, 1996) is used to determine the optimal lag structure along with the cointegration vector. We find a three lag-lead structure to be sufficient to eliminate the serial correlation. However, because we are conscious about Phillips and Loretan warning of over-fitting, we start reducing the number of leads by one. Every time, we check the serial correlation, the parameter estimates, and their significance. We find that a structure of three lags and two leads gives the same results as a three lag-lead structure without compromising the whiteness of the residuals. A further reduction in the leads introduces some serial correlation.

fails in tests involving the US currency. In the next section we seek possible explanations for excess returns (divergence from UIP).

## 6. Risk Premia and Rational Expectations

Figure 1 presents monthly excess returns, defined as the change in ex-post or expected exchange rates minus the interest rate differential, for exchange rates that involve the U.S. dollar. These plots can effectively be interpreted as the outcome of taking a short position in the currency. Clear patterns are evident in most cases. The expected payoffs are largely positive for the Canadian dollar and negative for the Yen, the Pound and the German Mark, whereas they appear largely negative, ex-post, for the Canadian dollar. This prima-facie evidence on inefficient or irrational market behaviour needs to be contextualised.

We note the following. Monthly returns that involve 3-month forecast horizons are known to be MA(2) processes. This aside, market efficiency does not require that returns on foreign currency positions are uncorrelated nor that are associated with zero expected profits (see Mark, 2001, p. 127). Fama (1991) has argued market efficiency should be assessed in conjunction with an equilibrium asset-pricing model. In other words, market efficiency requires that there are no unexploited excess profit opportunities and what we consider as excessive depends on the model of market equilibrium (see Mark, 2001).

Whether or not market efficiency can be reconciled within the context of a risk premium associated with a theory of asset pricing effectively requires that suitable economic forces can be identified to explain the volatility in excess returns. For example, in the use of the Lucas model to price forward exchange rate as in Hodrick (1987) or Engel (1992), variability in consumption growth should be sufficient to drive the variability in foreign currency excess returns. This seems unlikely to be the case from a visual inspection of the variability in excess returns shown in figure 1. In fact, there is ample empirical evidence to suggest that estimates of the coefficient of relative risk aversion in Lucas type of asset pricing models is uncomfortably large and imprecisely estimated (see Mark, 2001, p. 140).

Yet, Lucas' models provide a theoretical justification for some of the payoff patterns shown in figure 1. In the equilibrium solution of the Lucas model, the intertemporal marginal rate of substitution varies inversely with consumption growth. Risk averse investors are willing to buy an asset with an expected negative payoff to the extent that the negative payoff is regarded as an insurance premium. In other words, you want to buy an asset that pays well in bad (low consumption growth) states – when you really need it (see Mark, 2001).

Alternative explanations of UIP failure may rest on the expectations component of the efficiency hypothesis. Learning models involve rational behaviour in conjunction with systematic forecast errors during the transition period where agents learn about structural or policy changes. They typically show insufficient ability to model satisfactorily deviations from UIP as agents in financial markets are expected to learn quickly about single or infrequent changes in policy or other structural changes. A different explanation for UIP failure is provided through the effect of noise trading, that is, it can be argued that

some investors act irrationally and push asset prices away from their underlying fundamental values.

In Tables 3-6 we use survey expectations data to break down the joint efficiency hypothesis into separate tests for the existence of time varying risk premia and/or irrational expectations. The risk premium test, shown in the left side of Tables 3-6, is a test that the slope coefficient is equal to one in a regression of the expected exchange rate change on the interest rate differential. In other words, what we test is that the interest rate differential (or forward premium under covered interest parity) fully drives the expected change in the exchange rate. The rational expectations test is a test of no statistically significant relationship in a regression of the forecast error (the difference between expected and actual future spot rates) and information available at the current time period (the interest rate differential or forward premium). The results are mixed. The U.S.-Canadian dollar exchange rate passes the no risk premium test but fails the rationality test. The reverse is true for the Yen, whereas the Pound U.S. and the German Mark U.S. dollar exchange rates pass both tests. The form in which rationality is violated suggests that investors place excess weight on the interest differential which may be interpreted as noise trader over-reaction.

We intend to pursue further research to understand why UIP fails in situations involving the US dollar. This study will focus on (1) microstructure explanations either in the form of learning by rational agents or via an analysis of the impact of irrational trader behaviour in the foreign exchange market and (2) an analysis of the full effect of yield curve differentials as opposed to single interest rate differentials on the movement of the exchange rate.

