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ON TESTING THE HYPOTHESIS THAT PREMIUM IN US$/$A FORWARD EXCHANGE RATE IS CONSTANT: A SIGNAL EXTRACTION APPROACH

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ABSTRACT

In this paper, we implement the signal extraction approach, a methodology from engineering literature, to identify and measure premia in the pricing of 30-day and 90-day US$/$A forward foreign exchange. The estimated premium models indicate that the variance of the premium term is not equal to zero and reject the hypothesis that the premium is constant. It also lends support to some other findings that the variance of the premium term accounts for more that 70 per cent of the variation in forecast errors that results from the use of current forward rates as a predictor of future spot rates.

1 INTRODUCTION

In the past decade, there is an extensive literature testing the joint hypothesis of market efficiency and time-invariant risk premia in foreign exchange markets. The results of these studies are remarkably mixed. Hansen and Hodrick (1980, 1983), Bilson (1981), Fama (1984), Hsieh (1984) and Hidrick and Srivastave (1986), among others, reject the unbiased forward rate hypothesis. They find that a time-varying premia is present in several major foreign exchange markets. The implication of these empirical findings is that one cannot use the forward rate directly as a measure for the future spot rate. Frankel (1982), using a six-currencies test, fails to identify such premium, while Domowitz and Hakkio (1985) provide some evidence of non-zero constant risk premium for some, but not all currencies. Recently, Wolff (1987), using a signal extraction approach to analyse the three bilateral exchanges rates for the US dollar (US$/£, US$/Mark, and US$/Yen), indicates that premia, to some extent, do persist over time.

For Australia, the results are not much different. Levis (1982), and Turnovsky and Ball (1983), have found weak support for the unbiasedness of the US$/$A forward rate. Tease (1988) shows that the unbiasedness was holding only before the February depreciation in 1985 but became biased thereafter. In addition he identifies a behavioural change in the market after February 1985. Madsen (1990), however, finds the unbiasedness results quite in contrary to Tease's conclusion and a structural shift could exist on the 30-day forward market but not on the 90-day forward market.

The methodologies used in these studies usually involve estimation of equations in regression format. Conclusions about the behaviour of premia in the pricing of forward foreign exchange are thus drawn on the basis of the observed relationship between a regressand and a set of regressors. However, the choice of a set of regressors to be used in this context is fairly arbitrary. In this paper, the signal extraction approach, a technique from the engineering literature, will be used to identify and measure premia in the pricing of 30-day and 90-day US$/$A forward foreign exchange, using three weekly nominal exchange rates series (the spot US$/$A exchange rates, the 30-day forward rates and the 90-day forward rates) over the period from 21 May 1982 to 21 April 1989. The advantage of using the signal extraction approach is that an estimation will be updated once a new observation becomes available. In addition it allows the researcher to inspect the characteristics of the unobserved premium and see whether its presence, as suggested by some researches, only adds marginally to the variability of the forecast errors.

The paper is structured as follows: Section 2 will be used to model implementation. Preliminary discussion of the data will be in Section 3, while in Section 4 a sample correlogram will be examined for model identification. In Section 5, estimates of the unknown parameters which are computed by a maximum likelihood procedure based on the Kalman filter will be presented. Conclusion of the study will be in Section 6.

2 THE IMPLEMENTATION OF THE STATE SPACE MODEL

Conceptually, one can decompose the forward rate as the sum of the expected future spot rate and the premium term, that is

$$F(t,t+1)= E[S(t+1)/I(t)] + P(t) \qquad (1)$$

where F(t,t+1) is the natural logarithm of the forward rate at time t for the exchange rate at time t+1, E[S(t+1)/I(t)] is the rational or efficient forecast of the logarithm of the spot exchange rate at time t+1, given that information available at time t, and P(t) is the premium term. The main reason for using logarithms is to avoid the so-called Siegel paradox, see Siegel (1972).

