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Monotonicity and Currency Carry Trades: The Implications for Uncovered Interest Rate Parity and Expected Shortfall

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Abstract

Uncovered interest rate parity (UIP) implies that currency carry trades should not yield excess returns. We find that there is statistical evidence in favour of an increasing monotonic relation in the excess return from currency carry trades with respect to forward discounts. It follows that a larger interest rate differential indeed implies larger excess returns. Furthermore it seems that Expected Shortfall (ES) does not explain the excess return associated with currency carry trades adequately as we reject an increasing monotonic relation when we sort excess returns on their ES.

Keywords: Uncovered interest rate parity, currency carry trades, forward premium, expected shortfall, monotonicity tests

1 Introduction

In this paper we examine two factors associated with currency carry trades, namely the monotonicity implications by uncovered interest rate parity (UIP) and the relevance of Expected Shortfall (ES) as downside market risk factors for currency carry trades. Financial models and theories often imply increasing, decreasing or even flat patterns in expected returns with respect to certain risk factors or liquidity characteristics. A prime example is the covered interest rate parity (CIP) which states that there should be no arbitrage opportunities in borrowing in the domestic market to lend similar assets in a foreign market (and vice versa) while hedging the exchange rate exposure. We can hedge this exchange rate exposure through instruments such as a forward contracts. This theory is furthermore augmented by UIP which states that the expected future discount should be equal to the expected spot exchange rate appreciation and thus that currency carry trades should not yield excess returns.

Another example is the liquidity preference hypothesis (LPH) which implies that expected returns on treasury bills should monotonically increase with their time to maturity as investors prefer liquidity when all other things remain equal. More general, the no arbitrage pricing of assets implies a positive pricing kernel, or Stochastic Discount Factor (SDF), to discount future states of the world. This pricing kernel is implied to be monotonically decreasing with the quality of future states which is often measured by market returns. For example the discounting of future cash flows when pricing a bond.

In this paper we look at the monotonicity implications made by the aforementioned LPH as in Patton and Timmermann (2010), which serves as an illustration of correct implementation, but we focus on a popular hedge fund trading strategy called currency carry trades. Carry trades are based on exploiting interest rate differentials, in our case that means we short currencies with low interest rates, e.g. JPY and CHF and go long in currencies with high interest rates, e.g. AUD and NZD. Much is written about the profitability of these currency carry trades, Jylhä & Suominen (2010) for example state that it explains more than 16% of the risk adjusted returns in hedge funds and more than 33% of the fixed income arbitrage sub-index returns. However following the reasoning of UIP there should be no excess return in currency carry trades as the return we make should be the same as the expected currency appreciation. We formally test whether there is evidence for an increasing monotonic relation in excess returns with respect to their future discounts in order to test if UIP fails.

We furthermore examine if we can find evidence for an increasing monotonic relation in excess returns on one month currency carry trades when we sort on ES. Value-at-Risk and ES are closely related and well known risk measures in the financial industry and are used to quantify market risks. VaR, however, gives no indication on how large losses may be if the loss exceeds VaR, in other words VaR neglects so-called 'tail risk', and furthermore VaR is not subadditive (Acerbi and Tasche, 2002). These deficiencies gave rise to ES as a risk measure, which is the expected loss in case losses exceed VaR. ES deals with the problems associated with VaR as it accounts for tail risk and is subadditive. We focus on ES, however sorting on VaR or ES will not change the sorting relation.

The monotonic relations mentioned before can be tested through a framework developed by Patton and Timmermann (2010) and we will denote those monotonic relations tests as

MR tests for the remainder of this paper. The MR tests are straightforward in implementation and nonparametric as they adapt a bootstrap method to avoid estimating variance parameters directly. Contrary to the framework developed by Wolak (1987, 1989) and the Bonferroni bound test used by Fama (1984) a rejection of the null indicates that we find significant statistical evidence in favour of the hypothesized monotonic increasing relation. Both Wolak and Fama thus hold a monotonically increasing relation under the null whereas the MR test holds it under the alternative hypothesis. We include both these tests in our results.

We test the implications of UIP for carry trades short USD, JPY and CHF and long in other remaining G10 currencies using daily data from January 2000 to May 2017. We use the same data to test for an increasing relation in excess return sorted on ES. The increasing relation in yield to maturity implied by LPH is tested with monthly data on treasury securities in the period 1964 to 2001, similar to Patton and Timmermann (2010).

It follows that we cannot convincingly reject an increasing monotonic relation in log excess return from carry trades with respect to future discounts and thus the flat pattern implied by UIP seems to fail. It seems that bigger forward discounts or interest rate gaps indeed imply bigger excess returns. This is in line with Lustig et al. (2011) who conclude that common variation that can be found in assets after sorting on their interest rates can be a measure of differences in exposure to global risk and thus that more return can be expected from trades on a bigger interest rate gap.

Furthermore it follows that ES does not explain the risk associated with excess returns in currency carry trades adequately. We cannot reject an increasing monotonic relation with the Wolak test but we do not find statistical evidence in favour of the relation with the MR test either. We can therefore not draw convincing conclusions to support the hypothesized theory.

Our findings on UIP are in line with the evidence described in Engel (1996) and the conclusions drawn by Hansen and Hodrick (1980) who reject the hypothesis that the expected return of speculating in the forward markets is zero. Fama (1984) furthermore concludes that both the forecasting ability of forward exchange rates and the forward premiums are time varying. More recently the work of Lustig and Verdelhan (2007) and Lustig et al. (2011) shows that UIP also fails in the cross-section when they construct portfolios in forward contracts. They show that investors can earn large excess returns by holding bonds in currencies that currently have a high interest rate compared to other countries. Our intuition of a market risk factor to better describe excess returns in currency carry trades is followed by Dobrynskaya (2014). However they take a different approach as they perform regressions with a dummy variable for market risk to explain excess returns and conclude excess returns can be explained more adequately this way.

The remainder of this paper is structured as follows, we discuss the theory in section 2 followed by the data in section 3 and the methodology of the tests for monotonicity in section 4. Section 5 describes the results and finally section 6 concludes.

2 Theory

2.1 Covered and uncovered interest rate parity

A currency carry trade means to short a currency and go long in another (foreign) currency, which can be done through the fixed income market or the currency forward market. A popular example is to short JPY and go long AUD as Japan has low interest rates while Australia has considerably higher interest rates.

We define currency carry trades through the spot and forward currency market as in Lustig et al. (2011). We let s denote the log of the spot exchange rate S and f denote the log of the forward exchange rate F, both in terms of units of foreign currency per domestic currency, i.e. we go long in the foreign currency and short the domestic currency. We can now define the log excess return of a one-month currency carry trade as

$$rx_{t+21} = f_t - s_{t+21}, \tag{1}$$

where t is in days as we use daily data. We define the log forward discount (or forward premium) as the difference between the forward exchange rate and the spot exchange rate at time t, i.e. $f_t - s_t$. We can now rewrite equation 1 in terms of the forward discount as the log forward discount minus the appreciation of the spot exchange rate, that is

$$rx_{t+21} = f_t - s_t - (s_{t+21} - s_t) = f_t - s_t - \Delta_{21} s_{t+21}, \tag{2}$$

where $\Delta_{21} s_{t+21} = s_{t+21} - s_t$.

Covered interest rate parity (CIP) states that investors should not earn positive returns by borrowing domestic interest rate bearing assets to lend equal foreign interest bearing assets while covering for the exchange rate exposure through a forward contract of the same maturity and *vice versa*. Equal assets are defined as asset which have the same maturity and risk factors, such as liquidity, default and political risks. CIP thus states that the forward and spot rate differential must equal the interest rate differential, that is

$$\frac{F}{S} = \frac{1 + i^*}{1 + i}, \tag{3}$$

where i^* and i denote the foreign and domestic interest rates on similar assets respectively. In terms of logarithms 3 becomes

$$f - s \approx i^* - i. \tag{4}$$

We can use CIP to approximate the log excess returns as

$$rx_{t+1} \approx i_t^* - i_t - \Delta s_{t+1}. \tag{5}$$

Uncovered interest rate parity (UIP) states that the expected forward discount should equal the expected rate of depreciation, i.e. $E[f_t - s_t] = E[\Delta s_{t+1}]$. The currency carry trades specified above should thus yield no excess return and we would therefore expect a flat pattern when we sort the trades on their forward discounts.