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Table 1
Unit Root Tests

| Variable | Country | ADF | DFGLSu | Lags |
|---|---|---|---|---|
| 90-day interest rate | Germany | -0.89 | -0.38 | 1 |
| | UK | -0.96 | -1.23 | 4 |
| | Japan | -1.92 | -2.01 | 4 |
| | Canada | -0.06 | -0.25 | 1 |
| | New Zealand [A] | 0.56 | 0.19 | 1 |
| | Australia | -0.77 | -1.11 | 4 |
| | USA | -2.70 | -0.10 | 4 |
| $i_t^* ? ? E_t e_{t?1}$ | Germany + USDDM | -1.22 | -0.19 | 4 |
| | UK + USDGBP | -1.36 | -1.18 | 1 |
| | Japan + USDYEN | -1.36 | -1.37 | 4 |
| | Canada + USDCAD | -0.06 | -0.26 | 1 |
| | New Zealand + USDNZD [A] | 0.34 | 0.04 | 1 |
| | Australia + USDAUD | -0.15 | -1.04 | 1 |

? The sample size is June 1989-March 2000, except for Germany which is March 1997.
? Data for Germany exchange rate discontinued when the Euro system started.
? ADF is the Augmented Dickey-Fuller test and DFGLSu is the GLS corrected Elliott (1997) test for unit root.
? The 5% critical value for the ADF test with 25 observations is -2.997.
? The 5% critical value for DFGLSu is –2.73.
? The tests do not include time trends. They were dropped out because they were insignificant.
A. Because $? ? 1$ this series has an explosive root.

Table 2
The Phillips-Loretan Non -Linear Dynamic Least Square Estimator
1989:1 to 2000:1

$$y_{1t} = \alpha + \beta y_{2t} + \sum_{k=1}^{K} d_{1k} y_{1t-k} + \gamma y_{2t-k} + \sum_{i=p}^{p} d_{2i} y_{2t-i} + \varepsilon_{1t}$$

| | $\beta$ | $P$ value [B] | $\gamma$ | $P$ value [C] | DW | BKS [D] |
|---|---|---|---|---|---|---|
| USD-DM [A] | 2.96 | 0.000 | -0.50 | 0.0000 | 1.8 | 0.15 |
| USD-GBP | 4.08 | 0.8730 | -0.25 | 0.0000 | 1.7 | 0.14 |
| USD-YEN | 4.55 | 0.2935 | -1.85 | 0.0000 | 1.8 | 0.18 |
| USD-CAD | 0.74 | 0.3038 | 0.39 | 0.0000 | 1.8 | 0.22 |
| USD-NZD | -1.45 | 0.1840 | 1.00 | 0.9935 | 1.7 | 0.19 |
| USD-AUD | 0.28 | 0.3550 | 0.46 | 0.0000 | 1.7 | 0.20 |
| | | | | | | |
| AUD-NZD | -0.82 | 0.5318 | 1.04 | 0.8948 | 1.8 | 0.08 |
| CAD-AUD | -0.00 | 0.9941 | 0.79 | 0.1794 | 2.1 | 0.11 |
| CAD-NZD | -0.69 | 0.7022 | 1.07 | 0.8944 | 2.1 | 0.11 |
| GBP-CAD | 0.63 | 0.0187 | 0.95 | 0.6799 | 2.1 | 0.18 |
| GBP-NZD | 1.76 | 0.0204 | 0.52 | 0.0315 | 2.1 | 0.21 |
| GBP-AUD | 1.25 | 0.0008 | 0.58 | 0.0001 | 1.9 | 0.13 |

A. The sample is 89:1 to 97:1 for Germany, and 89:1 to 2000:1 for the rest.

B. The P value for the t-statistic to test $\beta = 0$.

C. The P value for the $\chi^2_1$ to test $\gamma = 1$.

D. The Kolmogrov-Smirnov cumulated periodogram test for white noise residuals. The 5% critical value is 0.24.

The regression is estimated using non-linear least squares and Hansen (1982) and White (1980) consistent estimators for the variance-covariance matrix.