By subtracting S(t+1), the logarithm of the subsequent spot rate at time t+1 from both sides of equation (1), we get:

$$F(t,t+1) - S(t+1) = P(t) + V(t+1) \qquad (2)$$

where $V(t+1) = E[S(t+1)/I(t)] - S(t+1)$

and

V(t) is an uncorrelated, zero-mean sequence.

Equation (2) states that the forecast error resulting from the forward rate as a predictor of the future spot rate consists of premium component and a white-noise error term due to the availability of new information between time t and t+1 concerning the spot exchange rate at time t+1. In the state space framework, the premium component P(t) is often referred to as the signal that we would like to characterise and measure and V(t+1) is the noise that contaminated the signal. The problem that we face thus involves extracting a signal from a noisy environment.

3 THE DATA

The 30-day and 90-day US$/$A forward exchange traded in the Australian Foreign exchange market will be used in our empirical testing. We chose this data set partly because of the convenient availability of the data[1], and partly because most of the previous studies have looked at the same problem for the similar time period. The US$/$A spot rates, the 30-day US$/$A forward rates and 90-day US$/$A forward rates are taken from the closing price of every Friday, for the period from 21 May 1982 to 21 April 1989; therefore, there are 362 observations. The realised 30-day US$/$A forward rate is taken from the subsequent US$/$A spot rates four weeks from that Friday, while that of the 90-day US$/$A forward rate is taken from the subsequent US$/$A spot rates thirteen weeks from the Friday in question.

Recently, Tease (1988) identified that there was a behavioural change on the forward market after the depreciation of February 1985. By avoiding a period of severe parameter instability from February to July 1985, Madsen (1990), conducting the same joint hypothesis, divided the sample period into two. The first period encompasses 21 May 1982 to 25 January 1985, and the second period encompasses 2 August 1985 to 21 April 1989. Figures 1 and 2 show the graph of F(t,t+1) - S(t+1) for 30-day forward rate and that for 90-day forward rate of the entire sample period. They are of particular interest. It is apparent that there is a wide fluctuation (instability) from February to July 1985. The correlograms of F(t,t+1) - S(t+1), which are shown in Table 1 and 2, for the two subperiods (as suggested in Madsen) and whole sample period, however, do not seem to support this. Instead, the correlograms for the two subperiods are very similar and they are very similar to the correlogram of the whole period for

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1 The data used in this study is kindly provided by Tony Hall.

the two series in question. Thus the arguments put forward by Madsen and Tease on the structural change in the forward market might not be valid. In order to compare the result with other studies and for testing whether there exists a behavioural change in the two forward exchange markets, the two series in question will therefore be split at the same time interval as in other studies.

4 MODEL IDENTIFICATION

The methodology for model identification of the unobserved component in this paper follows the approach pioneered by Box and Jenkins (1976: 121), in which the observations themselves are used to identify a suitably parsimonious model from the classes of autoregressive integrated moving average (ARIMA) process. From equation (2), we observe that the difference between the forward rate and the subsequent future spot rate, F(t,t+1) - S(t+1), is the premium p(t) plus the white noise V(t+1). Thus the autocorrelations of F(t,t+1) - S(t+1) must at the same time be the autocorrelations of the combined premium-plus-noise process P(t) + V(t+1). Once an appropriate model for F(t,t+1) - S(t+1) is identified based on the sample autocorrelation, an appropriate model for the premium can also be identified. This is based on the properties of aggregation of the time series.[2] Then the Kalman filter can be used to extract the unobserved premium component from the observed observations. Table 1 and 2 show the sample autocorrelations of F(t,t+1) - S(t+1) from lag one to fifteen for 30-day forward rates and 90-day forward rates for various periods. The sample autocorrelations of F(t,t+1) - S(t+1) decays slowly over time (with the exception of the 30-day forward rate which exhibits more rapid decay, as would be expected of a stationary process.) This indicates that differencing might be required for the 90-day forward series. The sample autocorrelations of the first difference, which are shown in Table 3, are generally small relative to their asymptotic standard error (equal to $1/\sqrt{n}$) with the exception of the seasonal autocorrelation at lag of 13 weeks for the 90-day forward rate. This is consistent with the random walk model. We can thus assume that the premium for the 90-day forward rate follows a random walk process, that is, P(t) = P(t-1) + a(t). For the 30-day forward rate, since the sample autocorrelations of

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2 For detail discussion of model identification, see Cheung (1986).

