Working Paper

INTERNATIONAL MONETARY FUND

WP/01/207

IMF Working Paper

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Authorized for distribution by Kenneth Rogoff

December 2001

Abstract

This paper tests for uncovered interest parity (UIP) using daily data for 23 developing and developed countries through the crisis-strewn 1990s. We find that UIP works better on average in the 1990s than in previous eras in the sense that the slope coefficient from a regression of exchange rate changes on interest differentials yields a positive coefficient (which is sometimes insignificantly different from unity). UIP works systematically worse for fixed and flexible exchange rate countries than for crisis countries, but we find no significant differences between rich and poor countries.

JEL Classification Numbers: F32, G15

Keywords: empirical, exchange rate, fixed, floating, developing, developed.

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I. INTRODUCTION

Uncovered interest parity (UIP) is a classic topic of international finance a critical building block of most theoretical models and a dismal empirical failure. UIP states that the interest differential is on average equal to the ex post exchange rate change. A strong consensus has developed in the literature that UIP works poorly; it predicts that countries with high interest rates should, on average, have depreciating currencies. Instead, such currencies tend to have appreciated. Surveys are provided by Hodrick (1987), Froot and Thaler (1990), and Lewis (1995). In this short paper, we use recent data for a wide variety of countries to re-examine the performance of UIP during the 1990s.

It is easy to motivate another look at UIP. The vast majority of literature on UIP uses data drawn from low-inflation floating exchange rate regimes (though our previous work also uses European fixed exchange rate observations; Flood and Rose (1996)). UIP may work differently for countries in crisis, where both exchange and interest rates display considerably more volatility. This volatility raises the stakes for financial markets and central banks; it also may provide a more statistically powerful test for the UIP hypothesis. UIP may also work differently over time as financial markets deepen; UIP deviations may also vary across countries for the same reason, as recently argued by Bansal and Dahlquist (2000). Finally, and as the proximate motivation for this paper, deviations from UIP are the basis for interest rate defenses of fixed exchange rates. Consider the actions of the monetary authority of a country under speculative pressure that is considering responding with an increase in interest rates—the classic interest rate defense. If UIP holds, the domestic interest rate increase is offset exactly by a larger expected currency depreciation. Investors see through the policy actions, so that no advantage is conferred to domestic securities. Policy exploitable deviations from UIP are, therefore, a necessary condition for an interest rate defense.

In this short piece, we test UIP using recent high-frequency data from a large number of countries. We use data from the 1990s, and include all the major currency crises. We find that the old consensual view needs updating. While UIP still does not work well, it works better than it used to, in the sense that high interest rate countries at least tend to have depreciating currencies (though not equal to the interest rate differential). There is a considerable amount of heterogeneity in our results, which differ wildly by country. Some of this is systematic; we find that UIP works worse for fixed rate countries. However, there is less heterogeneity by forecasting horizon, and almost none by country income.

In section II we lay out our methodology; the following section provides a discussion of our data set. Our main UIP results are presented in section IV. The paper ends with a brief summary.

II. METHODOLOGY

We use standard methods (summarized in Flood and Rose, 1996). The hypothesis of uncovered interest parity can be expressed as:

$$(1+i_t) = (1+i^*_t)E_t(S_{t+\Delta})/S_t \tag{1}$$

where: i_t represents the return on a domestic asset at time t of maturity Δ; i^* is the return on a comparable foreign asset; S is the domestic currency price of a unit of foreign exchange; and $E_t(.)$ represents the expectations operator conditional upon information available at t.

We follow the literature by taking natural logarithms and ignoring cross terms (most of the countries we consider have only low interest rates). Assuming rational expectations and rearranging, we derive:

$$\begin{aligned} & E_t(s_{t+\Delta} - s_t) \approx (i-i^*)_t \\ \Rightarrow \quad & (s_{t+\Delta} - s_t) = \alpha + \beta(i-i^*)_t + \varepsilon_t \end{aligned} \tag{2}$$

where: s is the natural logarithm of S; ε_t is (minus) the forecasting error realized at $t+\Delta$ from a forecast of the exchange rate made at time t; and α and β are regression coefficients. Equation (2) has been used as the workhorse for the UIP literature. The null hypothesis of UIP can be expressed as Ho: $\alpha=0$, $\beta=1$, though in practice almost all the focus in the literature has been on β. Since ε_t is a forecasting error, it is assumed to be stationary and orthogonal to information available at time t (including interest rates). Thus, OLS is a consistent estimator of β; it is the standard choice in the literature, and we follow this practice. Researchers have typically estimated β to be significantly negative (also, α is often found to be non-trivial).[2]

