IMF Working Paper

© 1997 International Monetary Fund

This is a *Working Paper* and the author(s) would welcome any comments on the present text. Citations should refer to a *Working Paper of the International Monetary Fund.* The views expressed are those of the author(s) and do not necessarily represent those of the Fund.

WP/97/39

INTERNATIONAL MONETARY FUND

Research Department

Foreign Exchange Risk Premium: Does Fiscal Policy Matter? Evidence from Italian Data

Prepared by Lorenzo Giorgianni*

Authorized for distribution by Peter Wickham

March 1997

Abstract

This paper challenges the conventional view that foreign exchange risk premiums are small, not volatile, and unrelated to macroeconomic variables. For the Italian lira (1987-94), unconditional risk premiums—constructed using survey data to measure exchange rate expectations—are found to be sizable (relative to the dimension of the forward premium), highly volatile (relative to the variability of the forward bias), and predictable. Estimation of structural models of the risk premium suggests that anticipated fiscal contractions in Italy and lower uncertainty about the future path of fiscal policy are associated with a lower risk premium on lira-denominated assets.

JEL Classification Numbers: F31, G1, H6, C42

Keywords: Risk Premium, Fiscal Policy, Survey Data, Italy

Author's E-Mail Address: lgiorgianni@imf.org

*I would like to thank Leonardo Bartolini, Peter Clark, Peter Isard, Lutz Kilian, Joaquim Levy, Karen Lewis, Ronald MacDonald, Richard Marston and Peter Wickham for their comments, Albert Ando, Charles Engel, Robert Flood and participants of the S.A.DI.BA.-Bank of Italy conference in Perugia (Italy) for helpful discussions, Mr. Alan Teck for permission to use the *Financial Times Currency Forecaster* data, and Susanna Mursula for technical assistance.

Table of Contents

Appendix Tables:

Appendix Figures:

SUMMARY

Economists and policymakers have long been interested in the effects of fiscal policies on exchange rates. In particular, for countries with large fiscal imbalances, it is often contended that fiscal policy may affect exchange rates through a "risk premium channel," whereby a fiscal contraction would reduce the risk carried by assets denominated in domestic currency, increase their demand, and ultimately appreciate the domestic currency. However, formal tests of this conjecture are notably absent from the literature. This shortcoming may, of course, reflect that most previous attempts to measure and model risk premiums alone have proven inconclusive. Risk premiums generated by rational expectations models or inferred ex post through linear projection equations are typically found to be implausibly small and stable and unresponsive to macroeconomic variables, let alone fiscal variables.

This paper challenges these observations and provides evidence that currency risk premiums are sizable and responsive to anticipated fiscal policies. It further suggests that the implausible results offered by the existing literature may reflect shortcomings in the measurement of risk premiums in a world with unobservable preferences. In contrast to most earlier studies, this paper employs survey data to overcome the unobservability of risk premiums.

For the Italian lira (1987-94), unconditional risk premiums are found to be sizable (relative to the dimension of the forward premium), highly volatile (relative to the variability of the forward bias), and predictable. Estimation of structural models of the risk premium suggests that anticipated fiscal contractions in Italy and lower uncertainty about the future path of fiscal policy are associated with a lower risk premium on lira-denominated assets.

I. Introduction

Economists and policymakers have long been interested in the effects of fiscal policies on the foreign exchange value of national currencies. Despite this long-standing interest, a recent study of the International Monetary Fund (1995), surveying the available evidence on this matter, notes the absence of clear predictions of the effects of fiscal policy on exchange rates. Both in that research and from anecdotal evidence, however, the view that for countries with large fiscal imbalances, like Belgium, Italy or Sweden, fiscal policy may affect exchange rates through a "risk premium channel" is strongly advocated. According to this view, a credible fiscal contraction in a country with a large stock of public debt, may produce two effects: first, it reduces the amount of government debt held by domestic and foreign investors and, second, it lowers uncertainty about future taxation and debt management policies, which render the domestic economy less vulnerable to external shocks. In a world populated by risk averse investors, both effects would produce an *increase* in the demand for home-currency denominated assets, thereby, easing or overturning the depreciating pressure on the domestic currency triggered by the initial fiscal contraction and arising from lower domestic interest rates.

Although the view that a fiscal stabilization should reduce currency risk is widely held in policy circles, formal tests of this conjecture are absent from the literature. This shortcoming may, of course, reflect that most previous attempts to measure and model risk premiums alone have proven inconclusive (see Engel (1995) for a comprehensive survey). Risk premiums generated by rational expectations asset pricing models, in particular, are typically implausibly small and stable.

This paper challenges both observations and provides evidence that currency risk premiums may be sizable and responsive to anticipated fiscal policies. It further suggests that the implausible results offered by the existing literature may reflect shortcomings in the measurement of risk premiums in a world with unobservable preferences. In contrast with most previous research, this study relies on a model-free measure of the risk premium, obtained by proxying the unobserved expectations of future spot rates with survey data. The directly constructed series for the risk premium are then used to assess the economic and statistical significance of risk premiums and, more importantly, to verify the significance of the link between fiscal policy and the risk premium.

Data for the empirical analysis are drawn from what seems an ideal laboratory to analyze the effects of fiscal policy on the currency risk premium, the market for the Italian lira from 1987 to 1994. During this period, Italy experienced a severe deterioration of its fiscal position and a fast growing government debt-to-GDP ratio (which stands as the highest among the G7 countries), while, at the same time, developing open domestic and offshore capital markets and experiencing a major exchange rate shock in 1992.

In anticipation of the findings of the present investigation, Italian lira risk premiums measured with survey data are found to be sizable --in excess of 1.5 percent per annum--, predictable and highly volatile. There is also evidence that some of the observed variability in the risk premium is systematically related to macroeconomic variables. In particular, anticipated fiscal contractions in Italy (fiscal expansions in the United States or Germany) and lower uncertainty on the future path of Italian fiscal policy (higher uncertainty on foreign fiscal policy) are found to reduce the risk premium on lira-denominated assets.

The paper is organized as follows. Section II reviews the notion of currency risk and illustrates the issues involved in the measurement of risk premiums (see also Appendix II). Section III provides empirical support for the hypothesis that Italian lira risk premiums are predictable (see also Appendix III). Motivated by this finding, the rest of the paper proposes and tests different structural models of the risk premium. Section IV describes existing theoretical approaches to modeling risk premiums, including a general equilibrium asset pricing model and the portfolio balance model. Empirical variants of these models are tested using Italian data and the results are presented in Section V, where the empirical significance of the link between the risk premium and fiscal policy is discussed. The main results are summarized and put into perspective in the closing section. Data description is relegated to Appendix I.

II. Measurement of the Foreign Exchange Risk Premium

A. Preliminary Definitions

The foreign exchange risk premium represents the compensation required by risk averse investors for holding an asset whose only risk depends on being issued in a particular currency. It equals the ex ante excess return from forward speculation, that is, the logarithmic difference between the time-t forward exchange rate for delivery at time $t+k$, $f_{t,k}$, and the k-step ahead expected future spot rate formed at the end of time t, $s^e_{t+k|t}$:[1]

$$er^e_{t,k} = f_{t,k} - s^e_{t+k|t}. \tag{1}$$

In equation (1) and in the remainder of the paper spot and forward rates are measured in units of domestic currency per foreign currency.

By Covered Interest Parity (CIP), equation (1) can be decomposed into an interest component and a capital gain. According to CIP,[2] the forward rate equalizes the sum of the

[1]In the remainder of the study the expressions "currency risk premium," "ex ante excess returns" and "expected excess returns" will be used interchangeably.

[2]Which holds in absence of capital restrictions or sizable transaction costs.

current spot rate *plus* the interest rate differential between comparable domestic and foreign offshore assets denominated in foreign currency:

$$f_{tk} = s_t + (r_{t,k} - r^*_{t,k}). \tag{2}$$

Here $r_{t,k}$ and $r^*_{t,k}$ are, respectively, the continuously compounded domestic and foreign Eurodeposit interest rates.[3]

Using equation (2), an equivalent and more intuitive definition of the ex ante excess return is:

$$er^e_{t,k} = (r_{t,k} - r^*_{t,k}) + [-(s^e_{t+k|t} - s_t)], \tag{3}$$

i.e., the negative of the residual from Uncovered Interest Parity. This definition highlights two components: the first, the interest rate differential, reflecting the interest gain, and the second, the expected appreciation rate, representing the expected capital gain. From equation (3), the expected excess return can be seen as the compensation demanded by foreign investors to be indifferent between investing one unit of foreign currency in assets denominated in foreign currency, earning $r^*_{t,k}$ with certainty, and one unit of foreign currency in assets denominated in home currency, expected to earn $r_{t,k} - (s^e_{t+k|t} - s_t)$.

B. Measurement Issues

A crucial difficulty involved in constructing risk premiums from either equation (1) or equation (3) is that expected future spot rates are unobservable. Previous literature has resolved this measurement problem either by assuming a model of formation of expectations, by which unobservable *ex ante* variables, e.g. expectations, are inferred from observable *ex post* data, or by proxying expectations with data extracted from opinion surveys. Under the usual assumption of rational expectations, the expected future exchange rate is equal to the mathematical expectation of the future spot rate conditional on information available at time t, i.e.:

$$s^e_{t+k|t} = \mathrm{E}_t(s_{t+k}). \tag{4}$$

Thus, the rational expectations risk premium can be expressed as:

$$er^{RE}_{t,k} = \mathrm{E}_t(f_{t,k} - s_{t+k}). \tag{5}$$

---

[3]Eurodeposit interest rates are used because CIP holds continuously in these markets. In line with the institutional features of Eurodeposit markets, interest rates are compounded continuously and are expressed in per-holding period, i.e., $r_{t,k} = \ln(1 + \frac{k}{12} R_{t,k})$, where $R_{t,k}$ is the annualized percentage nominal rate at the end of period t on Eurodeposits denominated in domestic currency maturing at the end of period $t+k$.

A data-based approach to estimate $er_{t,k}^{RE}$ consists in regressing ex post excess returns, the so-called forward bias, $f_{t,k} - s_{t+k}$, on a vector of variables belonging to investors' time-t information set, denoted by the vector $\mathbf{\Omega}_t$:[4]

$$f_{t,k} - s_{t+k} = \boldsymbol{\beta}\mathbf{\Omega}_t' + \varepsilon_{t+k}, \tag{6}$$

where $\boldsymbol{\beta}$ is a vector of regression coefficients. The predicted values from equation (6), provided that the estimated errors are well behaved, constitute *a* measure of the "rational expectations risk premium." (See Fama (1984), Cumby (1988), Canova and Marrinan (1993), and Lewis (1994), for different applications).

This approach has severe limitations: first, it relies on the econometrician having identified the correct exchange rate model,[5] as well as the correct information set available to investors at time t, $\mathbf{\Omega}_t$. And, second, it assumes that the underlying expectation model is stable, or that convergence to the rational expectations equilibrium in response to shocks (e.g., policy shocks) is instantaneous. Conversely, if agents learn only gradually about past policy switches, it is incorrect for the econometrician to assume this change to be part of investors' information set as soon as it occurs and, therefore, treat the regression coefficients $\boldsymbol{\beta}$ as time-invariant. It is also incorrect to assume stability of $\boldsymbol{\beta}$ in the presence of a "peso problem" in the data, i.e., the fact that agents assign a low probability to the possible occurrence of a large shift in the determinants of exchange rates, which does not materialize within the sample. The phenomena of learning or peso problems have been shown (see Kaminsky (1993), and Lewis (1994)) to produce small-sample systematic expectational errors. Failure to acknowledge their presence may bias the measure of the risk premium obtained in finite samples by regressions like equation (6) (this point is illustrated with an example in the Appendix II).

Also, when using ex post data to estimate equations like (6) it becomes impossible to test for the existence of the risk premium, because it is not clear how one would distinguish, say, the presence of predictable excess returns from an autocorrelated prediction error (see Engel (1995)).

