Uncovered Interest Parity in Crisis: The Interest Rate Defense in the 1990s*

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Abstract

This paper tests for uncovered interest parity (UIP) using daily data for twenty-three 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. Finally, we find evidence that varies considerably across countries and time, but is usually weakly consistent with an effective 'interest rate defense' of the exchange rate.

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1: 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. We also provide evidence on whether departures from UIP make viable an "interest rate defense" of a fixed exchange rate regime.

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 argues 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 2 we lay out our methodology; the following section provides a discussion of our data set. Our main UIP results are presented in section 4. Section 5 presents our evidence on the interest rate defense. The paper ends with a brief summary.

2: Methodology

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

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

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:

$$E_t(s_{t+\Delta} - s_t) \approx (i-i^*)_t$$

$$\Rightarrow \quad (s_{t+\Delta} - s_t) = \alpha + \beta(i-i^*)_t + \varepsilon_t \qquad (2)$$

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: $\alpha=0$, $\beta=1$. 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, and α to be non-trivial.[1]

In practice, we modify testing (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. 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 number of off-diagonal bands.

3: 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.[2] 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 UK and the US). 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 UK. We include also data for ten 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 euro market bid rate 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 UK), 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.

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 (non-German) 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).[3] 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: $\alpha=0$, $\beta=1$) $E_t er_{t+\Delta}=0$. Again, we use a monthly horizon as our default (so that we use one-month interest rates and set ? 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 US 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.

4. 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 (a) are not reported. 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 twenty-one estimates, twelve are negative and seven 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. 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).[4] Still, pooling is a dubious procedure given the heterogeneity manifest in Table 1, so we do not take these results too seriously.[5]

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.

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. 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).

5. The Interest Rate Defense

In this section we develop evidence on the efficacy of the interest rate defense.

The Framework

The model upon which we base our test is the one developed by Flood and Jeanne (2000) (FJ), itself an adaptation of Krugman (1979), and Flood and Garber (1984) that allows for a policy-exploitable wedge in UIP.[6] In FJ, defense efficacy is measured in terms of prolonging the fixed exchange rate regime. In other words, the defense works if raising the domestic-currency interest rate makes the fixed rate regime survive longer than it otherwise would without the rate increase. The UIP wedge in FJ is proportional to the worldwide privately held stock of domestic government issued domestic-currency denominated nominal debt.[7]

The main FJ results are: a) increasing the domestic-currency interest rate *prior* to a speculative attack will always *hasten* the onset of the speculative attack for fiscal reasons; and b) committing credibly to increase the domestic-currency interest rate *after* the speculative attack may obstruct the speculative attack. The most striking result is that it is the actions to be taken *after* the attack – like promising to hit back – that may deter the attack.

The key equation in FJ is:

$$i^*_t = i_t + E_t s_{t+\Delta} - s_t + \theta B_t/S_t \tag{3}$$

where: θ is a positive constant; and B_t is worldwide private holding of domestic-government issued domestic-currency bonds. The last term, $\theta B_t/S_t$, is the UIP wedge needed to analyze the interest rate defense.

FJ assume that all nominal bonds issued by the domestic government, N_t, are either held privately, B_t, or are held by the domestic monetary authority as domestic credit, D_t. FJ also assume that after the speculative attack, the exchange rate floats and domestic-monetary authority's international reserves are constant at zero. The wedge thus becomes

$$\theta[(N_t-D_t)/S_t] = \theta[(N_t-M_t)/S_t] = \theta(n_t-m_t)$$

where: $n \equiv N/S$; $m \equiv M/S$; M is high-powered money; and D=M because reserves are zero.

The state variable driving FJ to the attack precipice and beyond is N. During the fixed exchange rate regime that precedes the attack, the exchange rate stabilizes goods prices, and the government fixes the interest rate on its debt. Tracking N's growth is therefore an accounting exercise. In the post-attack floating rate epoch, FJ solve their model for n, the real value of government-issued debt.[8]

The Role of Excess Returns

We study the efficacy of the interest rate defense by first using the model to find the length of the fixed rate epoch, and then examining the data to find the direction in which interest rate increases change observable determinates of efficacy.

