

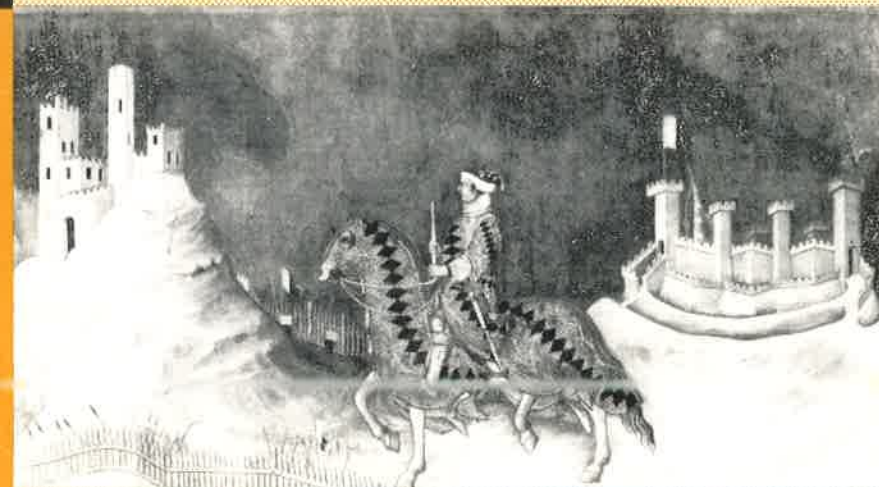
UNIVERSITA' DEGLI STUDI DI SIENA



QUADERNI DEL DIPARTIMENTO
DI ECONOMIA POLITICA

Giulio Cifarelli

**SYSTEMATIC FORWARD FORECAST
ERRORS, EX ANTE REAL INTEREST RATE
DIFFERENTIALS AND EX ANTE REAL
EXCHANGE RATE CHANGES.
AN EMPIRICAL ANALYSIS.**



**QUADERNI DEL DIPARTIMENTO
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Giulio Cifarelli

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REAL EXCHANGE RATE CHANGES.
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Siena, ottobre 1990

1. Introduction*

In this paper we consider the determinants of the (systematic) deviations from the unbiased efficiency hypothesis on forward exchange rate behaviour that have been observed in the 1980s. We expand an approach set out by R.A. Korajczyk (1985) to analyse the pricing of forward exchange rate in three ways:

- (i) we assess directly the relevance of ex ante real interest rate equality across currencies and of the efficient market version of purchasing power parity;
- (ii) we introduce explicitly ex ante real exchange rate changes in the estimation framework;
- (iii) we use a more flexible Generalised Method of Moments estimation approach, which is robust both to conditional heteroskedasticity and to serial correlation of the residuals.

Section 2 describes the economic rationale of the relationships that are to be estimated with the help of a simple model. Section 3 provides an empirical analysis of the behaviour of ex ante real interest rate differentials and of ex ante real exchange rate changes. Section 4 investigates the links between these factors and ex ante forward forecast errors.

The empirical findings suggest that the discrepancies

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between forward and future spot exchange rates are due to ex ante real exchange rate changes in the first half of the decade and - to a lesser extent - to risk premia on nominally riskless bonds in the second half. The determinants of systematic exchange rate forecast errors change over time and from one currency to the other.

2. Forecast Errors, Real Interest Rates and Real Exchange Rate Changes

The unbiased expectations hypothesis on forward exchange rate behaviour holds if

$$(1) E[(s_{t+1} - f_t) | \Phi_t] = 0$$

where s_{t+1} is the logarithm of the spot exchange rate at time $t+1$, f_t is the logarithm of the corresponding one period ahead forward exchange rate set at time t , Φ_t is the information set available at time t and $E(.)$ is the expectations operator.¹

Following R.A. Korajczyk (1985) we set out an estimable expression for equation (1) in terms of expected differences, across countries, in real interest rates and of expected deviations from relative purchasing power parity.

The covered interest parity condition implies that

$$(2) i_{t+1} - i_{t+1}^* = f_t - s_t$$

where i_{t+1} and i_{t+1}^* are the yields (from t to $t+1$) on one period nominally riskless bonds denominated in the home and foreign currencies respectively.

The logarithm of the deviations from purchasing power parity, d_t , can be expressed as

$$(3) d_t = s_t - P_t + P_t^*$$

where P_t and P_t^* are the logarithms of the domestic and foreign price levels. d_t can also be defined as the logarithm of the real exchange rate at time t . The nominal rate of interest from a one period bond can be decomposed as

$$(4) i_{t+1} = E(r_{t+1} | \Phi_t) + E(\Delta P_{t+1} | \Phi_t)$$

$$(5) i_{t+1}^* = E(r_{t+1}^* | \Phi_t) + E(\Delta P_{t+1}^* | \Phi_t)$$

where r_{t+1} and r_{t+1}^* are - respectively - the real returns on default free domestic and foreign bonds maturing at time $t+1$. ΔP_{t+1} and ΔP_{t+1}^* are the domestic and foreign inflation rates from time t to $t+1$.