In practice, we modify testing equation (2) in two slight ways. First, we pool data from a number of countries, an admissible way of increasing the sample under the null hypothesis.[3] Second, we use data of daily frequency for exchange rate forecasts of up to one-quarter (year) horizon. The fact that Δ is greater than unity induces ε to have a moving average "overlapping observation" structure. We account for this by estimating our covariance matrices with the Newey and West (1987) estimator, with an appropriate (Δ) number of off-diagonal bands.

[2] Many have tried to interpret deviations from UIP as risk premia; here we simply try to measure UIP deviations carefully and encourage others to link these deviations to other phenomena.

[3] It is likely that many of the countries are receiving correlated shocks, so that a SUR technique (that takes into account this cross-sectional dependence) would result in more efficient estimates; we did this in our 1996 paper. Nevertheless, we do not pursue this angle here, since to use SUR, one has to throw out observations when one or more countries are missing data; this results in a loss of efficiency. Further, the real problem with UIP, at least in our sample, is in the first moment of the data, not the precision of the slope estimates.

III. The Data Set

We are interested in studying how UIP performs of late in a variety of countries, especially those suffering from the currency crises that marked the 1990s. These crises were usually surprising events requiring quick policy responses.[4] In this spirit, we study the crises using a high-frequency cross-country data set. High-frequency data is of special importance to us given our focus on the interest rate defense of fixed exchange rates.

We gathered daily data for the interest and exchange rates of twenty-three countries during the 1990s. Our sample includes thirteen developed countries (Australia, Canada, Denmark, Finland, France, Germany, Italy, Japan, Norway, Sweden, Switzerland, the United Kingdom and the United States). We choose these countries to allow us to examine a variety of exchange rate regimes ranging from the floating Australian and Canadian dollars to countries like Denmark and France, European Monetary System (EMS) participants who joined European Economic and Monetary Union (EMU). A number of these countries also experienced currency crises in the 1990s, including Finland, Italy, Sweden, and the United Kingdom. We include also data for ten important and interesting developing countries (Argentina, Brazil, Czech Republic, Hong Kong, Indonesia, Korea, Malaysia, Mexico, Russia, and Thailand). The crises experienced by these countries account for most of the important action in the 1990s; we include all "the usual suspects." Indeed, it is difficult to think of an important emerging market that did not experience a crisis at some point during the 1990s. Nevertheless, there are considerable periods of tranquility through the period. These, together with the many successful and unsuccessful speculative attacks, lead us to believe that our estimates will not suffer from the "peso problem."

Our data are drawn from two sources. Whenever possible, we use the Bank for International Settlements (BIS) data set. Our default measure of exchange rates is QBCA, a representative dollar spot rate quoted at 2:15pm Brussels time. Our default measure of interest rates is JDBA, a one-month bid rate from the euro market quoted at about 10:00am Swiss time. However, a number of our countries do not have one or both of these series available. Accordingly, we supplement our BIS data with series drawn from Bloomberg. To check the sensitivity of our results with respect to the monthly forecast horizon, we include also interest rate data for three different maturities: one-day; one-week; and one-quarter. Further details (including mnemonics) and the data set itself are available online. The data set has been checked and corrected for errors.

We use the United States as the "center country" for all exchange rates (including Germany), except for nine European countries (Czech Republic, Denmark, Finland, France, Italy, Norway, Sweden, Switzerland, the United Kingdom), where we treat Germany as the anchor. We choose our center countries in this way to shed the maximum amount of light on the efficacy of the interest rate defense.

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[4] See e.g., Rose and Svensson (1994) and Boorman et al. (2000).

Figure 1 contains time-series plots of the exchange rates. The price of an American dollar rates is portrayed for all countries except for the nine European countries, which portray the price of a DM. (Scales vary across different plots, as they do in all the figures.) The breaks in series are usually associated with currency crises or other regime breaks. For instance, the Brazilian exchange rate shows clearly both the adoption of the real after the hyperinflation of the early 1990s, and the flotation of the real in January 1999. Similar breaks are apparent for many other countries, including: Indonesia, Italy, Korea, Malaysia, Mexico, Russia, and Thailand. The convergence of the EMS rates and the creation of the euro in 1999 are also apparent in the (nonGerman) EMU rates.