In light of these difficulties, the present work advocates the use of opinion survey data to resolve the unobservability of exchange rate expectations. The use of survey data in studies of exchange rate expectations can be traced back to Frankel and Froot (1987). In the context of studies of the risk premium, survey data have also been used by Froot and Frankel (1989),

---

[4]Note that estimation of the rational expectations risk premium requires, in general, knowledge of the model economy according to which agents form consistent expectations (see Section IV). The use of linear projections as in (6) is just a data-based shortcut to produce a measure of the rational risk premium. It does not require prior knowledge of a model economy (including functional forms and model parameters), and relies only on a guess of the relevant variables contained in agents' information sets and the assumption of linearity.

[5]In this case, by CIP, equal to $s_{t+k} = s_t + (r_{t,t+k} - r^*_{t,t+k}) - \boldsymbol{\beta}\mathbf{\Omega}_t + \varepsilon_{t+k}$.

Taylor (1989), MacDonald and Torrance (1990), Liu and Maddala (1992), Chinn and Frankel (1994b), Cavaglia, Verschoor and Wolff (1994) and MacDonald (1994). Consensus forecasts, obtained from opinion surveys, provide model-free measures of aggregate expectations and are immune to the problems discussed above. In particular, the use of survey data frees the econometrician from having to identify agents' information set, and produces valid measures of risk premiums even in presence of learning or peso problems. However, tests based on consensus forecasts are only as good as the data themselves; in particular, they might be biased because of measurement errors, or because of heterogeneity of forecasters (as in Ito (1990)).

III. Time Series Properties of the Lira Risk Premium

One of the tasks of the present investigation is to assess whether risk premiums are sizable and highly volatile relative to the forward premium, and whether they possess a predictable component. In the past, many studies have addressed this question. Svensson (1992) argues that, for floating currencies, annualized risk premiums are on average not greater than one percent, and that, for currencies joining a credible target zone agreement, they should be even smaller. Engel (1995), who surveys the recent literature on the rational expectations risk premium, finds that the unconditional mean of the risk premium is insignificantly different from zero and that its upper bound (in absolute value) is small relative to the size of the forward premium (i.e., the interest rate differential). Fama (1984) finds that, under rationality of expectations, the volatility of the risk premium is larger than the volatility of the expected depreciation and that these two series are negatively correlated. These findings are also confirmed in studies of the risk premium based on standard rational expectations general equilibrium models (see Engel (1995) and in the empirical literature generated by these models (see, for example, Domowitz and Hakkio (1985), Cumby (1988) Baillie and Bollerslev (1990), Cheung (1993) and MacDonald (1994)). As I will show shortly, a different characterization of the time-series properties of the risk premium emerges when using survey data to approximate expectations. This analysis is conducted using lira/dollar and lira/mark spot and forward exchange rates of maturity 3 and 12 months, and exchange rate expectations from the *Financial Times Currency Forecaster* survey.[6] All data are plotted in Figure 1.

The plots and correlograms of the 3- and 12-month lira/dollar and lira/mark risk premiums derived from survey data are presented in Figure 2, while the main statistical summary measures are reported in Table 1 and Table 2. A summary of the main features of these data follows.

---

[6]Forward rates are constructed from CIP using Eurodeposit interest rates. For a description of these and other data, see Appendix I.

Figure 1.a Spot Rates, Expectations and Forward Premiums: Lira/Dollar

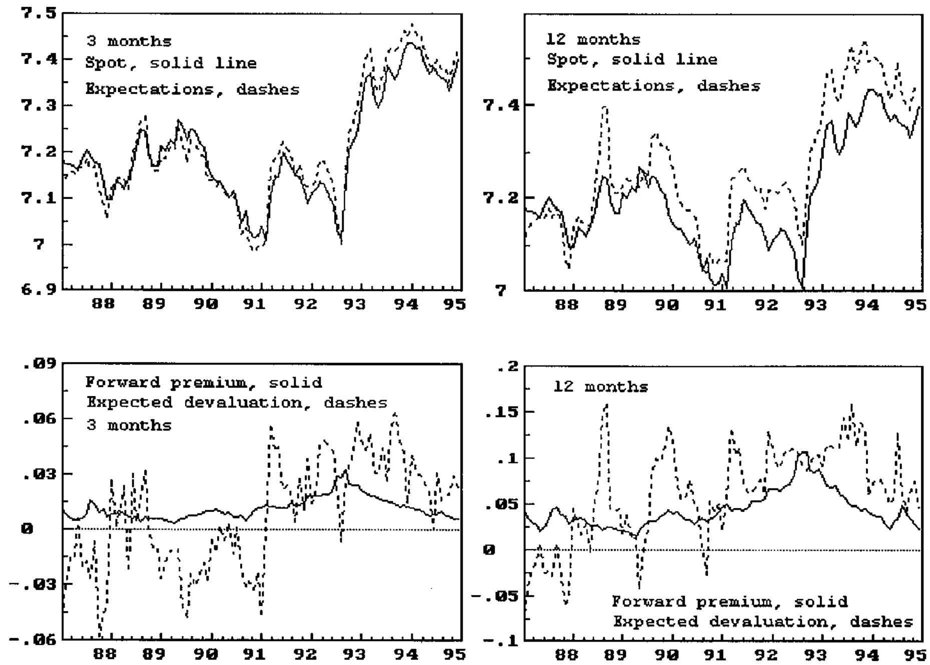
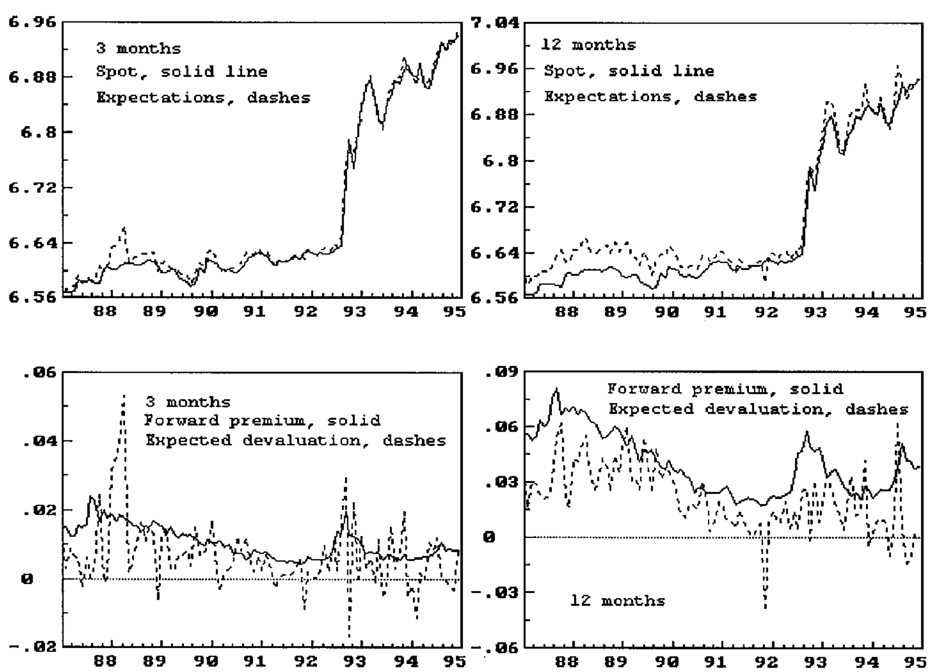


Figure 1.b Spot Rates, Expectations and Forward Premiums: Lira/Mark



Notes to figure:
In the top panels of Figures 1.a and 1.b, spot rates at time *t* (s_t) are plotted against expectations formed at time t for periods *t*+3 and *t*+12 ($s_{t+3|t}$ and $s_{t+12|t}$, respectively).

Figure 2.a   Survey-data Risk Premium and Correlograms: IL/US (1987-1994)

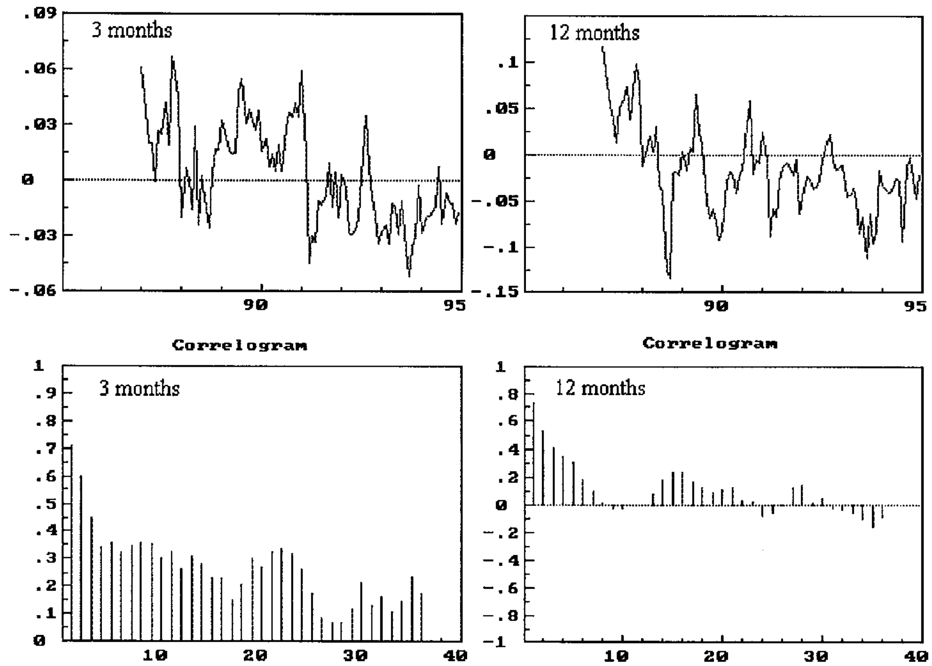


Figure 2.b   Survey-data Risk Premium and Correlograms: IL/GM (1987-1994)

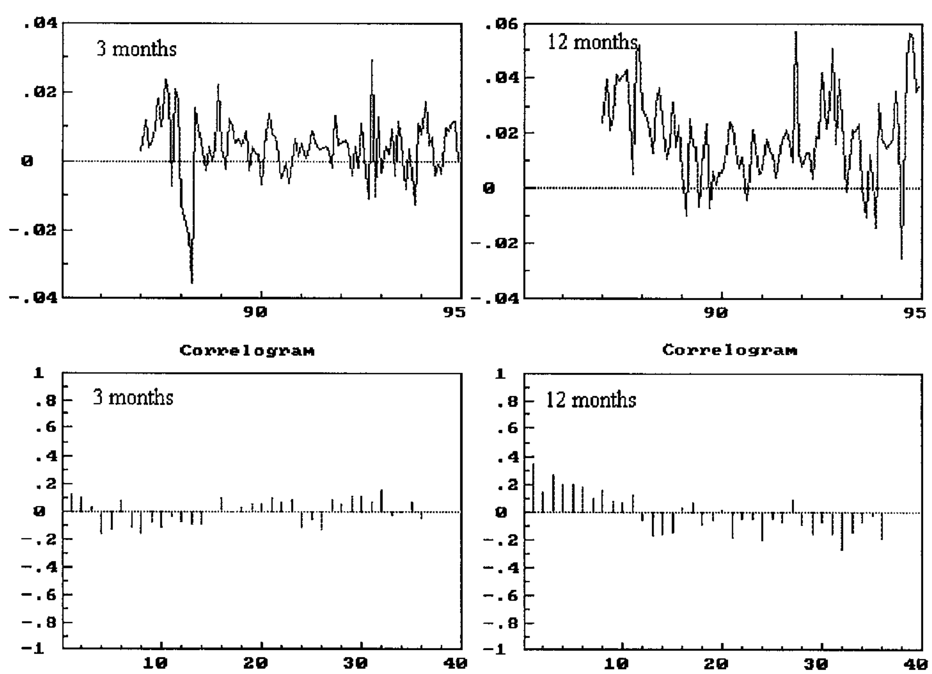


Table 1. Summary Statistics: Risk Premium ($er^{e}_{t,k}$), Annualized Percentage Values.

