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TESTS OF THE MARTINGALE HYPOTHESIS FOR FOREIGN CURRENCY FUTURES WITH TIME-VARYING VOLATILITY

Thomas H. McCURDY and Ieuan G. MORGAN *
Queen's University, Kingston, Ont., Canada K7L 3N6

Abstract: The martingale hypothesis for daily and weekly rates of change of futures prices for five currencies is tested in this paper. With daily data, we find some evidence against the null hypothesis for each currency. Although institutionally imposed limits on daily price changes were binding fairly often in the earlier years of the sample, the results are not substantially different when data affected by limit moves are removed. Trading day effects in foreign currency futures and spot prices introduce complicated day of the week patterns in futures price. For this reason, we retest the martingale hypothesis with weekly data and reject the null hypothesis for only one currency. For this currency, one interpretation of the evidence is that a time-varying risk premium exists.

Keywords: Foreign currency futures, Martingale, Time varying volatility, GARCH.

1. Introduction

The martingale hypothesis for daily and weekly rates of change of futures prices for five currencies is tested in this paper. With daily data we find some evidence against the null hypothesis for each currency, as do Hodrick and Srivastava (1985). When we retest the martingale hypothesis with weekly data, we obtain strikingly different results. We reject the null hypothesis for only one currency, the Deutschmark. For this currency, one interpretation of the evidence is that a time-varying risk premium exists.

Rates of change of futures prices for all five currencies show substantial time-varying volatility (see, for example, fig. 1 which plots the difference in logarithms of daily futures price data for the British pound over ten years). Such volatility is common to other financial price data. Mandelbrot (1963) observes that, for the prices of speculative assets, '...large changes tend to be followed by large changes – of either sign – and small changes tend to be followed by small changes...' Similarly, Engle (1982) quotes McNees (1979): 'the inherent uncertainty or randomness associated with different forecast periods seems to vary widely over time',..., 'large and small errors tend to cluster together (in contiguous time periods)'.

Given the time-varying volatility in foreign currency data, it is not surprising that a homoskedastic error structure has invariable been rejected for conditional distributions [see, for example, Cumby and Obstfeld (1983), or Gregory and McCurdy (1986)]. Hansen and Hodrick (1980, 1983), Hodrick

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and Srivastava (1984, 1985), and Hsieh (1984) have used estimation procedures that result in heteroskedasticity consistent covariance matrices. Alternatively, Diebold (1986), Domowitz and Hakkio (1985), Hsieh (1985), McCurdy and Morgan (1985), and Milhoj (1985) have modeled the time-varying volatility of foreign currency price changes as conditional heteroskedasticity that is a function of recent news or forecast errors, as in the autoregressive heteroskedasticity (ARCH) model of Engle (1982). [1]

Most previous studies have used monthly data. The highly organized and active markets for foreign currencies might imply that the persistence of the effects of shocks to volatility, and any associated effects on the mean, can be expected to be short lived. [2] Therefore, in addition to providing more information, daily data are more likely to allow detection and exploitation of any conditional heteroskedasticity.

Fine frequency data favour the use of futures data rather than forward market data. The fixed date of the futures contract maturity, as opposed to the fixed length of forward contracts, implies that daily futures price data refer to a sequence of expected values of a single future spot price, whereas daily forward data refer to the expected values of a sequence of future spot prices. Therefore, use of futures data avoids the moving average structure of the residuals arising from overlapping contracts in forward data. This allows straightforward identification and comparison of alternative specifications of the persistence of conditional heteroskedasticity.

In McCurdy and Morgan (1985), for daily rates of change in futures and spot prices for the Deutschmark, we evaluated the empirical performance of several specifications of the conditional heteroskedasticity and found the GARCH generalization of ARCH, due to Bollerslev (1986), to be a parsimonious model that represented the data well. As in Domowitz and Hakkio (1985), we also investigated any potential effect of the changing volatility of price on the level of the rate of change of price using the ARCH-M model. [3] Evidence of a time-varying risk premium of an ARCH-M type in daily data from nine Deutschmark/US dollar futures contracts (June 1983 to June 1985) appeared to be better interpreted as a weekend effect of the type discussed by Levi (1978) or French (1980). Otherwise, our test results were consistent with the martingale hypothesis for futures prices.

The daily data analysed in McCurdy and Morgan (1985) were collected from the Wall Street Journal and were necessarily for a relatively short period. The availability of data from the Center for Research in Futures Markets of the University of Chicago (CRFM) prompts us to attempt a more extensive evaluation of the martingal hypothesis for futures prices in five currencies over a much longer period. A recent analysis of foreign currency futures data is in Hodrick and Srivastava (1985). As mentioned above, they obtain standard errors that are robust to heteroskedasticity, by a generalized method of moments estimation procedure, instead of attempting to model the conditional heteroskedasticity explicitly. In addition, their transformation to induce stationarity differs from ours. In spite of these differences in analytic procedures, their results also allow the martingale hypothesis to be retained for the Deutschmark subsample which corresponds most closely to the time period we analysed. Nevertheless, the Hodrick and Srivastava tests reject the martingale hypothesis for the data in their complete sample in all currencies. So a further comparison of our method with theirs is warranted.

Section 2 integrates the theoretical results from equilibrium asset pricing theory necessary to derive an alternative hypothesis which involves a time-varying risk premium. Section 3 presents the particular forms of the test equations examined in this paper, including the specification of the

[1] See Engle and Bollerslev (1986) for a comprehensive survey of the various applications and extensions of the ARCH model.

[2] See Poterba and Summers (1984) who conclude, using weekly and monthly data, that the effect of changes in volatility of stock market prices on the level of return (and thus implicitly on a risk premium) is short lived.