Figure 2 is an analogue showing interest rates. Monthly interest rates are shown for all countries except for Russia (where weekly rates are shown since the monthly series is short), Finland and Korea (where quarterly rates are shown for the same reason).[5] Here the currency crises appear as spikes in interest rates. These spikes are particularly obvious during the EMS crisis of 1992-93 (for e.g., Denmark, France, Italy, Norway, and Sweden), the Mexico crisis of 1994-95 (for Argentina and Mexico), the Asian crisis of 1997 (for Hong Kong, Indonesia, Korea, Malaysia, and Thailand), and the Russian crisis of 1998.

Figure 3 combines the exchange and interest rate data into a single series, which we call "excess returns." Excess returns ("er") are defined as $[er_{t+\Delta} \equiv (s_{t+\Delta}-s_t)-(i-i^*)_t]$, annualized appropriately. Under the UIP null hypothesis (Ho: α=0, β=1) $E_t er_{t+\Delta}=0$. Again, we use a monthly horizon as our default (so that we use one-month interest rates and set Δ to one month); the only exceptions are Russia (we use weekly rates and horizon), Finland and Korea (quarterly rates and horizon are used).

In essence, the plots in Figure 3 show the results of taking a short position in the currency. For example, since Argentina, did not deviate from its peg with the U.S. dollar, the payoff from attacking the Argentine peso was consistently negative throughout the 1990s, dramatically so during the interest rate defense against the 'Tequila' attacks of early 1995. The successful attacks against the Korean won, Mexican peso, and the Russian ruble show up as large positive payoffs realized at the time of the flotations.

Where Figure 3 provides a look at a combination of exchange rate changes and interest differentials over time, Figure 4 graphs the exchange rate changes and interest rate differentials against each other. Instead of examining the time-series patterns on a country-by-country basis as in Figure 3, we pool the data across countries. Exchange rate changes (on the ordinate) are more volatile than interest rate differentials (on the abscissa) for each horizon. There is clearly no tight relationship between exchange rate changes and interest differentials. This is no surprise; interest differentials are not very useful in predicting exchange rate changes. Since the visual impression is unclear, we now proceed to more rigorous statistical analysis, which is essentially an analogue to the graphs of Figure 4.

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[5] We define a month as 22 business days, a week as five business days, and a quarter as 65 business days.