	Lira/Dollar		**Lira/Mark**	
1987.01-1994.12	**3 months**	**12 months**	**3 months**	**12 months**
Mean	1.48	–1.81	1.48	1.92
Standard Error for the Mean	*1.12*	*0.48*	*0.40*	*0.16*
Standard Deviation	10.84	4.71	3.36	1.59
Excess Kurtosis	–0.78	0.38	2.74	0.21
Skewness	0.28	0.26	–0.71	0.20
Normality Test (*M.S.L.*)	*0.053*	*0.32*	*0.00*	*0.50*
Autocorrelation(1)	0.71	0.73	0.12	0.34
Autocorrelation(2)	0.60	0.53	0.10	0.13
Autocorrelation(3)	0.44	0.40	0.03	0.26
Autocorrelation(4)	0.34	0.33	–0.16	0.20
Standard Error for Autocorrelation Coef.	*0.10*	*0.10*	*0.10*	*0.10*
Unit Root Test: *t*-ADF (5% critical value: -2.9)	-3.95	-4.76	-5.76	-7.19

Note: Figures in italics are either standard errors or marginal significance levels (M.S.L.)

Table 2a. Summary Statistics, Annualized Percentage Values: Lira/Dollar

Lira/Dollar					
3 months	$s_{t+k} - s_t$	$s^e_{t+k\|t} - s_t$	$r_{t,k} - r^*_{t,k}$	$er^e_{t,k}$	$f_{t,k} - s_{t+k}$
	1987.01 - 1994.09				
Mean	2.45	3.05	4.82	1.77	2.37
Standard Dev.	25.14	12.08	2.45	10.97	24.30
Corr($er^e_{t,k}$, •)		*–0.98*	*–0.36*		*0.04*
	1987.01 - 1992.08				
Mean	0.22	–1.00	4.27	5.28	4.05
Standard Dev.	25.44	11.17	2.16	10.38	24.98
Corr($er^e_{t,k}$, •)		*–0.98*	*–0.27*		*–0.03*
	1992.10 - 1994.09				
Mean	7.00	14.24	6.03	–8.21	–0.97
Standard Dev.	23.00	6.18	2.23	5.23	21.8
Corr($er^e_{t,k}$, •)		*–0.93*	*–0.25*		*0.15*
12 months					
	1987.01 - 1993.12				
Mean	3.09	5.99	4.42	–1.57	1.3
Standard Dev.	12.14	5.61	2.19	4.95	10.89
Corr($er^e_{t,k}$, •)		*–0.93*	*–0.12*		*–0.13*
	1987.01 - 1992.08				
Mean	2.25	4.80	3.93	–0.86	1.69
Standard Dev.	12.63	5.48	1.84	4.97	11.59
Corr($er^e_{t,k}$, •)		*–0.94*	*–0.10*		*–0.10*
	1992.10 - 1993.12				
Mean	5.69	11.27	6.23	–5.03	0.54
Standard Dev.	8.61	2.37	1.66	3.33	7.00
Corr($er^e_{t,k}$, •)		*0.74*	*–0.88*		*–0.72*

Note: $s_{t+k} - s_t$ is the realized devaluation rate;
$s^e_{t+k|t} - s_t$ is the expected (survey date) devaluation rate;
$r_{t,k} - r^*_{t,k}$ is the interest rate differential (or forward premium);
$er^e_{t,k}$ is the (survey data) risk premium;
$f_{t,k} - s_{t+k}$ is the forward bias.

All the variables are expressed in natural logarithms. Interest rates are continuously compounded.

Table 2b. Summary Statistics, Annualized Percentage Values: Lira/Mark

Lira/Mark					
3 months	$s_{t+k}-s_t$	$s^e_{t+k\|t}-s_t$	$r_{t,k}-r^*_{t,k}$	$er^e_{t,k}$	$f_{t,k}-s_{t+k}$
	1987.01 - 1994.09				
Mean	4.74	2.76	4.21	1.46	–0.53
Standard Dev.	12.75	4.13	1.86	3.82	12.78
Corr($er^e_{t,k}$, •)		*–0.89*	*0.07*		*0.12*
	1987.01 - 1992.08				
Mean	3.32	3.06	4.57	1.50	1.24
Standard Dev.	11.11	4.09	1.94	3.83	11.22
Corr($er^e_{t,k}$, •)		*–0.88*	*0.11*		*0.07*
	1992.10 - 1994.09				
Mean	7.76	1.53	3.06	1.52	–4.70
Standard Dev.	15.82	3.72	0.83	3.79	15.51
Corr($er^e_{t,k}$, •)		*–0.97*	*0.19*		*0.17*
12 months					
	1987.01 - 1993.12				
Mean	4.57	2.34	4.15	1.80	–0.44
Standard Dev.	6.87	1.67	1.72	1.47	7.64
Corr($er^e_{t,k}$, •)		*–0.40*	*0.46*		*0.08*
	1987.01 - 1992.08				
Mean	4.14	2.45	4.29	1.84	0.14
Standard Dev.	7.36	1.73	1.81	1.40	8.25
Corr($er^e_{t,k}$, •)		*–0.34*	*0.44*		*0.11*
	1992.10 - 1993.12				
Mean	5.93	1.83	3.38	1.54	–2.56
Standard Dev.	3.31	1.35	0.97	1.78	3.05
Corr($er^e_{t,k}$, •)		*–0.84*	*0.66*		*–0.09*

Note: $s_{t+k}-s_t$ is the realized devaluation rate;
$s^e_{t+k|t}-s_t$ is the expected (survey date) devaluation rate;
$r_{t,k}-r^*_{t,k}$ is the interest rate differential (or forward premium);
$er^e_{t,k}$ is the (survey data) risk premium;
$f_{t,k}-s_{t+k}$ is the forward bias.
All the variables are expressed in natural logarithms. Interest rates are continuously compounded.

Figure 3.a Correlograms: Lira/Dollar (1987-1994)

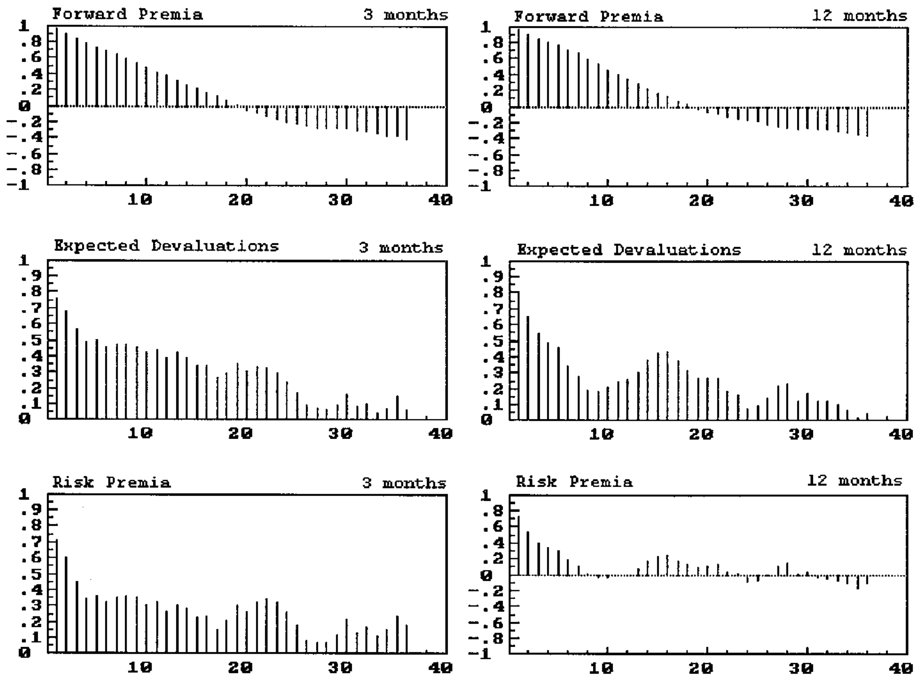
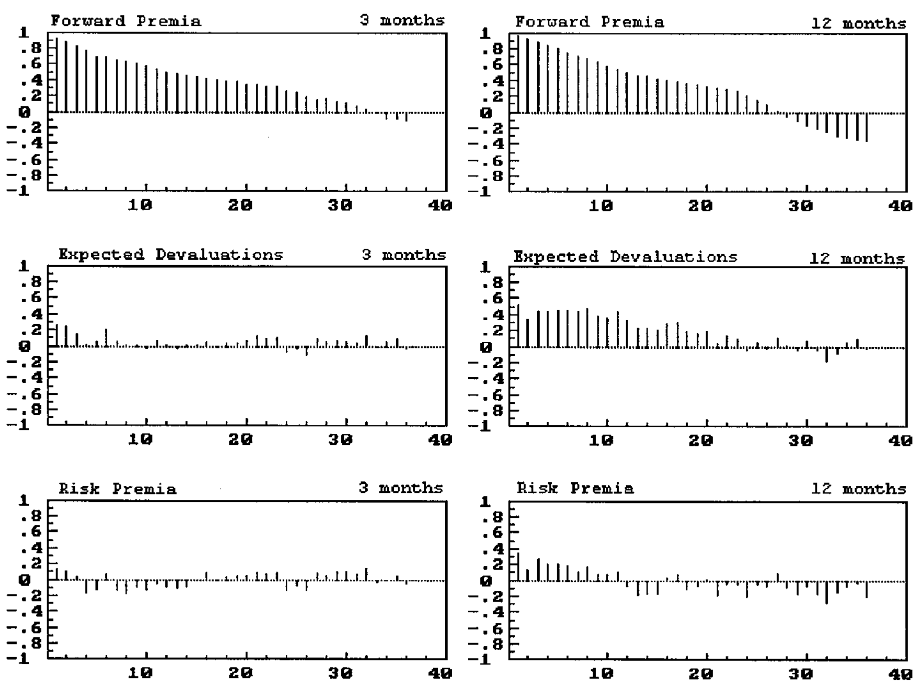


Figure 3.b Correlograms: Lira/Mark (1987-1994)



First, the unconditional mean of the lira/mark risk premiums at 3 and 12 months are, respectively, 1.5 percent and 1.9 percent per annum; and for the lira/dollar, 1.5 percent and –1.8 percent.[7,8] Contradicting the common wisdom and in particular Svensson's (1992) estimates, risk premiums are sizable, with an annualized average upper bound across maturities (in absolute terms) of 16 percent for the lira/dollar and 7.5 percent for the lira/mark. It is also interesting to notice that the unconditional mean of the lira/mark risk premium changes only slightly across the different institutional regimes; i.e., during the ERM (Exchange Rate Mechanism) period (1987.01-92.08), the 12-month risk premium has an annualized average of 1.8 percent (2.1 percent during the large band regime and 1.5 percent during the narrow band regime) and of 2.0 percent outside it (see Table 2.b). Also the lira/dollar risk premium displays significant unconditional mean nonconstancy.[9] For example, the 12-month lira/dollar premium is close to zero in the period preceding the end of the Gulf war and the associated mini oil crisis (March 1991), and becomes significantly negative thereafter (–3.6 percent).

The volatility of the risk premium is higher for the lira/dollar than the lira/mark and, for this last exchange rate, it is higher outside the ERM than inside (see Table 2.b). These observations are consistent with Svensson (1992), who suggests that the risk premium has mean and variance higher for floating currencies than for currencies belonging to a credible peg.

Second, although spot rates, expectations and interest rates taken individually appear to be nonstationary,[10] Augmented Dickey-Fuller (ADF) tests on $er^{e}_{t,k}$, a linear combination of the above variables, reject the unit root hypothesis at the usual 5 percent significance level for

---

[7]Notice that the sample means for the 3 month lira/dollar risk premium is insignificantly different from zero. The figures reported in Table 1 are similar to the ones obtained by Cavaglia *et al.*, (1993) with a different survey data set.