F(t,t+1) - S(t+1) die down quickly towards zero and the sample partial autocorrelations tail off towards zero, we can safely infer that this belongs to the class of ARMA(1,1) model. Using the theorem from the aggregation of time series, an ARMA(1,1) model for P(t) + V(t+1) is consistent with AR(1) model for the premium, and a white-noise error term, V(t+1). Thus we will employ AR(1) specification for the premium in the 30-day forward rate, that is, P(t) = ϕP(t-1) + a(t). Together with equation (2), this is precisely the state space format that is required to apply the Kalman filtering:

$$F(t,t+1) - S(t+1) = P(t) + V(t+1) \qquad (2)$$

$$P(t) = \phi P(t-1) + a(t) \qquad (3)$$

where a(t) is normally distributed with mean zero and variance σ_p^2.

5 THE EMPIRICAL RESULTS

Maximum-likelihood estimates of the state-space models for premia in the pricing of 30-day and 90-day US$/$A forward exchange rates are presented in Table 4. Most of the estimated parameters are statistically significant at a 95 per cent level. Since the maximum-likelihood estimate of ϕ in all cases is significantly different from zero, but not significantly different from one, this implies that the premium term follows a random walk rather that a first-order autoregressive process, AR(1). The estimated result confirms that the premium term do follows a random walk. The value of log-likelihood function has decreased by more than 2 in all cases. Based on the Akaike Information Criterion for model selection[3], this would be the better model of the two. The maximum-likelihood estimates of σ_p^2 (the variance of the premium term) are statistically significantly different from zero. They confirm that σ_p^2 is not equal to zero, and reject the hypothesis that the premium term is a constant.

The same pattern of results are observed for the 90-day UA$/$A forward rates series for the various periods. All the maximum-likelihood estimates of σ_p^2 are significantly different from

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3 An overall decision on the 'better' model can be obtained by using Akaike Information Criterion (AIC) (Akaike, 1973), that is, AIC + [-2 in Maximised likelihood + 2 (number of independent parameters estimated)].

zero. They further support the hypothesis that the variance of the premium is not equal to zero. The time-varying premium is observed in each period. These results indicate that premia do show a certain degree of persistence over time.

The estimated variance of the premium term σ_p^2 and the variance of the noise term, σ_v^2, reported in the Table 4, are of particular interest. In all cases, σ_p^2 is greater than σ_v^2 by a minimum of three times. This has an important implication that the variation in the premium terms accounts for more than 70 per cent of the variation in the forecast errors F(t,t+1) -S(t+1). Our result strongly supports the conclusion reached by Fama (1984) and Hodrich and Srivastave (1986) that the variability of the premium terms dominates the variance in the forecast errors that results from the use of the current forward rates as a predictor of future spot rates.

To test the validity of the filter model and assumption, we examine the randomness of the residuals or innovations (that is, the difference between the actual measurements and the estimated measurements). An optimal filter has the property that the estimated innovations sequence has a zero mean and is white with covariance v, if the model is adequate. Table 5 shows the autocorrelations of F(t,t+1) - S(t+1) - PS(t). They were calculated on the basis of estimates of PS(t) that result from Kalman filter[4]. We would like to test the whiteness or randomness of the residual left. None of the individual autocorrelations (with exception of lag 4 and lag 13 for the 30-day and the 90-day market respectively) of the residual in Table 5 is statistically significant. The significance levels associated with the Box and Ljung portmanteau test[5] of randomness at a 1 per cent level with 8 d.f. (with autocorrelation at lag 4 for 30-day neglected) also indicate that no significant residual autocorrelation is present. Thus the estimation of the premium model is successful in the sense that the extraction model appears to have captured the essence of the time-series properties of the premium terms.

