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QUADERNI DEL DIPARTIMENTO
DI ECONOMIA POLITICA

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ON TIME VARYING RISK PREMIA
IN THE FOREIGN EXCHANGE MARKET



***QUADERNI DEL DIPARTIMENTO
DI ECONOMIA POLITICA***

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Giulio Cifarelli

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Siena, Agosto 1995

Introduction

A substantial effort has recently been directed to the development of theoretical models of the forward risk premium, models that are based mostly on an extension to the foreign exchange market of financial market interpretations of the risk premia that may arise in the pricing of domestic assets.¹ The empirical evidence on these approaches is rather weak.² In this paper we provide an alternative analysis of the determinants of the forward risk premium that is explicitly set out in an international financial framework. Section 1 describes the economic rationale of the model of the forward risk premium; section 2 investigates the links between the risk premium, ex ante real interest rate differentials and ex ante real exchange rate changes. Section 3 analyses the properties of the first and second conditional moments of the forward risk premium distribution set in an ARCH model framework.

The estimates suggest that the determinants of the forward risk premium change over time. Risk premia on nominally riskless assets denominated in different currencies do not play the relevant role suggested by the theory. Inflation expectation differentials - that are not reflected in real interest rate differentials because of financial market efficiency - may well exert a direct influence on the conditional mean of the forward risk premium distribution. The volatility of this premium can be modelled as an ARCH process for some of the exchange rates of the sample.

1. The Risk Premium in the Foreign Exchange Market

The unbiased expectations hypothesis on the behaviour of the forward exchange rate implies that

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

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 at time t and $E(.)$ is the expectations operator. Under the maintained hypothesis of rational expectations the deviations of the expected forward forecast errors from zero can be seen as (minus) the exchange risk premium if we follow E.F. Fama (1984) and assume that the forward exchange rate be equal to the sum of the corresponding expected future spot rate and a risk premium.

With the help of the covered interest parity condition, of the Fisher interest rate relationships in the two countries of interest and of the relative purchasing power parity hypothesis, we can rewrite equation (1) in terms of expected differences, across countries, in the real interest rates and of expected real exchange rate changes.³ We obtain

$$E[(s_{t+1} - f_t) | \Phi_t] = E[(r_{t+1}^* - r_{t+1}) | \Phi_t] + E[\Delta d_{t+1} | \Phi_t] \quad (2)$$

where r_{t+1} and r_{t+1}^* are - respectively - the real returns on default free domestic and foreign bonds maturing at time $t+1$ and d_{t+1} is the logarithm of the real exchange rate at time $t+1$.

The right hand side is null and there is no risk premium (as the unbiased expectations hypothesis holds) if the ex ante real returns on default free assets are equal across countries and if the ex ante rate of change of the real exchange rate is zero.

If we assume a priori that $E[(r_{t+1}^* - r_{t+1}) | \Phi_t] = 0$, equation (2) becomes

$$E[(s_{t+1} - f_t) | \Phi_t] = E[\Delta d_{t+1} | \Phi_t]. \quad (3)$$

Expected deviations from the unbiased expectations hypothesis should reflect the expected change in the real exchange rate. The relevance of this hypothesis can be assessed by estimating the relationship

$$(s_{t+1} - f_t) = a_0 + a_1 E[\Delta d_{t+1} | \Phi_t] + u_{4t+1} \quad (4)$$

where $u_{4t+1} = (s_{t+1} - f_t) - E[(s_{t+1} - f_t) | \Phi_t]$.

The null hypothesis is that $a_0 = 0$, $a_1 = 1$ and $E[u_{4t+1} | \Phi_t] = 0$. Rejection of these restrictions is a rejection of the joint hypothesis that expected real exchange rate changes are associated with deviations from the unbiased expectations hypothesis (i.e. with a nonzero forward risk premium) and that expected real interest rates on default free bonds are equal across countries.

If we assume a priori that $E[\Delta d_{t+1} | \Phi_t] = 0$, equation (2) becomes

$$E[(s_{t+1} - f_t) | \Phi_t] = E[(r_{t+1}^* - r_{t+1}) | \Phi_t] \quad (5)$$

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

$$(s_{t+1} - f_t) = b_0 + b_1 E[(r_{t+1}^* - r_{t+1}) | \Phi_t] + u_{6t+1} \quad (6)$$

where $u_{6t+1} = (s_{t+1} - f_t) - E[(s_{t+1} - f_t) | \Phi_t]$.

The null hypothesis here is that $b_0 = 0$, $b_1 = 1$ and $E[u_{6t+1} | \Phi_t] = 0$. Rejection of these restrictions is a rejection of the joint hypothesis that risk premia on nominally riskless bonds are the cause of the failure of the unbiased expectations hypothesis (i.e. of a nonzero forward risk premium) and that $E[\Delta d_{t+1} | \Phi_t] = 0$.