2.2 Value-at-Risk and expected shorfall

In this subsection we define the general concepts of Value-at-Risk (VaR) and Expected Shortfall (ES). First, let Y_t denote the return on some asset at time t. We assume Y_t to be a stationary ergodic process. Furthermore, let $W_{t-1} = (Y_{t-1}, Z'_{t-1}, Y_{t-2}, Z'_{t-2}, \ldots)$ be the information set at time $t-1$, where Z_{t-1} denotes a vector of other relevant explanatory variables. We define the α-level VaR given the information set at time $t-1$ as the quantity $\text{VaR}_t(\alpha)$ such that

$$P(Y_t \leq \text{VaR}_t(\alpha)|W_{t-1}) = \alpha \tag{6}$$

for some coverage level $\alpha \in [0,1]$. ES is defined as the conditional expected loss given that the loss is larger than $\text{VaR}_t(\alpha)$, that is,

$$\text{ES}_t(\alpha) = E\left[Y_t|W_{t-1}, Y_t < \text{VaR}_t(\alpha)\right]. \tag{7}$$

We estimate a GARCH(1,1) model[1] to obtain estimates for the unconditional volatility, $\bar{\sigma}^2$, of the return series, that is we estimate

$$Y_t = v_t, \quad v_t = \sigma_t \varepsilon_t, \quad \varepsilon_t \sim t_v, \tag{8}$$

$$\sigma_t^2 = \omega + \alpha v_{t-1}^2 + \beta \sigma_{t-1}^2, \tag{9}$$

and calculate $\bar{\sigma}^2 = \frac{\omega}{1-\alpha-\beta}$. We can now calculate the values for VaR and ES as

$$\text{VaR}_t(\alpha) = \mu + \bar{\sigma} F_v^{-1}(\alpha), \tag{10}$$

$$\text{ES}_t(\alpha) = \mu - \frac{\bar{\sigma}}{\alpha} f_v(\text{VaR}_t(\alpha)). \tag{11}$$

where F is the CDF of a student's t distribution and v is the associated degrees of freedom.[2].

In order to test for an increasing relation in excess carry trade returns with respect to ES we sort the excess returns on ES in *descending* order. As ES is (generally) negative and we expect that currencies with more ES in absolute value should earn bigger excess returns.

2.3 Liquidity preference hypothesis

Liquidity preference hypothesis (LPH) states that investors prefer short-term investments over long-term investments and hence the yield curve should be upward sloping. LPH thus defines an increasing monotonic relation in term premium with respect to maturity. We define the term premium as in Richardson et al. (1992) and Boudoukh et al. (1999): the difference between the expected return on a T-bill with yield to maturity $\tau(j)$ and the expected return on a T-bill with 1 month yield to maturity, i.e. $E[r_t^{(\tau_j)} - r_t^{(1)}]$ where $r_t^{(\tau_j)}$ is the one-period return on a T-bill with maturity τ_j. For the application of the MR tests we now define $\Delta_i \equiv E[r_t^{(\tau_i)} - r_t^{(1)}] - E[r_t^{(\tau_{i-1})} - r_t^{(1)}]$ and test whether $\Delta_i > 0$ for $i = 2, \ldots, N$.

[1] We make use of Kevin Sheppard's MFE Toolbox to estimate the GARCH model using the function Tarch

[2] We use a student's t distribution to account for leptokurtic return distributions

3 Data

3.1 Currency carry trades

In order to test for a monotonic relation in excess returns associated with currency carry trades we use data on forward and spot exchange rates. We limit our attention ti the G10 currencies[3] and construct the trades using USD, JPY and CHF as domestic currencies, respectively. We use forward rate data from Datastream for most currencies, unfortunately not all currency forward rates are available through this Datastream so we construct the remaining forward rates by making use of CIP and one-month deposit rates available through Bloomberg. However, the data does contain some gaps, mostly around public holidays such as Christmas day and New Year's day, which we filled in by the average of the other observations. We use daily data and restrict our attention to data ranging from 3 January 2000 to 12 May 2017 (4530 observations).

We test for an increasing monotonic relation in excess returns in terms of their future discount. We apply the tests to the full sample but also to 2 subsamples. The first subsample dates from 3 January 2000 to 31 December 2008 (2348 observations) and describes the period before the global financial crisis with, in general, higher interest rates. The second subsample ranges from 1 January 2009 to 12 May 2017 (2182 observations) and describes the period after the crisis where interest rates are lower and even sub-zero in some countries. This drop in interest rates becomes clear from figure 1 which plots the one month deposit rates for the US, Japan, Swiss, New Zealand and Australia. We refer the reader to the appendix for a table with average interest rate differentials and forward discounts in the sample period and a plot with 1 month deposit rates for all G10 countries.

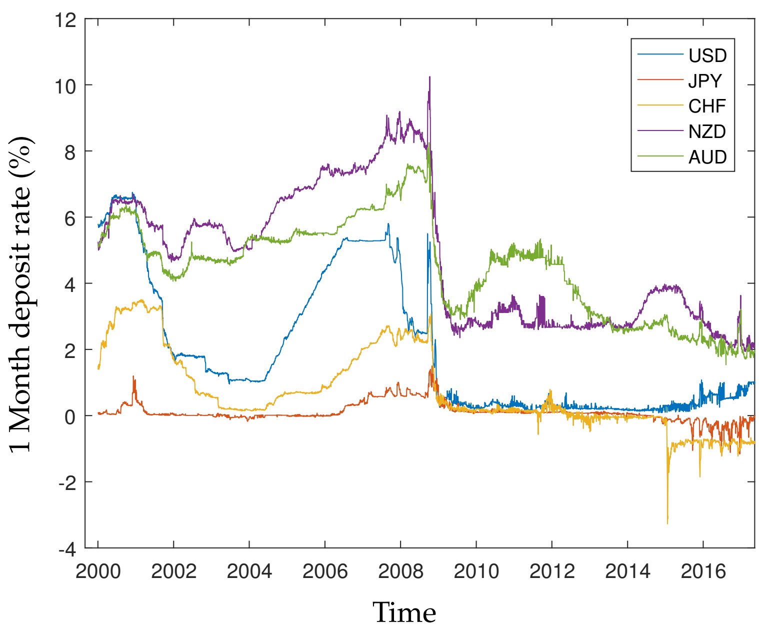


Figure 1: Daily one month deposit rates USD, JPY, CHF, NZD and AUD from January 2000 to May 2017

[3]Australia, Canada, the Euro zone, Japan, New Zealand, Norway, Sweden, Switzerland, UK and the US

Table 1 reports summary statistics on the log excess returns from currency carry trades short USD and long other G10 currencies. The returns distribution is leptokurtic, as the returns display skewness and excess kurtosis, indicating peakedness and fat tails. In the sample period after the global financial crisis the excess kurtosis seems to decline across the table as well as the skewness. The trades long CHF seem to be the exception to this observation as there is significant excess kurtosis and negative skew. The summary statistics for currency carry trades short JPY and CHF are given in the appendix.

Table 1: Summary statistics for the logarithms of the daily excess returns for currency carry trades short USD and long other G10 currencies

Sample Period	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
					2000-2017 (4030 obs)				
Mean	6.336	0.853	-2.449	0.108	-0.090	-4.616	3.032	1.038	0.156
Std	0.619	0.413	0.540	0.460	0.429	0.540	0.587	0.603	0.587
Skew	-0.589	-0.896	0.129	-0.082	-0.776	-0.201	-0.086	-0.376	-0.024
Kurtosis	5.561	10.267	4.165	4.035	6.085	2.734	3.925	4.537	3.505
Max	0.148	0.095	0.135	0.128	0.102	0.092	0.131	0.162	0.161
Min	-0.277	-0.230	-0.180	-0.151	-0.173	-0.123	-0.220	-0.222	-0.181
					2000-2008 (2348 obs)				
Mean	5.248	0.636	-3.812	0.613	0.034	-7.609	3.265	0.911	-0.284
Std	0.662	0.435	0.540	0.490	0.437	0.549	0.658	0.622	0.631
Skew	-0.938	-1.374	0.484	-0.125	-1.078	0.210	-0.128	-0.729	-0.148
Kurtosis	6.223	13.044	3.252	4.298	7.001	2.818	3.857	4.541	3.203
Max	0.125	0.095	0.135	0.128	0.076	0.092	0.119	0.118	0.111
Min	-0.277	-0.230	0.000	-0.151	-0.173	-0.123	-0.220	-0.222	-0.181
					2009-2017 (2182 obs)				
Mean	7.507	1.086	-0.983	-0.434	-0.223	-1.395	2.781	1.175	0.631
Std	0.566	0.372	0.000	0.432	0.412	0.447	0.505	0.568	0.538
Skew	0.136	-0.006	-0.247	-0.082	-0.394	-0.472	-0.045	0.133	0.254
Kurtosis	3.193	3.607	6.125	3.456	4.857	3.635	3.183	4.373	3.762
Max	0.148	0.093	0.135	0.090	0.102	0.082	0.131	0.162	0.161
Min	-0.090	-0.084	-0.180	-0.093	-0.122	0.000	-0.103	-0.144	-0.109

Mean and standard deviation are annualized using 252 trading days.