Combining equations (2), (3), (4) and (5) we can express the (ex post) exchange rate forecast error as

$$(6) (s_{t+1} - f_t) = E[(r_{t+1}^* - r_{t+1}) | \Phi_t] + [\Delta P_{t+1} - E(\Delta P_{t+1} | \Phi_t)] - [\Delta P_{t+1}^* - E(\Delta P_{t+1}^* | \Phi_t)] + \Delta d_{t+1}.$$

Taking expectations, we can define the expected deviations of s_{t+1} from f_t conditional on Φ_t as

$$(7) E[(s_{t+1} - f_t) | \Phi_t] = E[(r_{t+1}^* - r_{t+1}) | \Phi_t] + E[\Delta d_{t+1} | \Phi_t].$$

The right hand side is null and the unbiased expectations hypothesis holds if (i) the ex ante real returns on nominally riskless bonds are equal across countries and (ii) the ex ante rate of change of the real exchange rate is zero.

Condition (i) is generally believed not to hold (see e.g. R.J. Hodrick (1979), R.E. Cumby, M. Obstfeld (1984), F.S. Mishkin (1984.a, 1984.b), J.J. Merrick, A. Saunders (1986) and others). In order to hold the asset denominated in the more risky currency, an investor must be compensated with a risk premium, i.e. a higher expected real return.²

Condition (ii), the efficient market version of the purchasing power parity hypothesis, is usually believed to hold (see e.g. R.J. Rogalski, J.D. Vinso (1977), R. Roll (1979) and M. Adler, B. Lehmann (1983)). It implies that ex ante real returns are independent of the country of residence.³

If we assume a priori - following R.A. Korajczyk (1985), G.W. Evans (1986) and R. Levine (1989) - that condition (ii) holds and that $E[\Delta d_{t+1}|\Phi_t] = 0$, equation (7) becomes

$$(8) \quad E[(s_{t+1} - f_t)|\Phi_t] = E[(r_{t+1}^* - r_{t+1})|\Phi_t].$$

Forecastable deviations from unbiased expectations should be equal to the forecastable difference in real returns across bonds denominated in different currencies. The relevance of this assertion can be assessed empirically by estimating the following relationship

$$(9) \quad (s_{t+1} - f_t) = a_0 + a_1 E[(r_{t+1}^* - r_{t+1})|\Phi_t] + u_{9t+1}$$

where $u_{9t+1} = (s_{t+1} - f_t) - E[(s_{t+1} - f_t)|\Phi_t]$.

The theory would require that $a_0 = 0$, $a_1 = 1$ and $E[u_{9t+1}|\Phi_t] = 0$. Rejection of these restrictions is thus a rejection of the joint hypothesis that risk premia on nominally riskless bonds are the cause of deviations from the unbiased expectations hypothesis and that $E[\Delta d_{t+1}|\Phi_t] = 0$.

If we assume a priori that condition (i) holds and that $E[(r_{t+1}^* - r_{t+1})|\Phi_t] = 0$, equation (7) becomes

$$(10) \quad E[(s_{t+1} - f_t)|\Phi_t] = E[\Delta d_{t+1}|\Phi_t].$$

Expected deviations from unbiased expectations should reflect the forecastable change in the real exchange rate. The empirical relevance of this assertion can in turn be assessed by estimating the following relationship

$$(11) \quad (s_{t+1} - f_t) = b_0 + b_1 E[\Delta d_{t+1}|\Phi_t] + u_{11t+1}$$

where $u_{11t+1} = (s_{t+1} - f_t) - E[(s_{t+1} - f_t)|\Phi_t]$.

Here too the null hypothesis is that $b_0 = 0$, $b_1 = 1$ and $E[u_{11t+1}|\Phi_t] = 0$. Rejection of these restrictions is here a rejection of the joint hypothesis that ex ante real exchange rate changes cause deviations from the unbiased expectations hypothesis and that ex ante real interest rates on nominally riskless bonds are equal across countries.

Neither $E[(r_{t+1}^* - r_{t+1})|\Phi_t]$ nor $E[\Delta d_{t+1}|\Phi_t]$ are observable. However, consistent and asymptotically efficient parameter estimates may be obtained with the help of instrumental variables, following a methodology analysed, among others, by B.T. McCallum (1976), M.R. Wickens (1982) and M.H. Pesaran

(1987, pp. 164 - 170). The GMM estimator approach of L.P. Hansen (1982) becomes relevant when the residuals turn out to be conditionally heteroskedastic and/or serially correlated.