Figure 1. Exchange Rate Data

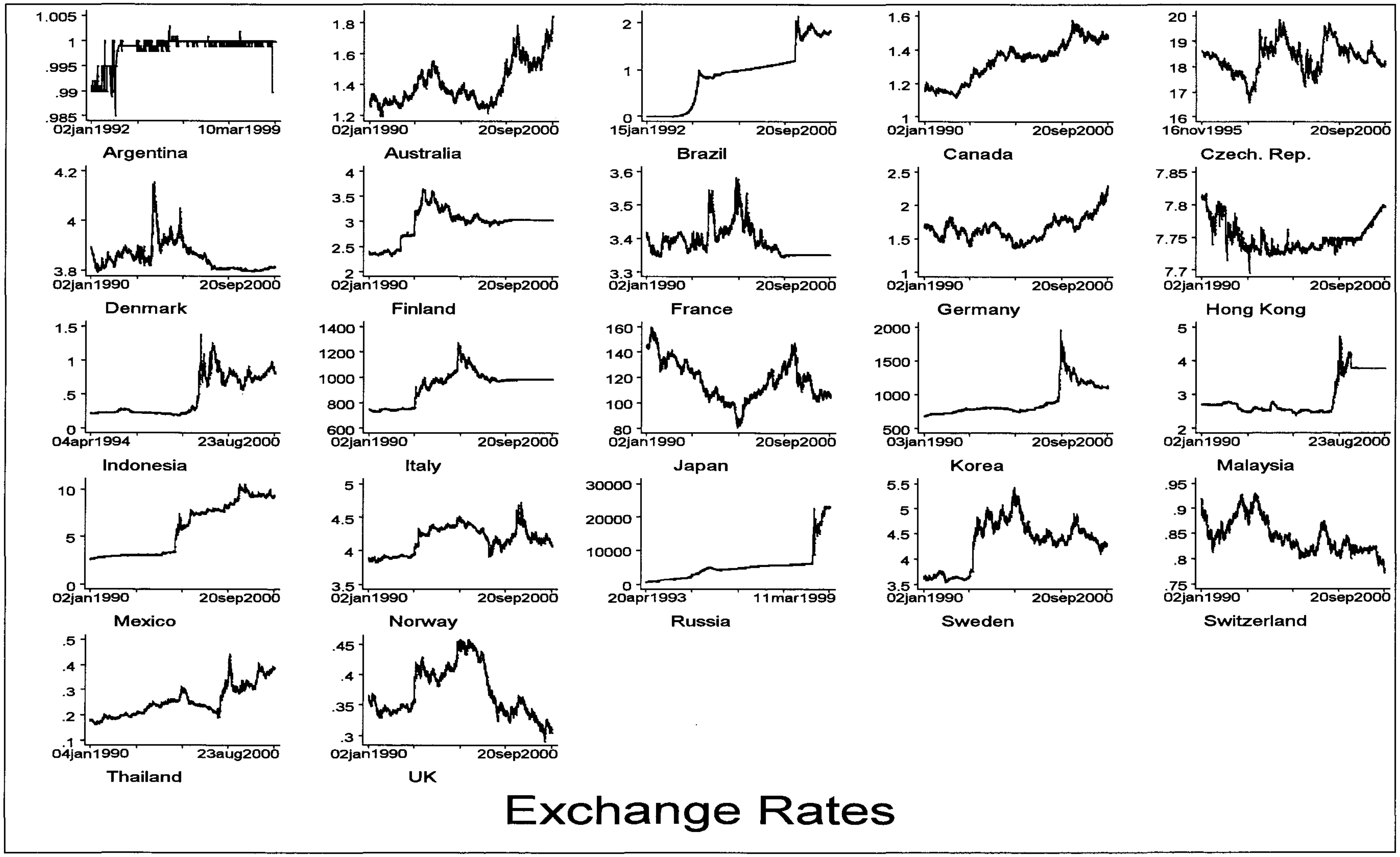


Figure 2. Interest