Table 3: Risk Premia and Rational Expectations (US/Canadian dollar)

```
Ordinary Least Squares Estimation
*******************************************************************
 Dependent variable is SCAN1
 48 observations used for estimation from 1989Q4 to 2001Q3
*******************************************************************
Regressor  Coefficient  Standard E  T-Ratio     [Prob]
C              3.5943     0.53234     6.7519 [.000]
RDCN           1.3595     0.24158     5.6275 [.000]
*******************************************************************
R-Squared     0.40774  R-Bar-Squared              .39486
S.E. of Re    3.3421  F-stat.   F( 1, 46)  31.6686[.000]
Mean of D     2.3274  S.D. of Dependent Variable    4.2963
Residual S   513.8029  Equation Log-likelihood    -125.0044
Akaike Info  -127.0044  Schwarz Bayesian Criterion  -128.8756
DW-statist    2.2824
*******************************************************************

                    Diagnostic Tests
*******************************************************************
*          Test Statis LM Versio F Version        *
*          *********** ********** ***************************
*          *           *          *
*          A:Serial C  CHSQ(  4)=  3.9845[.408]*      F(  4, 42)=  .95051[.445]*
*          *           *                          *
*          B:Function  CHSQ(  1)=  .086523[.769]*     F(  1, 45)=  .081262[.777]*
*          *           *                          *
*          C:Normalit  CHSQ(  2)=  3.2487[.197]*      Not applicable     *
*          *           *                          *
*          D:Heteros   CHSQ(  1)=  .32561[.568]*      F(  1, 46)=  .31418[.578]*
*******************************************************************
  A:Lagrange multiplier test of residual serial correlation
  B:Ramsey's RESET test using the square of the fitted values
  C:Based on a test of skewness and kurtosis of residuals
  D:Based on the regression of squared residuals on squared fitted values

a2=1
*******************************************************************
 Wald Statistic          CHSQ(1)=  2.2145[.137]
```

```
Ordinary Least Squares Estimation
*******************************************************************
 Dependent variable is SCAN2
 48 observations used for estimation from 1989Q4 to 2001Q3
*******************************************************************
Regressor  Coefficient  Standard E  T-Ratio[Prob]
C              6.6553     1.3451     4.9478 [.000]
RDCN           1.885     0.61044     3.088 [.003]
*******************************************************************
R-Squared     0.1717  R-Bar-Squared              .15369
S.E. of Re    8.4449  F-stat.   F( 1, 46)   9.5355[.003]
Mean of D     4.8987  S.D. of Dependent Variable    9.1797
Residual S    3280.5  Equation Log-likelihood    -169.4984
Akaike Info  -171.4984  Schwarz Bayesian Criterion  -173.3696
DW-statist    1.8581
*******************************************************************

                    Diagnostic Tests
*******************************************************************
*  Test Statistics *     LM Version      *      F Version       *
*******************************************************************
*            *              *              *
*          A:Serial C  *CHSQ(  4)=  6.9077[.141]      F(  4, 42)=  1.7651[.154]*
*                      *                          *                *
*          B:Function  *CHSQ(  1)=  1.0080[.315]      F(  1, 45)=  .96529[.331]*
*                      *                          *                *
*          C:Normali   *CHSQ(  2)=  6.6281[.036]      *    Not applicable     *
*                      *                          *                *
*          D:Heteros   *CHSQ(  1)=  1.4692[.225]      F(  1, 46)=  1.4524[.234]*
*******************************************************************
  A:Lagrange multiplier test of residual serial correlation
  B:Ramsey's RESET test using the square of the fitted values
  C:Based on a test of skewness and kurtosis of residuals
  D:Based on the regression of squared residuals on squared fitted values
```

# Table 4: Risk Premia and Rational Expectations (US dollar/Yen)

Ordinary Least Squares Estimation

Dependent variable is SJPN1
48 observations used for estimation from 1989Q4 to 2001Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| C | -0.59462 | 1.8597 | -0.31975 | [.751] |
| RDJP | -0.88189 | 0.48272 | -1.8269 | [.074] |

| | | | |
|---|---|---|---|
| R-Squared | 0.067649 | R-Bar-Squared | .047380 |
| S.E. of Re | 8.8197 | F-stat. F( 1, 46) | 3.3376[.074] |
| Mean of D | -3.0713 | S.D. of Dependent Variable | 9.0364 |
| Residual S | 3578.2 | Equation Log-likelihood | -171.5830 |
| Akaike Infc | -173.583 | Schwarz Bayesian Criterion | -175.4542 |
| DW-statist | 1.1673 | | |

Diagnostic Tests

| Test Statistics | LM Version | F Version |
|---|---|---|
| A:Serial C | CHSQ( 4)= 10.0203[.040]* | F( 4, 42)= 2.7703[.039]* |
| B:Function | CHSQ( 1)= 2.9439[.086]* | F( 1, 45)= 2.9402[.093]* |
| C:Normalit | CHSQ( 2)= .18848[.910]* | Not applicable |
| D:Heteros | CHSQ( 1)= .0054747[.941]* | F( 1, 46)= .0052472[.943]* |