3.2 Treasury Securities

To test the implications made by LPH we use monthly data on expected returns for T-bills with yields to maturity ranging from 1 month to 12 months. This data can be obtained through the Center for Research in Security Prices. For reasons of comparison with Patton & Timmermann (2010) we use data ranging from January 1964 to December 2001 and drop the data on 12 month T-bill returns. We apply the tests to the full sample (1964-2001) and two subsamples: a sample that consists of data ranging form January 1964 to December 1972 and a sample that ranges form January 1973 to December 2001.

4 Methods

4.1 Standard t-test

We first develop some general notation which we use through the remainder of this paper. As in Patton and Timmermann (2010) we consider the ranking of $N+1$ securities based on their expected return and take these as given. Let $\{r_{it}\}_{t=1}^{T}$ be the time series of returns of the i-th security where $i = 0, \ldots, N$. Furthermore let $\boldsymbol{\mu} = (\mu_0, \mu_1, \ldots, \mu_N)'$ be the expected returns of these securities.

We can now easily define a simple one-sided t-test to test for an increasing monotonic relation by testing if μ_N is significantly different from μ_0, that is we test using the following statistic

$$t = \frac{(\hat{\mu}_N - \hat{\mu}_0)}{\hat{\sigma}} \sim \mathcal{N}(0,1). \tag{12}$$

We obtain $\hat{\mu}_i$ by looking at the sample analog, i.e. $\hat{\mu}_i \equiv (1/T)\sum_{t=1}^{T} \mathrm{r}_{it}$. The standard errors are calculated as the square root of the variance of the spread of the returns, where the sample variance is calculated as $\hat{\sigma}^2 \equiv (1/T)\sum_{t=1}^{T} \left[(\mathrm{r}_{Nt} - \mathrm{r}_{1t}) - (\hat{\mu}_N - \hat{\mu}_1)\right]^2$. The test statistic in equation 12 rejects the increasing monotonic relation if the p-values of the t-statistic is not significant. We can easily adapt the test to a test for a downward relation by considering $\mu_1 - \mu_N$. In practice we will use Newey-West standard errors to account for possible heteroskedasticity and autocorrelation in the data.

4.2 Monotonic relation test

In this subsection we set forth the framework developed by Patton and Timmermann (2010) and hence we continue to adapt their notation. In essence we are testing inequality constraints on parameter estimates. We again consider the ranking of $N+1$ securities based on their expected return but now focus on their corresponding expected return differentials. We define these return differentials as $\Delta_i = \mu_i - \mu_{i-1}$ and perform the test by looking at (element by element) inequality constraints on the parameter vector $\boldsymbol{\Delta} \equiv [\Delta_1, \ldots, \Delta_N]'$. We specify a flat or weakly declining pattern under the null hypothesis and the hypothesized strictly increasing monotonic pattern under the alternative, that is

$$\begin{aligned} H_0 &: \ \boldsymbol{\Delta} \leq \mathbf{0} \\ H_1 &: \ \boldsymbol{\Delta} > \mathbf{0}. \end{aligned} \tag{13}$$

We can rewrite the above hypotheses as

$$\begin{aligned} H_0 &: \ \boldsymbol{\Delta} \leq \mathbf{0} \\ H_1 &: \ \min_{i=1,\ldots,N} \Delta_i > 0 \end{aligned} \tag{14}$$

because if the smallest value of $\Delta_i > 0$ then naturally $\Delta_i > 0 \; \forall \; i = 1, \ldots, N$. We estimate Δ_i as $\hat{\Delta}_i = \hat{\mu}_i - \hat{\mu}_{i-1}$, where $\hat{\mu}_i$ is again obtained by taking the sample mean. This gives rise to the following test statistic:

$$J_T = \min_{i=1,\ldots,N} \hat{\Delta}_i \tag{15}$$

Patton and Timmermann (2010) show that under standard conditions the estimated parameter vector $\hat{\boldsymbol{\Delta}} = [\hat{\Delta}_1, \ldots, \hat{\Delta}_N]'$ follows a normal distribution as $T \to \infty$,

$$\sqrt{T}(\hat{\boldsymbol{\Delta}} - \boldsymbol{\Delta}) \overset{a}{\sim} \mathcal{N}(\mathbf{0}, \Omega). \tag{16}$$

However we can avoid estimating Ω directly by adapting a bootstrap methodology, hence Patton and Timmermann (2010) propose to use the stationary bootstrap of Politis and Romano (1994) to randomly draw a new sample of returns $\{\tilde{r}^{(b)}_{i\tau(t)}, \tau(1), \ldots, \tau(T); i = 0, 1, \ldots, N\}$ with replacement. Here $\tau(t)$ is the new time index which we obtain by randomly drawing from the original set $\{1, \ldots, T\}$. Moreover to preserve any cross-sectional dependencies in returns we let $\tau(t)$ be equal across portfolios. The bootstrap indicator, b, ranges form b=1 to $b = B$ where B is chosen large enough to ensure the results do not depend on simulation errors. To further account for time series dependencies we will draw returns in blocks where the block length, L, is drawn from a geometric distribution such that we can control the average length of each block.

We thus perform the following routine:

```
for b = 1, ..., B do
    Draw random time index, τ(1)
    for t = 2, ..., T do
        Draw new random value, temp
        if temp > 1/L then
            Take next value of returns
        else
            Take new random time index
        end
    end
    return bootstrap time indices
end
```

In order to make valid statistical inferences we need to obtain the bootstrap distribution of $\hat{\mathbf{\Delta}}$ under the null and to do so Patton and Timmermann (2010) follow the reasoning of White (2000) and choose the point in the composite null space which is least favourable to the alternative, in our case that means we let $\mathbf{\Delta} = \mathbf{0}$ under the null. They impose this by subtracting the estimated parameter vector, $\mathbf{\Delta}$, from the bootstrapped parameter vector, $\mathbf{\Delta}^{(b)}$, such that our bootstrap test statistic becomes:

$$J_T^{(b)} = \min_{i=1,\ldots,N} (\hat{\Delta}_i^{(b)} - \hat{\Delta}_i), \quad b = 1, \ldots, B. \tag{17}$$

We obtain a p-value for the test by counting the number of times a pattern that is at least as unfavourable as the null in the real data appears, thus a higher value for the test statistic, and divide this by the number of bootstraps, B, i.e.

$$\hat{p} = \frac{1}{B} \sum_{b=1}^{B} \mathbb{1}(J_T^{(b)} > J_T), \tag{18}$$

where $\mathbb{1}(\cdot)$ is an indicator function. We reject the null when $\hat{p} < 0.05$. This test thus makes no assumptions on the functional form of the monotonic relation and in combination with the described stationary bootstrap method we also do not have to make any assumptions on the distribution of the underlying. The latter is an important difference with the Wolak test where normality of the returns is assumed (subsection 4.4). In practice we use a studentized version of the bootstrap method to account for cross-sectional heteroskedasticity in the returns, meaning we divide our test statistic by the bootstrap standard deviation which we can calculate analytically by using lemma 1 in Politis and Romano (1994).

4.3 Extensions of the monotonic relation test

We consider several logical extensions to the MR test specified above. First of all the test described in subsection 4.2 only tests for monotonicity in adjacent pairs, naturally we can also perform this test for all possible return pairs. That is we consider $E[r_{i,t}] - E[r_{j,t}]$, $\forall\, i > j$. This grows the size of the vector $\boldsymbol{\Delta}$ to $\frac{N(N+1)}{2}$ instead of N.