Rate Data

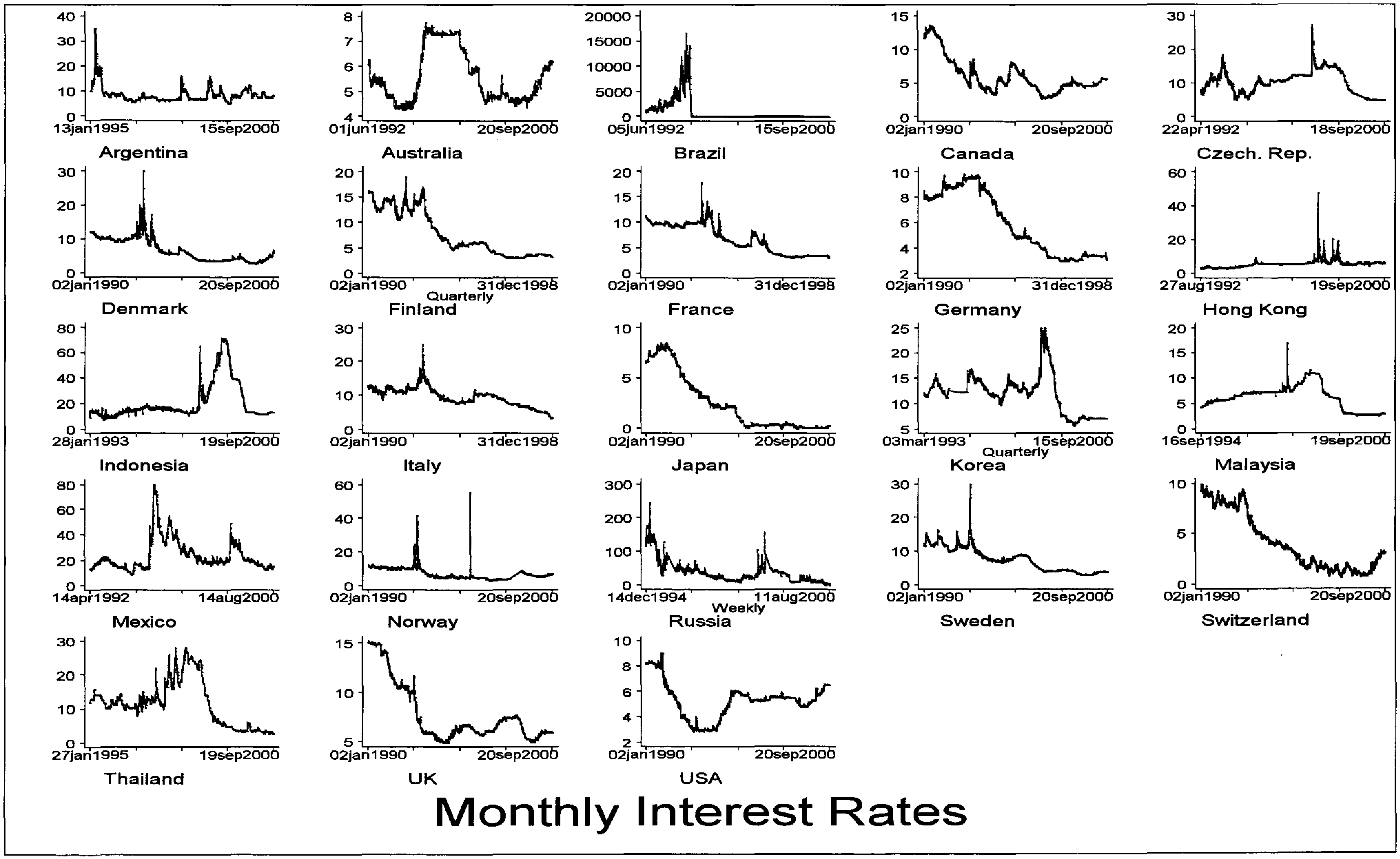


Figure 3. Monthly Excess Returns