3. Empirical Results: Ex Ante Real Interest Rates Equality and Ex Ante Relative Purchasing Power Parity

We have investigated the behaviour of the exchange rate between the U.S. dollar and four major currencies, the British pound, the German mark, the Japanese yen and the Italian lira over the periods January 1981 - February 1985 and March 1985 - October 1992.⁴ We have used - with the exception of Italy - Eurocurrency interest rates, which are less affected by capital controls.⁵ Consumer price indexes are used to derive the inflation rate series. Data are sampled monthly. The forward and spot exchange rate time series have been provided by the Bank of Italy. The consumer prices and the Italian interbank interest rate time series are published in the O.E.C.D. "Main Economic Indicators". The Eurocurrency time series (the interest rates on three-month deposits labelled in U.S. dollars, in Japanese yens, in German marks and in British pounds in London) are reported in the O.E.C.D. "Financial Statistics".

We investigate at first whether the conditions $E[(r_{t+1}^* - r_{t+1})|\Phi_t] = 0$ and $E[\Delta d_{t+1}|\Phi_t] = 0$ - which are crucial for the choice of unbiased expectations hypothesis model - hold over the two time periods of interest.

Ex ante real interest rate parity across countries (condition (i) of the previous section) implies, using equations (4) and (5)

$$(12) \quad i_{t+1} - E[\Delta P_{t+1}|\Phi_t] = i_{t+1}^* - E[\Delta P_{t+1}^*|\Phi_t].$$

Equation (12) can be expressed in regression format as

$$(13) \quad (\Delta P_{t+1} - \Delta P_{t+1}^*) = \alpha + \beta(i_{t+1} - i_{t+1}^*) + u_{13t+1}$$

where $u_{13t+1} = (\Delta P_{t+1} - \Delta P_{t+1}^*) - E[(\Delta P_{t+1} - \Delta P_{t+1}^*)|\Phi_t]$. A test of the null hypothesis $\alpha = 0, \beta = 1$ is a test of the hypothesis that expected real interest rates are equal across countries. The instrumental variables estimates of this

TABLE 1

$$(13) \quad (\Delta P_{t+1} - \Delta P_{t+1}^*) = \alpha + \beta(i_{t+1} - i_{t+1}^*) + u_{13t+1}$$

	1981:1-1985:2					1985:3-1992:10				
	α	β	R^2	H_0	n	α	β	R^2	H_0	n
U.S./U.K.	-0.002 (-2.70)	0.74 (2.34)	0.02	5.90	0	0.003 (1.46)	1.34 (2.15)	0.002	7.39*	0
U.S./GERMANY	-0.01 (-3.58)	2.55 (3.66)	0.31	11.12*	0	0.0001 (0.01)	0.93 (1.39)	0.002	24.95*	0
U.S./JAPAN	-0.001 (-0.60)	0.60 (1.56)	0.07	2.73	0	0.0002 (0.18)	0.81 (1.34)	0.0002	24.79*	0
U.S./ITALY	-0.003 (-1.72)	0.71 (2.44)	0.17	5.63	0	-0.0001 (-0.02)	0.34 (1.13)	0.001	28.09*	0

n : NUMBER OF AUTOCORRELATION TERMS USED IN COMPUTING THE COVARIANCE MATRIX OF ORTHOGONALITY CONDITIONS.

H_0 : TEST OF THE NULL HYPOTHESIS THAT $\alpha = 0, \beta = 1$. THE TEST STATISTIC IS DISTRIBUTED ASYMPTOTICALLY AS $\chi^2(2)$.

* : SIGNIFICANT AT THE 5% LEVEL.

relationship are set out in table 1. The instruments are the time t differentials between the nominal U.S. interest rate and the nominal interest rates of the four currencies of the sample. We have applied the GMM estimator procedure since conditional heteroskedasticity seems to affect the residuals.⁶ In the presence of serially correlated disturbances the standard GMM estimates will be consistent, but not asymptotically efficient if the weighting matrix is not appropriately modified. We have used therefore the positive semi-definite heteroskedasticity and autocorrelation consistent covariance matrix estimator of W.K. Newey, K.D. West (1987).⁷ The bound on the number of sample autocovariances used to compute the covariance matrix of orthogonality conditions (the weighting matrix) has been assessed with the help of the automatic bandwidth parameter selection procedure set out by D.W.K. Andrews (1991).⁸ (For an analogous approach, using however the prewhitening kernel method of D.W.K. Andrews, J.C. Monahan (1992), see B.E. Hansen (1992).)