4 PS(t) is obtained by the smoothing of the Kalman filter which adjusts each filtered estimate by a backward pass through all the subsequent measurements. Details of the smoothing algorithms can be found in Meditch (1969) and Anderson and Moore (1979).

5 The Box-Ljung portmantaeu test of randomness is fairly large for 30-day forward rates, but this is due primarily to the seasonal coefficient at lag 4.

Table 6 presents the break-point test of the structure of the two series in question. In this test, only the 'better' model[6] of each series will be used. Using the Chow test procedure, it is possible to establish such a break in 1985. Since the calculated F is greater than the critical F for the two series in question, we can reject the null hypothesis of parameter stability in the two periods at both the 1 per cent and 5 per cent significance level. Our result confirms the finding by Tease (1988) that there existed a behaviour change in both 30-day and 90-day US$/$A forward markets after February 1985.

6 CONCLUSION

In this article, we applied a methodology from engineering literature to identify and measure premia in the pricing of 30-day and 90-day US$/$A forward foreign exchange rates. The methodology was quite successful in capturing the essence of the time series properties of the premia in the two series in question. The model indicates that premia show a certain degree of persistence over time. Since the variance of the premium term is significantly greater than zero in each series, this rejects the hypothesis that the premium in forward exchange rates is constant. Furthermore, the empirical results have shown support to some of the findings by Fama (1984) and others that most of the variation in forward rates is due to variation in premia. The variance of the premium term has accounted for more than 70 per cent of the variance in the forecast error that results from the use of current forward rates as a predictor of future spot rates. In addition, the break-point test of the structure of the two series has shown that there existed such a break in 1985 for both the 30-day and 90-day US$/$A forward market.