[8]The presence of a negative lira/dollar risk premium, i.e., the perception that dollar assets are "riskier" than lira-denominated assets, may appear surprising. A closer look, however, reveals this anomaly to be coherent with the data. The presence of a very large and negative risk premium on the mark/dollar exchange rate (–5.6 percent for the same time period), pulls down the average lira/dollar risk premium, given that the lira/mark risk premium cannot be too high because these currencies belong (at least, for part of the sample) to a credible peg. This observation is also confirmed in a study by Favero *et al.* (1996), where it is contended that lira/mark excess returns may be driven not only by local factors, but also by international factors, i.e., by the presence of a highly negative mark/dollar risk premium.

[9]Giorgianni (1996) links the presence of instability in dollar ex ante excess returns with policy switches and shifts in expectations functions.

[10]Unit root test results on spot rates, expectations and interest rates are available from the author upon request. Standard Augmented Dickey-Fuller tests on these variables do not reject the unit root hypothesis at the usual 5 percent significance level.

both currencies and all maturities considered (see Table 1).[11] This result contradicts the findings of Liu and Maddala (1992), but is in accordance with MacDonald's (1994) results, and indicates that there exists a long-run (cointegrating) relationship between spot rates, expectations and domestic and foreign interest rates; that is, deviations from Uncovered Interest Parity, given by $er^e_{t,k}$, are only short lived.

Third, the persistence of $er^e_{t,k}$ (see the correlograms of Figure 2) is stronger for the lira/dollar than the lira/mark and it is weaker the shorter the maturity. Comparing the correlogram of $er^e_{t,k}$ to the correlograms of its two components, $(r_{t,k} - r^*_{t,k})$ and $(s^e_{t+k|t} - s_t)$, the difference across currencies is very noticeable (Figure 3). For the floating currency, i.e., the lira/dollar rate, it appears as if the shape of the correlogram of risk premiums is inherited from expected devaluations, while, for the pegged currency, i.e., the lira/mark, there seems to be no particular resemblance among the correlograms of risk premiums and its components.

Finally, corroborating Fama's (1984) calculations, risk premiums and expected exchange rate changes are strongly and negatively correlated. This correlation is stronger for the floating currency than for the pegged one (see Table 1). However, in contrast to another set of calculations by Fama, which are based on the assumption of absence of serial dependence in prediction errors (i.e., on the hypothesis of rationality of expectations), the volatility of the risk premium is found to be lower than the volatility of the process of expected depreciation for all currencies and horizons considered.

The presence of significant temporal dependence in $er^e_{t,k}$ (see also Appendix III) leaves open the question, to be answered in the remainder of the paper, of what determines the predictable components of risk premiums.

IV. Risk Premium and Fiscal Policy: Theoretical Models

In this section, I consider two different theoretical models that incorporate a link between fiscal policy variables and the risk premium. Empirical variants of these models are tested in the next section.

A. Asset Pricing Model

The first model considered is based on a general equilibrium asset pricing model. In this context, financial decisions are made according to an intertemporal optimization

---

[11] A top-down lag selection procedure was adopted in the choice of the maximal lag for the first difference terms to be included in the ADF auxiliary regressions (the parametric correction for the presence of serial dependence in the residuals). Two sets of regressions were considered, one including only a constant term and another including a constant term and a linear trend. Both regressions produced similar results. Only the first set of results is included in Table 1.

procedure and contemporaneously to consumption decisions. The economic importance of the risk premium within general equilibrium asset pricing models is of second order, because agents make economic decisions looking far into the future, assets are regarded as perfect substitutes and Ricardian equivalence holds. Here, a positiverisk premium on a specific asset arises only if the rate of return on this asset covaries positively with the return of a benchmark portfolio, or with the intertemporal marginal rate of substitution of consumption.

A popular asset pricing model, incorporating a link between risk premiums and macroeconomic variables is the one based on Lucas (1982), and successively refined by Svensson (1985a), (1985b) and Hodrick (1989). In this model, the statistical behavior of spot and forward exchange rates and, in turn, of risk premiums depends on the evolution of the exogenous home (h) and foreign (a) state variables: output, Y_{it}, money, M_{it}, and government expenditure, G_{it} (i=h, a). A general expression for the percentage k-period ex ante excess returns on forward speculation in the home currency, i.e., the currency risk premium, is given by:

$$\ln F_{t,k} - \mathrm{E}_t[\ln S_{t+k}] = er_{t,k}^{AP} \equiv \ln\left\{\frac{\mathrm{E}_t\left[U_{at+k}Y_{ht+k}(M_{ht+k})^{-1}\right]}{\mathrm{E}_t\left[U_{ht+k}Y_{at+k}(M_{at+k})^{-1}\right]}\right\} - \mathrm{E}_t\left[\ln\left\{\frac{\left[U_{at+k}Y_{ht+k}(M_{ht+k})^{-1}\right]}{\left[U_{ht+k}Y_{at+k}(M_{at+k})^{-1}\right]}\right\}\right], \tag{7}$$

where U_{it} (i=h, a) is the partial derivative of the instantaneous utility function with respect to country-i's consumption good, which depends, in turn, on output and government expenditure, given that in equilibrium consumption is $C_{it} = \frac{1}{2}(Y_{it} - G_{it})$. In equation (7), ex ante excess returns, $er_{t,k}^{AP}$, depend on expected future outputs, money supplies and, through the marginal utilities, government expenditures, as well as agents' degree of risk aversion, and other parameters of the utility function. Note that even with risk-neutral investors, $er_{t,k}^{AP}$ will differ from zero, due to the presence of the Jensens inequality term.[12] The only case in which $er_{t,k}^{AP}$ will be zero is when the exogenous stochastic processes appearing in equation (7) are constant or nonstochastic.

Equation (7) does not yield clear predictions on the effects of fiscal and monetary policy on risk premiums. To evaluate these effects one needs to derive a closed-form solution for $er_{t,k}^{AP}$. This can be achieved by specifying a preference structure and the stochastic properties of the exogenous state variables as, for example, in Canova and Marrinan (1993). In alternative, a general and empirically testable specification of the equation for the risk premium, can be obtained from a second order Taylor approximation of equation (7).[13] A second order approximation produces the following general expression:

$$er_{t,k}^{AP'} = \varphi\,\mathrm{Var}_t(x_{t+k}) + \xi_{t,k}, \tag{8}$$

---

[12]See Engel (1995) for a clarification of this point.

[13]See Engel (1992), among others.

where $x_t = \{x_{ht}, x_{at}\}$ is the state vector, with $x_{it} = [g_{it}, m_{it}, y_{it}]$ $(i = h, a)$, and φ is vector of coefficients measuring the contribution of the volatility of macroeconomic variables to explaining risk premiums. $\xi_{t,k}$, an error term, is the residual from the Taylor approximation.

Notice that equation (8) does not incorporate any direct supply effect. Thus, for instance, a change in the supply of government securities of the domestic country would not affect the home currency risk premium. The lack of a supply channel derives from the assumption of perfect substitutability among assets and of Ricardian equivalence. In this context, the only effect of fiscal policy on the risk premium derives from the volatility terms in equation (8): a higher anticipated volatility of future government expenditure in the home country should raise the risk premium on the home currency.

B. Portfolio Balance Model

A theoretical approach in which such supply channel is present is the simple static optimizing portfolio balance model. According to this approach, the composition of a portfolio among assets denominated in different currencies, regarded as imperfect substitutes, is chosen on the base of a mean-variance optimization process. In equilibrium, because of the imperfect substitutability among assets, an increase in the supply of a particular asset requires an increase in either the return or the price of risk associated with that asset. In particular, a simplified version of this model predicts that in equilibrium the risk premium (the ex ante excess return on assets denominated in domestic currency), $er_{t,k}^{PB}$, varies indirectly with the ratio of foreign assets over the total amount of assets in circulation, $(1-\xi_t)$, and directly with the volatility of the spot exchange rate, $\sigma_{t,k}^2(s)$, that is:

$$(r_{t,k} - r_{t,k}^*) - (\mathrm{E}_t\, s_{t+k|t} - s_t) = er_{t,k}^{PB} \equiv c_1[\sigma_{t,k}^2(s)] - c_2[\sigma_{t,k}^2(s)](1-\xi_t). \tag{9}$$

A higher volatility of the exchange rate, $\sigma_{t,k}^2(s)$, increases the uncertainty over the ex ante rate of return of assets denominated in the home currency, thus raising $er_{t,k}^{PB}$. Also, an increase in the supply of home currency denominated assets, that is a reduction in $(1-\xi_t)$, increases the required rate of return on domestic-currency denominated assets.

In this framework, given that Ricardian equivalence does not hold, an expansionary fiscal policy in the home country financed with outside bonds, by increasing the stock of home-currency denominated bonds in circulation, leads to a higher home currency risk premium.

V. Risk Premium and Fiscal Policy: Empirical Results

In this section, I present the estimation results of empirical versions of the asset pricing and the portfolio balance models discussed in Section IV (equations (8) and (9), respectively).

The goal of this section is to investigate whether there is a stable structural relationship between risk premiums and macroeconomic variables, and fiscal variables in particular.

Previous literature has been unable to identify this link. The inconclusiveness of these efforts is attributed by many (see Engel (1995), for example) to the presence of large and unsystematic forecast errors, and to the difficulty of measuring risk premiums in presence of a high volatility of the forward bias ($f_{t,k} - s_{t+k}$). Furthermore, shifts in expectations, a point well documented in this and other studies,[14] make it difficult to draw correct finite-sample inference from ex post data.

Because of these difficulties, the empirical tests of the risk premium carried out in the remainder of this investigation, will rely on survey data. However, it is important to stress that the goal of the present study is not to test *a particular* model of the risk premium. In fact, there is a conceptual difficulty in using survey data to test a behavioral equation of the risk premium derived under model-consistent expectations. For instance, if an equation like (7), which holds only in a world populated by rational and optimizing representative investors, is tested using "irrational" consensus forecasts, an unobservable extra term (i.e., the survey data prediction error) is introduced in the regression error. This term is typically correlated with information at time t, and may bias inference from general risk premium specifications like equation (8).

Hence, similarly to Dominguez and Frankel (1993a) and (1993b), who use survey data to test a version of the portfolio balance model, the goal of the present investigation is to test whether there is a link between the risk premium (which I measure accurately with survey data) and the variables that in the asset pricing model and the portfolio balance model are important determinants of the risk premium. In this sense, this approach justifies the estimation of more than one specification and the cross-validation of these specifications, with the hope to "learn" from the data, more than to "reject" the restrictions of a stylized model.

A. Asset Pricing Model

From equation (8), the risk premium depends on (anticipated) volatility and covariance terms of money, income and government expenditures of the home and foreign countries. Following a common assumption in the literature (see Canova and Marrinan (1993) and Engel (1992)), supply-side effects are ruled out a priori, that is, the variance and covariance terms of the two countries' outputs are restricted to be zero. This leaves fewer second moments to be estimated: the variances and covariances of the processes for the home (Italy) and foreign (United States and Germany, in turn) money growth rates and government expenditure to GDP ratios. With obvious notation, in the remainder of this section, I will refer to the

---

[14]See Section III.B, or the evidence in Giorgianni (1996) or, else, the results of Evans and Lewis (1995) and Goldberg and Frydman (1996a), (1996b).

volatility terms as $\sigma_t^e(g_h)$, $\sigma_t^e(g_a)$, $\sigma_t^e(m_h)$, and $\sigma_t^e(m_a)$, and the covariance terms as $\sigma_t^e(g_h,m_h)$, $\sigma_t^e(g_a,m_a)$, $\sigma_t^e(g_a,g_h)$, and $\sigma_t^e(m_h,m_a)$.[15]

Specification and Estimation of the Exogenous Processes

The conditional moments of the distribution of the state vectors are generated assuming that agents form forecasts with time-series models (typically, Vector AutoRegressions --VAR's--) and that the coefficients of the forecasting models are updated recursively, as in a real-time forecasting process. This procedure does not assume a priori the ergodicity of agents' forecast errors, or that the forecasts converge to a rational expectations solution.