The quality of fit declines significantly in the second half of the decade. The estimates suggest that ex ante real interest equality has to be rejected over the period March 1985 - October 1992 for all the countries of the sample. Over the period January 1981 - February 1985 the null hypothesis has to be rejected in the case of Germany only. A significant constant term suggests - however - that the ex ante U.S.-U.K. interest rate differential may have been constant but different from zero.

The hypothesis that $E[\Delta d_{t+1}|\Phi_t] = 0$ implies that the innovations in the real exchange rate, Δd_{t+1} , are serially uncorrelated and that $E[u_{t+1}|\Phi_t] = 0$, i.e. that the real

exchange rate follows a random walk process. We found, however, that Δd_{t+1} could be modelled as a stable function of a constant, of its own lagged value and of the time t forward premium in the period January 1981 - February 1985 (equation (14)), and as a function of its own lagged value and of the lagged forward forecast error in the period March 1985 - October 1992 (equation (15))

$$(14) \quad \Delta d_{t+1} = \gamma_0 + \gamma_1 \Delta d_t + \gamma_2 (f_t - s_t) + u_{14t+1}$$

$$(15) \quad \Delta d_{t+1} = \delta_0 \Delta d_t + \delta_1 (s_t - f_{t-1}) + u_{15t+1}.$$

The ex ante relative purchasing power parity hypothesis holds if $\gamma_0 = \gamma_1 = \gamma_2 = 0$ and u_{14t+1} is white noise in equation (14) and if $\delta_0 = \delta_1 = 0$ and u_{15t+1} is white noise in equation (15). As can be seen from the estimates of table 2, these hypotheses are always rejected. Expected real exchange rate changes do not seem to be unpredictable.

The validity of the specification of equation (14) is further corroborated by the dynamic simulations set forth in table 3. Both the R.M.S.E. and the correlation coefficients (between actual and simulated values of the dependent variable) suggest that this equation reproduces the effective behaviour of the real exchange rate better than the random walk model in the period January 1981 - February 1985.¹⁰ (Over the period March 1985 - October 1992 the results are less clearcut, since both the correlation coefficients and the R.M.S.E. statistics suggest that the random walk outperforms equation (15) in reproducing the effective real exchange rate behaviour.)

TABLE 2

1981:1-1985:2

$$(14) \Delta d_{t+1} = \gamma_0 + \gamma_1 \Delta d_t + \gamma_2 (f_t - s_t) + u_{14t+1}$$

	γ_0	γ_1	γ_2	R^2	S.E.	D.a.	F_{H_0}
L/\$	-0.01 (-1.81)	-0.27 (-1.53)	-7.86 (-2.38)	0.19	0.0338	-0.46	4.04*
DM/\$	-	-	-2.26 (-2.21)	0.02	0.0307	0.11	-
Y/\$	-	-	-1.66 (-1.68)	0.05	0.0329	-0.26	-
LIT/\$	-0.03 (-3.93)	-0.24 (-1.58)	-3.12 (-3.15)	0.15	0.0261	-1.59	3.12*

1985:3-1992:10

$$(15) \Delta d_{t+1} = \delta_0 \Delta d_t + \delta_1 (s_t - f_{t-1}) + u_{15t+1}$$

	δ_0	δ_1	R^2	S.E.	D.a.	F_{H_1}
L/\$	-0.31 (-1.54)	0.48 (2.22)	0.07	0.0412	-0.99	6.65*
DM/\$	-0.37 (-1.86)	0.43 (2.13)	0.06	0.0385	-1.60	4.35*
Y/\$	-0.58 (-2.37)	0.66 (2.54)	0.13	0.0396	-2.37*	11.64*
LIT/\$	-0.33 (-1.52)	0.47 (2.02)	0.06	0.0387	-1.65	5.45*

t RATIOS COMPUTED WITH HETEROSKEDASTICITY CONSISTENT STANDARD ERRORS.

D.a.: DURBIN'S ALTERNATIVE TEST FOR FIRST ORDER SERIAL CORRELATION.

 F_{H_0} : F TEST OF THE NULL HYPOTHESIS $H_0: \gamma_0 = \gamma_1 = \gamma_2 = 0$. F_{H_1} : F TEST OF THE NULL HYPOTHESIS $H_1: \delta_0 = \delta_1 = 0$.

* : SIGNIFICANT AT THE 5% LEVEL.