The regressors $E[(r_{t+1}^* - r_{t+1}) | \Phi_t]$ and $E[\Delta d_{t+1} | \Phi_t]$ are not observable. However, consistent and asymptotically efficient parameter estimates may be obtained with the help of instrumental variables, following a methodology analysed 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.

2. The Empirical Relevance of the Risk Premium

We have investigated the behaviour of the exchange rate between the U.S. dollar and four currencies, the British pound, the German mark, the Japanese yen and the Italian lira

in the periods January 1981 - February 1985 and March 1985 - October 1992.⁴ In a previous research (G. Cifarelli (1995)) we found that ex ante real interest rates on nominally riskless assets seemed to differ across the countries of interest only in the period March 1985 - October 1992. At the same time we found that the hypothesis that ex ante real exchange rate changes were nil (i.e. that the four real exchange rates under investigation followed a martingale process) had to be rejected only in the period January 1981 - February 1985. As a consequence we have analysed the behaviour of the risk premium with the help of equation (4) in the period January 1981 - February 1985 and of equation (6) in the period March 1985 - October 1992.⁵

As suggested by equation (4), systematic forward forecast errors may be due to anticipated real exchange rate changes. We have replaced the expected real exchange rate changes by their realized values Δd_{t+1} and have applied an errors in variables estimation method. We have obtained consistent and asymptotically efficient estimates of the resulting equation (7) with the help of the generalized instrumental variables estimation approach reported in the appendix.⁶ We have used as instruments a constant, the forward premium and the lagged real exchange rate change.

$$(s_{t+1} - f_t) = a_0 + a_1 \Delta d_{t+1} + u_{7t+1} \quad (7)$$

Instrumental variables:

constant, $(f_t - s_t)$, Δd_t .

The estimates indicate that in the period January 1981 - February 1985 systematic forward forecast errors for

equivalently the risk premium) can be attributed to the expectation of pound sterling, deutschemark, yen and lira real depreciation by financial market participants. The relevant coefficients have the appropriate signs. The Hansen tests of overidentifying restrictions (L.P. Hansen (1982), pp. 1049-1050) fail to reject these restrictions and the Andrews tests for parameter stability suggest that the estimates are stable.⁷ The null hypothesis that $a_0 = 0$ and $a_1 = 1$ is mostly rejected. This is due to the presence of constant terms that are significantly different from zero and negative.⁸ Indeed, with one exception, the hypothesis that $a_1 = 1$ is always accepted.

TABLE 1

$$(s_{t+1} - f_t) = a_0 + a_1 \Delta d_{t+1} + u_{7t+1} \quad (7)$$

1981:1-1985:2

	a_0	a_1	R^2	S.E.	D.W.	H_0	H_1	H_a	SupW	n
L/\$	-0.002 (-2.51)	1.06 (20.97)	0.98	0.0063	1.94	9.37*	1.33	1.22	2.93	0
DM/\$	-0.01 (-2.70)	0.79 (4.39)	0.99	0.0063	1.78	15.43*	1.28	0.01	1.41	0
Y/\$	-0.002 (-1.89)	1.10 (7.51)	0.96	0.0078	1.81	7.46*	0.44	1.53	0.92	0
LIT/\$	0.001 (0.84)	1.22 (13.14)	0.96	0.0087	1.71	6.70	5.43*	0.18	6.27	0

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

H_0 : WALD TEST OF THE NULL HYPOTHESIS $a_0 = 0, a_1 = 1$.

H_1 : WALD TEST OF THE NULL HYPOTHESIS $a_1 = 1$.

H_a : HANSEN'S TEST OF OVERIDENTIFYING RESTRICTIONS.

SupW : WALD TEST FOR PARAMETER STABILITY WITH UNKNOWN CHANGE POINT.

* : SIGNIFICANT AT THE 5% LEVEL.

In order to assess whether the deviations from the unbiased expectations hypothesis are caused only by ex ante real exchange rate changes we have estimated the following unrestricted version of equation (7)

$$(s_{t+1} - f_t) = a_0' + a_1' \Delta d_{t+1} + a_2'(r_t^* - r_t) + u_{8t+1} \quad (8)$$

with the help of the same methodology. Variables that enter the information set at time t, such as $(r_t^* - r_t)$, should not systematically affect the forward forecast error beyond their ability to influence Δd_{t+1} if the specification of equation (7) is correct. The estimates are reported in table 2.

TABLE 2

$$(s_{t+1} - f_t) = a_0' + a_1' \Delta d_{t+1} + a_2'(r_t^* - r_t) + u_{8t+1} \quad (8)$$

1981:1-1985:2

	a_0'	a_1'	a_2'	R^2	S.E.	D.W.	n
L/\$	-0.002 (-1.51)	1.15 (6.38)	-0.79 (-0.72)	0.96	0.0099	1.51	0
DM/\$	-0.004 (-0.54)	0.82 (2.17)	0.12 (0.08)	0.99	0.0056	1.76	0
Y/\$	0.001 (0.19)	0.99 (5.78)	1.25 (1.03)	0.91	0.0106	2.49	0
LIT/\$	0.001 (0.41)	1.21 (12.14)	-0.33 (-0.44)	0.95	0.0093	1.51	0

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

The additional regressor has no explanatory power and the quality of fit is not affected. We have assessed in the same way the influence of the forward premium and of the lagged forward forecast error and have obtained similar (disappointing) results. The forward risk premium seems to be affected by expected real exchange rate behaviour only and the role of additional explanatory variables seems to be irrelevant.