The estimation is conducted in two stages. First, I estimate recursively the conditional mean of x_t with a VAR of order p. Contrary to common practice (see, for example, Hodrick (1989), and Canova and Marrinan (1993)), this procedure acknowledges the presence of contemporaneous correlation between the shocks of the home and foreign state vectors.[16] Next, the estimated residuals from the VAR's are used to construct nonparametric measures of the conditional variances and covariances of the state vector.[17]

More specifically, using data for Italy, Germany and the United States (see Appendix I for a description of the macro data employed), I estimate recursively a VAR of order p for $x_t = \{x_{ht}, x_{at}\}$, with $x_{it} = \lfloor g_{it}, m_{it}, y_{it} \rfloor$, $i = h, a$. The lag-order of the VAR is chosen using the Schwartz and Hannan-Quinn information criteria, making sure that no serial correlation is left unaccounted for in the residuals.[18]

The dynamic behavior of the conditional second moments was approximated by smoothing the square and the cross-products of the one-step-ahead forecast errors from the

[15]A further a priori restriction is that $\sigma_t^e(g_h,m_a)$=0 and $\sigma_t^e(g_a,m_h)$=0.

[16]See Engel (1992) for an explanation of the advantages of using such approach.

[17]The two stage estimation strategy adopted here is close in spirit to Hodrick's (1989) and Cheung's (1993) tests of the Lucas model. A discussion of the statistical aspects of this procedure is provided in the first study (see page 455 of Hodrick (1989)).

[18]For Italy and the United States, this procedure identifies a VAR(3), while, for Italy and Germany a VAR(4). The results of the order selection procedure and of the tests for absence of autocorrelation, heteroskedasticity and nonnormality in the estimated errors of the chosen VAR's are available from the author.

estimated VAR's, denoted by $u_{it}(g)$, $u_{it}(m)$, $u_{it}(y)$ for $i = h, a$.[19] The smoothing was carried out employing the following constant-weight kernel:

$$\sigma_{it} = \frac{1}{\lambda}\sum_{\tau=0}^{\lambda-1} u_{it-\tau}^{2},$$

i.e., a noncentered moving average filter.[20]

Empirical Evidence

Two are the questions that motivate the empirical investigation: *(i)* are fiscal policy variables related to the risk premium? and if so, *(ii)* how much of the variability in the risk premium is explained by the variability in fiscal policy? To answer these questions, in a first set of tests, $er_{t,k}^{e}$ is regressed on a constant term and a set of fiscal variables, i.e., the volatility of the government expenditure to GDP (for Italy and the foreign country) plus their covariance term. The regression model is therefore:

$$er_{t,k}^{e} = \gamma_0 + \gamma_1 \sigma_t^{e}(g_h) + \gamma_2 \sigma_t^{e}(g_a) + \gamma_3 \sigma_t^{e}(g_a, g_h) + \xi_t .^{21} \tag{10}$$

Table 3 summarizes the results of the above regression for the lira/dollar and the lira/mark exchange rates and for two different forecasting horizons (3 and 12 months). All the lira/dollar regressions were augmented by the oil price, oil_t, to control for the effects the Gulf war, the "mini-oil crisis" of late 1990 and early 1991, and the subsequent sharp appreciation of the dollar (March 1991) --see Giorgianni (1996)--.

For both horizons, all the regressors are significantly different from zero at the 5 percent significance level, and enter with the correct sign: a higher volatility of future Italian fiscal policy, that is, a higher $\sigma_t^{e}(g_{IL})$ --or lower volatility of the future U.S. fiscal policy, i.e., a lower $\sigma_t^{e}(g_{US})$-- are associated with higher risk premiums on lira denominated assets. The

[19]See Hamilton (1994) on the differences between parametric and nonparametric estimates of conditional second moments. Ideally, a simultaneous and multivariate parametric model of first and second moments (i.e., a VAR+Multivariate GARCH) could be specified and estimated by Maximum Likelihood. However, besides its intrinsic computational difficulties, this strategy is not appealing because GARCH effects at monthly frequency are notoriously very weak.

[20]Different bandwidth parameters λ were experimented with and the corresponding second moments plotted to check the degree of smoothness. In the end, λ was fixed to 5 for all second moments. The results presented in the next sections were found robust to values of λ ranging between 2 and 11.

[21]The regressors in the risk premium equations are lagged once to account for the typical delays in the release of official statistics.

Table 3. OLS Estimation Results of Equation (14): Dependent Variable is $er^e_{t,k}$. 1987.02-1994.10

	Lira/Dollar				Lira/Mark	
	3 months (k=3)		12 months (k=12)		k=3	k=12
	[1]	[2]	[3]	[4]	[5]	[6]
Constant	–0.27	–0.12	–0.35	–0.12	0.01	0.02
S.E.	0.07	0.04	0.14	0.07	0.002	0.004
p-value	*0.00*	*0.01*	*0.01*	*0.12*	*0.00*	*0.00*
$\sigma^e_{IL,t-1}(g)$	0.02	0.01	0.03	0.02	–0.01	–0.00
S.E.	0.008	0.004	0.01	0.01	0.002	0.004
p-value	*0.02*	*0.00*	*0.03*	*0.02*	*0.00*	*0.95*
$\sigma^e_{*,t-1}(g)$	–1.67	–0.85	–2.87	–1.28	0.04	–0.08
S.E.	0.42	0.24	0.79	0.45	0.11	0.25
p-value	*0.00*	*0.00*	*0.00*	*0.00*	*0.70*	*0.73*
$\sigma^e_{t-1}(g_{IL}, g_*)$	–0.14	–0.07	–0.20	–0.06	0.06	0.07
S.E.	0.08	0.04	0.14	0.08	0.03	0.06
p-value	*0.06*	*0.10*	*0.18*	*0.41*	*0.04*	*0.25*
$er^e_{t-1,k}$		0.54		0.60		
S.E.		0.07		0.07		
p-value		*0.00*		*0.00*		
oil_{t-1}	0.07	0.03	0.09	0.03		
S.E.	0.02	0.01	0.03	0.02		
p-value	*0.00*	*0.01*	*0.01*	*0.15*		
R^2	0.35	0.56	0.27	0.58	0.15	0.02
Regr. S.E.	0.02	0.02	0.04	0.03	0.01	0.02
R.S.S.	0.04	0.03	0.14	0.08	0.01	0.02
D-W.	0.96	2.18	0.76	1.98	1.87	1.30
AR(6)	*0.00*	*0.42*	*0.00*	*0.82*	0.62	*0.00*
ARCH(6)	*0.40*	*0.89*	*0.05*	*0.56*	0.09	*0.57*
Normality	*0.22*	*0.14*	*0.45*	*0.18*	*0.00*	*0.74*
White	*0.38*	*0.86*	*0.13*	*0.77*	*0.45*	*0.28*
RESET	*0.25*	*0.73*	*0.05*	*0.26*	*0.51*	*0.07*
Var. Inst.	*<0.01*	*>0.10*	*<0.01*	*>0.10*	*<0.05*	*<0.05*
Coef. Inst.	*>0.10*	*>0.10*	*<0.01*	*>0.10*	*<0.01*	*<0.01*

Note: Figures in italics represent marginal significance levels. The S.E.'s and p-values are based on heteroskedastic-autocorrelation consistent residuals variance estimators. R.S.S. is the Regression Sum of Squares. The diagnostics reported are: D.-W. is the Durbin-Watson test for first order residuals autocorrelation; AR(6) is a Lagrange Multiplier (LM) test for residuals autocorrelation up to the sixth order; ARCH(6) is a LM test for the presence of ARCH effects in the residuals up to the sixth order; Normality is a test for gaussianity of the estimated residuals; White is a general test for residuals heteroskedasticity. RESET is Ramsey's functional mis-specification test. Var. Inst. and Coef. Inst. are Hansen's stability tests, respectively, for the residuals variance and the estimated coefficients (jointly).

covariance term, $\sigma_t^e(g_{IL}, g_{US})$, enters with a negative sign and is significant only in the 3-month regression.

Overall, the proportion of the variability of the risk premium explained by fiscal variables alone is very low: when the oil price is removed from the estimated equations the R^2 decreases sharply (e.g., from 0.35 to 0.15, in the 3-month-horizon equation, and from 0.27 to 0.17, in the 12-month-horizon regression). Furthermore, for all the estimated equations, the hypotheses of parameter constancy and lack of autocorrelation in the residuals are rejected at the usual significance levels, perhaps suggesting the presence of a mis-specification or omitted variable problems.

After the addition of the lagged dependent variable to equation (10), the overall performance of the estimated model improves, since all the parameters appear to be constant and residuals white-noise, while, at the same time, the significance of fiscal variables is preserved.

For the lira/mark exchange rate, the estimation of equation (10) produces unsatisfactory results. Irrespective of the prediction horizon, none of the fiscal variables enter with the correct sign or significantly. This outcome does not necessarily emerge as a surprise, for at least two reasons: first, as shown in Appendix III, the lira/mark risk premium is not very predictable (for example, for the 3-month horizon, standard statistical procedures cannot reject the absence of conditional mean dynamics); second, the presence of three different exchange rate regimes for the lira/mark (before and after the ERM breakdown --September 1992--, and, within the ERM period, before and after the introduction of the narrower oscillation band --January 1990--) may introduce non-linear dynamics in the data which are not captured by a simple linear specification.

To verify the robustness of the significance of fiscal variables for the lira/dollar risk premium, an alternative equation inclusive of monetary variables was estimated:

$$er_{t,k}^e = \gamma_0 + \gamma_1\sigma_t^e(g_h) + \gamma_2\sigma_t^e(g_a) + \gamma_3\sigma_t^e(g_a, g_h) + \gamma_4\sigma_t^e(g_h, m_h) + \gamma_5\sigma_t^e(g_a, m_a) + \xi_t^*, \qquad (11)$$

where all variables are defined above.

The introduction of monetary variables for the lira/dollar risk premium did not change substantially the previous finding of a significant effect of U.S. and Italian fiscal policy variables on the risk premium (to economize in space these results are not included in the tables). Further, the only monetary variable that resulted significant was the covariance between U.S. monetary and fiscal shocks, $\sigma_t^e(g_{US}, m_{US})$. It appears that a higher degree of comovement between these shocks reduces the Italian lira risk premium relative to U.S. dollar-denominated assets.

For the lira/mark case, the introduction of monetary variables did not improve significantly the previous unsatisfactory findings.

Summing-up, in line with the predictions of the asset pricing model illustrated in Section IV, the lira/dollar risk premium is found to be significantly influenced by the volatility of the home and foreign state variables. In particular, higher uncertainty about the future fiscal policy in Italy relative to the United States is associated with higher expected returns in lire, and vice versa. Finally, the link between risk premiums and fiscal variables is stronger during the floating exchange rate regime than during the pegged one, and is stronger the shorter the forecasting horizon. However, the explanatory power of this model is low, even with respect to a simple first-order autoregressive process (see Appendix III). This result, together with the disappointing findings for the lira/mark rate, confirm the inability of the asset pricing model to capture the dynamic behavior of risk premiums previously documented in the literature. These results motivate the search for a more data-driven empirical specification for the survey-based measure of the risk premium tackled in the next section.

B. Portfolio Balance Model

In the asset pricing model financial instruments are perfectly substitutable, investors are forward-looking and Ricardian equivalence holds. Under these circumstances, an increase in the supply of government securities does not alter their expected return. As a consequence, it is impossible to test whether an increase in the budget deficit in the home country (financed with emission of new bonds) affects the risk premium on assets denominated in the domestic currency.

To fill this gap, in this section I specify and test an empirical version of equation (9), the risk premium equation implied by the portfolio balance model. According to this model, an anticipated fiscal expansion in the home country increases the home government borrowing requirement and, for unchanged monetary and debt management policies, the supply of government bonds denominated in the home currency. With risk aversion and low degree of substitutability among assets denominated in different currencies, foreign investors would absorb the increase in the supply of home denominated assets only if these offered higher expected excess returns, that is, only in presence of a higher risk premium on the home currency.