TABLE 3

1981:1-1985:2

	EQUATION (14)		RANDOM WALK ¹¹	
	$\rho_{d_{t+1}\hat{d}_{t+1}}$	R.M.S.E.	$\rho_{d_{t+1}\hat{d}_{t+1}}$	R.M.S.E.
L/\$	0.948	0.063	0.899	0.069
DM/\$	0.932	0.044	0.930	0.070
Y/\$	0.724	0.078	0.611	0.098
LIT/\$	0.870	0.098	0.753	0.055

1985:3-1992:10

	EQUATION (15)		RANDOM WALK	
	$\rho_{d_{t+1}\hat{d}_{t+1}}$	R.M.S.E.	$\rho_{d_{t+1}\hat{d}_{t+1}}$	R.M.S.E.
L/\$	0.868	0.222	0.779	0.081
DM/\$	0.714	0.414	0.886	0.067
Y/\$	0.740	0.430	0.940	0.050
LIT/\$	0.917	0.366	0.917	0.071

$\rho_{x_t \hat{x}_t}$: CORRELATION COEFFICIENT BETWEEN THE HISTORICAL AND THE SIMULATED VALUES OF THE DEPENDENT VARIABLE x_t .

$$\text{R.M.S.E. (ROOT MEAN SQUARE ERROR)} : \sqrt{\frac{1}{T} \sum_{t=1}^T (\hat{x}_t - x_t)^2}$$

WHERE T IS THE NUMBER OF OBSERVATIONS, $\hat{x}_t$ AND x_t ARE THE SIMULATED AND HISTORICAL VALUES OF THE VARIABLES OF INTEREST.

4. Empirical Results: the Unbiased Efficiency Hypothesis, the Real Interest Rates Differentials and the Real Exchange Rate Changes

The results of the previous section contradict the hypotheses that are accepted a priori in the studies mentioned above, viz. that condition (ii) holds and that condition (i) is always rejected. Ex ante real interest rates on nominally riskless bonds differ across countries, violating condition (i), in the March 1985 - October 1992 period only, and condition (ii) is clearly violated in the period January 1981 - February 1985. (In the period March 1985 - October 1992 we obtain mixed results: the random walk restrictions on equation (15) are rejected but, at the same time, the random walk outperforms the alternative specification in the dynamic simulations.)

We want to assess whether these results can be used to explain the (observed) systematic discrepancies between forward and future spot exchange rates with the help of the model set out in section 2 above. Indeed, we want to check whether the forecastable deviations from forward exchange rate unbiasedness are equal to the forecastable change in the real exchange rate in the period January 1981 - February 1985 and whether they are equal to the forecastable difference in real returns on nominally riskless bonds across currencies in the period March 1985 - October 1992.¹²

We have replaced the ex ante real interest rate differences of equation (9) by their realised values $(r_{t+1}^* - r_{t+1})$ and have applied an errors in variables estimation method. We have obtained consistent and asymptotically efficient estimates of

the resulting equation (16) below with the help of the generalised instrumental variables estimation approach. (We use as instruments in deriving the orthogonality conditions a constant, the lagged real interest rates differential across the two countries of interest, the lagged inflation rate differential, the forward premium and the average value - over the previous twelve months - of the real interest rate differential.) We have introduced the GMM methodology since a preliminary analysis suggested that the estimates might be affected by conditional heteroskedasticity.¹³ Here too we have applied the bandwidth selection procedure of D.W.K. Andrews and have adjusted accordingly the covariance matrix of orthogonality conditions.

$$(16) (s_{t+1} - f_t) = c_0 + c_1(r_{t+1}^* - r_{t+1}) + u_{16t+1}$$

Instrumental variables:

$$\text{constant}, (r_t^* - r_t), (\Delta P_t^* - \Delta P_t), (f_t - s_t), \left[\frac{1}{12} \sum_{i=1}^{12} (r_{t+1-i}^* - r_{t+1-i}) \right].$$

The estimates set out in table 4 indicate that in the period March 1985 - October 1992 (in which condition (i) does not hold) some of the forecastable deviations in the unbiased expectations hypothesis can indeed be explained by ex ante real interest rate differentials in the case of the deutschmark-dollar and lira-dollar exchange rates. The hypothesis that $c_0 = 0$ and $c_1 = 1$ in equation (16) is never rejected. The power of the test is low, however, in the case of the yen-dollar exchange rate given the low value of the multiple correlation coefficient. In the period January 1981 - February 1985 condition (i) seems to hold and, as expected,

TABLE 4

$$(16) (s_{t+1} - f_t) = c_0 + c_1(r_{t+1}^* - r_{t+1}) + u_{16t+1}$$

1981:1-1985:2

	c_0	c_1	R^2	S.E.	D.W.	H_0	n
$\text{L}/\text{\$}$	-0.01	2.74	0.01	0.0396	1.86	7.52*	0
	(-0.70)	(0.62)					
$\text{DM}/\text{\$}$	-0.02	-1.75	0.02	0.0297	1.79	3.52	0
	(-1.78)	(-0.61)					
$\text{Y}/\text{\$}$	-0.01	-0.45	0.004	0.0332	1.89	1.58	0
	(-0.75)	(-0.10)					
$\text{LIT}/\text{\$}$	-0.01	-1.91	0.02	0.0275	1.90	6.80*	0
	(-2.16)	(-1.31)					