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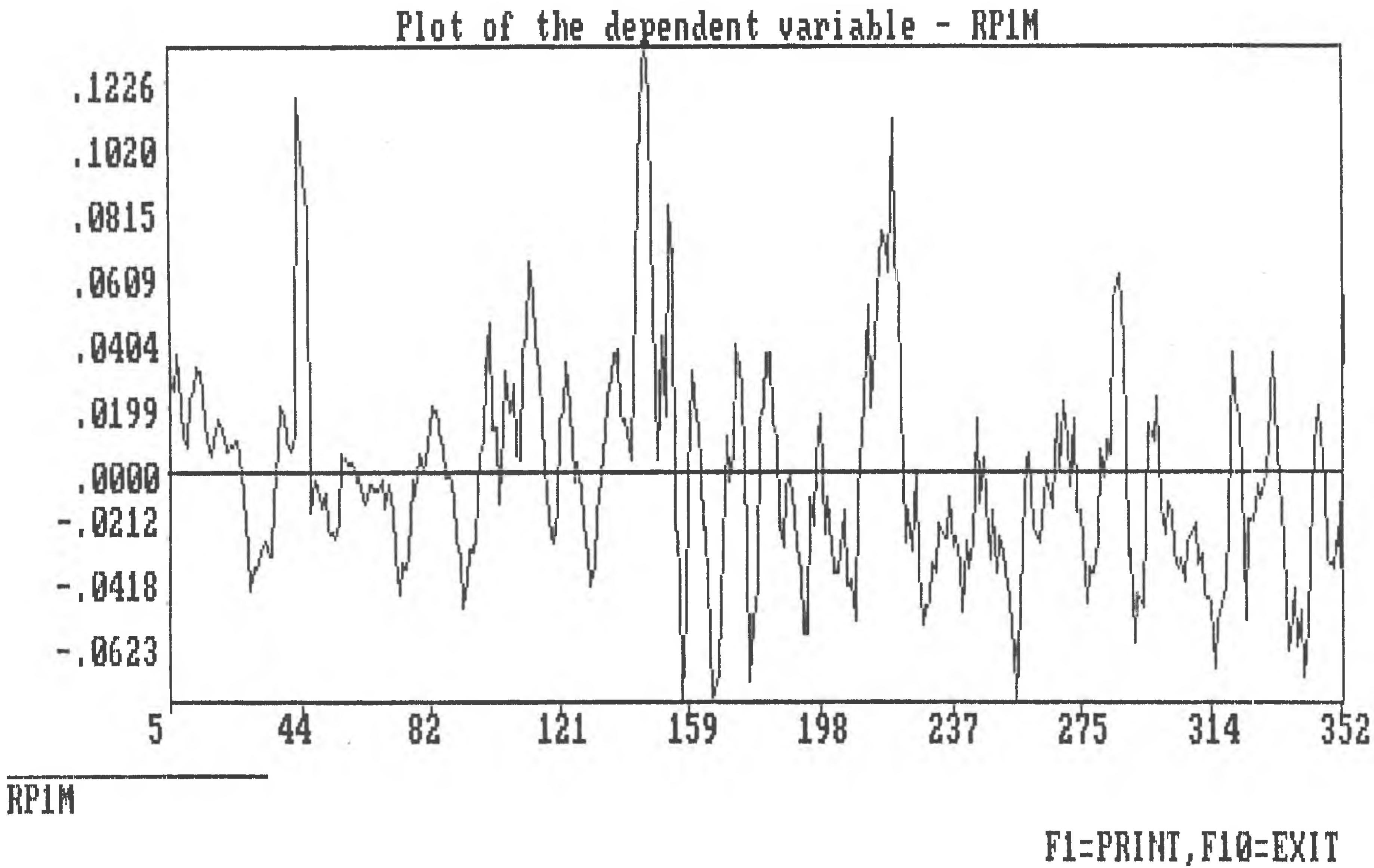
6 As defined footnote 3.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.0815
.0609
.0404
.0199
.0000
-.0212
-.0418
-.0623
5
44
82
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198
237
275
314
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RP1M
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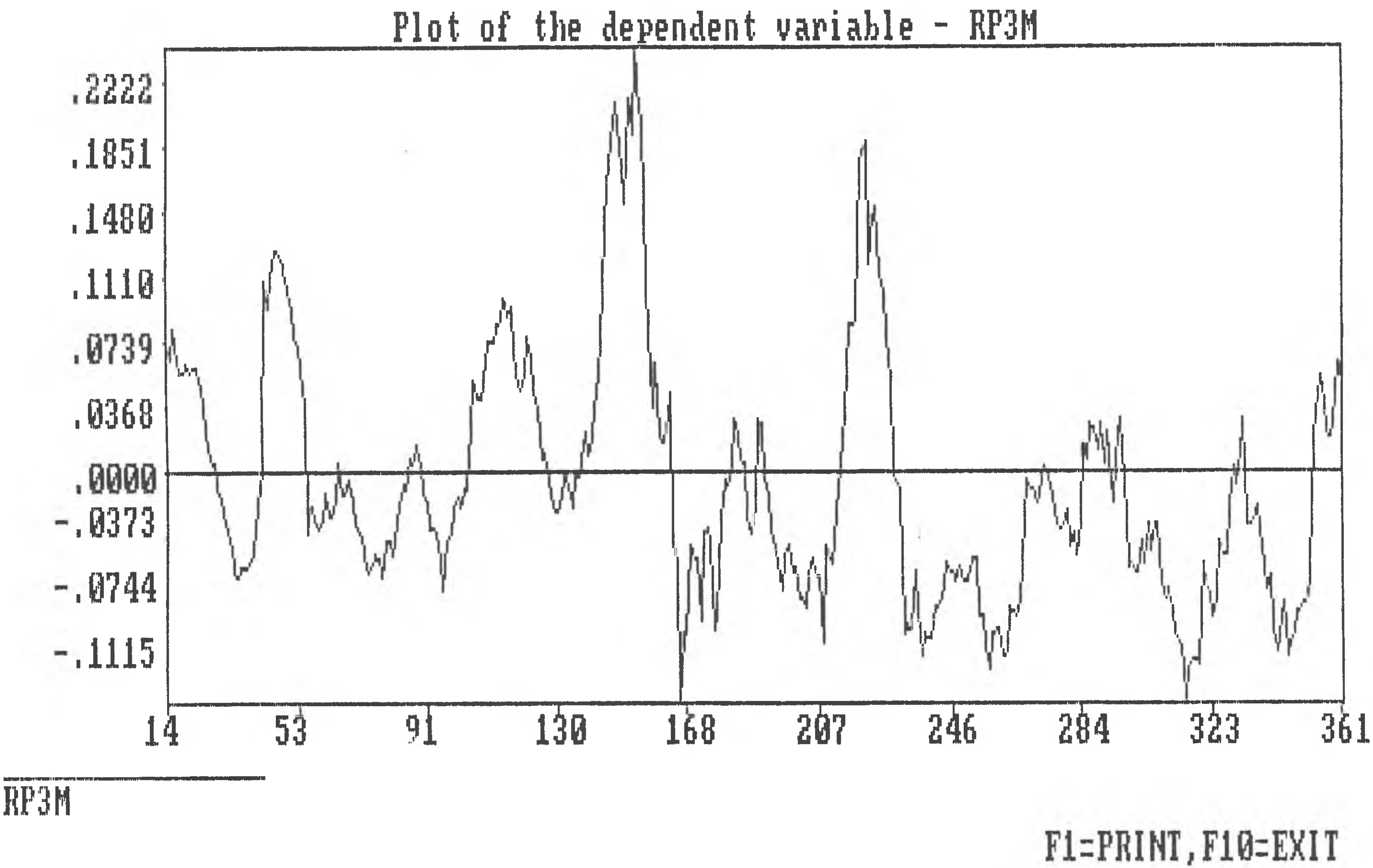