Furthermore to account for possible low power of the MR test in small samples we will also calculate an Up and Down test statistic. These test statistics capture the direction of deviations from a flat pattern, but also the frequency and magnitude of these deviations. The Up test can be performed by testing the following null hypothesis:

$$\begin{aligned} H_0 &: \ \boldsymbol{\Delta} = \mathbf{0}, \text{ against} \\ H_1^+ &: \ \sum_{i=1}^{N} |\Delta_i| \mathbb{1}[\Delta_i > 0] > 0, \end{aligned} \tag{19}$$

where $\mathbb{1}(\cdot)$ is an indicator function. The null again defines no monotonic relation, but conversely to the MR test in 4.2 the null now describes a flat pattern as opposed to a weakly decreasing pattern. The test statistic now logically becomes:

$$J_T^+ = \sum_{i=1}^{N} |\hat{\Delta}_i| \mathbb{1}[\hat{\Delta}_i > 0]. \tag{20}$$

The Down test can be defined in corresponding fashion by flipping the sign of the indicator function, i.e.

$$\begin{aligned} H_0 &: \ \boldsymbol{\Delta} = \mathbf{0} \\ H_1^- &: \ \sum_{i=1}^{N} |\Delta_i| \mathbb{1}[\Delta_i < 0] > 0, \end{aligned} \tag{21}$$

with associated test statistic

$$J_T^- = \sum_{i=1}^{N} |\hat{\Delta}_i| \mathbb{1}[\hat{\Delta}_i < 0]. \tag{22}$$

The p-values for these tests are again obtained by adapting the bootstrap methodology described in subsection 4.2.

4.4 Wolak test

Wolak (1989) develops a framework to test inequality constraints in econometric models. A special case of the test can be used in the application of monotonicity tests, which is closely related to a likelihood ratio test. In contrast to the MR test the null now holds the (weak) monotonic relation and the alternative hypothesis specifies the nonexistence of a monotonic relation. That is:

$$\begin{aligned} H_0 &: \ \boldsymbol{\Delta} \geq \mathbf{0} \\ H_1 &: \ \boldsymbol{\Delta} \text{ unrestricted (or } \boldsymbol{\Delta} \in \mathbb{R}^N). \end{aligned} \tag{23}$$

Rejecting the null thus indicates that there is no evidence for an increasing monotonic relation. Low power due to noisy data or small samples for example could make it difficult to reject the null, however as we have the hypothesized theory under the null this consequently leads to less confidence in the statistical inferences made from this test. Moreover the null now also includes the case $\mathbf{\Delta} = \mathbf{0}$, a flat pattern, and thus accepting the null could also simply be due to the nonexistence of a monotonic relation between expected returns and the associated sorting variable. The MR test holds the hypothesized relation under the alternative and accounts for these problems, furthermore the Up and Down test statistics could help to asses whether low power leads to failure to reject the null.

We perform the Wolak test by solving the following quadratic problem (QP):

$$\begin{aligned} &\min_{\mathbf{\Delta}} (\hat{\mathbf{\Delta}} - \mathbf{\Delta})'\Omega^{-1}(\hat{\mathbf{\Delta}} - \mathbf{\Delta}) \\ &\text{s.t. } \mathbf{\Delta} \geq \mathbf{0}, \end{aligned} \tag{24}$$

where Ω is the covariance matrix of the data. Let $\tilde{\mathbf{\Delta}}$ be the solution of this QP then we define the test statistic IU as follows:

$$IU = (\hat{\mathbf{\Delta}} - \tilde{\mathbf{\Delta}})'\Omega^{-1}(\hat{\mathbf{\Delta}} - \tilde{\mathbf{\Delta}}). \tag{25}$$

In practice we again use Newey-West standard errors. This test statistic tests inequality constraints under the null versus unrestricted parameters under the alternative[4].

Wolak (1989) shows that the distribution of the test statistic follows a weighted sum of independent chi-squared variables:

$$Pr[IU \geq c] = \sum_{i=1}^{N} Pr[\chi^2(i) \geq c]\omega(N, N-i), \tag{26}$$

where $\omega(N, N-i)$ are weights and $\chi^2(i)$ is a chi-squared variable with i degrees of freedom. The weights are generally unknown but we can calculate them through Monte Carlo simulation[5]. That is we perform the following simulation:

for $ns = 1, \ldots, nS$ **do**
- Simulate $\mathbf{\Delta}$, say $\mathbf{\Delta}^s$, from a normal distribution with zero mean and variance $\hat{\Omega}$
- Use $\mathbf{\Delta}^s$ to solve the QP in equation 24 for $\tilde{\mathbf{\Delta}}^s$
- Calculate the number of elements in $\tilde{\mathbf{\Delta}}^s > 0$, say m
- Update the weights vector accordingly $\omega(N, m) = \omega(N, m) + \frac{1}{ns}$

end

Monte Carlo simulation can however be a limitation towards financial applications as it is computationally costly and harder to implement when the number of constraints increases. Furthermore we have to make assumptions about the underlying of the distribution, namely that the returns follow a normal distribution, which we do not in the MR test.

[4] Wolak (1989) also provides a test statistic to test equality constraints under the null against inequality constraints under the alternative in the form of $EI = \tilde{\mathbf{\Delta}}'\Omega^{-1}\tilde{\mathbf{\Delta}}$. This test statistic seems closer to the MR test but simulation studies have shown that the test behavior is unknown when the relation is not monotonic as the test was probably not intended to work when (weak) monotonicity is violated (Patton & Timmermann (2010))

[5] Wolak (1989) mentions that for $N \leq 4$ there exist closed form solutions (Kudo, 1963) and for $5 \leq N \leq 10$ some advanced numerical methods but that these methods become expensive and thus unattractive for $N \geq 8$

4.5 Bonferroni bounds test

Fama (1984) uses a Bonferroni bound in an attempt to summarize the information of a t-test for the expected return differentials $\hat{\Delta}_i \geq 0 \; \forall i = 1, \ldots, N$ in their application to term premia. The test is based on the following hypotheses:

$$\begin{aligned} &H_0: \; \hat{\Delta}_1 \geq 0, \ldots, \hat{\Delta}_N \geq 0 \; \text{against} \\ &H_1: \; \hat{\Delta}_j < 0 \; \text{for some } j \in \{1, \ldots, N\}. \end{aligned} \tag{27}$$

Let t_i, $i = 1, \ldots, N$, be the result of a t-test on Δ_i. We determine *Bonferroni p-values* (p_B) by checking whether the smallest t-statistic falls below a Bonferroni bound on Type 1 errors. We can calculate these p-values as follows: $p_B = N\,F(\min_i\{t_i\})$ where $F(\cdot)$ is the CDF of a normal distribution. We reject the null of an increasing pattern in the data if $p_B < \alpha$, where α is the significance level of the test. Note that these p-values can in fact be larger than 1 as we multiply by N.

This test can easily be adapted to test for a decreasing monotonic pattern by testing the following hypotheses:

$$\begin{aligned} &H_0: \; \hat{\Delta}_1 \leq 0, \ldots, \hat{\Delta}_N \leq 0 \; \text{against} \\ &H_1: \; \hat{\Delta}_j > 0 \; \text{for some } j \in \{1, \ldots, N\}. \end{aligned} \tag{28}$$

and calculate the Bonferroni p-values as $p_B = N\,F(\max_i\{-t_i\})$ where $F(\cdot)$ is again the CDF of a normal distribution.

5 Results

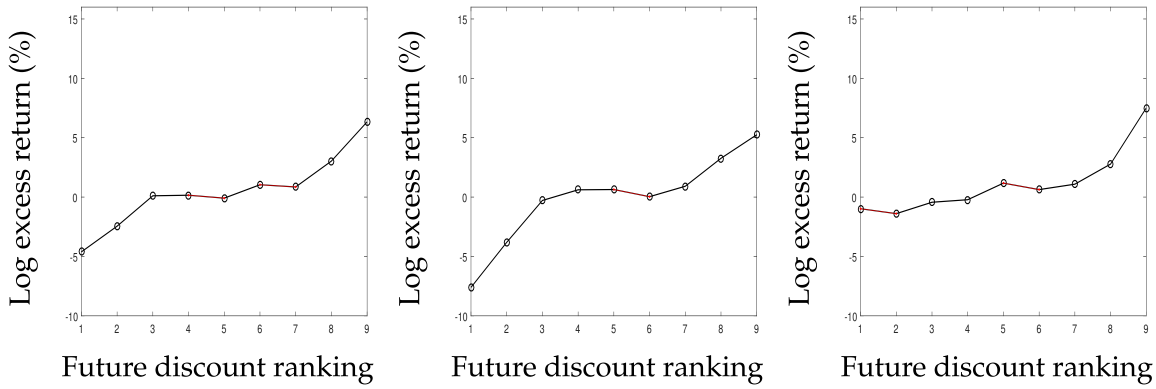
5.1 Currency Carry trades

5.1.1 Short US Dollar

Table 2: Annualized average daily log returns on 1-month currency carry trades short USD and long in other G10 currencies from January 2000 to may 2017

Sample Period	1	2	3	4	5	6	7	8	9
2000-2017	-4.536	-2.520	0.000	0.252	0.000	1.008	0.756	3.024	6.300
2000-2008	-7.560	-3.780	-0.252	0.504	0.756	0.000	1.008	3.276	5.292
2009-2017	-1.008	-1.512	-0.504	-0.252	1.260	0.756	1.008	2.772	7.560

Returns are in percentage points and are sorted in increasing order with respect to future discounts. Returns are annualized using 252 trading days per year. Other G10 currencies are AUD, CAD, CHF, EUR, GBP, JPY, NOK, NZD and SEK.