[3] We use the ARCH-M terminology from Engle, Lilien and Robins (1986)

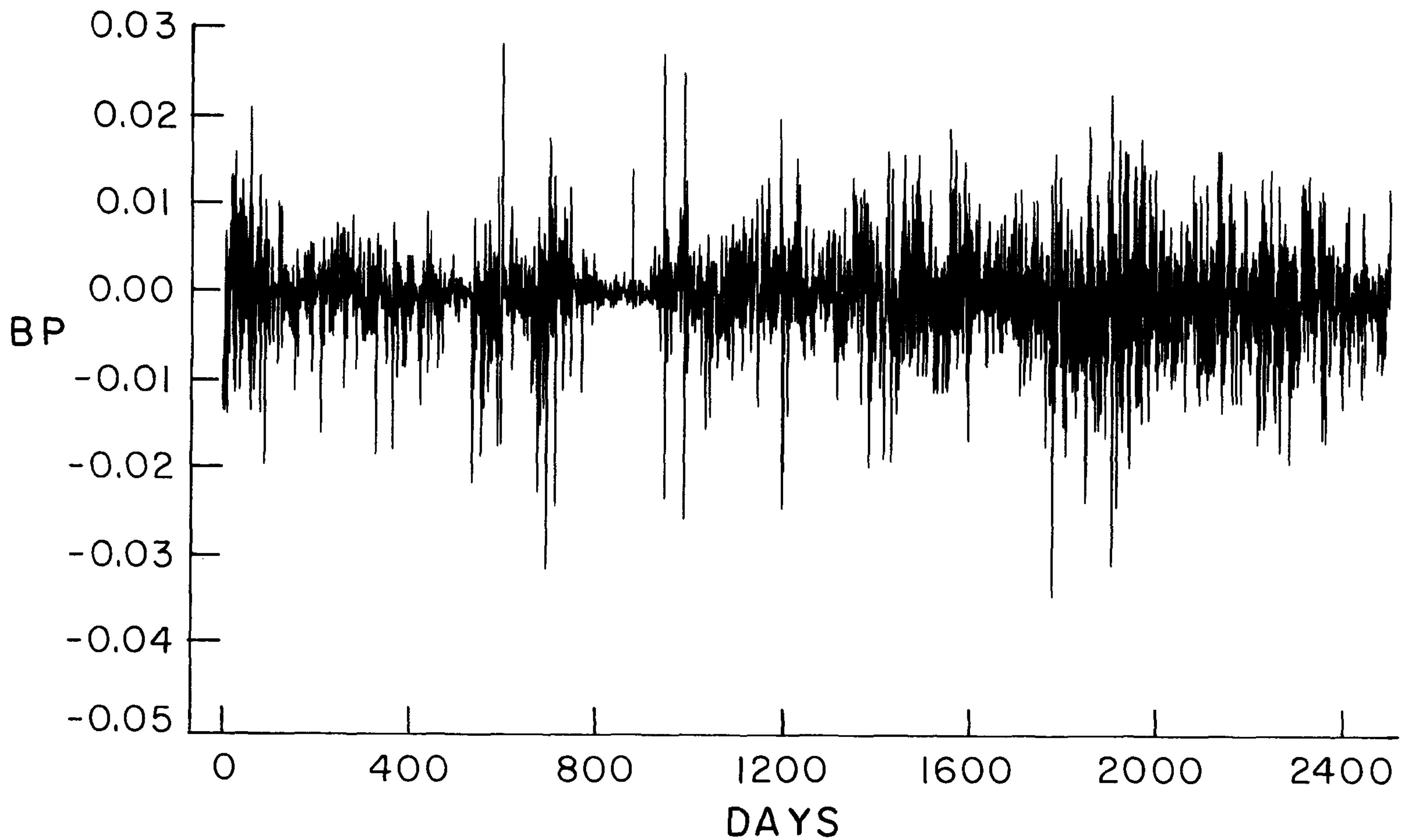

Fig. 1. Daily rates of change of futures prices for the British pound.

time-varying volatility as generalized autoregressive conditional heteroskedasticity (GARCH) under the null hypothesis and as GARCH-M under the alternative hypothesis. Section 4 describes the diagnostic checks and tests for omitted variables with which we evaluate the statistical adequacy of our test equations. Sections 5 and 6 summarize the results for the daily and weekly data respectively, and section 7 contains concluding comments.

2. Futures prices and time dependent risk premia

Chicago Mercantile Exchange futures contracts for foreign currencies (in terms of US dollars) are delivered on the third Wednesday of March, June, September, and December. These contracts expire two business days earlier on the data we define to be T. Since there is also a two day delivery lag for spot contracts, a spot settlement data corresponding to delivery of the futures contract requires that the futures price predict the spot price at the expiration date T. [4] First, to define notation, let

F_t = futures price at t for contract expiring at T,
S_T = spot price for underlying commodity at T,
R_t = 1 plus the riskless rate of interest from t to $t+1$,
C_t = consumption at t,
ρ = discount factor in additive, separable multiperiod utility function U_t,
$u(C_t)$ = utility of consumption at t, with greed, $u_t' > 0$ and risk aversion, $u_t'' < 0$,
v_t = market value at t of a random quantity q_T of commodity to be delivered at T.

A convenient and meaningful route to the definition of testable hypotheses for futures prices is as follows. First, Proposition 2 of Cox, Ingersoll and Ross (1981) defines the futures price in terms of a contract which will pay amounts that depend on a sequence of random one period interest rates, as follows. The futures price at t, F_t, is the value at t of a contract which will pay at T the amount $S_T \prod_{k=t}^{T-1} R_k$. While this proposition defines the futures price explicitly, it does so in terms of ex post interest rates. The next step is to introduce a present value form of the futures price based on expectations at t. What follows is a single commodity discrete time analogue of a continuous time analysis given by Richard and Sundaresan (1981).

Given

$$U_t = E_t \sum_{k=t}^{\infty} \rho^k u(C_k), \tag{1}$$

Theorem 1 of Richard and Sundaresan states that

$$v_t = E_t \left[\rho^{T-t} \frac{u'(C_T)}{u'(C_t)} S_T q_T \right]. \tag{2}$$

First order conditions for maximum expected utility lead to

$$1 = E_t \left[\rho^{T-t} \frac{u'(C_T)}{u'(C_t)} \prod_{k=t}^{T-1} R_k \right], \tag{3}$$

[4] The Canadian dollar spot delivery lag is one day and for contracts up to and including June 1980 the futures contracts expired on the Tuesday, before a Wednesday delivery date, instead of the Monday.

which is the present value of one dollar rolled over in one period bonds, Application of the present value operator in (3) to futures gives

$$F_t = E_t\left[\rho^{T-t}\frac{u'(C_T)}{u'(C_t)}S_T\prod_{k=t}^{T-1}R_k\right]. \tag{4}$$