We have assessed the influence of ex ante real interest rate differentials on forward forecast errors in the period March 1985 - October 1992. We report in table 3 the generalized instrumental variables estimates of equation (9) in which we have replaced the ex ante real interest rate differences of equation (6) by their realized values ($r_{t+1}^* - r_{t+1}$). We use as instruments in deriving the orthogonality conditions a constant, the lagged real interest rate differential between the two countries of interest, the lagged inflation rates differential, the forward premium and the average value - over the previous twelve months - of the real interest rate differential.

$$(s_{t+1} - f_t) = b_0 + b_1(r_{t+1}^* - r_{t+1}) + u_{9t+1} \quad (9)$$

Instrumental variables:

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

The estimates set out in table 3 indicate that ex ante real interest rate differentials (i.e. risk premia) have a systematic impact on forward forecast errors in the case of the deutschmark dollar and lira dollar exchange rates. The

TABLE 3

$$(s_{t+1} - f_t) = b_0 + b_1(r_{t+1}^* - r_{t+1}) + u_{9t+1} \quad (9)$$

1985:3-1992:10

	b_0	b_1	R^2	S.E.	D.W.	H_0	H_a	SupW	n
$\text{L}/\$$	0.01 (1.63)	0.36 (1.21)	0.04	0.0388	1.64	5.85	5.62	2.06	1
$\text{DM}/\$$	0.01 (2.02)	0.54 (2.33)	0.07	0.0365	1.90	7.24*	7.07	5.19	1
$\text{Y}/\$$	0.01 (2.20)	-0.06 (-0.49)	0.004	0.0352	1.91	85.84*	2.14	4.10	0
$\text{LIT}/\$$	0.01 (2.12)	0.48 (2.11)	0.04	0.0361	1.76	8.10*	6.02	4.25	1

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

H_0 : WALT TEST OF THE NULL HYPOTHESIS $b_0 = 0, b_1 = 1$.

SupW : WALT TEST FOR PARAMETER STABILITY WITH UNKNOWN CHANGE POINT UNDER THE ALTERNATIVE.

H_a : HANSEN'S TEST OF OVERIDENTIFYING RESTRICTIONS.

* : SIGNIFICANT AT THE 5% LEVEL.

hypothesis that $b_0 = 0$ and $b_1 = 1$ is always rejected, however, with the exception of the sterling-dollar exchange rate. There is no strong evidence of misspecification as the Hansen tests of overidentifying restrictions fail to reject these restrictions. The Andrews Wald tests for parameter stability suggest that the estimates are stable.

We have estimated also the following unrestricted version of equation (9) with the help of the same methodology

$$(s_{t+1} - f_t) = b_0' + b_1'(r_{t+1}^* - r_{t+1}) + b_2'\Delta d_t + b_3'(s_t - f_{t-1}) + u_{10t+1} \quad (10)$$

and have tested whether $b_2' = b_3' = 0$. Information available at time t (such as Δd_t and $(s_t - f_{t-1})$) should not systematically affect $(s_{t+1} - f_t)$ beyond its ability to influence $(r_{t+1}^* - r_{t+1})$ if the specification of equation (9) is correct. The estimates are set out in table 4. The hypothesis that $b_2' = b_3' = 0$ cannot be rejected, with the exception of the lira-dollar exchange rate.

TABLE 4

$$(s_{t+1} - f_t) = b_0' + b_1'(r_{t+1}^* - r_{t+1}) + b_2'\Delta d_t + b_3'(s_t - f_{t-1}) + u_{10t+1} \quad (10)$$

1985:3-1992:10

	b_0'	b_1'	b_2'	b_3'	R^2	S.E.	H_0	n
$L/\$$	-0.003	-0.53	0.81	0.82	0.02	0.0744	1.53	1
	(-0.32)(-0.16) (0.38) (1.18)							
$DM/\$$	0.01	0.53	0.12	-0.78	0.07	0.0373	0.38	1
	(1.66) (1.45) (0.29)(-0.57)							
$Y/\$$	0.01	0.27	-0.17	0.54	0.008	0.0371	1.36	0
	(0.68) (0.79) (-0.97) (0.80)							
$LIT/\$$	0.001	0.58	0.06	0.66	0.04	0.0426	13.81*	1
	(0.16) (2.23) (0.28) (3.36)							

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

H_0 : WALD TEST OF THE NULL HYPOTHESIS $b_2' = b_3' = 0$.