The connection to excess returns proceeds in three steps. First, we solve for n noting that at the instant of the attack we must have $n=N/\overline{S}$, where $\overline{S}$ is the pre-attack fixed exchange rate.[9] Second, since $\overline{S}$ is fixed and N grows in lockstep with the mechanical pre-attack deficit, anything (and only those things) that increases n must increase the length of the fixed rate epoch also. Third, we have no daily data on N or, therefore, on n but we do have daily excess returns. According to the above model:

$$er_{t+\Delta} = \theta(m_t - n_t) + s_{t+\Delta} - E_t s_{t+\Delta} \tag{4}$$

Since neither money nor debt is available at a daily frequency, our investigation of the efficacy of the interest rate defense involves regressing $er_{t+\Delta}$ on i_i. If we estimate the following OLS regression:

$$er_{t+\Delta} = \lambda + \gamma i_t + v_t, \tag{5}$$

the question then is what can be learned?[10]

The FJ model helps but still does not allow a straightforward interpretation of the regression results. We measure $\hat{g} = \theta(\Delta m/\Delta i - \Delta n/\Delta i)$; thus, even if $q > 0$ we do not measure the sign of $\Delta n/\Delta i$ directly. Instead, we measure it combined with $\Delta m/\Delta i$. We assume, therefore, that m is negatively related to i through substitution in money demand.

Thus if $\hat{g} > 0$, we conclude $\Delta n/\Delta i<0$, so that the interest rate defense is *ineffective*. If however, $\hat{g} < 0$, the test is *inconclusive* but *consistent* with the efficacy of the interest rate defense. When $\hat{g} = 0$, either $q = 0$ so that the interest rate defense is *ineffective* because the UIP wedge is not exploitable, or $\Delta m/\Delta i = \Delta n/\Delta I$, making the interest rate defense *effective* when $\Delta m/\Delta i <0$. Our only possibility for strong results requires $\hat{g} > 0$.

Model-specific considerations make our test sound narrow. But it is also possible to put a more positive spin on our evidence. What policymakers are trying to accomplish with an active interest rate defense is to decrease the expected excess return to (short) positions against the domestic currency. That is, by increasing the domestic interest rate the authorities are trying to increase the expected excess return to holding domestic-currency debt. Our empirical work simply asks: Does this strategy usually work?

In Tables 3 through 5 we provide a number of estimates of γ. The results tabulated in Table 3 are analogues to those in Table 1 for UIP; these estimates of γ use time-series data on a country-by-country basis. Table 4 uses data that is pooled across countries on a year-by-year basis. Finally, Table 5 is the analogue to Table 2, and provides estimates of γ that use data which is pooled across both countries and time.

The estimates in Tables 3-5 show that γ is typically negative, but vary wildly. The country-specific time series evidence of Table 3 shows that γ varies substantially across countries and even across horizons within countries. The negative estimates for Argentina are striking but intuitive, since Argentina successfully used the interest rate defense to support the peso through the 1990s; results for Hong Kong are similar. But a number of countries such as Italy and Malaysia provide positive estimates of γ. There is also an interesting lack of strong results for Korea, Mexico, Thailand and the UK, all victims of highly visible and successful speculative attacks.

The heterogeneity of results also characterizes the results in Table 4 that pool data across countries within specific years. Perhaps the most striking results are the positive coefficients that characterize 1997 (the year of the Asian crisis) for all maturities. Manifestly an effective interest rate defense did not characterize that crucial year.

The results in Table 5 pool observations across countries and time. The typical estimate of γ is negative, significantly so at the key monthly horizon. This is consistent with the efficacy of the interest rate defense. However, the lower panels of the table show that we are unable to find a link between the efficacy of this strategy and either the exchange rate regime or income.

Conclusion

Uncovered interest parity works better than it used to, in the sense that interest rate differentials seem typically 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. Nevertheless, there are still massive departures from uncovered interest parity. There is enormous heterogeneity in the UIP relationship across countries, though we have been unable to find a close relationship between UIP departures and either the exchange rate regime and country income.

We also presented evidence on the efficacy of the 'interest rate defense' of a fixed exchange rate. Our evidence on the interest rate defense is both model-specific and loose in the sense that data limitations prevent a direct test of the model. Nevertheless, we think it is suggestive. We cannot establish the effectiveness of this strategy; but neither has our empirical work been able to unequivocally rule it out; so far as we are concerned, the door is open. However, the evidence is murky, and we provide only slightly more evidence consistent with the interest rate defense than we do for the complete absence of any effect from the domestic interest rate on UIP deviations. We think of this as an intriguing place to pass on the baton.