1985:3-1992:10

	c_0	c_1	R^2	S.E.	D.W.	H_0	n
$\text{L}/\text{\$}$	0.01	0.36	0.04	0.0384	1.64	1.20	1
	(1.63)	(1.21)					
$\text{DM}/\text{\$}$	0.01	0.54	0.07	0.0362	1.90	0.83	1
	(2.02)	(2.33)					
$\text{Y}/\text{\$}$	0.01	-0.06	0.004	0.0348	1.91	0.87	0
	(2.20)	(-0.49)					
$\text{LIT}/\text{\$}$	0.01	0.48	0.04	0.0357	1.76	1.46	1
	(2.12)	(2.11)					

n : NUMBER OF AUTOCORRELATION TERMS USED IN COMPUTING THE COVARIANCE MATRIX OF THE ORTHOGONALITY CONDITIONS.

H_0 : LIKELIHOOD RATIO-LIKE TEST OF THE NULL HYPOTHESIS $c_0 = 0, c_1 = 1$.

* : SIGNIFICANT AT THE 5% LEVEL.

the overall quality of fit of the estimates declines as ex ante real interest rate differentials exert no effect on forward forecast errors.

As pointed out in section 2, systematic forward forecast errors may also be due to anticipated real exchange rate changes, an hypothesis rejected a priori in the empirical analyses mentioned above. We report in table 5 the generalised instrumental variables estimates of equation (17) below, in which we have replaced the ex ante real exchange rate changes of equation (11) by their realised values. (We use as instruments the regressors of equations (14) and (15) in the two relevant time periods.)

$$(17) (s_{t+1} - f_t) = e_0 + e_1 \Delta d_{t+1} + u_{17t+1}^{14}$$

Instrumental variables:

constant, Δd_t , X_t , where

$X_t = (f_t - s_t)$ in the period January 1981 - February 1985

$X_t = (s_t - f_{t-1})$ in the period March 1985 - October 1992.

The GMM estimates indicate that in the period January 1981 - February 1985 (in which condition (ii) is violated) a relevant fraction of the deviation from the unbiased expectations hypothesis can be attributed to the expectation of deutschemark, pound sterling, yen and lira real depreciation by financial market participants.¹⁵ The overall quality of fit is good and the relevant coefficients have the appropriate signs. The null hypothesis that $e_0 = 0$ and $e_1 = 1$ is rejected. This is due to the presence of constant terms that are significantly different from zero. Indeed, the null hypothesis that $e_1 = 1$ is always accepted, with the exception of the lira-

TABLE 6

$$(17) (s_{t+1} - f_t) = \alpha_0 + \alpha_1 \Delta d_{t+1} + u_{17t+1}$$

1981:1-1985:2

	α_0	α_1	R^2	S.E.	D.W.	H_0	H_1	n
L/\$	-0.002	1.06	0.98	0.0062	1.94	8.85*	2.45	0
	(-2.51)	(20.97)						
DM/\$	-0.01	0.74	0.99	0.0078	1.79	26.62*	3.63	0
	(-5.77)	(7.11)						
Y/\$	-0.001	1.23	0.96	0.0107	1.79	9.62*	2.72	0
	(-1.13)	(9.39)						
LIT/\$	0.001	1.22	0.96	0.0085	1.71	7.16*	5.74*	0
	(0.04)	(13.14)						

1985:3-1992:10

	α_0	α_1	R^2	S.E.	D.W.	H_0		n
L/\$	0.004	0.25	0.65	0.0312	1.80	3.24		2
	(2.08)	(0.68)						
DM/\$	0.001	-0.24	0.61	0.0456	1.82	0.54		3
	(0.71)	(-1.05)						
Y/\$	0.001	0.16	0.63	0.0300	1.89	1.46		3
	(0.64)	(1.31)						
LIT/\$	0.005	-0.22	0.61	0.0434	1.73	9.53*		2
	(2.26)	(-0.79)						

n : NUMBER OF AUTOCORRELATION TERMS USED IN COMPUTING THE COVARIANCE

MATRIX OF THE ORTHOGONALITY CONDITIONS.

 H_0 : LIKELIHOOD RATIO-LIKE TEST OF THE NULL HYPOTHESIS $\alpha_0 = 0$, $\alpha_1 = 1$. H_1 : LIKELIHOOD RATIO-LIKE TEST OF THE NULL HYPOTHESIS $\alpha_1 = 1$.

* : SIGNIFICANT AT THE 5% LEVEL.

dollar exchange rate. Deviations from the unbiased expectations hypothesis are not caused exclusively by ex ante real exchange rate changes.