In line with this argument, the equation of the risk premium tested is:

$$er^e_{t,k} = \mu + \beta_{IL}(bd^e_{IL,t-1}) + \beta_*(bd^e_{*,t-1}) + \omega_{t,k}, \tag{12}$$

where β_{IL} and β_* measure the effect on $er^e_{t,k}$ of, respectively, home and foreign expected future ratios of government net borrowing to GDP ($bd^e_{IL,t-1}$ and $bd^e_{*,t-1}$). The variables $bd^e_{IL,t-1}$ and $bd^e_{*,t-1}$ are constructed by recursively estimating high-order autoregressive processes on the historical ratios.[22]

[22]The one-step ahead predicted values from these regressions were smoothed to emphasize the low frequency components of the data.

Table 4. OLS Estimation Results of Equation (16):
Dependent Variable is $er^{e}_{t,k}$

	Lira/Dollar 87.02-94.10				Lira/Mark 87.02-94.07	
	3 months (k=3)		12 months (k=12)		k=3	k=12
	[1]	[2]	[3]	[4]	[5]	[6]
Constant	–0.37	–0.15	–0.76	–0.30	–0.02	–0.09
S.E.	0.11	0.06	0.19	0.12	0.03	0.05
p-value	*0.00*	*0.02*	*0.00*	*0.02*	*0.42*	*0.07*
$bd^{e}_{IL,t-1}$	0.19	0.09	0.48	0.21	0.02	0.11
S.E.	0.07	0.04	0.12	0.08	0.02	0.03
p-value	*0.01*	*0.02*	*0.00*	*0.01*	*0.43*	*0.00*
$bd^{e}_{*,t-1}$	–0.24	–0.12	0.04	0.002	0.09	–0.01
S.E.	0.07	0.04	0.12	0.07	0.17	0.26
p-value	*0.00*	*0.00*	*0.73*	*0.97*	*0.58*	*0.96*
$er^{e}_{t-1,k}$		0.56		0.58		
S.E.		0.07		0.08		
p-value		*0.00*		*0.00*		
oil_{t-1}	0.07	0.03	0.06	0.02		
S.E.	0.02	0.01	0.03	0.02		
p-value	*0.00*	*0.02*	*0.05*	*0.32*		
R^2	0.30	0.54	0.31	0.57	0.01	0.13
Regr. S.E.	0.02	0.02	0.03	0.03	0.01	0.01
R.S.S.	0.04	0.03	0.13	0.08	0.01	0.02
D-W.	0.95	2.19	0.91	1.98	1.79	1.55
AR(6)	*0.00*	*0.53*	*0.00*	*0.69*	*0.44*	*0.15*
ARCH(6)	*0.04*	*0.68*	*0.05*	*0.20*	*0.01*	*0.96*
Normality	*0.14*	*0.12*	*0.09*	*0.05*	*0.00*	*0.08*
White	*0.01*	*0.08*	*0.69*	*0.33*	*0.06*	*0.84*
RESET	*0.29*	*0.53*	*0.01*	*0.46*	*0.30*	*0.30*
Var. Inst.	*<0.01*	*>0.10*	*>0.10*	*>0.10*	*<0.05*	*>0.10*
Coef. Inst.	*<0.01*	*>0.10*	*<0.01*	*>0.10*	*>0.10*	*>0.10*

Note: Figures in italics represent marginal significance levels. The S.E.'s and p-values are based on heteroskedastic-autocorrelation consistent residuals variance estimators. R.S.S. is the Regression Sum of Squares. The diagnostics reported are: D.-W. is the Durbin-Watson test for first order residuals autocorrelation; AR(6) is a Lagrange Multiplier (LM) test for residuals autocorrelation up to the sixth order; ARCH(6) is a LM test for the presence of ARCH effects in the residuals up to the sixth order; Normality is a test for gaussianity of the estimated residuals; White is a general test for residuals heteroskedasticity. RESET is Ramsey's functional mis-specification test. Var. Inst. and Coef. Inst. are Hansen's stability tests, respectively, for the residuals variance and the estimated coefficients (jointly).

Table 4 presents the estimation results of equation (12) for both the lira/dollar and the lira/mark exchange rates. The Italian net government borrowing requirement enters always significantly and with the positive sign in the lira/dollar regressions, suggesting that an expected higher government deficit, which can be thought of being accompanied by a higher supply of government liabilities, is associated with a higher risk premiums on Italian lira assets. This effect is stronger the longer is the maturity. The expected U.S. federal government borrowing requirement enters significantly and with the negative sign only in the 3-month regression.

In the case of the lira/mark, a significant portfolio effect is observed for the 12-month maturity and pertains only to the Italian fiscal variables. This finding is encouraging, especially in light of the unsatisfactory results of the previous section obtained for this exchange rate. On the negative side, the explanatory power of this model is not much greater than the asset pricing specification of the previous section. Furthermore, as in the previous case, the presence of parameter instability and of residuals autocorrelation is eliminated only after the inclusion of the lagged-dependent variable among the regressors.

VI. Concluding Remarks

This paper uses survey data to produce novel and model-free evidence on the magnitude and predictability of the Italian lira risk premium.

In the existing literature, risk premiums simulated from rational expectations asset pricing models, or inferred using regression equations with ex post data are found to be small relative to the size of the forward premium and surprisingly too stable given the observed high volatility of the forward bias. In contrast, this paper shows that risk premiums in excess of 1.5 percent per annum (a third of the average forward premium) may be the norm, and so may be a significant degree of volatility (approximately, in the order of 40 percent of the volatility of the forward bias).

The main contribution of this study is the finding that some of the observed variability in the survey-based measure of the risk premium may be systematically related to macroeconomic variables. Two structural models of the risk premium are tested. The estimation results of the first model, an empirical version of the Lucas two-country asset pricing model, suggest that anticipated uncertainty surrounding the future path of fiscal policy has a significant impact on currency risk. When fiscal policy is measured by the ratio of government expenditure to GDP, higher uncertainty about Italian (U.S.) fiscal policy are shown to be associated with higher (lower) lira risk premiums, that is, lower (higher) dollar risk premiums.

In another set of regressions, inspired by a portfolio balance model, anticipated fiscal expansions in Italy vis-à-vis the United States and Germany, whereby the fiscal stance is measured with the ratio of budget deficits to GDP, appear to be broadly correlated with the risk premium on lira assets. Other things being equal, an increase in the ratio of Italian government deficit to GDP induces an increase in the lira/dollar and lira/mark risk premiums.

In light of policymakers' constant concern with the external effects of domestic policies, this study provides encouraging evidence of a significant empirical linkage between fiscal policy and exchange rates through a currency risk premium channel, and suggests that a credible fiscal contraction in Italy would reduce the degree of macroeconomic uncertainty and allow for lower risk premiums in the Italian lira. Still, the evidence provided does not necessarily carry the implication that a credible fiscal contraction in the home country would appreciate the domestic currency as discussed in the Introduction. This is because, theoretically, a strong response of exchange rate expectations to anticipated fiscal policy, namely, when the home currency is expected to depreciate in reaction to an anticipated domestic fiscal contraction, might neutralize or offset the risk premium effect. To better understand the effects of fiscal policy on exchange rate movements, future research should be devoted to gather evidence on the presence of a direct link between fiscal variables and expectations of devaluations.

DATA DESCRIPTION AND SOURCES

A. Survey Data

Expected values of the spot rate are taken from the *Financial Times Currency Forecaster* (FTCF).[23] The FTCF constitutes an improvement over previous surveys, because of the dimension of the panel, the stability over time of the survey respondents, the clarity of the aggregation rules and the knowledge of the timing of the survey.[24] Since January 1987, point forecasts of 45 individual forecasters have been collected every month for currencies of the G7 countries, plus some Asian, Latin American and Eastern European economies. Four forecasting horizons are available: 1, 3, 6 and 12 months. Forecasters are separated into two groups: the first includes 30 multinational firms (mainly based in the United States); the second, 15 forecasting services firms. Only aggregate forecasts are available from the survey. In this paper, the "consensus" forecast, i.e., the geometric mean of the individual forecasts, is used. The geometric mean differs from the arithmetic average because it downplays the role of outliers.

The forecast of the spot rate at time t (i.e., the end of month t) for time $t+k$ (i.e., the end of month $t+k$), with $k=1, 3, 6, 12$, is typically formed during the third week of month t, i.e., at time $t-\eta$, with η a fraction of month. (To simplify notation, in the body of the paper the information set at time $t-\eta$ has been denoted with t). Forecasts are collected by fax on the third Friday of each month.[25]

B. Other Data

Spot rates are the 10:00 a.m. (Swiss time) Bank of International Settlements' (BIS) bid-ask average quotes. The monthly series are extracted from the daily data-base by choosing the quote of the day in which the exchange rate forecasts are formed. Interest rates are BIS Eurodeposit rates quoted at the same time as the spot rates.

---

[23]The FTCF, formerly known as the *Currency Forecasters Digest*, has been studied in Frankel and Chinn (1993) and Chinn and Frankel (1994a), (1994b). The samples considered in the first two studies were very short (from February 1988 to February 1991, 36 observations overall). In the third study the authors updated their results to June 1994. However, their data set carries three missing observations (Chinn and Frankel (1994b: 23). This paper employs a larger sample (from January 1987 to December 1994) and has no missing observations.

[24]A list of the surveys on exchange rate expectations previously employed can be found in Takagi (1991, Table 1). For completeness, two surveys should be added to that list: the *U.K. Money Market Services* and the *Business International Corporation* (respectively, MacDonald (1990) and Cavaglia *et al.* (1993a) and(1993b).

[25]Exceptions are the months of November and December, when the publication date is generally one week earlier because of Thanksgiving and Christmas.

The monthly series of macroeconomic data are from the *International Financial Statistics* (IFS) and the WEFA data-bank. The gross rate of growth of money is constructed using seasonally adjusted (end of period) figures for M1 (line *34b* of IFS for Germany and Italy, and WEFA *vmm1a* series for the United States)

Output growth is approximated by the rate of growth of the seasonally adjusted industrial production index (for Germany and the United States, IFS *66i*, and for Italy, WEFA *qjtta*.)

The monthly ratios of government expenditure to GDP and of government deficit to GDP are constructed using, respectively, the WEFA series of total government expenditure on a cash basis (series *vget*, for Italy and Germany, and series *vgbo* --federal government net outlays-- for the United States), and the series of government net borrowing requirement (IFS *84* for Germany, and WEFA *vgb* for the United States and *vgbr* for Italy). Before these ratios are computed, fiscal data are converted into 1990 prices using (seasonally adjusted) Consumer Price Index series (CPI, line *64* of IFS). The monthly measure of output used to compute these ratios is obtained applying the rate of growth of industrial production to the 1990 (end of period) figure of GDP (taken from line *99b.c* of the IFS). And residual seasonality was removed by means of seasonal dummy regressions. The oil price series is the monthly 'crude petroleum' producer price index.

Data is generally available from January 1980 to December 1994. The only exceptions are Italian government expenditure, available until September 1994 and German net borrowing requirement, until July 1994. German data refers only to West Germany until July 1990, and to unified Germany afterwards.

"REGRESSION-BASED" VERSUS SURVEY DATA RISK PREMIUMS

The goal of this appendix is to analyze the shortcomings that arise when inferring risk premiums from ex post data using linear projections like equation (6), which I will refer to as the "regression-based" risk premium. In particular, it will be showed that, in presence of an unstable economic environment, the regression-based risk premium is biased, and differs significantly from its bias-free survey-data counterpart.

To estimate the equation (6), the econometrician must chose the variables appearing in agents' information set, Ω_t. An obvious and simple option, frequently adopted in the literature and maintained here, is to include only a constant term and the forward premium (i.e., the interest rate differential).[26] Although richer specifications for equation (6) have been proposed in previous studies, the explanatory power of variables other than the forward premium has been generally found to be low.[27] The predicted values from OLS estimation of such regression,[28] i.e., the regression-based measure of the risk premium, $er_{t,k}^{FB}$, are plotted in Figure A1.a, along with the two-standard error bands.[29]

The large error bands suggest that the regression-based risk premium is measured rather imprecisely. In line with previous findings (see Engel (1995)), the linear projection of the forward bias is very persistent and rarely changes sign.