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

	ß (se)	**Num. Obs.**
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

	ß (se)	**FIX*ß (se)**	**FLOAT*ß (se)**	**Num. 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

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

Table 3: Excess Return/Domestic Interest Rate Relationship by Country

OLS Estimates of γ from $er_{t+\Delta} = \lambda + \gamma i_t + v_t$
Newey-West standard errors in parentheses.

Horizon:	Daily	Weekly	Monthly	Quarterly
Argentina	-.96 (.11)		-.96 (.01)	-.96 (.01)
Australia			-1.78 (2.16)	
Brazil	-65 (94)		-.81 (.01)	
Canada			-1.56 (.41)	
Czech Rep.	-.28 (1.14)		-2.41 (.92)	-2.51 (1.15)
Denmark			-.27 (.25)	
Finland	1.10 (1.22)		2.98 (1.68)	1.01 (.60)
France			-.22 (.16)	
Germany	-2.39 (1.54)		-1.56 (1.34)	-1.76 (1.42)
Hong Kong	-.54 (.17)	-1.14 (.06)	-.71 (.09)	-.69 (.10)
Indonesia	-.76 (2.06)		-2.17 (1.14)	
Italy	.86 (1.04)		1.36 (1.42)	1.16 (.96)
Japan	-1.50 (1.41)	-8.56 (3.54)	-2.22 (1.21)	-2.49 (1.29)
Korea	2.64 (4.30)	.41 (2.06)		-1.22 (1.72)
Malaysia			1.51 (2.25)	1.26 (2.13)
Mexico	-1.26 (.97)	-1.46 (.64)	-1.60 (.66)	
Norway			.25 (.38)	
Russia	.48 (1.45)	.29 (.57)	-.78 (.11)	
Sweden	-.83 (.08)		.24 (.61)	.80 (.95)
Switzerland			-.32 (.47)	
Thailand	-.54 (1.93)	-2.51 (1.65)	-1.79 (1.81)	
UK	.31 (.80)		.01 (.71)	.13 (.85)

Table 4: Excess Return/Domestic Interest Rate Relationship by Year

OLS Estimates of γ from $er_{it+\Delta} = \lambda + \gamma i_{it} + v_{it}$
Newey-West standard errors in parentheses.

Horizon:	Daily	Weekly	Monthly	Quarterly
1990	-.43 (.93)		-.40 (.37)	.52 (.27)
1991	1.20 (2.72)		-.65 (.34)	.52 (.33)
1992	-.78 (.14)	-2.19 (.45)	-.75 (.004)	3.00 (.16)
1993	-.57 (.60)	-.85 (.54)	-.78 (.003)	.12 (.16)
1994	.33 (1.15)	.80 (.54)	-.84 (.002)	.61 (.19)
1995	-.58 (.21)	-.49 (.09)	-.73 (.02)	-2.04 (.18)
1996	-.73 (.15)	-.73 (.09)	-.70 (.02)	-1.85 (.12)
1997	.19 (1.30)	.37 (1.05)	1.10 (.34)	1.03 (.37)
1998	1.35 (2.51)	3.01 (1.81)	-.52 (.36)	-1.80 (.19)
1999	-1.52 (.88)	-.89 (.32)	-1.29 (.31)	-.25 (.36)

Table 5: Pooled Excess Return/Domestic Interest Rate Relationship

OLS Estimates of γ from $er_{it+\Delta} = \lambda + \gamma i_{it} + v_{it}$
Newey-West standard errors in parentheses.