Richard and Sundaresan decompose (4) in a way which turns out to be more complicated than is necessary here. Hodrick and Srivastava (1985) proceed more directly from (4) to obtain

$$F_t = (E_tS_T)E_t\left[\rho^{T-t}\frac{u'(C_T)}{u'(C_t)}\prod_{k=t}^{T-1}R_k\right] + cov_t\left[S_T,\rho^{T-t}\frac{u'(C_T)}{u'(C_t)}\prod_{k=t}^{T-1}R_k\right], \tag{5}$$

$$F_t = E_tS_T + cov_t\left[S_T,\rho^{T-t}\frac{u'(C_T)}{u'(C_t)}\prod_{k=t}^{T-1}R_k\right], \tag{6}$$

by use of (3). The covariance term in eq. (5) may be defined to be a risk premium P_t, giving

$$F_t = E_tS_T + P_t. \tag{7}$$

Although (6) relates the current futures price to the expected spot price at T and a risk premium, its form does not allow exploitation of the main advantage of futures price data over forward price data, as discussed in the previous section. Full use of the sequence of daily futures prices is better tied to Proposition 1 of Hodrick and Srivastava, which says that the futures price at t is equal to the expected futures price at $t+1$ plus a risk premium. To obtain this result, first rewrite (4) as

$$F_t = E_t\left[\rho\frac{u'(C_{t+1})}{u'(C_t)}R_t\rho^{T-t-1}\frac{u'(C_T)}{u'(C_{t+1})}S_T\prod_{k=t+1}^{T-1}R_k\right], \tag{8}$$

and then decompose it into

$$\begin{aligned} F_t = {} & E_t\left[\rho\frac{u'(C_{t+1})}{u'(C_t)}R_t\right]E_t\left[\rho^{T-t-1}\frac{u'(C_T)}{u'(C_{t+1})}S_T\prod_{k=t+1}^{T-1}R_k\right] \\ & + cov_t\left[\rho\frac{u'(C_{t+1})}{u'(C_t)}R_t,\rho^{T-t-1}\frac{u'(C_T)}{u'(C_{t+1})}S_T\prod_{k=t+1}^{T-1}R_k\right]. \end{aligned} \tag{9}$$

Application of iterative expectations gives

$$F_t = E_tF_{t+1} + cov_t\left[\rho\frac{u'(C_{t+1})}{u'(C_t)}R_t,\rho^{T-t-1}\frac{u'(C_T)}{u'(C_{t+1})}S_T\prod_{k=t+1}^{T-1}R_k\right]. \tag{10}$$

If the covariance term in (10) is non-zero, it represents a risk premium p_t. If, instead, this covariance is zero, the martingale hypothesis [Samuelson (1965)] holds. Therefore, the two hypotheses are

$H_0: F_t = E_tF_{t+1}$,

$H_1: F_t = E_tF_{t+1} + p_t$,

3. A test equation for daily changes in futures prices

3.1. The martingale hypothesis

Under the null hypothesis of unbiasedness, futures prices follow a martingale process which implies that changes in those prices have the property of a fair game. That is, the changes are forecast errors (innovations or news) and as such they are orthogonal to information available when the forecast was made. [5]

To obtain a test equation for the martingale hypothesis, define f_t as the logarithm of F_t and ϵ_t as the forecast error with respect to the the one day ahead futures price, that is,

$$f_t - E_{t-1} f_t = \epsilon_t. \tag{11}$$

With H_0 expressed in its fair game form, $E_{t-1} f_t - f_{t-1} = 0$, the test equation is

$$f_t - f_{t-1} = x_{t-1}\gamma + \epsilon_t, \tag{12}$$

in which x_{t-1} is a vector of variables from the information set at time $t-1$, and ϵ_t is the error term. Under the null hypothesis that the futures price follows a martingale process the change in the futures price from $t-1$ to t should be orthogonal to the information set I_{t-1} and therefore to any subset of I_{t-1}. One such subset, the lagged rate of change in the futures price, corresponds to what Roberts (1967) referred to as a weak form efficiency test. In this case, the test equation can be written as

$$f_t - f_{t-1} = \gamma_0 + \gamma_1 (f_{t-1} - f_{t-2}) + \epsilon_t. \tag{13}$$

Under the null hypothesis, the regression coefficients γ_0 and γ_1 equal zero while the error ϵ_t has mean zero and is not autocorrelated. Note that the martingale hypothesis does not imply that the errors are homoskedastic.

On occasion, in semi-strong form tests, other variables from the available information set, in addition to $f_{t-1} - f_{t-2}$, will be included in the vector x_{t-1} in (12). In particular, variables proxying trading day effects will be included in the test equations used for the daily data. Day of the week dummy variables, chosen on the basis of the analysis in McCurdy and Morgan (1986), are as follows: pre and post 1981 10 1 for Mondays, M_1 and M_2; and pre 1981 10 1 for Wednesdays, W_1. In this case, the test equation (13) becomes

$$f_t - f_{t-1} = \gamma_0 + \gamma_1 (f_{t-1} - f_{t-2}) + \gamma_2 M_1 + \gamma_3 M_2 + \gamma_4 W_1 + \epsilon_t. \tag{14}$$

Further variables, such as $(\log S_{t-1} - \log S_{t-2})$, are included in x_{t-1} during diagnostic testing for omitted variables.

Hodrick and Srivastava (1985) use a different transformation to induce stationarity. That is, their dependent variable is $(F_t - S_{t-1})/S_{t-1}$ and their independent variable is $(F_{t-1} - S_{t-1})/S_{t-1}$. We extend this to take account of the day of the week patterns in price, as in

$$(F_t - S_{t-1})/S_{t-1} = \gamma_0 + \gamma_1 (F_{t-1} - S_{t-1})/S_{t-1} + \gamma_2 M_1 + \gamma_3 M_2 + \gamma_4 W_1 + \epsilon_t. \tag{15}$$

In section 5 below we compare the results from (14) and (15).

[5] As mentioned above, this assumes that any institutionally imposed limits on daily price changes are not effective.

As discussed above, time-varying volatility is a feature of the changes in price data that cannot be ignored. Given the results of McCurdy and Morgan (1985), in this paper we model the forecast errors ϵ_t as a GARCH process. Then, from (12), $f_t - f_{t-1}$, has conditional mean $x_{t-1}\gamma$ and conditional variance

$$h_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i \epsilon_{t-i}^2 + \sum_{j=1}^{p} \beta_j h_{t-j} + \phi z_{t-1}^2, \tag{16}$$

where z_{t-1} is an element of x_{t-1}. The conditional variance is modeled as a linear function of the last q innovations or forecast errors, and the last p conditional variances. In the optimization, nonnegativity constraints are imposed on the parameters of (16). As described in the empirical results below, the appropriate choice for the values of p and q is 1. The final term in (16) is used only in tests for omitted variables.