A:Lagrange multiplier test of residual serial correlation
B:Ramsey's RESET test using the square of the fitted values
C:Based on a test of skewness and kurtosis of residuals
D:Based on the regression of squared residuals on squared fitted values

a2=1

Wald Statistic CHSQ( 1)= 15.1984[.000]

Ordinary Least Squares Estimation

Dependent variable is SJPN2
48 observations used for estimation from 1989Q4 to 2001Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| C | -7.8158 | 5.5399 | -1.4108 | [.165] |
| RDJP | 1.3078 | 1.438 | 0.90942 | [.368] |

| | | | |
|---|---|---|---|
| R-Squared | 0.017662 | R-Bar-Squared | -.0036934 |
| S.E. of Re | 26.2738 | F-stat. F( 1, 46) | .82705[.368] |
| Mean of D | -4.1431 | S.D. of Dependent Variable | 26.2254 |
| Residual S | 31754.4 | Equation Log-likelihood | -223.9791 |
| Akaike Infc | -225.9791 | Schwarz Bayesian Criterion | -227.8503 |
| DW-statist | 1.6062 | | |

Diagnostic Tests

| Test Statistics | LM Version | F Version |
|---|---|---|
| A:Serial C | CHSQ( 4)= 9.6545[.047]* | F( 4, 42)= 2.6436[.047]* |
| B:Function | CHSQ( 1)= 1.7189[.190]* | F( 1, 45)= 1.6713[.203]* |
| C:Normalit | CHSQ( 2)= .46151[.794]* | Not applicable |
| D:Heteros | CHSQ( 1)= 2.2798[.131]* | F( 1, 46)= 2.2938[.137]* |

A:Lagrange multiplier test of residual serial correlation
B:Ramsey's RESET test using the square of the fitted values
C:Based on a test of skewness and kurtosis of residuals
D:Based on the regression of squared residuals on squared fitted values

# Table 5: Risk Premia and Rational Expectations (US dollar/Pound)

Ordinary Least Squares Estimation

Dependent variable is SGBP1
48 observations used for estimation from 1989Q4 to 2001Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| CONST | 0.61295 | 1.2559 | 0.48807 | [.628] |
| DRGB | -0.38059 | 0.36183 | -1.0518 | [.298] |

| | | | |
|---|---|---|---|
| R-Squared | 0.023487 | R-Bar-Squared | .0022580 |
| S.E. of Re | 6.2519 | F-stat. F( 1, 46) | 1.1064[.298] |
| Mean of D | 1.5317 | S.D. of Dependent Variable | 6.2590 |
| Residual S | 1798 | Equation Log-likelihood | -155.0663 |
| Akaike Inf | -157.0663 | Schwarz Bayesian Criterion | -158.9375 |
| DW-statist | 1.0462 | | |

Diagnostic Tests

| * Test Statistics * | LM Version | * | F Version | * |
|---|---|---|---|---|
| * A:Serial C | CHSQ( 4)= | 15.8754[.003]* | F( 4, 42)= | 5.1889[.002]* |
| * B:Functio | CHSQ( 1)= | 6.4176[.011]* | F( 1, 45)= | 6.9451[.011]* |
| * C:Normal | CHSQ( 2)= | 6.2566[.044]* | Not applicable | * |
| * D:Hetero | CHSQ( 1)= | .40080[.527]* | F( 1, 46)= | .38733[.537]* |

A:Lagrange multiplier test of residual serial correlation
B:Ramsey's RESET test using the square of the fitted values
C:Based on a test of skewness and kurtosis of residuals
D:Based on the regression of squared residuals on squared fitted values

Ordinary Least Squares Estimation
Based on Newey-West adjusted S.E.'s Parzen weights, truncation lag= 6

Dependent variable is DGBP3
48 observations used for estimation from 1989Q4 to 2001Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| CONST | 0.61295 | 2.0263 | 0.3025 | [.764] |
| DRGB | -0.38059 | 0.67542 | -0.56349 | [.576] |

a2=1 Based on Newey-West adjusted S.E.'s

Wald Statistic CHSQ( 1)= .84103[.359]