(a) Full sample 2000-2017 (b) Subsample 2000-2008 (c) Subsample 2009-2017

Figure 2: Annualized average daily log excess return on 1-month currency carry trades short USD and long in other G10 currencies from January 2000 to May 2017

UIP implies a flat pattern as forward discounts should capture the risk associated with currency carry trade. Table 2 reports annualized average excess returns in the different sample periods and it clearly follows that the excess returns are non-zero. Figure 2 shows plots of these excess returns when sorted on future discounts in the different (sub)samples and there seems to be an increasing pattern in log excess return with respect to forward discounts, especially in the full and pre-crisis sample. This increasing pattern seems to flatten out in the sample period after the global financial crisis as interest rate gaps shrink, but there is still significant positive excess return when carry trading the currency with the biggest future discount (AUD, see table 12 in the appendix). We observe some negative excess return as we also performed trades going long in currencies with lower one month deposit rates than in the US, e.g. JPY and CHF.

Table 3 reports the results of testing for an increasing monotonic relation in the data, we use $B = 10{,}000$ bootstrap simulations in addition to a block length $L = 200$[6] for the MR-tests and $nS = 1{,}000$ Monte Carlo simulations for the weights in the Wolak test. We found that increasing any of these numbers does not change the results a lot.

Table 3: p-values of tests applied to log excess returns from currency carry trades short USD sorted on forward discounts

Sample Period	top-bottom	t-stat	t-$p$val	MR	MR^{all}	Up	Down	Wolak	Bonferroni
2000-2017	0.043	23.523	0.000	0.007	0.004	0.000	0.995	0.969	1.000
2000-2008	0.051	20.324	0.000	0.054	0.060	0.000	0.996	0.586	0.550
2009-2017	0.034	17.904	0.000	0.077	0.027	0.000	0.960	0.845	0.926

Data ranges from January 2000 to May 2017. We use 10,000 bootstrap simulations for the MR-tests and 1,000 Monte Carlo simulations for the weights in the Wolak test.

[6] There is significantly more autocorrelation in this data compared to the bonds hence we choose a larger block length

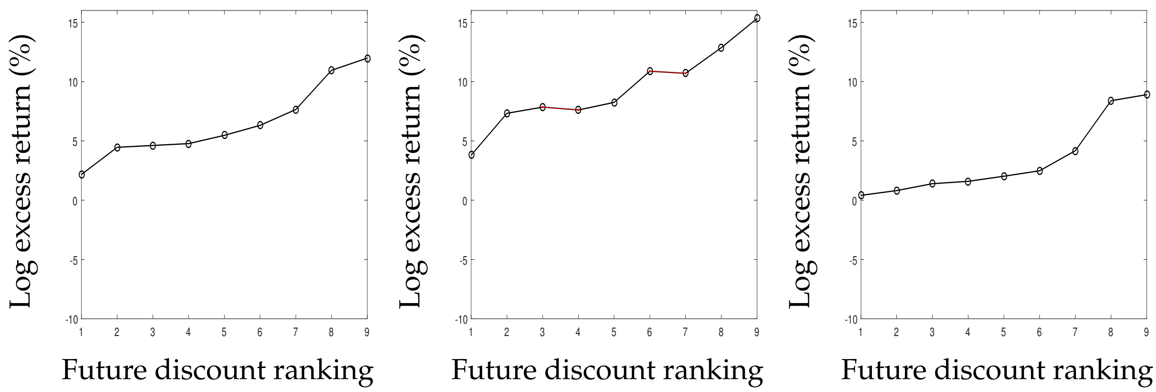
Both the Wolak and Bonferroni tests do not reject an increasing monotonic relation in log excess return with respect to future discounts for all sample periods. The MR test furthermore seems to find no statistical evidence in favour of a weakly declining monotonic pattern. This is strengthened by the conclusions of both the MR^{all} test, which also rejects a weakly declining pattern, and the Up test which rejects a flat pattern versus an increasing pattern. We thus find no significant evidence in favour of the flat pattern implied by UIP in the data.

5.1.2 Short Japanese Yen

Table 4: Annualized average daily log excess returns on 1-month currency carry trades short JPY and long in other G10 currencies from January 2000 to may 2017

Sample Period	1	2	3	4	5	6	7	8	9
2000-2017	2.268	4.536	4.536	4.788	5.544	6.300	7.560	10.836	12.096
2000-2008	3.780	7.308	7.812	7.560	8.316	10.836	10.584	12.852	15.372
2009-2017	0.504	0.756	1.512	1.512	2.016	2.520	4.284	8.316	8.820

Returns are in percentage points and are sorted in increasing order with respect to future discounts. Returns are annualized using 252 trading days per year. Other G10 currencies are AUD, CAD, CHF, EUR, GBP, NOK, NZD, SEK and USD



(a) Full sample 2000-2017 (b) Subsample 2000-2008 (c) Subsample 2009-2017

Figure 3: Annualized average daily log excess return on 1-month currency carry trades short JPY and long in other G10 currencies from January 2000 to May 2017

Table 4 reports the annualized average log excess returns on carry trades short JPY and long in other G10 currencies. Noticeably all returns are positive, indicating why this is a popular trading strategy. Figure 3 shows plots of these excess returns against their future discounts in the different sample periods and we again recognize an increasing pattern. In the period after the financial crisis the excess returns have come down as the interest rates decreased throughout the G10 currencies, however there still seems to be positive returns and an increasing pattern in the data.

Table 5 reports the results of the statistical tests for an increasing monotonic relation. We again use $B = 10,000$ bootstrap replication with a block length of $L = 200$ and use $nS = 1,000$ simulations for the weights in the Wolak test.

Table 5: p-values of tests applied to log excess returns of currency carry trades short JPY sorted on forward discounts

Sample Period	top-bottom	t-stat	t-$p$val	MR	MR^{all}	Up	Down	Wolak	Bonferroni
2000-2017	0.039	25.321	0.000	0.000	0.000	0.000	1.000	0.999	1.000
2000-2008	0.046	22.093	0.000	0.003	0.001	0.000	0.998	0.997	1.000
2009-2017	0.034	17.846	0.000	0.000	0.000	0.000	1.000	1.000	1.000

Short JPY and long other G10 currencies, data ranges from January 2000 to May 2017. We use 10,000 bootstrap simulations for the MR-tests and 1,000 Monte Carlo simulations for the weights in the Wolak test.

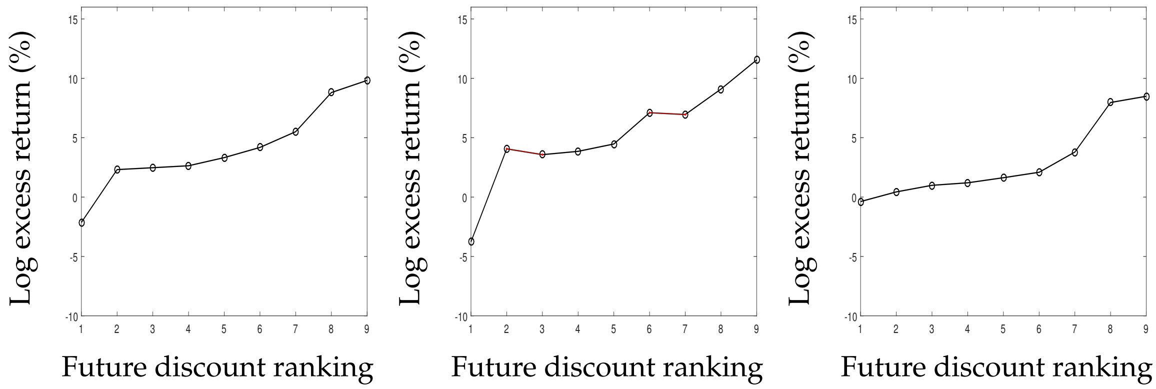
The Wolak and Bonferroni bounds tests clearly find no statistical evidence to reject the null of an increasing monotonic relation in excess returns with respect to forward premia as the p-values approach 1. Moreover it follows that the MR test finds sufficient statistical evidence to reject a flat or weakly declining pattern. This is again strengthened by the result of the MR^{all} and UP tests. We again find little evidence in favour of the flat pattern implied by UIP in the data.