3.2. *An alternative hypothesis: time-varying risk premia*

Of course, other regressors from the information set at $t-1$ could be included in (14). However, it should be clear that empirical evidence that some coefficient γ is statistically different from zero does not imply rejection of the rational expectations or efficient market hypothesis since this is just the first part of a joint hypothesis, and there are other possible choices for the second part, such as a model including an explicit time-varying risk premium. We allow for a possibility of this sort in the alternative hypothesis H_1.

Empirical testing of H_1 must be indirect because the covariance term in (10) contains unobservable utility function derivatives. Nevertheless, if the alternative hypothesis is true, it should be possible to find a regressor which proxies the premium so that its estimated coefficient, in a test equation such as (12), would be non-zero. To avoid inclusion of confounding variables which may introduce spurious correlations, it is useful to specify a highly restrictive version of a model of the premium. One such restrictive choice with this purpose is the GARCH-M specification, in which $f_t - f_{t-1}$ given I_{t-1} has conditional mean $x_{t-1}\gamma + \theta h_t^{1/2}$, with h_t defined by (16). If there is a positive association between systematic or covariance risk and variance, it is possible that the GARCH-M specification proxies changes in systematic risk. However, the strength of the association for a particular asset or, in this case, currency, will determine the usefulness of the proxy.

4. Diagnostic tests

Finding a test equation which is not rejected by the data, in the sense that it survives a comprehensive battery of diagnostic checks, reduces the danger of making inferences that are invalid because of inconsistent estimates of parameters or standard errors. The test equations (14) and (15) are subjected to two types of tests of misspecification. Table 2 reports the results of several diagnostic checks on the standardized residuals, while table 3 reports the outer product of the gradient (OPG) variant [see Davidson and MacKinnon (1983)] of the Lagrange multiplier (LM) test statistics for omitted variables [proposed by Godfrey and Wickens (1981), (1982)]. The test statistic and the marginal level of significance (p-value) for its asymptotic distribution under the null hypothesis, the chi-square, are reported in each case.

In particular, the diagnostic tests include

(i) a first order autocorrelation test, as in Godfrey (1978), with the test statistic G computed as the product of the number of observations, N, and R^2 from the artificial regression;

Table 1
Coefficient estimates for daily futures data, 1974–1983. [a]

Currency data [b]	γ_0	γ_1	M_1 γ_2	M_2 γ_3	W_1 γ_4	α_0	α_1	β_1
BP	−0.0306	0.015	0.084	−0.104	0.144	0.0060	0.105	0.886
	(0.0119)	(0.021)	(0.024)	(0.053)	(0.023)	(0.0014)	(0.014)	(0.014)
BP H&S [c]	−0.0587	0.924	0.078	−0.016	0.136	0.0059	0.105	0.887
	(0.141)	(0.017)	(0.025)	(0.055)	(0.025)	(0.0013)	(0.014)	(0.013)
CD	−0.0108	0.054				0.0003	0.078	0.922
	(0.0048)	(0.025)				(0.0002)	(0.012)	(0.012)
CD H&S	−0.0127	0.931				0.0003	0.078	0.921
	(0.0050)	(0.027)				(0.0002)	(0.013)	(0.013)
DM	−0.0310	−0.044		−0.263	0.150	0.0148	0.118	0.867
	(0.0165)	(0.021)		(0.070)	(0.035)	(0.0033)	(0.015)	(0.015)
DM H&S	0.0202	0.927		−0.289	0.143	0.0139	0.119	0.867
	(0.0204)	(0.022)		(0.083)	(0.035)	(0.0030)	(0.015)	(0.015)
JY	−0.0030	−0.006	0.112	−0.127		0.0040	0.064	0.930
	(0.0163)	(0.026)	(0.043)	(0.060)		(0.0019)	(0.012)	(0.014)
JY H&S	0.0383	0.922	0.105	−0.123		0.0053	0.074	0.918
	(0.0204)	(0.025)	(0.045)	(0.060)		(0.0022)	(0.013)	(0.015)
SF	−0.0159	−0.008		−0.121	0.100	0.0032	0.079	0.917
	(0.0130)	(0.020)		(0.060)	(0.029)	(0.0012)	(0.012)	(0.012)
SH H&S	0.0253	0.939		−0.114	0.102	0.0033	0.079	0.917
	(0.0175)	(0.019)		(0.063)	(0.029)	(0.0013)	(0.012)	(0.012)

[a] Standard errors are shown in parenthesis.
[b] For the Canadian dollar and the Japanese yen the time period is 1977–1983.
[c] The H&S equations use the Hodrick and Srivastava (1985) transformation of the futures prices, that is $(F_t - S_{t-1})/S_{t-1}$ and $(F_{t-1} - S_{t-1})/S_{t-1}$ rather than $(\log F_t - \log F_{t-1})$ and $(\log F_{t-1} - \log F_{t-2})$ for the dependent variable and the regressor corresponding to γ_1 respectively.

(*ii*) the Ljung–Box (1978) version of the portmanteau test on the first 10 lags of the autocorrelation function, $Q(10)$;
(*iii*) and (*iv*) the Kiefer–Salmon (1983) tests for nonnormality with respect to skewness (SK) and kurtosis (KU);
(*v*) White's (1982) information matrix test (calculated with the method suggested by Chesher (1983)) which may have power against nonnormality [see, for example, Newey (1985)] and/or time-varying parameters [Chesher (1984)];and
(*vi*) The Ljung–Box test on the first 10 lags of the autocorrelation function of the squared standardized residuals [see, for example, McLeod and Li (1983)].

The p-values are reported for each diagnostic test statistic without adjustments for the fact that several tests are being jointly considered. That is, tests designed to detect a particular problem may also pick up other statistical deficiencies [see, for example, Davidson and MacKinnon (1985)]. For example, the Ljung–Box and the Godfrey test statistics may reflect remaining heteroskedasticity as well as autocorrelation. Pagan and Hall (1983) discuss the necessary conditions under which diagnostic tests, such as those used in this paper, would be additive. Since our test equations generally involve a lagged dependent variable, those conditions would not be satisfied here. Given the complexity of determining the joint probabilities, each diagnostic is treated as if it were calculated in isolation. Clearly, the p-values should be interpreted as being suggestive rather than taken literally.