Ordinary Least Squares Estimation

Dependent variable is SGBP2
48 observations used for estimation from 1989Q4 to 2001Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| CONST | 0.13325 | 4.5514 | 0.029277 | [.977] |
| DRGB | -0.28954 | 1.3113 | -0.2208 | [.826] |

| | | | |
|---|---|---|---|
| R-Squared | 0.001059 | R-Bar-Squared | -.020657 |
| S.E. of Re | 22.6578 | F-stat. F( 1, 46) | .048752[.826] |
| Mean of D | 0.83218 | S.D. of Dependent Variable | 22.4273 |
| Residual S | 23615.3 | Equation Log-likelihood | -216.8718 |
| Akaike Inf | -218.8718 | Schwarz Bayesian Criterion | -220.7430 |
| DW-statist | 2.0101 | | |

Diagnostic Tests

| * Test Statistics * | LM Version | * | F Version | * |
|---|---|---|---|---|
| * A:Serial C | CHSQ( 4)= | 3.4219[.490]* | F( 4, 42)= | .80601[.528]* |
| * B:Functio | CHSQ( 1)= | 1.8403[.175]* | F( 1, 45)= | 1.7940[.187]* |
| * C:Normal | CHSQ( 2)= | 74.6235[.000]* | Not applicable | * |
| * D:Hetero | CHSQ( 1)= | 9.4272[.002]* | F( 1, 46)= | 11.2425[.002]* |

A:Lagrange multiplier test of residual serial correlation
B:Ramsey's RESET test using the square of the fitted values
C:Based on a test of skewness and kurtosis of residuals
D:Based on the regression of squared residuals on squared fitted values

Ordinary Least Squares Estimation
Based on White's Heteroscedasticity adjusted S.E.'s

Dependent variable is DGBP2
48 observations used for estimation from 1989Q4 to 2001Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| CONST | 0.13325 | 3.4353 | 0.038788 | [.969] |
| DRGB | -0.28954 | 2.0104 | -0.14402 | [.886] |

# Table 6: Risk Premia and Rational Expectations (US dollar/German Mark)

Ordinary Least Squares Estimation

Dependent variable is SDM1
36 observations used for estimation from 1989Q4 to 1998Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| C | -2.7218 | 1.181 | -2.3046 | [.027] |
| RDDM | 0.80003 | 0.39174 | 2.0422 | [.049] |

| | | | |
|---|---|---|---|
| R-Squared | 0.10927 | R-Bar-Squared | .083068 |
| S.E. of Re | 6.8378 | F-stat. F( 1, 34) | 4.1707[.049] |
| Mean of D | -3.355 | S.D. of Dependent Variable | 7.1408 |
| Residual S | 1589.7 | Equation Log-likelihood | -119.2616 |
| Akaike Infc | -121.2616 | Schwarz Bayesian Criterion | -122.8451 |
| DW-statist | 0.56464 | | |

Diagnostic Tests

| * Test Statistics * | LM Version | * | F Version | * |
|---|---|---|---|---|
| A:Serial Cc | CHSQ( 4)= | 20.8406[.000]* | F( 4, 30)= | 10.3107[.000]* |
| B:Function | CHSQ( 1)= | 2.9557[.086]* | F( 1, 33)= | 2.9518[.095]* |
| C:Normalit | CHSQ( 2)= | 1.8170[.403]* | Not applicable | * |
| D:Heterosc | CHSQ( 1)= | .020928[.885]* | F( 1, 34)= | .019777[.889]* |

A:Lagrange multiplier test of residual serial correlation
B:Ramsey's RESET test using the square of the fitted values
C:Based on a test of skewness and kurtosis of residuals
D:Based on the regression of squared residuals on squared fitted values

Ordinary Least Squares Estimation
Based on Newey-West adjusted S.E.'s Parzen weights, truncation lag= 6

Dependent variable is SDM1
36 observations used for estimation from 1989Q4 to 1998Q3

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| C | -2.7218 | 2.0522 | -1.3263 | [.194] |
| RDDM | 0.80003 | 0.43925 | 1.8213 | [.077] |

a2=1

Wald Statistic CHSQ( 1)= .20726[.649]