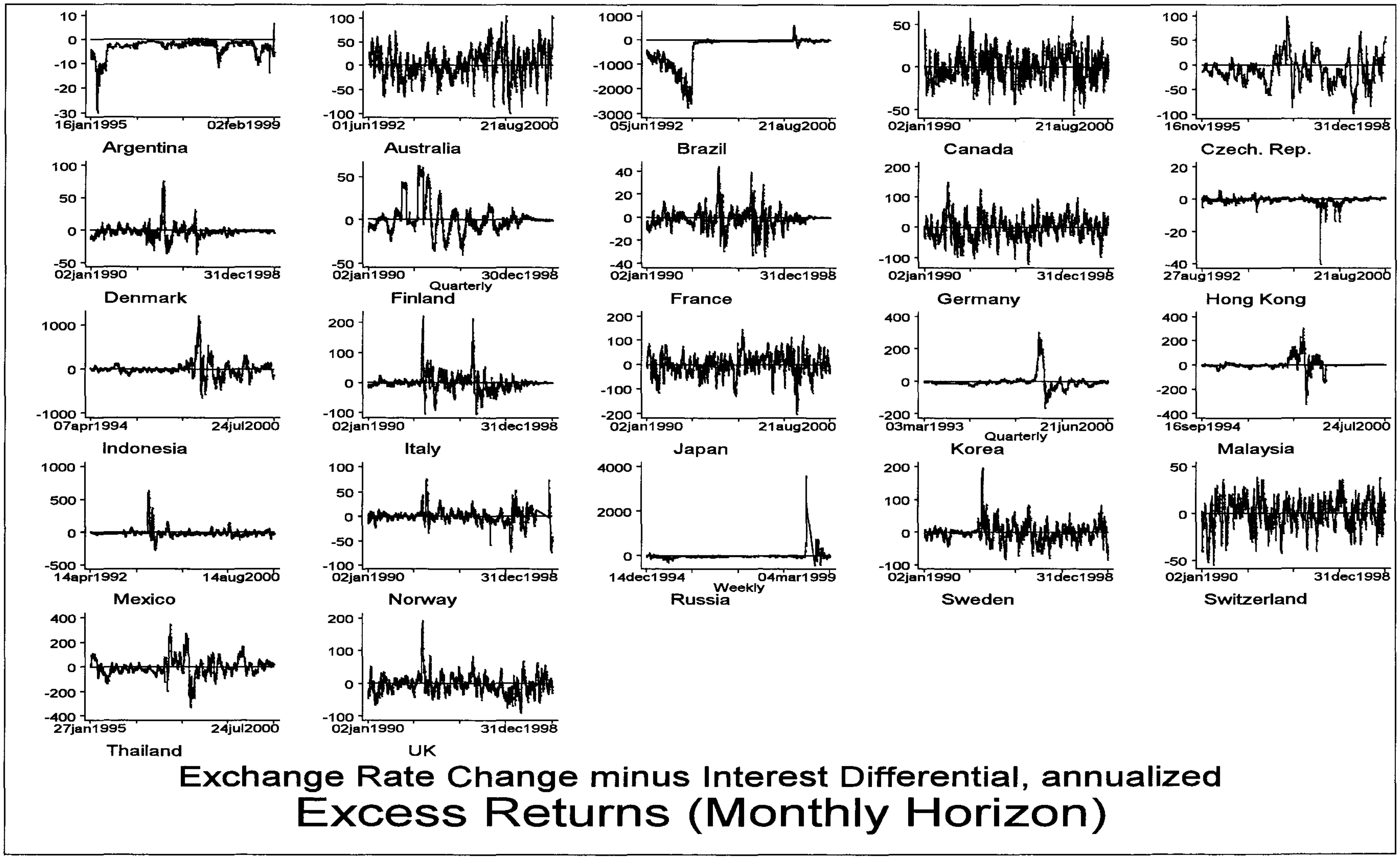
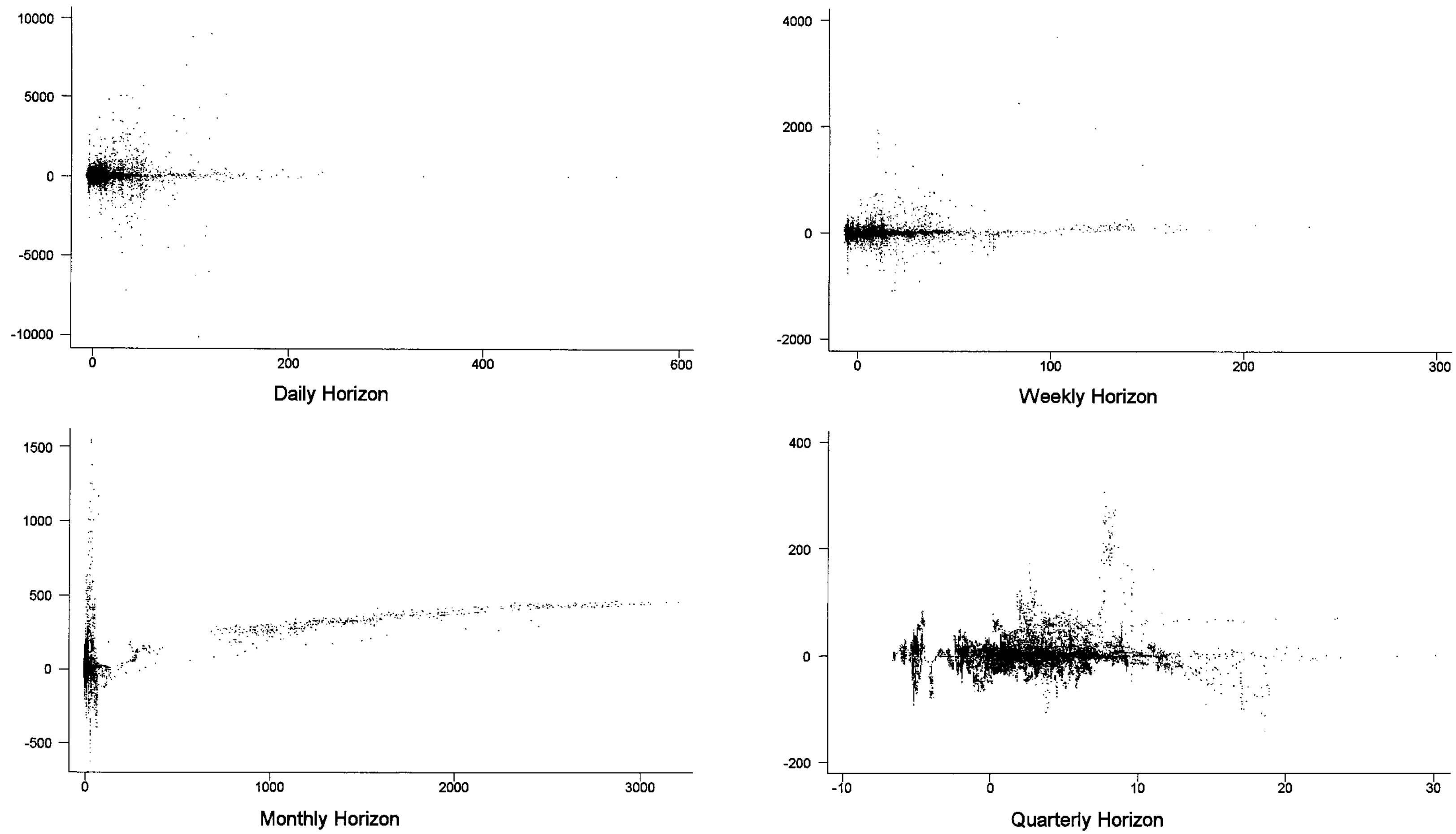