Table 2
Diagnostic checks for the daily test equations. [a]

Currency data	L	G	$Q(10)$	SK	KU	I	$Q^*(10)$
BP	247.6	7.8	10.3	11.4	2115.	24.4 [b]	6.6
		(0.005)	(0.415)	(0.003)	(0.000)	(0.001)	(0.763)
BP H&S	265.5	7.5	15.4	6.8	2260.	20.8	6.1
		(0.006)	(0.118)	(0.009)	(0.000)	(0.004)	(0.807)
CD	1661.6	1.3	11.2	0.1	60.	0.7	12.5
		(0.254)	(0.342)	(0.752)	(0.000)	(0.873)	(0.253)
CD H&S	1662.9	9.9	19.5	0.0	72.	0.9	12.3
		(0.002)	(0.034)	(0.999)	(0.000)	(0.825)	(0.265)
DM	−689.5	7.3	45.5	14.0	102.	17.7	12.6
		(0.007)	(0.000)	(0.000)	(0.000)	(0.013)	(0.247)
DM H&S	−673.4	0.4	33.8	19.5	112.	32.8	9.9
		(0.527)	(0.000)	(0.000)	(0.000)	(0.000)	(0.449)
JY	−157.9	3.9	24.8	1.7	35.	19.0	18.2
		(0.048)	(0.006)	(0.192)	(0.000)	(0.008)	(0.052)
JY H&S	−172.0	1.5	20.1	3.8	47.	19.5	16.5
		(0.221)	(0.028)	(0.051)	(0.000)	(0.002)	(0.086)
SF	−191.4	2.5	12.1	9.4	67.	7.2	11.0
		(0.114)	(0.278)	(0.002)	(0.000)	(0.408)	(0.358)
SF H&S	−212.1	4.1	10.2	15.5	80.	14.9	11.3
		(0.043)	(0.423)	(0.000)	(0.000)	(0.037)	(0.335)

[a] L is the log of the likelihood function; G, the Godfrey (1978) test for (first order) serial correlation; $Q(10)$, the Ljung–Box (1978) portmanteau test on the first 10 lags of the autocorrelation function; SK and KU, the Kiefer-Salmon (1983) tests for skewness and excess kurtosis; I, the White (1982) information test; and $Q^*(10)$, the Ljung–Box test on the squared standardized residuals. p values, for the chi-square distribution, are shown in parenthesis.

[b] This test statistic corresponds to a BP test equation without the M_1 regressor due to size limitations for the information matrix test routine.

The OPG Lagrange multiplier tests are particularly useful for evaluating the specification against particular variables excluded from the mean or variance functions. These test statistics, reported in tables 3 and 5, are calculated as NR^2 where R^2 is the explained sum of squares from a regression of a vector of ones on the matrix of scores for the locally equivalent alternative model evaluated under the null hypothesis [see, for example, Engle, Lilien and Robins (1986) for a similar application]. The OPG LM tests are also useful in determining whether the lower bound of zero on the parameters of the variance function is a restriction which is incompatible with the data.

5. Daily data analysis

5.1. Daily data definition

In this section we report the results of GARCH estimation of (14) with daily data for five currencies; the British pound, the Canadian dollar, the Deutschmark, the Japanese yen, and the Swiss franc. For the British pound, the Deutschmark, and the Swiss franc, the data for the dependent variable, the rate of change of futures price, extend from 1974 to 1983. Although the CRFM price data are available for some months in 1973 the starting dates vary from currency to currency, and we decided to use a common starting data of 1974 1 2 for the first observation. For the Canadian dollar

Table 3
Lagrange multiplier test statistics for omitted variables: daily test equations. [a]

Currency data	$h_t^{1/2}$	$f_{t-2} - f_{t-3}$	$s_{t-1} - s_{t-2}$	W_2	W_1	M_1	M_2
in mean:							
BP	1.36	0.72	12.14	0.01			
	(0.244)	(0.396)	(0.001)	(0.920)			
CD	3.64	1.09	2.50	0.67	0.16	0.25	0.38
	(0.056)	(0.296)	(0.114)	(0.413)	(0.689)	(0.617)	(0.538)
DM	2.35	7.89	13.60	0.03		0.02	
	(0.125)	(0.005)	(0.000)	(0.862)		(0.888)	
JY	2.44	0.09	2.06	1.12	1.37		
	(0.118)	(0.764)	(0.151)	(0.290)	(0.242)		
SF	0.75	1.92	8.53	0.05		0.86	
	(0.386)	(0.166)	(0.003)	(0.823)		(0.354)	
in variance:							
	ϵ^2_{t-5}	M_2	W_1	$(f_{t-1} - f_{t-2})^2$			
BP	0.23	1.43	1.02	0.75			
	(0.632)	(0.232)	(0.313)	(0.386)			
CD	0.14						
	(0.708)						
DM	3.00	2.03	2.78	1.92			
	(0.083)	(0.154)	(0.095)	(0.166)			
JY	3.75	0.30	1.48 [b]	3.04			
	(0.053)	(0.584)	(0.224)	(0.081)			
SF	3.44	0.52	0.66	0.33			
	(0.064)	(0.471	(0.471)	(0.566)			

[a] p-values, for the chi-square distribution with one degree of freedom, are in parenthesis below the test statistics.
[b] Implies that the variable is M_1 rather than W_1.

and the Japanese yen we chose the common starting date of 1977 1 7. The common ending date is 1983 12 30. We always take the date for the outstanding contract with the shortest time to maturity of the contracts maturing in March, June, September, and December. Data for the first and second lagged observations of the change in log futures price and for the first lag of the change in log spot price are computed. Normally this is a straightforward computation, but around the time of expiration of a contract care is needed to ensure that each of the variables $f_t - f_{t-1}$, $f_{t-1} - f_{t-2}$, and $f_{t-2} - f_{t-3}$ are calculated from prices corresponding to the appropriate, common, contract. The dependent variable and the regressor are scaled before the maximum likelihood estimation is undertaken. This scaling takes the form of multiplication by a factor, typically 150 in these analyses, and it is necessary if certain conditions recommended for the Numerical Algorithms Group (1983) system of numerical optimisation are to be satisfied. A gradient method is used.

For the spot price we use the series Spot2 given in the CRFM term structure file. In the Hodrick and Srivastava transformation, if no spot price is given in the file for $t-1$ we substitute S_{t-2} for S_{t-1} in (15).

Limit moves in futures price are a potential problem. In the earlier years of the sample period the limits imposed for settlement purposes were quite tight for certain currencies and the futures price often moved the maximum allowed. The price is then an administered rather than an equilibrium price and the price change computed from such a price reflects only part of the new information received during the day. Problems that could arise include positive autocorrelation induced by the partial adjustment of price. Although this is potentially serious in tests of the martingale hypothesis, when we purge the data of observations affected by limit moves the results are not substantially different. This subsection deals with the data for which the limit move problems are ignored and the next subsection deals with the subsample remaining after action has been taken to avoid the limit moves.