* : SIGNIFICANT AT THE 5% LEVEL.

We have assessed in the same way the influence of the

forward premium $(f_t - s_t)$ and of the lagged forward forecast error $(s_t - f_{t-1})$ and have obtained analogous (negative) results.⁹ In the second half of the sample period the specification of equation (9) does not seem to be rejected by the data and at least some of the deviations from forward exchange rate unbiasedness can be explained by expected differences in real interest rates on default free assets. The impact of additional explanatory variables is marginal.

W.K. Newey (1985) has shown that optimal tests of the violation of specific moment conditions can be used to formulate tests for the validity of instrumental variables. His approach becomes relevant if - as is the case here - the theory provides no definite guidelines for the choice of instrumental variables. Moreover, as pointed out by G. Tauchen (1986), a large set of instruments may result in biased GMM estimates. We have tested with the help of the Newey optimal GMM test statistic the relevance of model misspecification for various possible choices of X_{2t} , the vector of (possibly) contaminated s additional instrumental variables, where X_{1t} is the vector of instruments used in the GMM estimates of equations (7) and (9) above. This statistic has an asymptotic chi-square distribution with s degrees of freedom.¹⁰ The results of table 5 suggest that - with one exception for the lira-dollar exchange rate estimates of equation (7) in the period January 1981 - February 1985 - the null hypothesis that the introduction of the additional instruments leads to model misspecification cannot be rejected at the 5% level of significance. The parsimonious choice of instruments adopted in the estimates above does not seem to be contradicted by the data.

TABLE 5

X_{2t}	EQUATION (7)			EQUATION (9)		
	1981:1-1985:2			1985:3-1992:10		
	Δd_t^j	$(r_t^* - r_t)^j$	$(s_t - f_{t-1})$	$(r_t^* - r_t)^j$	Δd_t	$(s_t - f_{t-1})$
L/\$	6.352	1.936	1.979	2.825	0.196	0.198
DM/\$	3.721	3.656	1.387	3.838	0.009	0.066
Y/\$	0.713	0.014	0.011	0.200	0.320	0.340
LIT/\$	1.608	5.790*	3.708	3.917	0.002	0.001

$(r_t^* - r_t)^j$: VECTOR OF LAGGED DIFFERENTIALS BETWEEN THE U.S. REAL INTEREST RATE AND THE REAL INTEREST RATES OF THE THREE OTHER COUNTRIES OF THE SAMPLE.

Δd_t^j : VECTOR OF THE LAGGED REAL EXCHANGE RATE CHANGES OF THE THREE OTHER CURRENCIES OF THE SAMPLE.

* : SIGNIFICANT AT THE 5% LEVEL.

3. A Joint Parameterization of the Conditional Mean and of the Conditional Variance of the Risk Premium

The principal features of the distributions of the residuals of the GMM estimates of equations (7) and (9) in the first and in the second half of the time interval of interest are presented in table 6. A preliminary inspection of the results suggests that the volatility of the residuals - roughly measured by the standard errors - is much larger in the second half of the sample than in the first. The residuals of equation (7) fail the Jarque Bera normality test in the case of the sterling-dollar and deutschmark-dollar exchange rates and

TABLE 6

	$\hat{u}_{7t+1}$				$\hat{u}_{9t+1}$			
	1981:1-1985:2				1985:3-1992:10			
	L/\$	DM/\$	Y/\$	LIT/\$	L/\$	DM/\$	Y/\$	LIT/\$
MEAN	0.0001	0.0001	-0.0003	-0.0001	-0.0016	-0.0019	-0.0037	-0.0010
S.D.	0.0062	0.0063	0.0078	0.0086	0.0385	0.0376	0.0350	0.0359
SKEW.	-1.2900	0.8882	0.1338	0.1181	-0.6630	-0.4749	0.0623	-0.9586
KURT.	4.7608	3.4096	2.2423	3.1763	4.1343	3.5067	2.5335	5.0758
J.B.	20.327*	6.9232*	1.3442	0.1809	11.6730*	4.4432	0.8936	30.6062*

S.D. : STANDARD DEVIATION.

SKEW. : COEFFICIENT OF SKEWNESS.

KURT. : COEFFICIENT OF KURTOSIS.

J.B. : JARQUE BERA NORMALITY TEST.

* : SIGNIFICANT AT THE 5% LEVEL.

the residuals of equation (9) fail this test in the case of the sterling-dollar and lira-dollar exchange rates. These findings are probably due to a high degree of kurtosis (compounded by some skewness in the first half of the sample). One possible source of leptokurtosis lies in the ARCH structure of the residuals.