Figure 4. Exchange Rate Changes and Interest Rate Differentials



Annualized Percentage Changes of Daily Data, 1990s

Exchange Rate Changes (y-axis) against Interest Differentials (x)

IV. UIP Regression Analysis

Table 1 provides estimates of β when equation (2) is estimated on a country-by-country basis; that is, the regressions are estimated for an individual country over time. Newey-West standard errors that are robust to both heteroskedasticity and autocorrelation (induced by the overlapping observation problem) are recorded in parentheses below. Estimates of the intercept (α) are not reported; they are of less interest, and are usually insignificantly different from zero at conventional confidence levels. We focus on the monthly horizon results, but tabulate the results for the three other forecasting horizons as a sensitivity check.

The most striking thing about the estimates of β is their heterogeneity. Of the 21 estimates, 12 are negative and 7 are positive (two are essentially zero). This in itself is interesting, since virtually all estimates in the literature are negative. Further, all but one of the negative estimates are insignificantly so, while three of the positive coefficients are significant. At conventional significance levels, only nine of the slopes are insignificantly different from the hypothesize value of unity. However, this is frequently because of large standard errors rather than point estimates close to unity, so even this evidence is weak.[6] Finally, the point estimates vary across forecast horizon, often switching signs across horizons.

Table 2 pools the data across countries, so that a single β is estimated for all countries and periods of time. Here too, the results are striking. In particular, the top panel shows that the pooled estimate is positive at all four horizons. At the monthly horizon, β is significantly positive, though at .19 it is far below its theoretical value of unity. At the other horizons, β is even higher and insignificantly different from unity (and strikingly close to unity at the daily and weekly horizons, though with large standard errors).[7] Still, pooling is a dubious procedure given the heterogeneity manifest in Table 1, so we do not take these results too seriously.[8]

The other panels of Table 2 add interactions between dummy variables and the interest differential. Panel B includes an interaction with the exchange rate regime. We consider Argentina, Denmark, France and Hong Kong to have fixed their exchange rates throughout the sample, while we classify Australia, Canada, Germany, Japan, Norway, and Switzerland as floaters. The other ("crisis") countries experienced at least one regime switch and are omitted as our control group.

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[6] Some of the standard errors are very low however; they may be biased because of nonnormalities associated with jumps at currency crises. Hence we recommend that readers not take our covariance estimates too literally.

[7] Chinn and Meredith (2000) find even more positive results using long-maturity data.

[8] This is especially true since the Hildreth-Houck random-coefficients method delivers slope coefficients, which are economically and statistically insignificant on our pooled data.

Table 1. Uncovered Interest Parity Tests by Country
OLS Estimates of β from $(s_{t+\Delta} - s_t) = \alpha + \beta(i\text{-}i^*)_t + \varepsilon_t$
Newey-West standard errors in parentheses.