5.2. Daily data including limit moves in futures prices

In table 1 the regression coefficient estimates for eqs. (14) and (15) are summarised for each of the currencies. Two sets of estimates are shown for each currency; the first set is for eq. (14) in which the stationarity transformation is the difference in log price, and the second set is for eq. (15) in which the Hodrick and Srivastava stationarity transformation is adopted.

The choice of which dummy variables to include is determined in an analysis of day of the week patterns in futures price described by McCurdy and Morgan (1986). It is important to distinguish between data before and after 1981 10 1, which was the date of an acceleration of the clearing process for U.S. dollar denominated cheques for purchases of foreign currency. As explained in McCurdy and Morgan (1986), and demonstrated in table 3, some of the dummy variables are unimportant for certain currencies.

Under the martingale hypothesis, γ_0 is 0, while γ_1 is 0 in eq. (14) and 1 in the Hodrick and Srivastava eq. (15). This null hypothesis is rejected in all five currencies with the Hodrick and Srivastava equation, consistent with their own empirical results [Hodrick and Srivastava (1985, Table 6)]. For eq. (14), the null hypothesis for γ_1 is retained in three currencies, the British pound, the Japanese yen, and the Swiss franc. It seems possible that the Hodrick and Srivastava form of the equation is rejecting the null hypothesis too frequently on account of a confounding influence of the spot price in the construction of the dependent variable and regressor.

Table 2 shows that it cannot be claimed that eq. (14) represents the data adequately, since the null hypothesis of no autocorrelation is rejected by either the Godfrey test or the Ljung–Box test for the British pound, the Deutschmark, and the Japanese yen. To what extent these results reflect the departures from normality detected by the Kiefer–Salmon statistics, particularly with respect to kurtosis, is unknown. Note that the information matrix test rejects eq. (14) in the same three currencies. The test for remaining heteroskedasticity, as given by the portmanteau statistic for autocorrelation of the standardised squared residuals in the final column of table 2 tends to support the view that the problem is with eq. (14) itself rather than with the specification of the variance function. This idea is best explored with the OPG LM test for omitted variables, summarised in table 3.

The two panels of table 3 give the OPG LM test statistics and p-values for additional variables constrained to have zero coefficients. The first panel examines such constraints for variables potentially influencing the mean and the second panel evaluates more elaborate forms of the variance function. For the mean, the second lag in the rate of change of futures price and the first lag in the rate of change of the spot price are subsets of the information set I_{t-1}. Under H_0 they should be irrelevant. Table 3 suggests the contrary for the British pound, the Deutschmark and the Swiss franc and H_0 is rejected for these currencies. Nor can the deficiencies of eq. (14) be attributed to inadequate specification of the variance function. As mentioned above, the McLeod–Li test results in table 2 allowed the homoskedasticity hypothesis to be retained for the standardized GARCH residuals. The second panel of table 3 provides support for the view that the variance function specification is sound, since all omitted variables examined turn out to have insignificant OPG LM test statistics. In other words, an improved form of the variance function has not been found. In

Table 4
Results when observations for which daily price change limits were effective are removed. [a]

Currency data	coefficient γ_1	test statistics G	$Q(10)$	SK	KU	OPG LM $s_{t-1}-s_{t-2}$
BP	0.0276	9.3	11.9	71.4	1438.	0.01
	(0.0216)	(0.002)	(0.292)	(0.000)	(0.000)	(0.920)
CD	0.0501	0.59	9.7	1.9	39.	0.95
	(0.0238)	(0.442)	(0.467)	(0.168)	(0.000)	(0.330)
DM	−0.0711	6.45	45.4	5.3	46.	5.83
	(0.0214)	(0.011)	(0.000)	(0.021)	(0.000)	(0.016)
JY	−0.0239	1.88	30.2	7.6	15.	0.01
	(0.0244)	(0.170)	(0.001)	(0.006)	(0.000)	(0.920)
SF	−0.0417	0.81	14.4	4.2	29.	0.22
	(0.0217)	(0.368)	(0.156)	(0.040)	(0.000)	(0.639)

[a] Standard errors are shown in parenthesis below the coefficient; p-values, for the chi-square distribution with one degree of freedom are in parenthesis below the test statistics.

summary, the martingale hypothesis is rejected with daily data in all five currencies: by the weak form test for the CD and DM; by the tests for autocorrelation for the BP, DM and JY; and, in semistrong form tests, by correlation of the rate of change in futures price with variables in the available information set for the BP, DM and SF. In the next subsection we examine the possibility that the rejection could be attributed to problems caused by limit moves in futures prices.

5.3. Daily data avoiding limit moves in futures prices

Other things being equal, a limit move will tend to be followed by another price change in the same direction. There are many ways of treating limit moves in tests of H_0 for futures prices. One approach is to throw out any observation for which either $F_t - F_{t-1}$ or $F_{t-1} - F_{t-2}$ is a limit move. This method will avoid the problem of the resulting contribution of positive autocorrelation. Table 4 gives selected statistics for the analyses corresponding to the above approach. In the periods analysed, the number of observations lost because of limit moves was greatest for the Swiss franc (199 out of 2517), and the Deutschmark (109 out of 2511). The Japanese yen lost 62 out of 1750, the British pound 30 out of 2499 and the Canadian dollar 22 out of 1760.

The estimate of γ_1 in table 4 is more negative than in table 1 for the two currencies with the most limit moves removed, the Swiss franc and the Deutschmark. This is as expected: avoiding limit moves will remove some positive dependence. However, one or both of the autocorrelation diagnostics, the Godfrey test and the Ljung–Box test, still suggest rejection of the null hypothesis for the British pound, the Deutschmark, and the Japanese yen. Since this is the same group of currencies showing evidence of autocorrelation in table 2, removal of the limit moves does not change the overall conclusion about autocorrelation. It does, of course, reduce the kurtosis.

Removal of the limit moves sheds light on the role of the lagged rate of change of spot price in the OPG LM tests. In contrast to table 3, where the null hypothesis was rejected in these tests for three of the five currencies, the OPG LM test statistics for the lagged rate of change of spot price in table 4 are not significant except for the Deutschmark. This shows that the rejection of the martingale hypothesis in table 3 is partly attributable to problems associated with limit moves. When the lagged change of the futures price, $F_{t-1} - F_{t-2}$, is constrained by the limit rules, S_{t-1} and therefore the lagged rate of change of the spot price, $\log S_{t-1} - \log S_{t-2}$, contains information F_{t-1} cannot fully reflect.