5.1.3 Short Swiss Franc

Table 6: Annualized average daily log excess returns on 1-month currency carry trades short CHF and long in other G10 currencies from January 2000 to may 2017

Sample Period	1	2	3	4	5	6	7	8	9
2000-2017	-2.016	2.268	2.520	2.520	3.276	4.284	5.544	8.820	9.828
2000-2008	-3.780	4.032	3.528	3.780	4.536	7.056	7.056	9.072	11.592
2009-2017	-0.252	0.504	1.008	1.260	1.512	2.016	3.780	8.064	8.568

Returns are in percentage points and are sorted in increasing order with respect to future discounts. Returns are annualized using 252 trading days per year. Other G10 currencies are AUD, CAD, EUR, GBP, JPY, NOK, NZD, SEK and USD.



(a) Full sample 2000-2017 (b) Subsample 2000-2008 (c) Subsample 2009-2017

Figure 4: Annualized average daily log excess return on 1-month currency carry trades short CHF and long in other G10 currencies from January 2000 to May 2017

Table 6 reports the annualized average returns on carry trades short CHF and long other G10 currencies and these returns are plotted in figure 4. Similar to the results for carry trades short JPY in subsection 5.1.2 we see an increasing pattern in returns, however the absolute value of the excess returns now is lower. Furthermore the increasing pattern again seems to flatten out in the period after the crisis. Table 7 reports the results after testing for an increasing monotonic relation. We use $B = 10,000$ bootstrap replication with a block length of $L = 200$ and use $nS = 1,000$ simulations for the weights in the Wolak test.

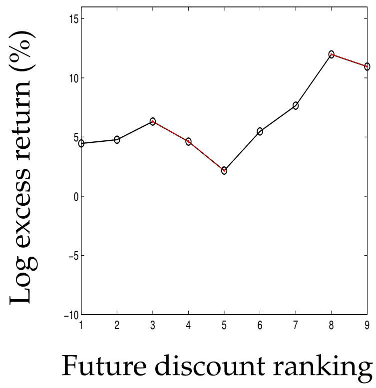
Table 7: p-values of tests applied to log excess returns currency carry trades short CHF sorted on future discounts

Sample Period	top-bottom	t-stat	t-$p$val	MR	MR$^{\text{all}}$	Up	Down	Wolak	Bonferroni
2000-2017	0.047	24.613	0.000	0.000	0.000	0.000	1.000	1.000	1.000
2000-2008	0.061	24.245	0.000	0.155	0.084	0.000	0.963	0.492	0.150
2009-2017	0.035	15.887	0.000	0.000	0.000	0.000	1.000	1.000	1.000

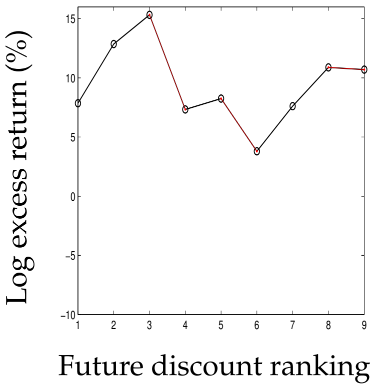
Short CHF and long other G10 currencies, data ranges from January 2000 to May 2017. We use 10,000 bootstrap simulations for the MR-tests and 1,000 simulations for the weights in the Wolak test.

The Wolak and Bonferroni bounds tests find no significant statistical evidence to reject an increasing monotonic relation in all samples. The MR test rejects a weakly declining relation in the full sample and the sample after the financial crisis, however it cannot strongly reject a weakly declining relation for the sample period preceding the financial crisis. Moreover, the MR^{all} test cannot convincingly reject a weakly declining relation as well. However, the UP test still rejects a flat pattern against an increasing relation indicating the possibility of low power of the MR tests in this sample period. There is still substantial evidence in favour of an increasing monotonic relation in excess return and UIP thus again seems to fail.

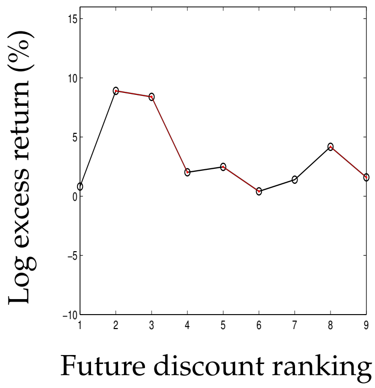
5.2 Expected Shortfall



(a) Full sample 2000-2017



(b) Subsample 2000-2008



(c) Subsample 2009-2017

Figure 5: Annualized average daily log excess return on 1-month currency carry trades short JPY and long in another G10 currency from January 2000 to May 2017 when sorted on expected shortfall

We hypothesize an increasing monotonic relation with respect to the absolute value of ES, however we can already infer, with inspection from figure 5, that this does not seem to be the case. For the two subsample periods it is hard to even recognize a clear pattern. ES does not seem to explain the risks associated with the excess returns of carry trades short JPY and long other G10 currencies well. We formalise this intuition by applying the tests for an increasing monotonic relation to the data. Table 8 reports the results of these tests, we use $B = 50{,}000$ bootstrap replication with a block length of $L = 200$ and use $nS = 2{,}500$ simulations for the weights in the Wolak test.

Table 8: p-values of tests applied to log excess returns from currency carry trades short JPY and long other G10 currencies sorted on ES

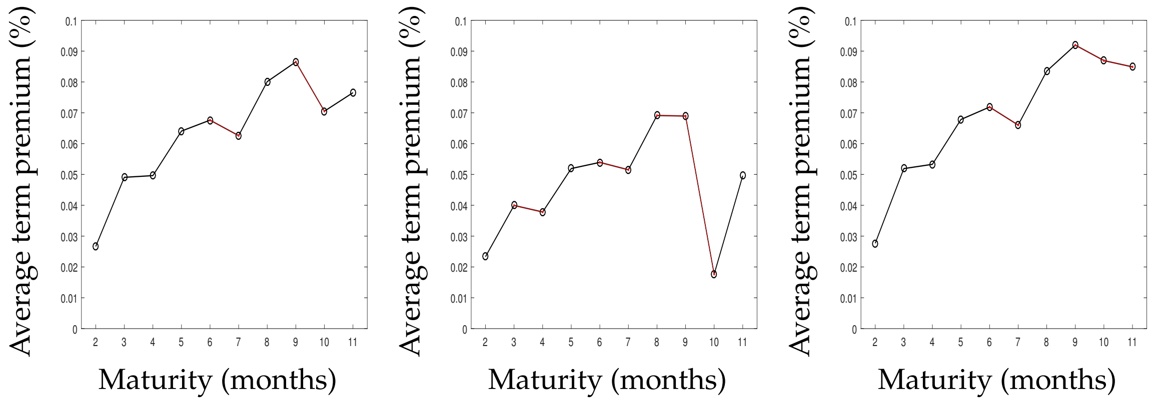
Sample Period	top-bottom	t-stat	t-$p$val	MR	MR^{all}	Up	Down	Wolak	Bonferroni
2000-2017	0.026	19.428	0.000	0.990	1.000	0.000	0.014	1.000	0.000
2000-2008	0.011	8.366	0.000	1.000	1.000	0.000	0.000	1.000	0.000
2009-2017	0.003	2.369	0.009	1.000	1.000	0.000	0.000	1.000	0.000

Short JPY and long other G10 currencies, data ranges from January 2000 to May 2017. We use 10,000 bootstrap simulations for the MR-tests and 1,000 simulations for the weights in the Wolak test.

The MR tests clearly find no statistical evidence in favour of an increasing monotonic relation. Noticeably both the Up and Down test reject a flat pattern but find evidence in favour of an increasing and decreasing relation, indicating that we might not draw valid conclusion from this test. The Wolak test however cannot reject an increasing relation, while the bonferroni test does.