We have therefore fitted the following ARCH model to the data

$$y_{t+1} = Z_{t+1}b + u_{11t+1} \quad (11)$$

(equation (11) represents either equation (7) or equation (9) above)

$$E[u_{11t+1}|\bar{w}_t] = 0, E[u_{11t+1}^2|\bar{w}_t] = h_{t+1}$$

$$h_{t+1} = \alpha_0 + \sum_{i=1}^p \alpha_i u_{11t+1-i}^2 \quad (12)$$

$$\sum_{i=1}^p \alpha_i < 1, \alpha_i \geq 0, i = 1, \dots, p.$$

y_{t+1} is the dependent variable, Z_{t+1} is a vector of explanatory variables included in Φ_t and b is a vector of coefficients. R.W. Rich, J. Raymond and J.S. Butler (1991) suggest that the ARCH model be written in innovation form as follows

$$u_{11t+1} = y_{t+1} - Z_{t+1}b \quad (11')$$

$$\eta_{t+1} = (y_{t+1} - Z_{t+1}b)^2 - \alpha_0 - \sum_{i=1}^p \alpha_i (y_{t+1-i} - Z_{t+1-i}b)^2 \quad (13)$$

$$\eta_{t+1} = u_{11t+1}^2 - h_{t+1}, E[\eta_{t+1}|\Phi_t] = 0,$$

and estimated with the help of a generalized method of moments estimation procedure. The innovations to the ARCH model at time $t+1$ are uncorrelated with the information available at time t . Thus, if s_{1t} and s_{2t} denote respectively a $1 \times k_1$ and a $1 \times k_2$ vector of observable variables that enter the agents' information set at time t , we obtain the following (k_1+k_2) orthogonality conditions implied by the model

$$E[u_{11t+1}s_{1t} \quad \eta_{t+1}s_{2t}] = 0.$$

The use of this procedure has three advantages: (i) it is compatible with the estimation of the coefficients of the unobservable variables which appear on the right hand side of equations (7) and (9), which describe the behaviour of the

conditional mean of the distribution of the risk premium; (ii) it does not require a distributional assumption for the innovations to the ARCH model; (iii) it allows for the implementation of the specification tests developed by L.P. Hansen (1982) and W.K. Newey (1985) and of the stability tests developed by D.W.K. Andrews (1993).¹¹ The use of the GMM estimation procedure in an ARCH context can be found also in N.C. Mark (1988), W.E. Ferson (1989), D.P. Simon (1989) and J.N. Bodurtha, N.C. Mark (1991), among others.

In the estimation below we assume that s_{1t} includes the instruments used in the GMM estimates of the (conditional mean) equations (7) and (9) above and that s_{2t} includes the lagged squared residuals which enter the specification of the conditional variance. s_{1t} and s_{2t} are used as instruments for the regressors in the ARCH model. Following I. Domowitz, C.S. Hakkio (1985) we use squares to impose the non-negativity constraints on the parameters of the variance equations and estimate equations of the form

$$u_{11t+1}^2 = \alpha_0^2 + \sum_{i=1}^p \alpha_i^2 u_{11t+1-i}^2 + \eta_{t+1} \quad (14)$$

Table 7 reports the ARCH model description of the distributions of the forward forecast errors $(s_{t+1} - f_t)$ in the period January 1981 - February 1985, conditional on the information set Φ_t , with mean $a_0 + a_1 \Delta d_{t+1}$ and variance $h_{t+1} = E[u_{11t+1}^2|\Phi_t]$. Parameter estimates for the variance equations are reported in terms of the underlying unsquared parameters. The quality of fit of the conditional mean of the distribution of $(s_{t+1} - f_t)$ is good: the estimates of a_0 and a_1 are similar to those reported in table 1 above and can be given the same

TABLE 7

$$(s_{t+1} - f_t) = a_0 + a_1 d_{t+1} + u_{7t+1} \quad (7)$$

$$u_{7t+1}^2 = \alpha_0^2 + \alpha_1^2 u_{7t}^2 + \alpha_2^2 u_{7t-1}^2 + \eta_{t+1} \quad (7')$$

1981:1-1985:2

a_0	a_1	α_0	α_1	α_2	Q(6)	H _a	H ₀	SupW	n	SK.	KURT.	J.B.
L/\$												
-0.002	1.04	0.004	0.67	-	6.80	1.47	-	3.08	0	-0.76	3.17	4.76
(-2.00)(21.44) (3.08) (4.35) - (5.64)												
DM/\$												
-0.003	0.97	0.003	-0.0001	-0.41	6.62	3.79	4.25	2.82	0	0.11	3.50	0.60
(-7.23)(33.46) (6.96) (-1.1x10 ⁻⁷)(-2.07) (6.09)												
Y/\$												
-0.002	1.07	0.01	-0.001	0.37	13.17*	2.22	2.16	2.80	0	0.16	2.25	1.35
(-1.97) (8.22) (2.64) (-1.2x10 ⁻⁶)(1.42) (9.91)												
LIT/\$												
0.001	1.15	0.01	0.0002	-0.02	11.42	0.69	0.01	6.40	0	0.14	3.01	0.15
(0.66)(11.66) (3.12) (0.4x10 ⁻⁶)(-0.01) (1.66)												

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

H_a : HANSEN'S TEST OF OVERIDENTIFYING RESTRICTIONS.