Plot of the dependent variable - RP3M
.2222
.1851
.1480
.1110
.0739
.0368
.0000
-.0373
-.0744
-.1115
14
53
91
130
168
207
246
284
323
361
RP3M
F1=PRINT,F10=EXIT

TABLE 1

Autocorrelations of F(t,t+1) - S(t+1) for 30-day Forward Market

Lag	21/5/1982-21/4/1989	21/5/1982-25/1/1985	2/8/1985-21/4/1989
1	0.80*	0.76*	0.76*
2	0.60*	0.51*	0.56*
3	0.36*	0.28*	0.31*
4	0.16*	0.11	0.09
5	0.13	0.09	0.04
6	0.10	0.05	-0.02
7	0.10	0.00	-0.02
8	0.08	-0.07	-0.06
9	0.06	-0.13	-0.07
10	0.04	-0.12	-0.06
11	0.01	-0.10	-0.09
12	-0.02	-0.10	-0.06
13	-0.03	-0.11	-0.05
14	-0.05	-0.16	-0.04
15	-0.07	-0.19	-0.04

Table 2

Autocorrelations of F(t,t+1) - S(t+1) for 90-day Forward Market

Lag	21/5/1982 - 21/4/1989	21/5/1982 - 25/1/1985	2/8/1985 - 21/4/1989
1	0.94*	0.94*	0.92*
2	0.88*	0.85*	0.84*
3	0.80*	0.76*	0.76*
4	0.74*	0.66*	0.67*
5	0.65*	0.57*	0.58*
6	0.56*	0.47*	0.47*
7	0.47*	0.36*	0.37*
8	0.38*	0.24*	0.24*
9	0.27*	0.13	0.11
10	0.19*	0.03	0.01
11	0.09	-0.06	-0.09
12	0.01	-0.13	-0.17
13	-0.07	-0.17	-0.24
14	-0.10	-0.19	-0.26
15	-0.12	-0.20	-0.27

Table 3

Autocorrelations of the First Difference of F(t,t+1) - S(t+1) for 90-day Forward Market