Horizon:	**Daily**	**Weekly**	**Monthly**	**Quarterly**
Argentina	.03		.00	-.003
	(.11)		(.01)	(.002)
Australia			-3.58	
			(2.55)	
Brazil	15.3		.19	
	(15.9)		(.01)	
Canada			-.58	
			(.54)	
Czech Rep.	.73		-1.27	-1.41
	(1.13)		(.85)	(1.14)
Denmark			-.03	
			(.70)	
Finland	2.50		7.06	2.56
	(2.20)		(3.80)	(1.21)
France			-1.42	
			(.62)	
Germany	-.60		.13	-.11
	(1.32)		(1.11)	(1.16)
Hong Kong	-.35	-.20	.00	-.00
	(.18)	(.06)	(.03)	(.02)
Indonesia	.22		-1.19	
	(2.05)		(1.13)	
Italy	1.66		.29	-.75
	(1.87)		(2.55)	(1.92)
Japan	-.82	-3.14	-1.71	-1.84
	(1.36)	(1.83)	(1.11)	(1.19)
Korea	3.41	1.42		-.31
	(4.12)	(2.08)		(1.57)
Malaysia			2.24	2.07
			(2.08)	(1.95)
Mexico	-.37	-.60	-.77	
	(1.00)	(.66)	(.70)	
Norway			.59	
			(.75)	
Russia	1.48	1.29	.22	
	(1.46)	(.58)	(.11)	
Sweden	.08		-.44	1.28
	(.03)		(.95)	(2.03)
Switzerland			-2.08	
			(1.40)	
Thailand	.52	-1.29	-.83	
	(1.86)	(1.57)	(1.80)	
UK	-1.15		-1.26	-1.42
	(1.06)		(.97)	(.98)

Table 2. Pooled UIP Tests

OLS Estimates of β from $(s_{it+\Delta} - s_{it}) = \alpha + \beta(i\text{-}i^*)_{it} + \varepsilon_{it}$

Newey-West standard errors in parentheses.

Panel A: No interactions

	β (se)	**Number Observations**
Daily	.86 (.65)	26,972
Weekly	.87 (.34)	8,033
Monthly	.19 (.01)	37,992
Quarterly	.29 (.39)	18,942

Panel B: Exchange Rate Regime Interactions

	β (se)	**FIX*β (se)**	**FLOAT*β (se)**	**Number Obs.**	**P-value: Interactions=0**
Daily	.87 (.67)	-.94 (.58)	-.71 (1.23)	26,972	.21
Weekly	.92 (.37)	-.87 (.29)	-1.26 (1.40)	8,033	.00
Monthly	.19 (.01)	-.93 (.32)	-.20 (.48)	37,992	.01
Quarterly	.43 (.49)	-.54 (.42)	-.47 (.94)	18,942	.44

Panel C: Country Income Interactions

	β (se)	**OECD*β (se)**
Daily	.97 (.75)	-.80 (.48)
Weekly	.92 (.37)	-1.28 (1.40)
Monthly	.19 (.01)	-.31 (.36)
Quarterly	.27 (.54)	.06 (.68)

Panel D: High Inflation Interactions

	β (se)	**High Inflation*β (se)**
Daily	.38 (.47)	,89 (1.22)
Weekly	.32 (.33)	.71 (.50)
Monthly	.00 (.42)	.19 (.42)
Quarterly	.31 (.40)	-.45 (.37)

We find that both fixers and floaters have significantly lower estimates of β, in contrast to Flood and Rose (1996) who use data from late 1970s through the early 1990s. Thus the marginally better UIP results that stem from pooling across countries must be largely due to the inclusion of countries that were successfully attacked.

When we interact the interest rate differential with a dummy variable that is unity for countries that were members of the OECD at the beginning of the decade, we find insignificantly different results. This result stands in contrast to the estimates provided by Bansal and Dahlquist (2000).

Finally, we dummy out the three countries, which experienced high inflation at some point during the sample period (Argentina, Brazil and Russia). When we do so, we find in Panel D that some of our positive results stem from high inflation countries; the interaction terms are typically positive and economically large (especially at shorter horizons). However, our point estimates for β are still positive or zero, unlike those in most of the literature (though our standard errors are large).

V. Conclusion

Uncovered interest parity works better than it used to, in the sense that interest rate differentials seem often to be followed by subsequent exchange rate depreciation. The fact that this relationship has been positive on average during the 1990s contrasts sharply with the typically negative estimates of the past. At the daily and weekly horizons, this relationship even seems to be proportionate if one includes high-inflation countries. Nevertheless, there are still massive departures from uncovered interest parity. The enormous cross-country heterogeneity in the UIP is relationship uncorrelated with either the exchange rate regime or country income.

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