We have seen in the previous section that the random walk outperforms the alternative specification in reproducing real exchange rate behaviour in the period March 1985 - October 1992. This finding is reflected in the estimates of equation (17): the quality of fit declines and ex ante real exchange rate changes have no systematic effect on ex ante forward forecast errors.

5. Conclusion

In this paper we have shown that systematic forward forecast errors have to be associated with ex ante real exchange rate changes in the period January 1981 - February 1985. A possible interpretation of this finding is that the four currencies under investigation paid a negative risk premium vis a' vis the dollar due to expected (national) real depreciation. The failure of the unbiased expectations hypothesis seems to be the result of goods market inefficiency insofar as the association of non zero ex ante real exchange rate changes with goods market inefficiency holds.

In the period March 1985 - October 1992 ex ante real interest differentials explain a fraction of the systematic forward forecast errors of the dollar exchange rate with the E.M.S. currencies of the sample, the deutschemmark and the lira. (Britain joined the E.M.S. only in 1990.) This finding is probably connected with the rigidities introduced into

financial markets by the E.M.S. itself. Ex ante real exchange rate differentials are not reflected in systematic forward forecast errors in the case of the yen-dollar and sterling-dollar exchange rates, a result in line with the growing efficiency of international capital markets. On the whole the empirical findings suggest that the link of systematic forward forecast errors with financial risk premia is rather weak.

notes

- (1) Under the maintained hypothesis of rational expectations the deviations of $E[(s_{t+1} - f_t) | \Phi_t]$ from zero are to be interpreted as (minus) the exchange risk premium.
- (2) A differential in ex ante real returns may also be due to barriers to capital movements across countries, which prevent complete arbitrage and - temporarily - to systematic exchange rate manipulation by central banks.
- (3) R.E. Cumby, M. Obstfeld (1984), D. Longworth (1986) and R.D. Huang (1987), however, found evidence contrary to the ex ante relative purchasing power parity hypothesis. In this case expected real returns on nominally riskless bonds will depend upon the currency in which they are evaluated. This will affect the forward exchange rate, which will differ from the future spot rate expected by the market.
- (4) As shown in a previous paper (G. Cifarelli (1992)) the dollar behaviour seems to have been affected by a regime shift in February 1985. In the period 1981 - 1985 market forces seem to have played a dominant role in causing its

appreciation. In the period 1985 - 1992 central banks have influenced its behaviour, essentially with the help of more or less coordinated interest rate manipulation.

(5) F. Giavazzi, M. Pagano (1985) and L.E. Mohlo (1990) have found that, in the case of Italy, capital controls have not been binding on average. The behaviour of the Eurolira interest rates does not differ in a significant way from the behaviour of equivalent short run onshore interest rates.

(6) Conditional heteroskedasticity seems to affect the residuals over the period March 1985 - October 1992. Conditional homoskedasticity implies that $E(u_{13t+1}^2 | \Phi_t) = \text{constant}$. It requires that in the regression $u_{13t+1}^2 = \theta_0 + \theta_1 u_{13t} + \theta_2 u_{13t}^2 + e_{t+1}$ the estimates of the coefficients θ_1 and θ_2 be insignificantly different from zero. We obtain the following F test results, where u_{13t+1} is set equal to $(\Delta P_{t+1} - \Delta P_{t+1}^*) - (i_{t+1} - i_{t+1}^*)$, under the assumption that the null hypothesis of ex ante real interest rate equality be valid in equation (13) above:

$F_{\theta_1=\theta_2=0}$	U.S./U.K.	U.S./GERMANY	U.S./JAPAN	U.S./ITALY
1981:1-1985:2	2.42	0.26	0.24	0.24
1985:3-1992:10	31.50*	38.36*	33.99*	36.74*

* : SIGNIFICANT AT THE 5% LEVEL.

The conditional homoskedasticity restrictions are rejected in the period March 1985 - October 1992.

(7) L.P. Hansen, K.J. Singleton (1982) proposed a modified weighting matrix for the case of serially correlated disturbances pointing out, however, that it was no longer positive definite by construction. This shortcoming might bias asymptotic confidence interval formation and hypothesis testing. Heteroskedasticity and autocorrelation consistent

covariance matrix estimators have been set out by H. White (1984, pp. 147 - 161), A.R. Gallant (1987, p. 533), W.K. Newey, K.D. West (1987) and others. They introduce the use of various kernels to guarantee positive semi-definiteness in finite samples. We apply in this paper the Bartlett kernel (suggested by Newey and West) to smooth the sample autocovariance function.