H₀ : WALD TEST OF THE NULL HYPOTHESIS $\alpha_1 = \alpha_2 = 0$.

SupW : WALD TEST FOR PARAMETER STABILITY WITH UNKNOWN CHANGE POINT.

Q(6) : BOX-PIERCE Q-STATISTIC. SKEW. : COEFFICIENT OF SKEWNESS.

KURT. : COEFFICIENT OF KURTOSIS. J.B. : JARQUE BERA NORMALITY TEST.

AR(4) : TEST FOR FOURTH ORDER ARCH. * : SIGNIFICANT AT THE 5% LEVEL.

economic interpretation. The conditional variance seems to be compatible with an ARCH parameterization in the case of the sterling-dollar and (possibly) of the deutschmark-dollar and

yen-dollar exchange rates (even if the t-ratios of the coefficients of the lagged squared residuals of the latter estimates are low). We cannot reject the hypothesis that the standardized residuals are normally distributed as the Jarque Bera statistics are no longer significant at the 5% level. The

TABLE 8

$$(s_{t+1} - f_t) = b_0 + b_1(r_{t+1}^2 - r_{t+1}) + u_{9t+1} \quad (9)$$

$$u_{9t+1}^2 = \alpha_0^2 + \alpha_1^2 u_{9t}^2 + \alpha_2^2 u_{9t-1}^2 + \alpha_3^2 u_{9t-2}^2 + \eta_{t+1} \quad (9')$$

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b_0	b_1	α_0	α_1	α_2	α_3	Q(6)	H ₀ (H ₁)	SupW	n	SK.	KURT.	J.B.
L/\$												
0.004	0.52	0.04	0.34	-	-	3.19	-	1.68	1	-0.47	3.59	4.73
(1.05) (1.73) (7.56) (2.55) - - (5.13) (2.73)												
DM/\$												
0.006	0.60	0.04	-6.6x10 ⁻⁶	-	-	4.09	-	2.52	1	-0.48	3.48	4.35
(1.99) (2.54)(10.13) (-6.1x10 ⁻⁸) - - (8.10*) (2.72)												
Y/\$												
0.006	0.01	0.04	-1.1x10 ⁻⁶	0.28	-	4.79	2.18	4.04	0	0.03	2.50	0.94
(1.72) (0.03)(10.14)(-2.4x10 ⁻⁹)(1.45) - (7.53) (0.85)												
LIT/\$												
0.006	0.50	0.03	0.22	0.17	0.44	4.28	7.88*	11.89	1	-0.89	4.91	25.16*
(1.64) (2.15) (4.84) (0.90) (0.55) (2.27) (7.26) (4.09)												

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

H_a : HANSEN'S TEST OF OVERIDENTIFYING RESTRICTIONS.

H₀ (H₁) : WALD TEST OF THE NULL HYPOTHESIS $\alpha_1 = \alpha_2 (= \alpha_3) = 0$.

SupW : WALD TEST FOR PARAMETER STABILITY WITH UNKNOWN CHANGE POINT.

AR(4) : TEST FOR FOURTH ORDER ARCH. * : SIGNIFICANT AT THE 5% LEVEL.

time dependent heteroskedasticity is successful in explaining the nonnormality that affected the residuals of the sterling-dollar and deutschmark-dollar estimates of equation (7) set out in table 1 above.

Table 8 reports the ARCH model representation of the distribution of $(s_{t+1} - f_t)$ in the period March 1985 - October 1992 conditional on the information set Φ_t , with mean $b_0 + b_1(r_{t+1}^* - r_{t+1})$ and variance $h_{t+1} = E[u_{t+1}^2 | \Phi_t]$. The quality of fit of the conditional mean has improved, if compared to the estimates of table 3 above, especially in the case of the sterling-dollar exchange rate. The link between the exchange risk premium and the financial risk premium seems to be stronger. The conditional variance has an ARCH structure in the case of the sterling-dollar and of the lira-dollar exchange rates. The nonnormality that affected the residuals of the estimates of table 3 above has been eliminated in the case of the sterling-dollar exchange rate and only reduced in the case of the lira-dollar exchange rate (even if the standardized residuals exhibit no ARCH), a finding that does not affect the consistency of the GMM estimator.