In summary, after removal of the limit moves in futures price, the martingale hypothesis is rejected with daily data in four of the five currencies: by the weak form test for the CD and DM; by the tests for autocorrelation for the BP, DM and JY; and, in semi-strong form tests, by correlation of the rate of change in futures price with the lagged rate of change of spot price for the DM. The Swiss franc is the one currency which allows the null hypothesis to be retained. Of course, a strict interpretation of the martingale hypothesis also rules out systematic patterns in futures price related to the day of the week. Given the evidence of these patterns, it is clear that further testing of the martingale hypothesis with weekly rather than daily data might be informative.

6. Weekly data analysis

6.1. Weekly data definition

GARCH estimation of (13) with weekly data is described in this section. The sample period is the same as in section 5; for the British pound, the Deutschmark, and the Swiss franc, the data for the dependent variable, the rate of change of futures price, extend from the first week of 1974 to the last of 1983. For the Canadian dollar and the Japanese yen the data extend from the second week of 1977 to the last of 1983. The rate of change of price for a given week is normally calculated from Wednesday prices. If Wednesday data are missing on account of a holiday, the week is extended so that it ends on a Thursday, or even a Friday if necessary, and the following week then generally covers less than seven calendar days.

The dependent variable and the regressor are scaled by multiplication by a factor of 10 before the maximum likelihood estimation is undertaken.

Although limit moves in futures price again constitute a potential problem for tests of the martingale hypothesis with weekly data, they are less important than in daily data because they are relevant only if they occur at the end of a given week. When we remove the influence of limit moves from the weekly data the results are generally changed only slightly. This subsection deals with the data for which the limit moves are ignored. The next subsection deals with the subsample remaining after action has been taken to avoid the limit moves by ending the week on Thursday, or Friday if necessary, when the Wednesday settlement price corresponds to a limit move.

6.2. Weeks defined without reference to limit moves in futures prices

Table 5 summarises the regression coefficient estimates obtained from weekly data for the five currencies. Under the martingale hypothesis, γ_0 and γ_1 are 0 in eq. (13). The null hypothesis is retained in four out of the five currencies and rejected for the Deutschmark only. With the exception of the Deutschmark, these results are strikingly different from those obtained with daily data. Tables 5 and 6 suggest that eq. (13) represents the data adequately, except with respect to the distributional assumption of conditional normality. The null hypothesis of no autocorrelation is retained by both the Godfrey test and the Ljung–Box test for all currencies. It is not known to what extent the departures from nonnormality detected by the Kiefer–Salmon statistics, particularly with respect to kurtosis, should qualify the inferences made about H_0. Note that the information matrix test for model misspecification, which should have some power against departures from normality, does not reject eq. (13) for any currency. The results of the test for remaining heteroskedasticity, as given by the portmanteau statistic for autocorrelation of the squared standardised residuals, in the final column of table 6, are all consistent with the hypothesis of homoskedasticity. This suggests that the distributional problems detected by the Kiefer–Salmon tests are not associated with the specification

Table 5
Coefficient estimates for weekly futures data, 1974–1983. [a]

Currency data [b]	γ_0	γ_1	α_0	α_1	β_1
BP	−0.0014	0.071	0.0015	0.089	0.811
	(0.0048)	(0.045)	(0.0006)	(0.031)	(0.060)
CD	−0.0048	0.059	0.0001	0.075	0.907
	(0.0028)	(0.054)	(0.0001)	(0.027)	(0.030)
DM	0.0004	0.147	0.0003	0.092	0.892
	(0.0047)	(0.045)	(0.0002)	(0.028)	(0.034)
JY	−0.0002	0.080	0.0021	0.096	0.814
	(0.0089)	(0.054)	(0.0028)	(0.062)	(0.174)
SF	−0.0012	0.079	0.0003	0.125	0.868
	(0.0057)	(0.044)	(0.0002)	(0.030)	(0.028)

[a] Standard errors are shown in parenthesis.
[b] For the Canadian dollar and the Japanese yen the time period is 1977–1983.

Table 6
Diagnostic checks for the weekly test equations. [a]

Currency	L	G	$Q(10)$	SK	KU	I	$Q^*(10)$
BP	836.4	0.6	16.7	26.1	152.	1.9	4.5
		(0.439)	(0.081)	(0.000)	(0.000)	(0.593)	(0.922)
CD	865.6	3.6	9.2	3.9	29.	4.3	4.3
		(0.058)	(0.513)	(0.048)	(0.000)	(0.231)	(0.933)
DM	851.3	1.3	11.0	0.0	14.	1.5	12.8
		(0.254)	(0.358)	(0.999)	(0.000)	(0.682)	(0.235)
JY	507.2	0.8	13.5	6.3	3.	2.8	8.8
		(0.371)	(0.197)	(0.012)	(0.083)	(0.424)	(0.551)
SF	740.4	0.7	14.3	1.9	13.	2.4	5.5
		(0.403)	(0.160)	(0.168)	(0.000)	(0.494)	(0.855)

[a] p-values, for the chi-square distribution, are shown in parenthesis. See table 2 for the key to the tests.

of the variance function. [6] In summary, with weekly data, if it is assumed that the departures from the distributional assumption of conditional normality detected by the Kiefer–Salmon tests are not sufficiently serious to interfere with the inferences based on estimates obtained by maximum likelihood estimation, the Deutschmark is the only currency that is not well described by the martingale hypothesis applied to eq. (13). The next subsection reevaluates these conclusions after care has been taken to avoid defining the end of a week to coincide with a limit move.

6.3. *Weeks defined to avoid ending on a limit move in futures price*

In this subsection we evaluate the role of limit moves in weekly data. This we do by defining the weeks so that they do not end on a limit move. Tables 7 and 8 summarise the analysis of the new series. In general, the contents of these tables are very similar to those of tables 5 and 6 respectively, since it is only for the Deutschmark and the Swiss franc that the limit moves are at all frequent. The

[6] Although Bollerslev (1985) allows for conditionally Student distributed errors by including an error term in the conditional variance equation.