During the period under study (1987-1994) important institutional changes took place, possibly introducing instability in the probability distribution of the data employed. While the lira was always floating freely vis-à-vis the U.S. dollar, it experienced different regimes vis-à-vis the German mark: until September 1992, the lira was part of the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS) --with a large fluctuation band until January 1990 and a smaller band thereafter--; since the speculative attack of September 1992, the lira has been floating freely also vis-à-vis the German mark. In presence of discrete shifts, like the ones just mentioned, even rational investors might find it optimal to temporarily produce biased forecasts, which would naturally occur if, for example, investors were engaged in a learning process. The presence of learning also implies that over time agents update the parameters of the expectational model. This updating introduces an instability in the estimated coefficients of equations (6).

---

[26]See, for all, Fama (1984), Canova and Marrinan (1993) and Lewis (1994). Lewis (1994) justifies the adoption of this specification on the ground of parsimony, although she acknowledges the fact that other authors have found significant effects of other variables in explaining the forward bias.

[27]See, Hodrick and Srivastava (1984) and Cumby (1988).

[28]The estimation sample is 1986.01-94.12.

[29]These error bands are based on consistently estimated residuals' standard errors. Detailed regression results are available from the author.

Figure A1.a Regression-based Risk Premium and 2*SE Bands (1986-1994)

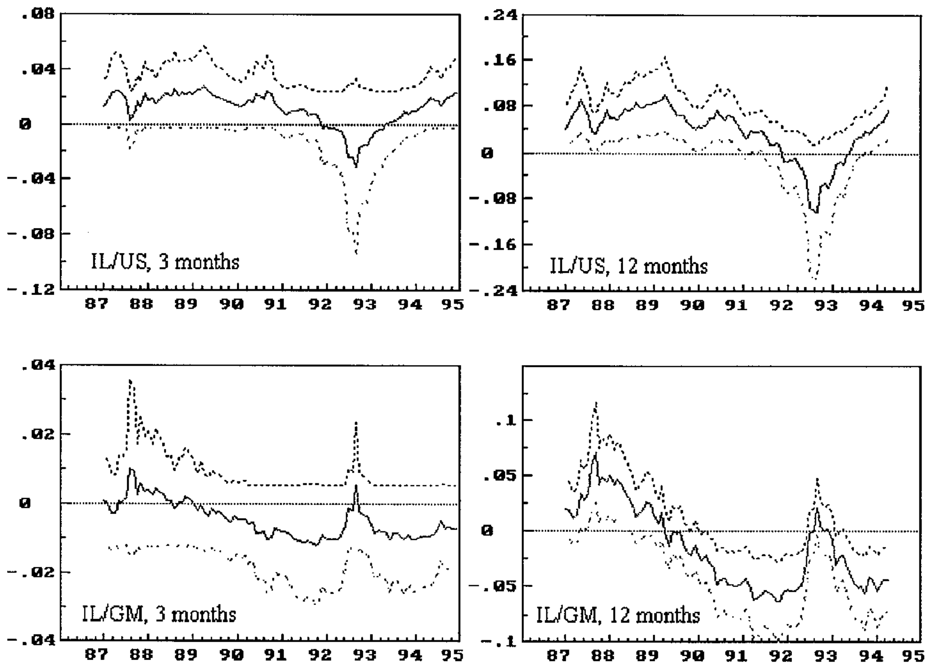


Figure A1.b Survey-data Risk Premium and 2*SE Bands from Regression-based Risk Premium (1986-1994)

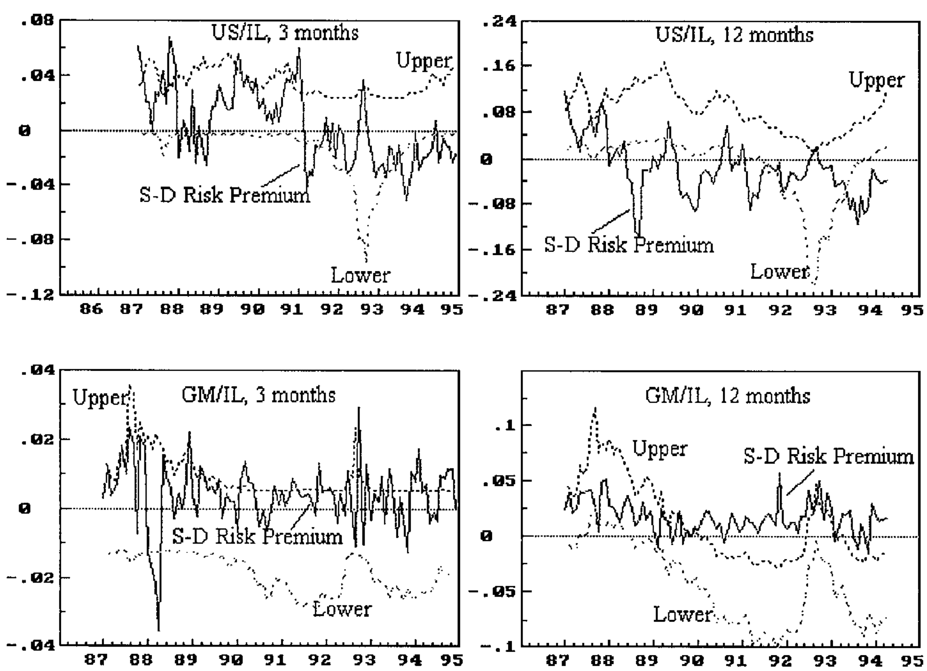
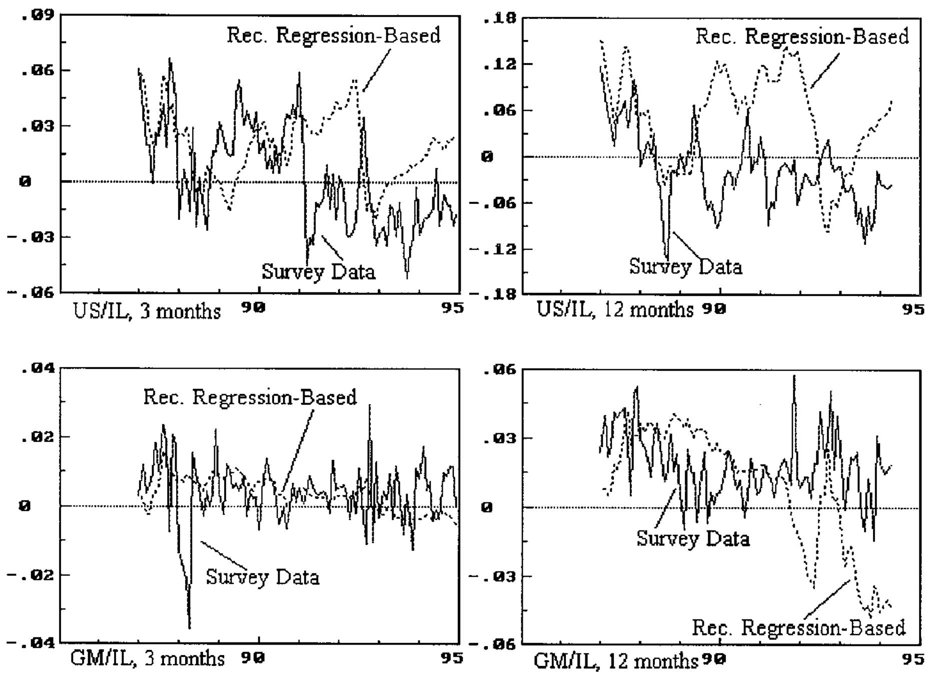


Figure A1.c Survey-data Risk Premium and Recursive Regression-based Risk Premium (1986-1994)



To verify this conjecture, equation (6) is estimated recursively and the sequence of the estimated projection coefficients, along with the one- and multi-step prediction failure tests, are plotted in Figure A2.[30] These plots clearly reject the hypothesis of time invariance. The lack of parameter constancy introduces a bias in the regression-based measure of the risk premium, $er_{t,k}^{FB}$, obtained employing full-sample information (i.e., erroneously imposing stability of the estimated coefficients). To assess the magnitude and significance of this bias, I compare $er_{t,k}^{FB}$ with a bias-free measure of the risk premium, i.e. the survey-based risk premium, denoted by $er_{t,k}^{e}$. Figure A1.b plots $er_{t,k}^{e}$ together with the two-standard error bands of Figure A1.a, e.g., the 95 percent prediction interval associated with $er_{t,k}^{FB}$, the regression-based risk premium. Whenever the survey-based risk premium falls outside the two-standard error bands, the bias in the regression-based risk premium is statistically significant (at the 5 percent confidence level). As Figure A1.b makes clear, the size of the bias becomes larger the closer one gets to the breakdown of the ERM (September 1992) --i.e., the main source of instability in this period, especially for the lira/mark case-- and becomes more significant the longer the horizon.

In presence of significant data nonconstancy, one can expect that the ex post prediction exercise, implicit in the whole sample estimation of equation (6), may differ from a recursive "real-time" approach. To explore this possibility, I compare the survey-based measure of the risk premium with the recursively generated regression-based risk premium, denoted by $\vec{er}_{t,k}^{FB}$. The point-wise comparison of these two series (see Figure A1.c) is very insightful: first, there is a closer match for the lira/mark than for the lira/dollar and for shorter horizons versus long; second, even in presence of a similar behavior in the two series, there are sub-periods in which the pattern of the two series is markedly dissimilar. This is not unexpected, since the recursive procedure is backward looking and does not incorporate investors' anticipation of future shifts, i.e., a peso problem.

Summing-up, the survey-based measure of the risk premium differs substantially from its whole-sample regression-based counterpart, mainly because the former is more volatile and less persistent than the latter. Also, the correlation between the two series is low and decreases as the forecasting horizon increases. In addition, the presence of nonconstancy introduces significant biases in the regression-based risk premium, which are only partly mitigated if the coefficients of the linear projection are updated sequentially with the arrival of new information.

---

[30]Notice that Figure A2 presents only the recursions for the 3-month case (for both currencies). This is a more stringent test, because the 12-month estimated equations (not included in the figures to save space) display higher instability.

Figure A2.a Recursive Estimation of Regression-based Risk Premium: IL/US, 3 months (1986-1994)