Conclusion

The model originally set out by R.A. Korajczyk (1985) assumes that the rejection of the unbiased expectations hypothesis is due to the risk premia reflected in the real interest rate differentials on riskless nominal assets denominated in different currencies, in line with the findings of the international asset pricing models of P.J.K. Kouri (1977), E.F. Fama, A. Farber (1979), R.J. Hodrick (1981) and

others. This approach predicts that the time-varying forward risk premium (i.e. the forecastable difference between forward and future spot exchange rate) should be equal to the risk premium differential caused by different levels of purchasing power risk across currencies. It is not strongly supported by the data as ex ante real interest rate differentials are not reflected in systematic forward forecast errors in the case of the yen-dollar and (possibly) the sterling-dollar exchange rates. In the period March 1985 - October 1992 the conditional means of the lira-dollar and deutschmark-dollar forward forecast errors are affected by ex ante real interest rate differentials and the conditional variances of the sterling-dollar and lira-dollar forecast errors can be modelled as ARCH processes. We have also 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 the estimates is that the four currencies under investigation paid a risk premium vis a' vis the dollar because of the expected real appreciation of the latter. Indeed, the expected dollar appreciation - in excess of the inflation rate differentials - brought about an increase of the risk premium (and a decrease in $E[(s_{t+1} - f_t) | \Phi_t]$), in line with the negative covariation between the risk premium and the expected spot exchange rate depreciation identified by E.F. Fama (1984) and justified by R.J. Hodrick, S. Srivastava (1986) in the context of the Lucas intertemporal asset pricing framework. The relevance of the ARCH parameterization of the conditional variance of the expected forward forecast errors seems to be weaker than in the second half of the sample.

Appendix

Generalized Method of Moments Estimation Methodology

Consider the linear model

$$y_t = Z_t b + u_t \quad (A.1)$$

where y_t is the dependent variable, Z_t is a $1 \times q$ vector of explanatory variables, b is a $q \times 1$ vector of coefficients and assume that X_t is a $1 \times r$ vector of instruments. The orthogonality conditions can be written as

$$E[X_t'(y_t - Z_t b)] = 0. \quad (A.2)$$

$\hat{b}_T$, the GMM estimate of b , is the value of b that minimizes the objective function

$$g_T(b)' \hat{Q}_T^{-1} g_T(b) \quad (A.3)$$

where $g_T(b) = \frac{1}{T} \sum_{t=1}^T X_t'(y_t - Z_t b)$ and $\hat{Q}_T$ is a consistent estimate of the matrix of orthogonality conditions (or weighting matrix)

$$Q = \lim_{T \rightarrow \infty} \frac{1}{T} \sum_{t=1}^T \sum_{v=-\infty}^{\infty} E[X_t'(y_t - Z_t b) X_{t-v}'(y_{t-v} - Z_{t-v} b)']$$

$\hat{b}_T$ can be estimated as

$$\hat{b}_T = [(\sum_{t=1}^T Z_t' X_t) \hat{Q}_T^{-1} (\sum_{t=1}^T X_t' Z_t)]^{-1} [(\sum_{t=1}^T Z_t' X_t) \hat{Q}_T^{-1} (\sum_{t=1}^T X_t' y_t)] \quad (A.4)$$

and the corresponding asymptotic variance - covariance matrix

can be estimated as

$$\hat{V}_T = T^2 [(\sum_{t=1}^T Z_t' X_t)^{-1} \hat{Q}_T (\sum_{t=1}^T X_t' Z_t)^{-1}]. \quad (A.5)$$

In the presence of serially correlated disturbances the orthogonality conditions will be affected by a moving average structure. The standard GMM estimates will be consistent but not asymptotically efficient if the weighting matrix is not appropriately modified (and asymptotic confidence interval formation and hypothesis testing will be biased). We use therefore the positive semi-definite heteroskedasticity and autocorrelation consistent covariance matrix estimator of W.K. Newey, K.D. West (1987)

$$\hat{Q}_T = \Lambda_{0T} + \sum_{v=1}^h \left(1 - \left(\frac{v}{h+1}\right)\right) [\Lambda_{vT} + \Lambda_{vT}'], \quad (A.6)$$

where $\Lambda_{vT} = \frac{1}{T} \sum_{t=v+1}^T u_t u_{t-v} X_t' X_{t-v}$ and $u_t = y_t - Z_t b$.

The bound on the number of sample autocovariances h has been assessed with the help of the automatic bandwidth parameter selection procedure set out by D.W.K. Andrews (1991).¹² Alternative heteroskedasticity and autocorrelation consistent covariance matrix estimators have been provided by R.E. Cumby, J. Huizinga, M. Obstfeld (1983), H. White (1984, pp. 147-161), A.R. Gallant (1987, p. 533) and D.W.K. Andrews, J.C. Monahan (1992).

notes

(1) See, among others, the international equilibrium asset

pricing models of F.L.A. Grauer, R.H. Litzenberger, R.E. Stehle (1976), P.J.K. Kouri (1977), E.F. Fama, A. Farber (1979), R. Roll, B.H. Solnik (1979), J.A. Frankel (1979), R.J. Hodrick (1981), R.M. Stulz (1981) and R.E. Lucas (1982).

(2) Previous empirical investigations, following a pathbreaking study by L.P. Hansen, R.J. Hodrick (1983), can be found in R.J. Hodrick, S. Srivastava (1984), N.C. Mark (1985, 1988), I. Domowitz, C.S. Hakkio (1985) and R.D. Huang (1989). Additional analyses - that are not based, however, on the international asset pricing framework - have been performed by J.A. Frankel (1982) and M.P. Dooley, P. Isard (1983), among others.