Ordinary Least Squares Estimation

Dependent variable is SDM2
35 observations used for estimation from 1989Q4 to 1998Q2

| Regressor | Coefficient | Standard E | T-Ratio | [Prob] |
|---|---|---|---|---|
| C | -3.0823 | 4.4048 | -0.69976 | [.489] |
| RDDM | 0.87995 | 1.4497 | 0.60701 | [.548] |

| | | | |
|---|---|---|---|
| R-Squared | 0.011042 | R-Bar-Squared | -.018926 |
| S.E. of Re | 24.9638 | F-stat. F( 1, 33) | .36846[.548] |
| Mean of D | -3.8494 | S.D. of Dependent Variable | 24.7308 |
| Residual S | 20565.2 | Equation Log-likelihood | -161.2430 |
| Akaike Infc | -163.243 | Schwarz Bayesian Criterion | -164.7984 |
| DW-statist | 1.9036 | | |

Diagnostic Tests

| * Test Statistics * | LM Version | * | F Version | * |
|---|---|---|---|---|
| A:Serial Cc | CHSQ( 4)= | 4.1871[.381]* | F( 4, 29)= | .98518[.431]* |
| B:Function | CHSQ( 1)= | .62149[.430]* | F( 1, 32)= | .57849[.452]* |
| C:Normalit | CHSQ( 2)= | 1.9461[.378]* | Not applicable | * |
| D:Heterosc | CHSQ( 1)= | 2.5649[.109]* | F( 1, 33)= | 2.6096[.116]* |

A:Lagrange multiplier test of residual serial correlation
B:Ramsey's RESET test using the square of the fitted values
C:Based on a test of skewness and kurtosis of residuals
D:Based on the regression of squared residuals on squared fitted values
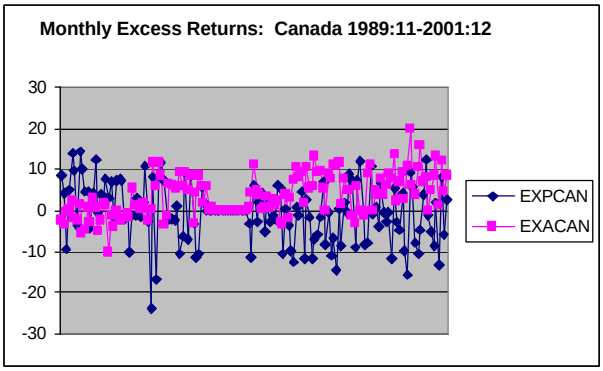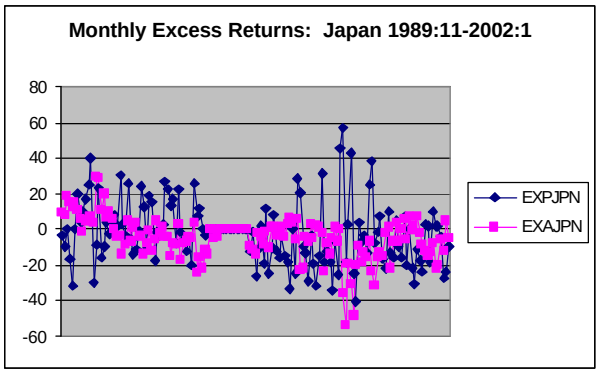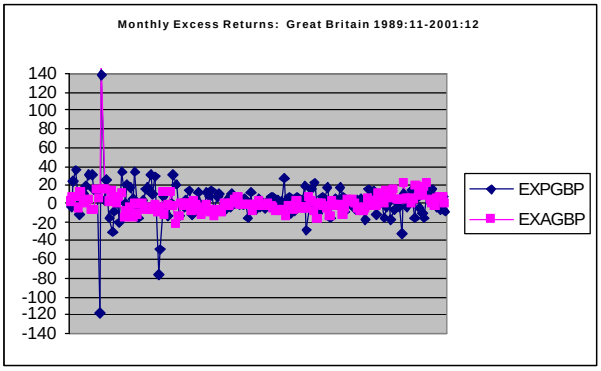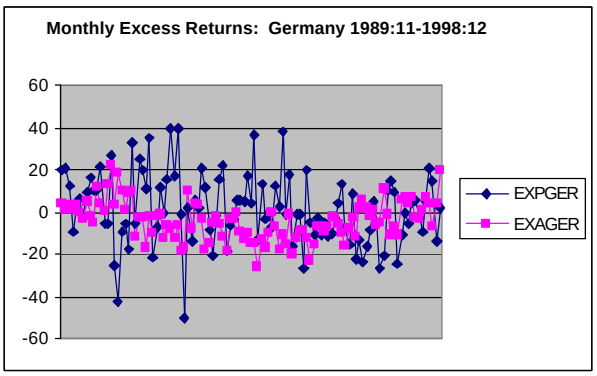
Figure 1: Excess Returns of US dollar Exchange Rates