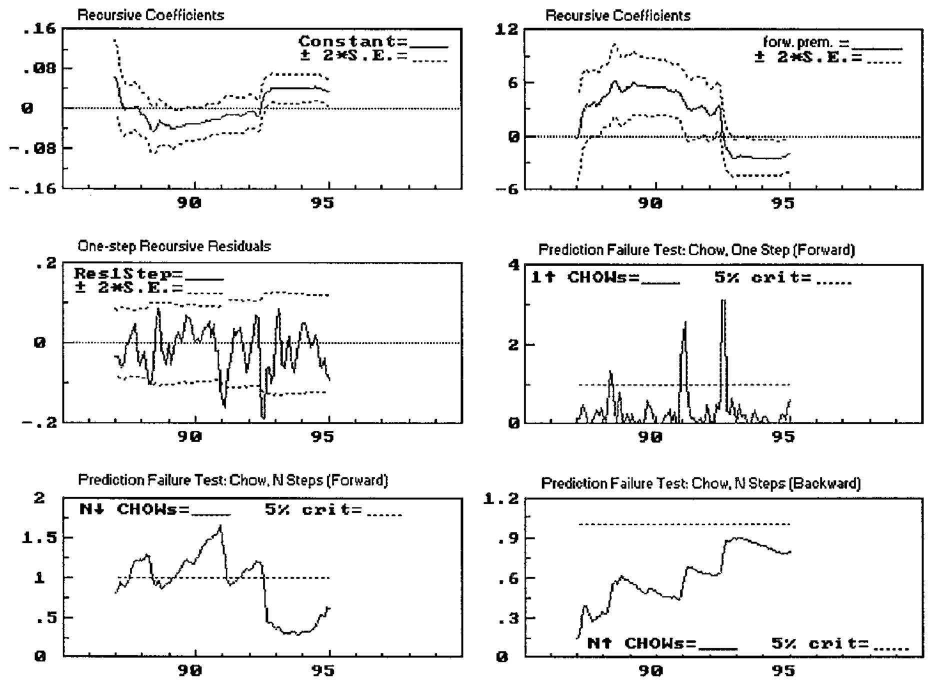
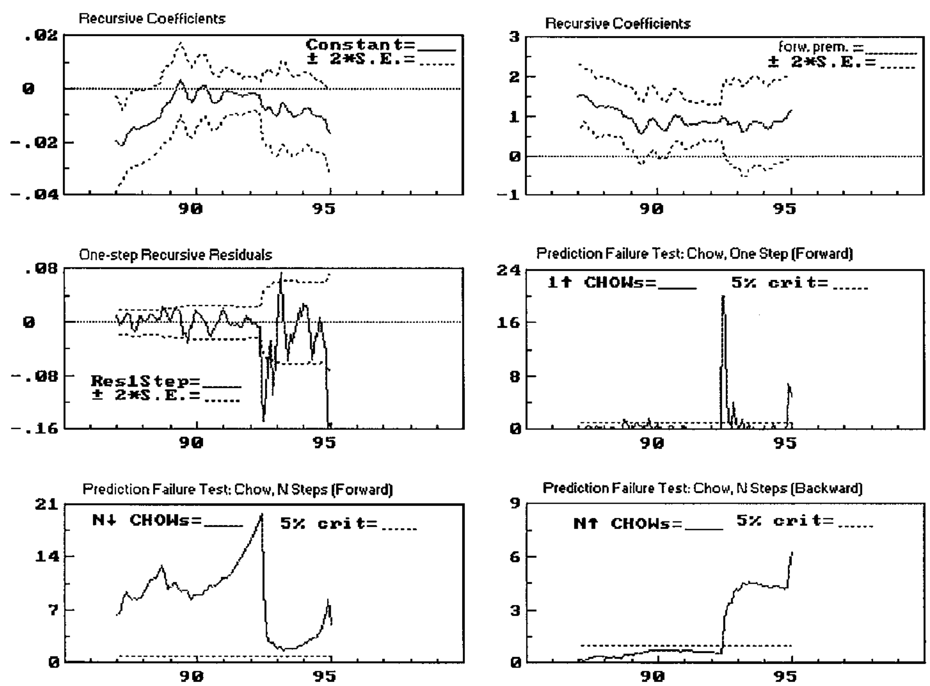


Figure A2.b Recursive Estimation of Regression-based Risk Premium: IL/GM, 3 months (1986-1994)



A TIME SERIES REPRESENTATION FOR THE LIRA RISK PREMIUM

The objective of this appendix is to verify one conjecture stemming from all the general equilibrium models of the risk premium, namely, that risk premiums are well approximated by low order stationary autoregressive processes.[31] The search for a suitable time series process for $er^{e}_{t,k}$ is restricted to the class of Auto-Regressive Moving Average plus Trend, ARMA(p,q)+Tr(r), specifications. Standard Schwartz's Bayesian Information Criteria (BIC) were employed to select the orders p, q and r.

For the lira/dollar rate, the 3- and 12-month risk premiums are well described by a simple AR(1) process without deterministic components (see estimation results of Table A1). The estimates of the autoregressive root across maturities are around 0.75 and highly significant. The adequacy of this specification was checked by a series of diagnostic tests, which suggest that the estimated residuals are well behaved (see Table A1).

Differently from the lira/dollar case, for the lira/mark risk premiums the optimized BIC's do not suggest a uniform specification for the whole sample and across maturities. For the 3-month maturity the chosen model is just the sample mean, i.e., conditional mean dynamics are absent. For the 12-month case a MA(1) plus constant, with a very small but significant MA coefficient of 0.02, is selected. Undoubtedly, the presence of different institutional arrangements involving the lira/mark rate complicates the dynamics of risk premiums in such a way that it becomes difficult to retrieve from the data a whole-sample and time-invariant linear time series model. Excluding the ERM-crisis period and the following floating of the lira/mark (92.09 to 94.12), and restricting the identification procedure to search only within the AR family, the chosen models were: an AR(1) without constant for the 3-month maturity and an AR(1) with constant for the 12-month case. Both AR(1) models produced low but significant estimates of the autoregressive parameters, near to 0.38 (see Table A1). However, the low explanatory power (e.g., the R^2 is close to 0.14) and the nonnormality and heteroskedasticity of the estimated errors reveal the inadequacy of a pure linear time series model for the lira/mark risk premium.

---

[31]See also Nijman *et al.* (1993) and MacDonald (1994).

Table A1. OLS Estimation Results of the Time Series Model for the Risk Premium: $er^e_{t,k}$.

	Lira/Dollar 1987.02-1994.12		Lira/Mark 1987.02-1992.08	
	k=3	k=12	k=3	k=12
Constant				0.01
S.E.				0.003
p-value				*0.00*
$er^e_{t-1,k}$	0.69	0.74	0.39	0.38
S.E.	0.07	0.06	0.11	0.11
p-value	*0.00*	*0.00*	*0.00*	*0.00*
R^2	0.51	0.59	0.15	0.14
Regr. S.E.	0.02	0.03	0.01	0.01
R.S.S.	0.03	0.09	0.006	0.01
D.-W.	2.28	2.05	2.02	2.02
AR(6)	*0.41*	*0.39*	*0.17*	*0.19*
ARCH(6)	*0.80*	*0.09*	*0.03*	*0.82*
Normality	*0.20*	*0.13*	*0.00*	*0.05*
White	*0.09*	*0.28*	*0.00*	*0.97*
RESET	*0.86*	*0.99*	*0.20*	*0.64*
Var. Instab.	*>0.10*	*>0.10*	*<0.01*	*>0.10*
Coef. Instab.	*>0.10*	*>0.10*	*<0.01*	*>0.10*

Note: Figures in italics represent marginal significance levels. The S.E.'s and p-values are based on heteroskedastic-autocorrelation consistent residuals variance estimators. R.S.S. is the Regression Sum of Squares. The diagnostics reported are: D.-W. is the Durbin-Watson test for first order residuals autocorrelation; AR(6) is a Lagrange Multiplier (LM) test for residuals autocorrelation up to the sixth order; ARCH(6) is a LM test for the presence of ARCH effects in the residuals up to the sixth order; Normality is a test for gaussianity of the estimated residuals; White is a general test for residuals heteroskedasticity. RESET is Ramsey's functional mis-specification test. Var. Inst. and Coef. Inst. are Hansen's stability tests, respectively, for the residuals variance and the estimated coefficients (jointly).

REFERENCES

Baillie, R.T. and T. Bollerslev, 1990, "A Multivariate Generalized ARCH Approach to Modeling Risk Premia in Forward Foreign Exchange Rate Markets," *Journal of International Money and Finance*, No. 9, pp. 309-24.

Canova, F. and J. Marrinan, 1993, "Profits, Risk and Uncertainty in Foreign Exchange Markets," *Journal of Monetary Economics*, No. 32, pp. 259-86.

Cavaglia S., W.F.C. Verschoor and C.C.P. Wolff, 1994, "On the Biasedness of Forward Foreign Exchange Rates: Irrationality or Risk Premia?" *Journal of Business*, Vol. 67, No. 3, pp. 321-43.

Cheung, Y., 1993, "Exchange Rate Risk Premiums," *Journal of International Money and Finance*, No. 12, pp. 182-94.

Chinn, M. and J.A. Frankel, 1994a, "Patterns in Exchange Rate Forecasts for Twenty-five Currencies," *Journal of Money Credit and Banking*, Vol. 26, No. 4, pp. 759-770.

——— and ———, 1994b, "More Survey Data on Exchange Rate Expectations: More Currencies, More Horizons, More Tests." Presented at the 1995 *AEA* meetings.

Cumby, R.E., 1988, "Is it Risk? Explaining Deviations from Uncovered Interest Parity," *Journal of Monetary Economics*, No. 22, pp. 279-99.

Dominguez, K.M. and J.A. Frankel, 1993a, *Does Foreign Exchange Intervention Work?* Washington, DC, Institute for International Economics.

——— and ———, 1993b, "Does Foreign Exchange Intervention Matter? The Portfolio Effect," *American Economic Review*, Vol. 83, No. 5, pp. 1356-69.

Domowitz, I. and C.S. Hakkio, 1985, "Conditional Variance and the Risk Premium in the Foreign Exchange Market," *Journal of International Economics*, No. 19, pp. 47-66.

Engel, C., 1992, "On the Foreign Exchange Risk Premium in a General Equilibrium Model," *Journal of International Economics*, No. 32, pp. 305-19.

———, 1995, "Why is the Forward Exchange Rate Forecast Biased? A Survey of Recent Evidence," NBER *Working Paper Series*, No. 5312, October.

Evans, M.D.D. and K.K. Lewis, 1995, "Do Long-Term Swings in the Dollar Affect Estimates of the Risk Premia?" *Review of Financial Studies*, Vol. 8, No. 3, pp. 709-42.

Fama, E.F., 1984, "Forward and Spot Exchange Rates," *Journal of Monetary Economics*, No. 14, pp. 319-38.

Frankel, J.A. and M. Chinn, 1993, "Exchange Rate Expectations and the Risk Premium: Tests for a Cross Section of 17 Currencies," *Review of International Economics*, Vol. 1, No. 2, p. 136.

——— and A.K. Froot, 1987, "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations," *American Economic Review*, Vol. 77, No. 1.

——— and ———, 1989, "Forward Discount Bias: Is It an Exchange Risk Premium?" *Quarterly Journal of Economics*, No. 54, February, pp. 139-61.

Giorgianni, L., 1996, "On Expectations and Risk Premia in Foreign Exchange Markets. Evidence from Survey Data," *University of Pennsylvania*, Ph.D. dissertation.

Hamilton, J.D., 1994, *Time Series Analysis*, Princeton, Princeton University Press.

Hodrick, R.J., 1989, "Risk, Uncertainty and Exchange Rates," *Journal of Monetary Economics*, No. 23, pp. 433-59.

International Monetary Fund, 1995, "The Exchange Rate Effects of Fiscal Consolidation," *World Economic Outlook*, Annex I. October.

Lewis, K.K., 1994, "International Financial Markets." Forthcoming in *Handbook of International Economics*, Grossman, G. and K. Rogoff. (eds.), North-Holland.

Lucas, R.E., 1982, "Interest Rates and Currency Prices in a Two-Country World," *Journal of Monetary Economics*, No. 16, pp. 335-60.

Liu, P.C. and G.S. Maddala, 1992, "Rationality of Survey Data and Tests for Market Efficiency in the Foreign Exchange Market," *Journal of International Money and Finance*, No. 11, pp. 366-81.

MacDonald, R., 1994, "Is the Foreign Exchange Market 'Risky'? Some New Survey-Based Results," *University of Strathclyde*, Working Paper.

——— and T.S. Torrance, 1990, "Expectations Formation and Risk in Four Foreign Exchange Markets," *Oxford Economic Papers*, No. 42, pp. 544-61.

Nijman, T.E., Palm, F.C. and C.C.P. Wolff, 1993, "Premia in Forward Exchange as Unobserved Components," *Journal of Business and Economic Statistics*, No. 11, pp. 361-65.

Svensson, L.E.O., 1985a, "Currency Prices, Terms of Trade and Interest Rates: A General Equilibrium Asset-Pricing Cash-in-Advance Approach," *Journal of International Economics*, No. 18, pp. 17-41.

———, 1985b, "Money and Asset Prices in a Cash-in-Advance Economy," *Journal of Political Economy*, No. 93, pp. 919-44.

———, 1992, "The Foreign Exchange Risk Premium in a Target Zone with Devaluation Risk," *Journal of International Economics*, No. 33, pp. 21-40.

Taylor, M.P., 1989, "Expectations, Risk and Uncertainty in the Foreign Exchange Market: Some Results Based on Survey Data," *The Manchester School*, Vol. 57, No. 2, pp. 142-53.