(3) The above mentioned relationships

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

$$d_t = s_t - P_t + P_t^*$$

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

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

where i_{t+1} (i_{t+1}^*) is the yield - from t to $t+1$ - on one period domestic (foreign) riskless assets, P_t (P_t^*) is the logarithm of the domestic (foreign) price level and d_t is the logarithm of the real exchange rate, can be combined to express the ex post forward forecast error as

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

Taking expectations we obtain equation (2) of the text.

(4) Data are sampled monthly. The forward and spot exchange rate time series have been provided by the Bank of Italy. We have used - with the exception of Italy - Eurocurrency interest rates, which are less affected by capital controls. These 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". The consumer prices and the Italian interbank interest rate time series are published in the O.E.C.D. "Main Economic Indicators". Consumer prices are used to derive the inflation rates series.

(5) We could not assess the joint effect of ex ante real interest rate differentials and of ex ante real exchange rate changes on ex ante forward forecast errors because of a severe multicollinearity problem.

(6) Conditional heteroskedasticity may well be present in monthly exchange rate time series. A feature of the GMM approach is that no auxiliary assumption of conditional homoskedasticity need be made in constructing the covariance matrix of parameter estimates.

(7) The alternative hypothesis is that a break has occurred at some time t such that $t/T = \pi \in \Pi$, where Π is some pre-specified subset of $[0, 1]$ whose closure lies in $(0, 1)$ and T is the sample size. $\text{Sup}W = \sup_{\pi \in \Pi} W(\pi)$, where $W(\pi)$ is a Wald statistic for known structural break. When $\Pi = [.15, .85]$, the null hypothesis of parameter stability is rejected if the statistic exceeds the 5% critical values of 11.79, 14.15, 16.45 and 18.35, respectively with two, three, four and five degrees of freedom, provided by D.W.K. Andrews (1993), page 840.

(8) This finding reflects the systematic underprediction of the dollar appreciation from 1981 to 1985. It has been investigated - among others - by G.W. Evans (1986), J.A. Frankel, K.A. Froot (1987) and K.K. Lewis (1988, 1989).

(9) These variables had been shown by L.P. Hansen, R.J. Hodrick (1980, 1982), J.E.O. Bilson (1981), R.D. Huang (1989)

and others to affect the forward forecast error.

(10) We want to test the hypothesis that when the model is misspecified $E(X_{1t}'u_t) = 0$ and $E(X_{2t}'u_t) = c$, where X_{1t} is a $1 \times (r - s)$ vector of uncontaminated instruments, X_{2t} is a $1 \times s$ vector of contaminated instruments such that $X_t = (X_{1t}, X_{2t})$ is $1 \times r$. u_t denotes the residuals of the GMM estimates of $y_t = Z_t b + u_t$ where y_t is the dependent variable and Z_t is a $1 \times q$ vector of explanatory variables.

The Newey optimal GMM statistic for misspecification of this form is $m_T = T(L_T g_T(b))' Q_T^{-1} (L_T g_T(b))$, where $L_T = (-V_{T21} V_{T11}^{-1}, I_s)$ and $g_T(b) = (\frac{1}{T} \sum_{t=1}^T u_t X_{1t}', \frac{1}{T} \sum_{t=1}^T u_t X_{2t}')'$. $Q_T = V_{T22} - V_{T21} V_{T11}^{-1} V_{T12} - B_T (H_T' V_T^{-1} H_T)^{-1} B_T'$, where $H_T = (H_{T1}', H_{T2}')' = (-\frac{1}{T} \sum_{t=1}^T Z_t' X_{1t}, -\frac{1}{T} \sum_{t=1}^T Z_t' X_{2t})'$ and $B_T = (H_{T2} - V_{T21} V_{T11}^{-1} H_{T1})$. V_{T11} , V_{T12} , V_{T21} and V_{T22} are obtained by partitioning the asymptotic covariance matrix of the orthogonality conditions V_T conformably with the vector of moments function $g_T(b)$. A modified version of this test has been developed by J.N. Bodurtha and N.C. Mark (1991).

(11) Usually the maximum likelihood estimation of ARCH models is performed under the assumption that the conditional distribution of the residual is either normal or Student t. T. Bollerslev and J.M. Wooldridge (1992) suggest that in the case of nonnormality the QML estimator is generally consistent and has a limiting normal distribution. They point out, however, that the GMM approach can produce an instrumental variables estimator that is asymptotically more efficient.

(12) D.W.K. Andrews (1991) suggests the following procedure for the selection of a bandwidth (lag truncation) parameter. First, one estimates m univariate first order autoregressive AR(1) models for the m elements of the (Kronecker) product

of the residuals of the GMM model estimates with the vector of instruments (m is the number of orthogonality conditions). Then one computes

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

where $(\hat{\rho}_a, \hat{\sigma}_a^2)$, $a=1, \dots, m$ 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.

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