(8) D.W.K. Andrews (1991) suggests the following guidelines for the selection of a bandwidth (lag truncation) parameter. First, one estimates p univariate first order autoregressive AR(1) models for the p elements of the Kronecker product of the residuals of the GMM model estimates with the vector of instruments (p is the number of orthogonality conditions). Then one computes

$$\hat{\alpha} = \sum_{a=1}^p \frac{4\hat{\rho}_a^2 \hat{\sigma}_a^4}{(1-\hat{\rho}_a)^6 (1+\hat{\rho}_a)^2} \bigg/ \sum_{a=1}^p \frac{\hat{\sigma}_a^4}{(1-\hat{\rho}_a)^4}$$

where $((\hat{\rho}_a, \hat{\sigma}_a^2), a=1, \dots, p)$ denote the autoregressive and innovation variance estimates for the a^{th} element.

The automatic bandwidth parameter in the case of the Newey-West approach is given by $\hat{M} = 1.1447(\hat{\alpha}T)^{1/3} - 1$, where T is the sample size. In this paper we have used integer-valued lag truncation parameters only and have replaced $\hat{M}$ by the corresponding closest integer.

(9) Analogous attempts to model real exchange rate changes behaviour can be found in J.A. Frankel, A.T. MacArthur (1988) and R. Levine (1989). A nonzero expected real exchange rate change is interpreted by W.H. Branson (1988) as representing a lack of goods market integration. L.E. Mohlo (1990) suggests that serial correlation in the real exchange

rate changes may be due to incomplete credibility on the nonaccommodating exchange rate policy (to the inflation differentials) of the monetary authorities.

(10) The estimates of equations (14) and (15) explain the behaviour of the change in the real exchange rate. We have thus replaced Δd_{t+1} by $d_{t+1} - d_t$ and have used d_t as a predetermined variable.

(11) Random Walk models used in the simulations:

$$d_{t+1} = \tau + \psi d_t + e_{t+1}$$

	1981:1-1985:2					1985:3-1992:10				
	τ	ψ	R^2	S.E.	D.h.	τ	ψ	R^2	S.E.	D.h.
L/\$	0.02 (0.77)	0.94 (21.94)	0.95	0.0358	-0.46	0.06 (3.07)	0.91 (27.99)	0.89	0.0410	0.61
DM/\$	-	1.01 (220.66)	0.93	0.0311	0.25	-0.05 (-2.50)	0.92 (33.09)	0.93	0.0377	-0.72
Y/\$	-	1.001 (1175.18)	0.91	0.0337	0.41	-0.42 (-3.21)	0.92 (36.56)	0.92	0.0393	-0.13
Lit/\$	-0.59 (-1.84)	0.92 (20.18)	0.88	0.0271	0.31	-0.46 (-2.84)	0.93 (38.87)	0.94	0.0381	-0.27

t RATIOS COMPUTED WITH HETEROSKEDASTICITY CONSISTENT STANDARD ERRORS.

(12) We could not assess the joint effect of ex ante real interest differentials and of ex ante real exchange rate changes on forward forecast errors because of a severe multicollinearity problem. For an economic interpretation of the correlation between these two regressors, see J.Y. Campbell, R.H. Clarida (1987).

(13) Conditional heteroskedasticity seems to affect the residuals of equation (16) in the period March 1985 - October 1992. We have applied the test set out in note (6) above to the residuals of equation (16), $u_{1985:1-1992:10}$, and have obtained the

following F statistics, where $u_{16t+1} = (s_{t+1} - f_t) - (r_{t+1}^* - r_{t+1})$.

	L/\$	DM/\$	Y/\$	LIT/\$
1981:1-1985:2	0.05	0.26	1.28	0.29
1985:3-1992:10	3.52*	1.03	9.60*	1.70

* : SIGNIFICANT AT THE 5% LEVEL.

(14) We have applied the conditional homoskedasticity tests also to the residuals of equation (17), u_{17t+1} . We obtain the following F statistics, where $u_{17t+1} = (s_{t+1} - f_t) - \Delta d_{t+1}$.

	L/\$	DM/\$	Y/\$	LIT/\$
1981:1-1985:2	2.25	0.03	0.19	0.30
1985:3-1992:10	31.70*	38.56*	0.36	1.62

* : SIGNIFICANT AT THE 5% LEVEL.

(15) We provide in this way a possible interpretation of a phenomenon - the systematic underprediction of the dollar strength from 1981 to 1985 - that has been intensively investigated. G.W. Evans (1986) labels it a "bubble" but provides no definite explanation of its causes. J.A. Frankel, K.A. Froot (1987) interpret the systematic forward forecast errors as evidence of market irrationality. K.K. Lewis (1989) attributes these errors to rational (bayesian) learning by market participants about the increase in the U.S. money demand of the early 1980s.

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