Table 7
Coefficient estimates for weekly futures data: variable week to avoid daily price limit changes, 1974–1983. [a]

Currency data [b]	γ_0	γ_1	α_0	α_1	β_1
BP	−0.0012	0.064	0.0015	0.089	0.811
	(0.0048)	(0.046)	(0.0006)	(0.031)	(0.059)
CD	−0.0048	0.059	0.0001	0.075	0.907
	(0.0028)	(0.054)	(0.0001)	(0.027)	(0.030)
DM	−0.0005	0.085	0.0006	0.143	0.824
	(0.0049)	(0.047)	(0.0004)	(0.037)	(0.051)
JY	0.0010	0.049	0.0009	0.071	0.894
	(0.0080)	(0.056)	(0.0008)	(0.027)	(0.046)
SF	−0.0006	0.076	0.0003	0.138	0.856
	(0.0057)	(0.047)	(0.0002)	(0.032)	(0.029)

[a] Standard errors are shown in parenthesis.
[b] For the Canadian dollar and the Japanese yen the period is 1977–1983.

main conclusions to be drawn from tables 7 and 8 is that the martingale hypothesis is retained for all five currencies and that eq. (13) also survives all the diagnostic tests except the Kiefer–Salmon tests for nonnormality. These conclusions are very different from those drawn from the analysis of daily data and it seems probable that the patterns in price by day of the week account for the contrasting inferences of sections 5 and 6.

The remaining task is to examine the model for the possible relevance of omitted variables. Table 9 shows the OPG LM test statistics and associated p-values for two additional variables constrained to have zero coefficients. The first variable is the lagged rate of change of the spot price which, as a subset of the information set I_{t-1}, should be irrelevant under H_0. Table 9 confirms that this variable is not relevant to the futures pricing equation, since the OPG LM test statistics are insignificant for all five currencies. This is a useful result, since in section 5 the corresponding tests with daily data rejected the null hypothesis for the BP, the DM and the SF before limit moves were removed and for only the DM afterwards. Since limit moves in weekly data are generally less important than in daily data, retention of the null hypothesis in all five currencies strengthens the view that the lagged rate of change of spot price is irrelevant except when the limit rules prevent the futures price from reflecting the information available at $t-1$.

Table 8
Diagnostic checks for the variable week test equation. [a]

Currency	L	G	$Q(10)$	SK	KU	I	$Q^*(10)$
BP	836.6	0.5	17.3	25.5	152.	1.4	4.7
		(0.480)	(0.068)	(0.000)	(0.000)	(0.706)	(0.910)
CD	865.6	3.6	9.2	3.9	29.	4.3	4.3
		(0.058)	(0.513)	(0.048)	(0.000)	(0.231)	(0.933)
DM	839.1	3.8	16.9	0.9	23.	1.0	8.8
		(0.051)	(0.077)	(0.343)	(0.000)	(0.801)	(0.551)
JY	500.9	0.0	14.1	3.7	8.	0.9	5.5
		(0.999)	(0.168)	(0.054)	(0.005)	(0.993)	(0.855)
SF	734.3	2.8	17.9	4.9	18.	2.1	8.9
		(0.094)	(0.057)	(0.027)	(0.000)	(0.552)	(0.542)

[a] p-values, for the chi-square distribution, are shown in parenthesis. See table 2 for the key to the tests.

Table 9
Lagrange multiplier test statistics for omitted variables: weekly test equations. [a]

Currency data	For table 5		For table 7	
	$s_{t-1} - s_{t-2}$	$h_t^{1/2}$	$s_{t-1} - s_{t-2}$	$h_t^{1/2}$
BP	0.10	0.51	0.01	0.46
	(0.752)	(0.475)	(0.920)	(0.498)
CD	1.09	2.85	1.09	2.85
	(0.296)	(0.091)	(0.296)	(0.091)
DM	1.61	5.65	0.98	7.22
	(0.204)	(0.017)	(0.322)	(0.007)
JY	0.02	0.00	0.01	1.25
	(0.888)	(0.999)	(0.920)	(0.264)
SF	3.41	1.84	1.99	2.30
	(0.065)	(0.175)	(0.158)	(0.129)

[a] p-values, for the chi-square distribution with one degree of freedom, are shown below the test statistics.

The second variable examined in table 9 corresponds to the GARCH-M specification instead of GARCH. GARCH-M allows the conditional mean to be an ex ante function of the variance. A significant statistic for the OPG LM test for omitting $h_t^{1/2}$ from the conditional mean is consistent with a particular version of the risk premium hypothesis. These test statistics in table 9 are consistent with the martingale hypothesis except for the Deutschmark. It is not obvious why this result is obtained for the DM for the weekly data. Nevertheless, this is the currency which has rejected the null hypothesis throughout the whole paper. Also, as is clear from the contrast with the quite different results for the same currency in the period 1983–1985 in McCurdy and Morgan (1985), the source of the rejection is in the earlier part of the sample.

In summary, the martingale hypothesis is rejected with weekly data in only one of five currencies. One interpretation of the rejection of the null hypothesis for the Deutschmark in the OPG LM test for omission of $h_t^{1/2}$ from the conditional mean is that there is a time-dependent risk premium, as under H_1.

7. Conclusion

In this paper, we have tested the martingale hypothesis for foreign exchange futures prices. A comprehensive series of diagnostic tests has been carried out. These tests included checks on model specification, residual properties, and tests for omitted variables. The GARCH model representation of conditional heteroscedasticity survived these tests in almost all cases.

It is shown that the martingale hypothesis is rejected for daily data, generally due to the complex pattern in futures price related to day of the week, but is retained in tests with weekly data in four of the five currencies in the sample. In the fifth currency, the Deutschmark, there is evidence which is consistent with the existence of a time-varying risk premium. The conclusions we draw with respect to the martingale hypothesis from the daily data do not differ substantially from those of Hodrick and Srivastava (1985). On the other hand, from our weekly data tests our overall conclusions about the relative merits of the martingale and alternative hypotheses are quite different.

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Biography: Thomas McCURDY is an Assistant Professor in the Department of Economics, Queen's University, Kingston, Ontario, Canada. He has a Ph.D. from the London School of Economics. His research interests include: topics in international finance, particularly intertemporal asset pricing models with time-varying volatility and risk premia; and computable general equilibrium modeling of adjustment to structural shocks such as technical change.

Ieuan MORGAN is Associate Professor of Finance in the School of Business at Queen's University, Kingston, Ontario, Canada. He has a Ph.D. from the University of Chicago. His research interests are empirical testing of models of capital asset pricing and hypotheses about the time series behaviour of financial and speculative securities.