The tests thus leave us with contradictory conclusions and hence we cannot convincingly reject the increasing relation nor do we find convincing evidence in favour of the relation. It does seem that ES is not an adequate measure to capture market risk associated with excess returns in currency carry trades. We obtain similar results when we look at carry trades short CHF or USD and we report those results in the appendix.

5.3 Liquidity preference hypothesis



(a) Full sample 1964-2001 (b) 1964-1972 (c) 1973-2001

Figure 6: Average monthly term premia for US T-bills with one month to maturity in the different (sub) samples

This section replicates the part in Patton and Timmerman (2010)[7] and serves as an illustration of correct implementation of the tests described in section 4. We test for an increasing monotonic relation in term premia with respect to yield to maturity as is implied by the liquidity preference hypothesis. Figure 6 shows plots of the average monthly term premia on US treasury bills, for all subsamples. It seems that nine-month treasury bills get a relatively high average term premium, which is in line with Fama (1984) who draws similar conclusions for the sample 1964 to 1982. Furthermore it is evident that a test for an increasing monotonic relation is expected to reject the increasing relation in the subsample 1964 to 1972, as the drop after the nine-month treasury bill is significant, with the ten-month premium even dropping below the average term premium for bonds with 1 month to maturity. McCulloch (1987) argues this unconventional behaviour is caused by the behaviour of the bid-ask spread for nine-month bills as it was half the size of the spread for eight- or ten-month T-bills in the subsample 1964 to 1972.

Table 9 shows the results for the different tests, we use $B = 1{,}000$ bootstrap simulations in addition to a block length of $L = 10$ for the MR-tests and $nS = 1{,}000$ Monte Carlo simulations for the weights in the Wolak test as we found that increasing those numbers did not change the outcome significantly.

[7] 4.3 Testing monotonicity of the term premium

Table 9: p-values of tests applied tot the term premium of treasury securities

Sample Period	top-bottom	t-stat	t-$p$val	MR	MR^{all}	Up	Down	Wolak	Bonferroni
1964-2001	0.050	2.416	0.008	0.953	0.906	0.000	0.369	0.035	0.020
1964-1972	0.026	0.908	0.182	0.991	0.991	0.003	0.375	0.009	0.004
1973-2001	0.057	2.246	0.012	0.617	0.617	0.002	0.474	0.328	0.704

Data ranges from January 1964 to December 2001. Results are obtained with 1,000 bootstrap simulations for the MR-tests and 1,000 Monte Carlo simulations for the weights used in the Wolak test.

The MR test cannot reject the null of an weakly declining pattern in all subsamples and thus finds little statistical evidence to support an increasing monotonic relation as implied by the LPH. Furthermore both the Wolak and Bonferroni bounds test reject an increasing relation in the full sample and, unsurprisingly, the subsample from 1964 to 1972. Remarkably monotonicity is not rejected by the Wolak and Bonferroni tests in the last subsample, while the MR test finds no evidence in support of the mononotonic relation. The Up-test however does not reject an increasing pattern in this subsamples which might indicate low power of the MR test.

The increasing monotonic relation implied by LPH thus seems questionable at best. We find little evidence to support it in the sample January 1964 to December 2001 and its subsample January 1964 to December 1972. However the theory may indeed hold in the subsample January 1973 to December 2001.

6 Conclusion

Firstly we found significant statistical evidence in favour of an increasing monotonic relation in excess returns with respect to forward discounts indicating that UIP fails in our data. A higher interest rate differential thus seems to earn higher excess return and these unconditional deviations from UIP seem to indicate its failure. Secondly it seems that ES does not capture the risk associated with excess returns of currency carry trades as we found no indication that larger ES (in absolute value) can explain larger excess return in currency carry trades.

To test UIP We have looked at carry trades of single currencies only, further extensions could be made by looking at baskets of currencies constructed through forward contracts as in Lustig and Verdelhan (2007). We have furthermore looked at ES as a risk factor for the excess returns in currency carry trades, but there are different risk factors of which the monotonicity implications may be explored in further research. We could explore the risk factor proposed by Dobrynskaya (2014) who states that high return currency carry rates can be explained through a downside market risk factor different from ES and thus that excess returns are justifiable. Furthermore a model as in Bansal and Dahlquist (2000) could be explored. They conclude that the interest rate differential as a risk factor is limited to developed countries and that differences can be explained through GDP per capita and average inflation rates with associated volatility.

Appendix

Summary statistics

Table 10: Summary statistics for the logarithms of the daily excess returns for currency carry trades short JPY and Long other G10 currencies

Sample Period	AUD	CAD	CHF	EUR	GBP	NOK	NZD	SEK	USD
					2000-2017				
Mean	10.952	5.469	2.166	4.461	6.312	7.648	11.982	4.772	4.616
Std	0.687	0.615	0.000	0.591	0.646	0.682	0.727	0.639	0.538
Skew	-1.170	-0.915	-0.363	-0.340	-0.716	-0.452	-0.509	-0.529	0.202
Kurtosis	9.536	7.694	5.264	4.976	4.802	5.367	4.942	5.499	2.733
Max	0.174	0.144	0.148	0.171	0.146	0.177	0.213	0.187	0.122
Min	-0.391	-0.333	-0.178	0.000	-0.255	-0.282	-0.279	-0.288	-0.099
					2000-2008				
Mean	12.858	8.247	7.851	3.797	10.698	10.878	15.320	7.328	7.611
Std	0.703	0.617	0.549	0.000	0.575	0.687	0.722	0.612	0.548
Skew	-2.074	-1.830	-0.833	-0.212	-1.652	-1.051	-1.117	-1.290	-0.210
Kurtosis	15.295	13.737	9.084	6.039	10.064	8.437	7.306	9.391	2.829
Max	0.174	0.144	0.171	0.148	0.146	0.177	0.212	0.135	0.122
Min	-0.391	-0.333	-0.251	-0.175	-0.255	-0.282	-0.279	0.000	-0.099
					2009-2017				
Mean	8.901	2.481	0.411	0.812	1.592	4.172	8.391	2.022	1.393
Std	0.645	0.553	0.000	0.547	0.585	0.604	0.663	0.622	0.446
Skew	-0.158	-0.117	-0.398	0.016	-0.252	0.003	-0.029	0.161	0.478
Kurtosis	3.781	3.229	4.735	3.500	4.411	3.395	4.269	3.900	3.620
Max	0.174	0.121	0.129	0.121	0.128	0.147	0.213	0.187	0.107
Min	-0.126	-0.106	-0.178	0.000	-0.170	-0.136	-0.165	-0.144	-0.087

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Table 11: Summary statistics for the logarithms of the daily excess returns for currency carry trades short CHF and Long other G10 currencies

Sample Period	AUD	CAD	EUR	GBP	JPY	NOK	NZD	SEK	USD
					2000-2017				
Mean	8.804	3.324	2.315	4.182	-2.133	5.508	9.826	2.638	2.474
Std	0.554	0.514	0.331	0.473	0.000	0.452	0.568	0.427	0.523
Skew	-0.755	-0.814	-2.195	-0.857	0.331	-0.306	-0.457	-0.737	-0.168
Kurtosis	6.748	6.780	24.771	6.527	5.163	7.936	4.890	10.893	4.317
Max	0.258	0.187	0.172	0.162	0.170	0.223	0.223	0.221	0.181
Min	-0.193	-0.184	0.000	-0.183	-0.139	-0.167	-0.124	-0.204	-0.152
					2000-2008				
Mean	9.097	4.471	4.061	6.952	-3.766	7.104	11.555	3.573	3.852
Std	0.570	0.516	0.207	0.394	0.497	0.422	0.568	0.355	0.528
Skew	-1.116	-0.694	-0.336	-1.524	0.208	-0.459	-0.743	-0.587	-0.479
Kurtosis	6.517	4.172	6.979	9.377	5.899	5.976	4.314	4.959	3.324
Max	0.156	0.103	0.092	0.090	0.170	0.162	0.147	0.081	0.105
Min	-0.193	-0.145	0.000	-0.153	-0.139	-0.142	-0.122	-0.097	-0.122
					2009-2017				
Mean	8.489	2.090	0.436	1.202	-0.377	3.792	7.965	1.631	0.991
Std	0.536	0.501	0.394	0.479	0.552	0.459	0.544	0.486	0.502
Skew	-0.303	-1.040	-2.007	-0.414	0.346	-0.115	-0.227	-0.613	0.145
Kurtosis	7.181	10.109	22.176	7.412	4.701	10.602	6.509	11.679	6.320
Max	0.258	0.187	0.172	0.162	0.170	0.223	0.223	0.221	0.181
Min	-0.129	-0.184	0.000	-0.183	-0.133	-0.167	-0.124	-0.204	-0.152