Lag	21/5/1982 - 21/4/1989	21/5/1982 - 25/1/1985	2/8/1985 - 21/4/1989
1	0.05	0.12	0.04
2	0.14	0.26	-0.01
3	-0.05	0.07	0.05
4	0.16	0.12	0.09
5	0.01	0.03	0.15
6	0.03	0.07	-0.11
7	0.04	0.04	0.19
8	0.07	-0.01	-0.07
9	-0.12	-0.15	-0.09
10	0.10	-0.05	0.03
11	-0.10	-0.07	-0.09
12	-0.03	-0.09	-0.01
13	-0.49*	-0.31*	-0.43*
14	-0.05	-0.09	-0.01
15	-0.01	-0.06	0.03

Table 4

Maximum-Likelihood Estimates of State Space Models for Premia in the Pricing of Forward Foreign Exchange

Series	ϕ	σ_p^2	σ_ε^2	Log L	σ_{pe}^2	H^a	Q^b	R_D^2
		(X10^4)	(X10^4)		(X10^4)		Statistic	
rp1m	0.834	3.473	0.0	1196.1	3.800	1.401		0.136
	(34.76)	(12.72)	(0.0)					
		3.029	0.264	1169.9	4.390	1.415	16.10	0.002
		(7.620)	(1.421)					
rp1m1	0.894	2.453	0.0	508.7	2.685	0.697		0.108
	(30.94)	(8.206)	(0.0)					
		3.014	0.0	500.3	2.438	0.653	7.08	0.190
		(8.426)	(0.0)					
rp1m2	0.830	3.245	0.164	625.9	3.781	0.849		0.147
	(17.45)	(3.771)	(0.377)					
		2.549	0.554	612.9	4.341	0.882	9.75	0.021
		(5.326)	(2.149)					
rp3m		3.820	0.342	1148.8	4.948	1.431	16.01	0.048
		(7.271)	(1.420)					
rp3m1		1.543	0.499	496.5	2.637	1.001	12.37	0.255
		(4.267)	(2.398)					
rp3m2		3.651	0.364	604.7	4.609	0.646	16.60	0.063
		(5.239)	(1.130)					

Notes:

t ratios are in parentheses;

σ_{pe}^2 (the prediction error variance): The steady state variance of the one-step ahead prediction error.

[a]: Heteroscedasticity statistic: asterisks indicate significant values at the 10 per cent level for a 2-sided test based on the F distribution of (m,m) d.f. where $m=T^*/3$ or the nearest integer to it.

[b]: Box-Ljung Q Statistic (autocorrelation at lag 4 has been ignored in 30-day market.)

R_D^2: defined in Harvey (1984) as a yardstick to measure goodness of fit.

Table 5

Autocorrelations of the Residual of F(t,t+1) - (S(t+1) - PS(t) for 30-day and 90-day Forward Market.

Lag	RP1M	RP3M
1	-0.005	-0.000
2	0.043	0.031
3	-0.052	-0.148
4	-0.387*	0.097
5	0.106	-0.031
6	0.034	0.013
7	0.104	0.052
8	0.070	0.093
9	0.040	-0.073
10	0.103	-0.175
11	0.016	-0.008
12	-0.013	0.040
13	0.050	-0.430*
14	0.010	0.047
15	0.078	0.120

Table 6

Chow Test Results for Structural Stability*
(for Pre-February 1985 and Post-August 1985)

Series	Calculated F	Critical F
RP1M	$F^{2}_{345}= 15.73$	$F^{2}_{345} = 4.61$ (1%)
		$= 4.00$ (5%)
RP3M	$F^{2}_{345} = 16.75$	

Note: The F-statistic is calculated by

$$F^{k}_{T^* - 2k} = \frac{[SSRa - (SSR_1 + SSR_2)]/k}{(SSR_1 + SSR_2/T^*_1 + T^*_2 - 2k)}$$

where

SSR_a = Sum of square residuals from entire period.
SSR_1 = Sum of square residuals from first period.
SSR_2 = Sum of square residuals from second period.
k = number of parameters estimated,

and T^*_i = number of observations (usually $T_i = 5$).

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