Panel A: No interactions

	g **(se)**	**Num. Obs.**
Daily	-.11 (.64)	26,972
Weekly	-.12 (.33)	8,033
Monthly	-.81 (.01)	37,992
Quarterly	-.21 (.35)	18,942

Panel B: Exchange Rate Regime Interactions

	g **(se)**	**FIX***g **(se)**	**FLOAT***g **(se)**	**P-value: Interactions=0**
Daily	-.12 (.64)	-.38 (.26)	-.41 (.62)	.33
Weekly	-.16 (.35)	-.24 (.53)	-4.17 (3.23)	.39
Monthly	-.81 (.01)	-.50 (.21)	-.57 (.28)	.06
Quarterly	-.25 (.35)	-.37 (.18)	-.76 (.52)	.09

Panel C: Country Income Interactions

	g **(se)**	**OECD***g **(se)**
Daily	-.48 (.47)	-.05 (.68)
Weekly	-4.02 (3.04)	-.16 (.35)
Monthly	-.21 (.28)	-.81 (.01)
Quarterly	.20 (.35)	-.31 (.44)

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Figure 1: Exchange Rate Data

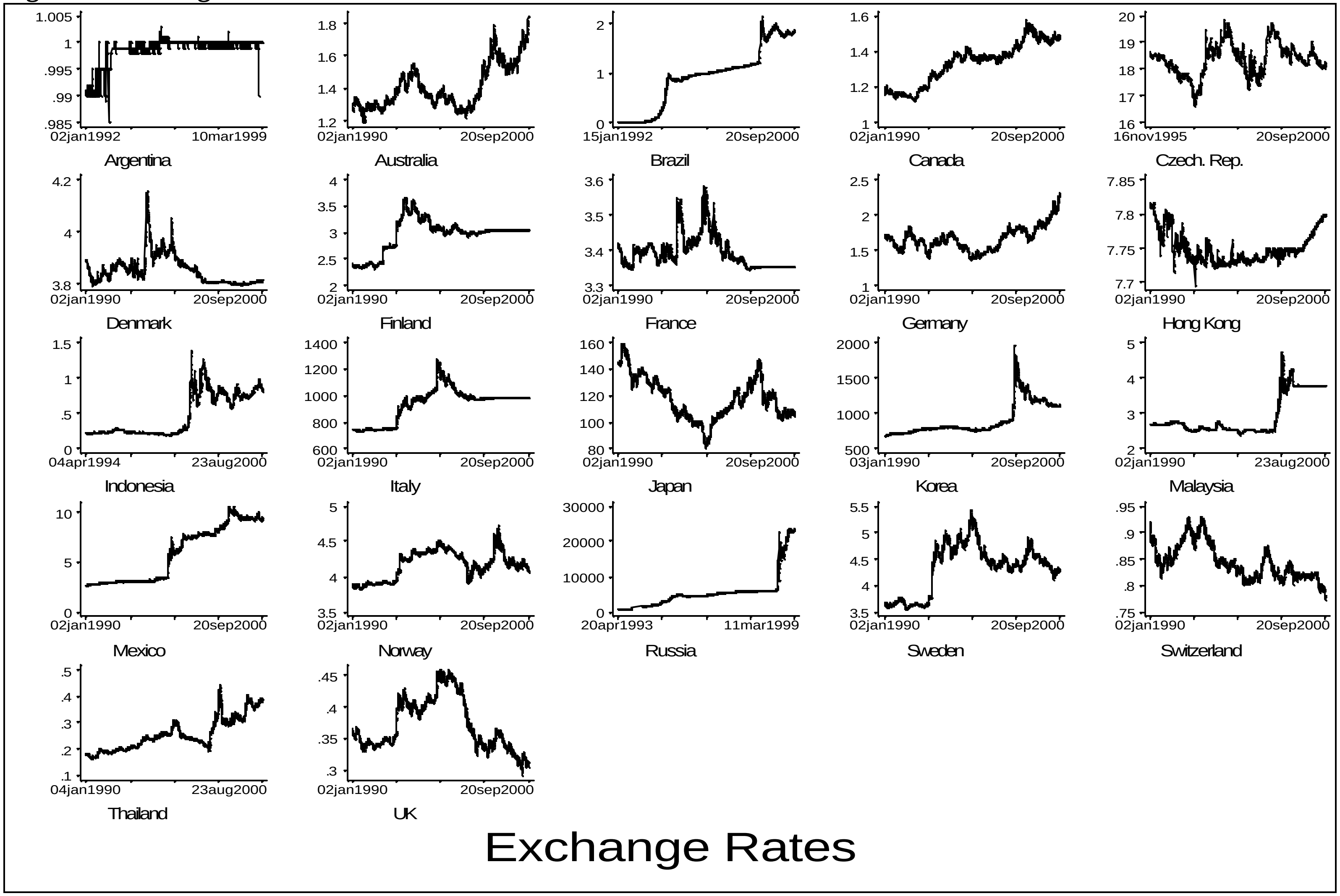


Figure 2: Interest Rate Data

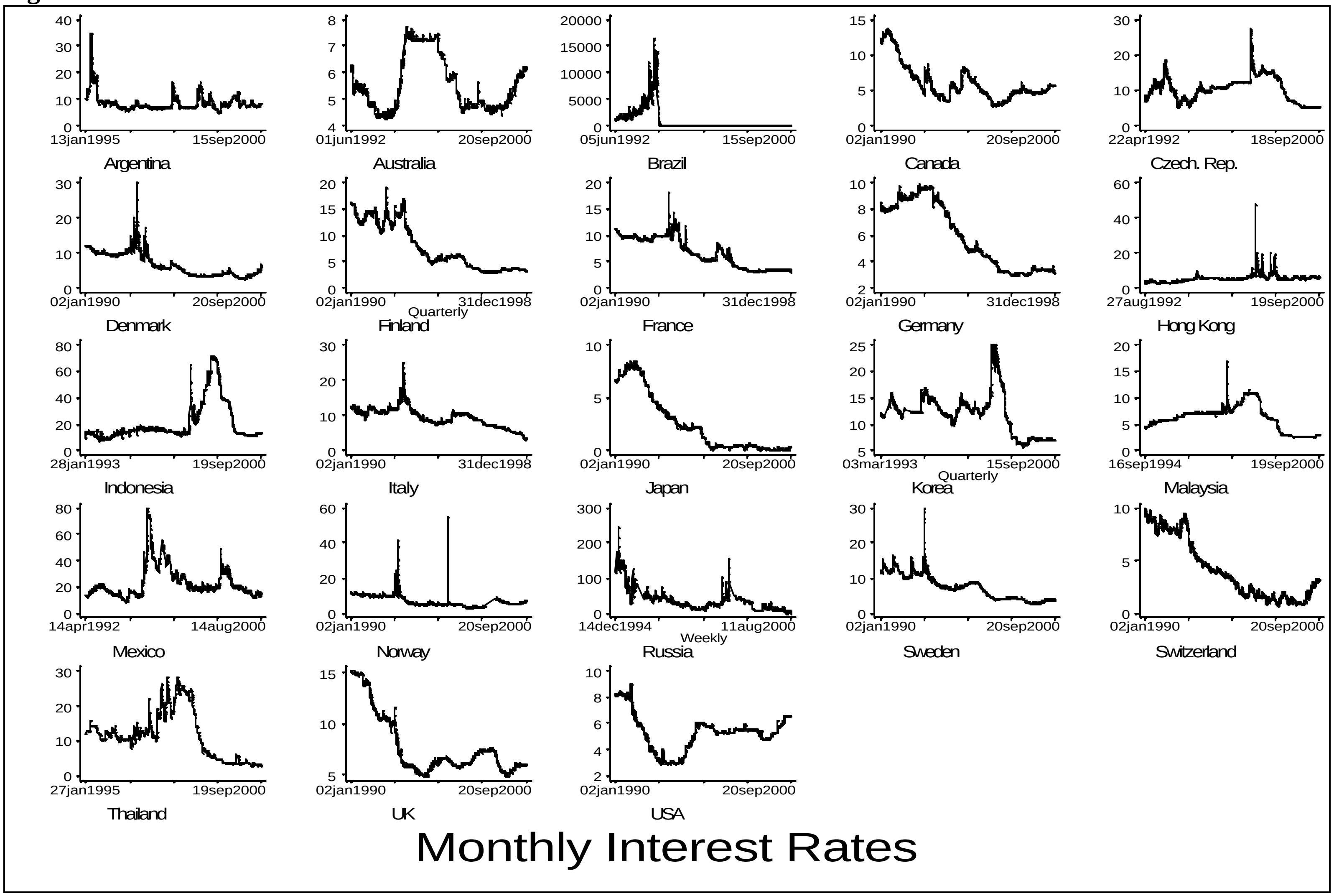


Figure 3: Monthly Excess Returns

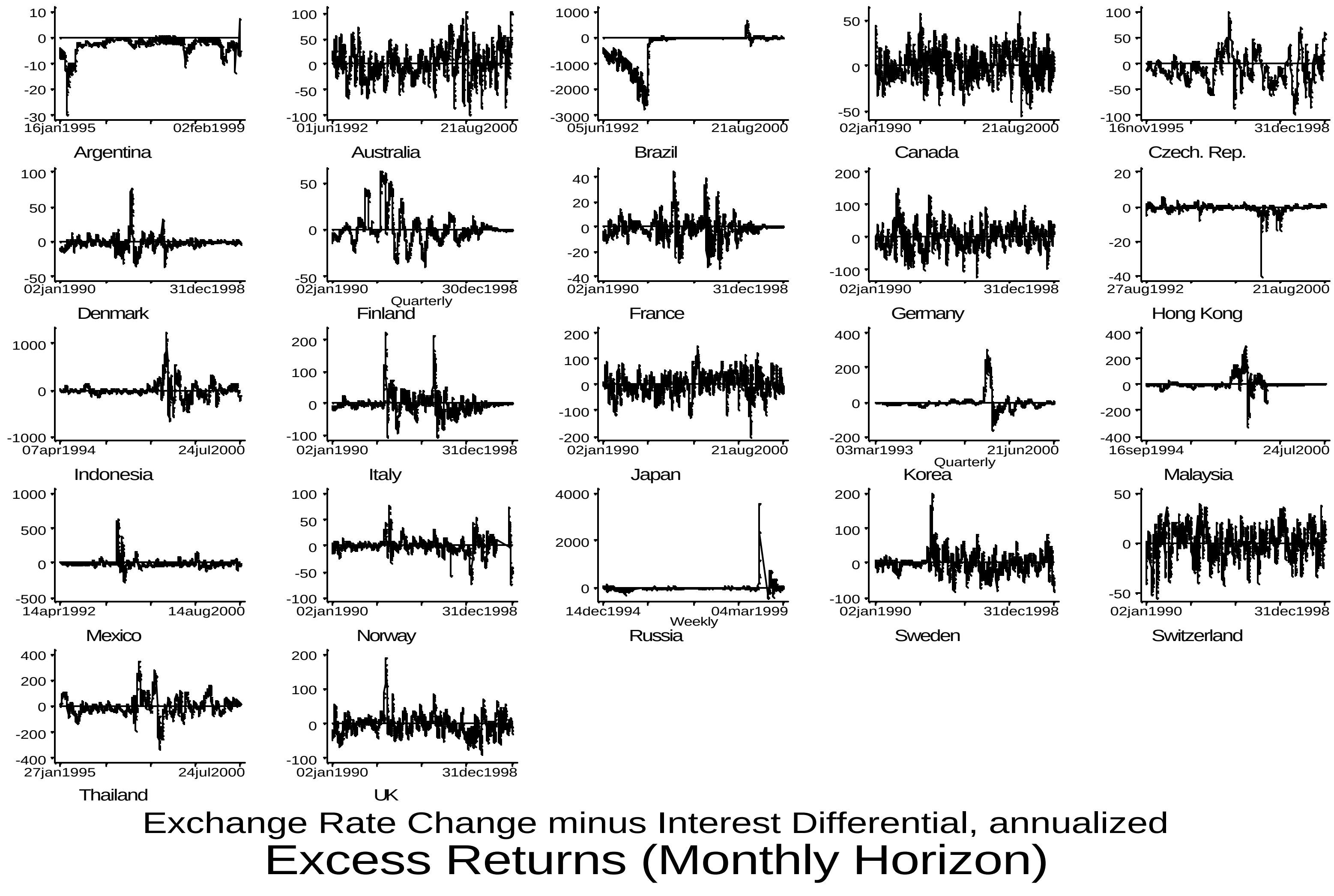
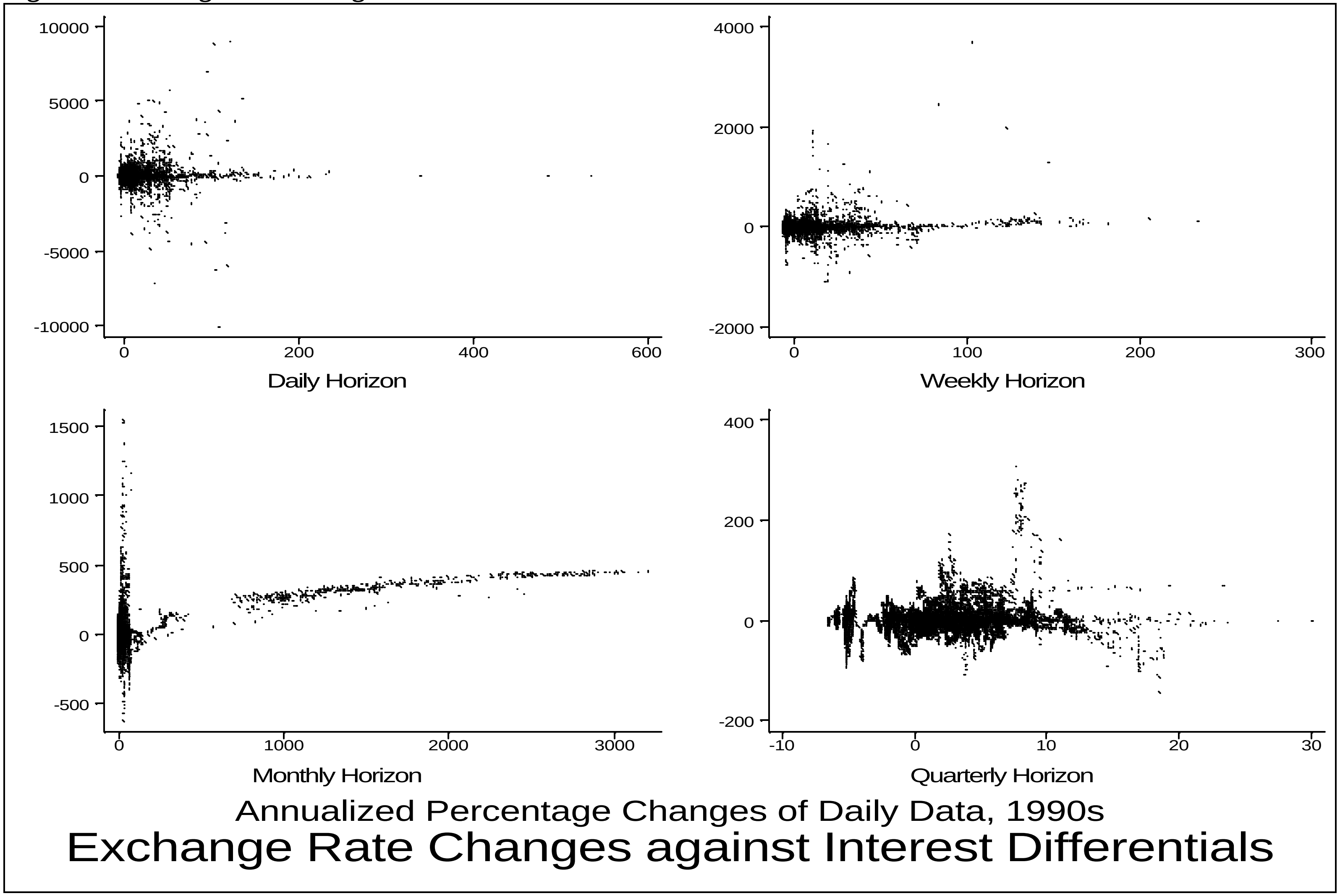


Figure 4: Exchange Rate Changes and Interest Rate Differentials



Endnotes

[1] 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.

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

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

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

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

[6] Other interest rate defense models include Bensaid and Jeanne (1997), Drazen (1999), and Lahiri and Végh (1999, 2000).

[7] This functional form is derived in Jeanne and Rose (1999) and is discussed more in FJ.

[8] FJ is a perfect foresight model. The translation of their results to real-world data requires us to refer to the permanent component of disturbances.

[9] This terminal condition would be altered slightly in a stochastic setup. See, e.g., Flood and Garber (1984).

[10] Viewing equation (5) as the linearization of equation (4) turns v_t into an error composed of an exchange rate prediction error plus a linearization error. Arbitrary exclusion restrictions are required for the model-specific interpretation that follows.