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Table 12: Average interest rate differentials and forward discounts from a USD, JPY an CHF investor

Sample Period	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
					USD Interest rate differential					
2000-2017	2.523	0.304	-1.228	-0.116	0.806	-1.845	1.259	2.868	0.068	-
2000-2008	2.132	0.094	-1.880	-0.165	1.378	-3.230	1.218	3.131	-0.175	-
2009-2017	2.944	0.530	-0.527	-0.063	0.191	-0.355	1.304	2.585	0.329	-
					USD Forward discount					
2000-2017	2.455	0.306	-1.199	-0.003	0.076	-1.807	1.226	0.264	0.074	-
2000-2008	2.053	0.101	-1.824	-0.001	0.125	-3.157	1.168	0.294	-0.158	-
2009-2017	3.034	0.579	-0.432	0.006	0.030	-0.293	1.390	0.244	0.473	-
					JPY Interest rate differential					
2000-2017	4.368	2.149	0.617	1.729	2.651	-	3.104	4.713	1.913	1.845
2000-2008	5.362	3.324	1.350	3.065	4.608	-	4.448	6.361	3.055	3.230
2009-2017	3.299	0.884	-0.172	0.292	0.546	-	1.658	2.940	0.684	0.355
					JPY Forward discount					
2000-2017	4.262	2.113	0.607	1.699	2.591	-	3.033	4.583	1.881	1.807
2000-2008	5.210	3.258	1.333	3.009	4.493	-	4.325	6.149	2.999	3.157
2009-2017	3.327	0.871	-0.139	0.327	0.538	-	1.683	2.932	0.766	0.293
					CHF Interest rate differential					
2000-2017	3.752	1.532	-	1.112	2.035	-0.617	2.488	4.097	1.296	1.228
2000-2008	4.012	1.974	-	1.715	3.258	-1.350	3.098	5.011	1.705	1.880
2009-2017	3.471	1.057	-	0.464	0.718	0.172	1.831	3.113	0.856	0.527
					CHF Forward dicount					
2000-2017	3.661	1.514	-	1.100	1.998	-0.594	2.437	3.979	1.286	1.209
2000-2008	3.891	1.933	-	1.678	3.180	-1.321	3.002	4.828	1.683	1.839
2009-2017	3.466	1.018	-	0.481	0.686	0.150	1.835	3.063	0.914	0.437

Differentials and discounts are reported in percentage points.

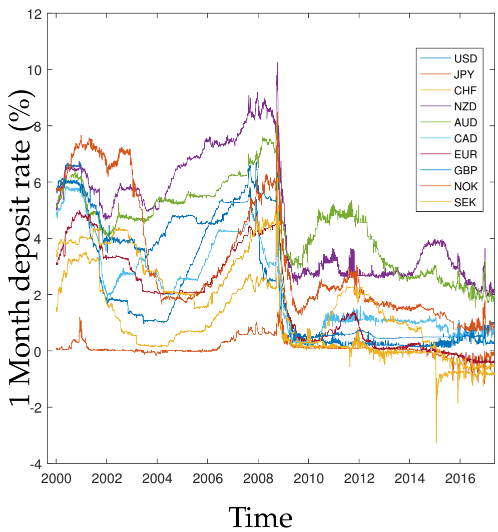
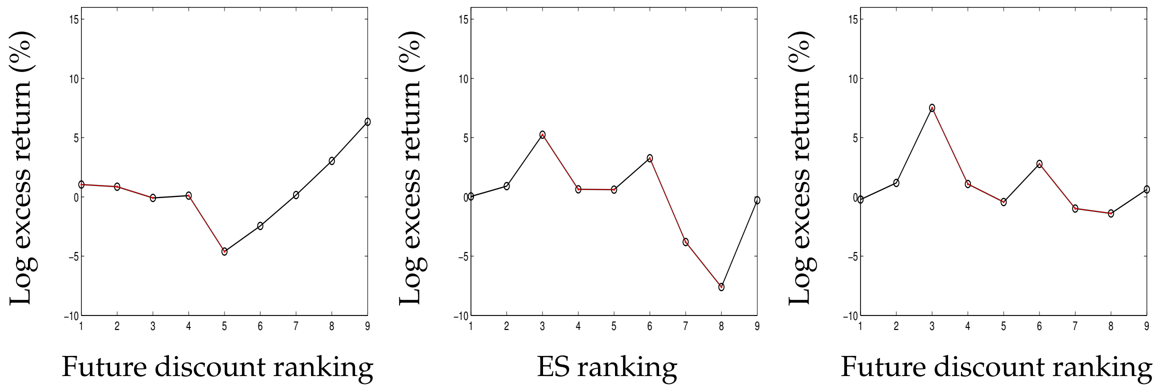


Figure 7: Daily one month deposit rates USD, JPY, CHF, NZD and AUD from January 2000 to May 2017

Results ES



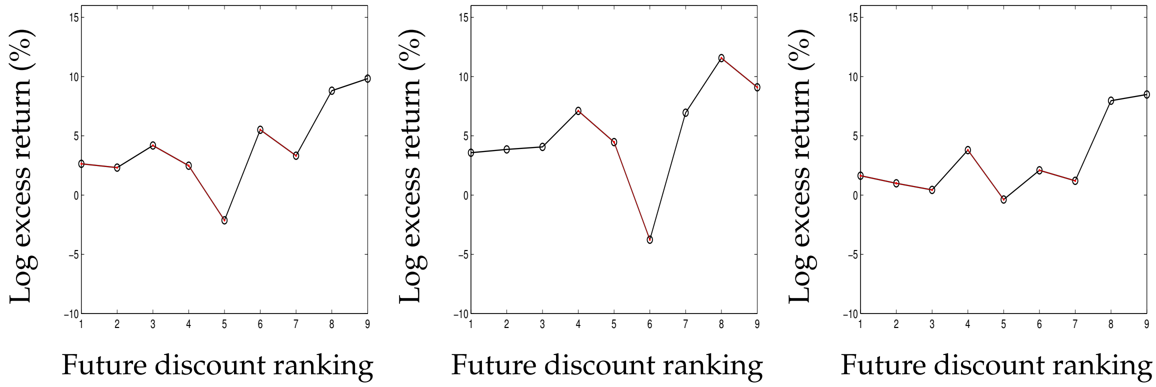
(a) Full sample 2000-2017 (b) Subsample 2000-2008 (c) Subsample 2009-2017

Figure 8: Annualized average daily log excess return on 1-month currency carry trades short USD and long in another G10 currency from January 2000 to May 2017 when sorted on expected shortfall

Table 13: *p*-values of tests applied to log excess returns from currency carry trades short USD and long other G10 currencies sorted on ES

Sample Period	top-bottom	*t*-stat	*t*-*p*val	MR	MR^{all}	Up	Down	Wolak	Bonferroni
2000-2017	0.021	18.389	0.000	1.000	1.000	0.000	0.000	1.000	0.000
2000-2008	-0.001	-0.708	0.761	1.000	1.000	0.000	0.000	1.000	0.000
2009-2017	0.003	2.025	0.021	1.000	1.000	0.000	0.000	0.397	0.000

Short USD and long other G10 currencies, data ranges from January 2000 to May 2017. We use 10,000 bootstrap simulations for the MR-tests and 1,000 simulations for the weights in the Wolak test.



(a) Full sample 2000-2017 (b) Subsample 2000-2008 (c) Subsample 2009-2017

Figure 9: Annualized average daily log excess return on 1-month currency carry trades short CHF and long in another G10 currency from January 2000 to May 2017 when sorted on expected shortfall

Table 14: p-values of tests applied to log excess returns from currency carry trades short USD and long other G10 currencies sorted on ES

Sample Period	top-bottom	t-stat	t-$p$val	MR	MR^{all}	Up	Down	Wolak	Bonferroni
2000-2017	0.029	21.620	0.000	1.000	1.000	0.001	0.000	0.765	0.000
2000-2008	0.022	13.586	0.000	1.000	1.000	0.000	0.000	0.764	0.000
2009-2017	0.027	17.719	0.000	1.000	1.000	0.009	0.000	0.000	0.000

Short CHF and long other G10 currencies, data ranges from January 2000 to May 2017. We use 10,000 bootstrap simulations for the MR-tests and 1,000 simulations for the weights in the Wolak test.

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