Carry trade attractiveness:
A time-varying currency risk premium approach

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אטרקטיביות של מסחר בתשואה:
גישת פרמיית סיכון מטבעי משתנה בזמן

תקציר

עבודה זו בוחנת את האטרקטיביות של אסטרטגיות מסחר בתשואה (Carry Trade) בשוק המט״ח הישראלי בתקופה שבין ינואר 2003 לדצמבר 2014. אנו בוחנים מספר מדדים פופולריים של תשואה לסיכון (Carry To Risk) ומציעים מדד חדש המבוסס על פרמיית סיכון מטבעי משתנה בזמן. בנוסף, אנו מעריכים את יכולת הניבוי של מדדי תשואה לסיכון ביחס לאחזקות של אג״ח ממשלתיות בידי משקיעים זרים תוך שימוש במאגר נתונים קנייני. כדי לאמוד את פרמיית סיכון המטבע, אנו בוחנים גם את שקלויות פער הריביות המכוסה (CIP) ופער הריביות הלא מכוסה (UIP). אנו מוצאים כי ה-CIP לא התקיים במחצית השנייה של תקופת המדגם, ואילו ה-UIP לא התקיים לאורך כלל תקופת המדגם באופק של חודש אחד, אך התקיים באופק של שנים-עשר חודשים. אנו מוצאים, באמצעות מתודולוגיות שונות, כי מדד התשואה לסיכון שהצענו ניבא פוזיציות של זרים באג״ח ממשלתיות בתקופת תת-המדגם השני. ניבוי זה היה איתן (Robust) יותר ולמשך זמן ארוך יותר, יחסית למדדים פופולריים אחרים של תשואה לסיכון.

Carry trade attractiveness:
A time-varying currency risk premium approach

Abstract

This study investigates the attractiveness of carry trade strategies in the Israeli FX market during the period 1/2003 - 12/2014. We examine several Carry To Risk (CTR) popular measures and propose a new measure which is based on a time-varying currency risk premium. The predictive capability of the examined CTRs is examined on foreigners' holdings in Israeli government bonds using a proprietary data set. In order to estimate the currency risk premium we also assess the Covered Interest rate Parity (CIP) and the Uncovered Interest rate Parity (UIP) hypotheses. We find that the CIP hypothesis did not hold during the second half of the sample period while the UIP hypothesis did not hold along the entire sample period for a one-month horizon, but prevailed for the twelve-month horizon. We find, using various methodologies, that our proposed CTR measure predicted foreigners' positions in government bonds during the second sub-sample. Such a prediction was more robust and of longer duration than other popular CTR measures.

1 Introduction

Carry trading (hereinafter CT) is one of the most profitable strategies in FX, though it carries a substantial risk (see Doskov and Swinkels (2015) and Burnside et al. (2007)). The most basic strategy involves borrowing in a low-interest rate currency and simultaneously depositing that amount in a high-interest currency. The former currency is referred to as the 'funding currency' while the latter is referred to as the 'target currency'[1]. The gain from the positive interest rate differential (hereinafter IRD) is exposed to a depreciation in the target currency such that the loss due to a depreciation may slash the profit from the positive IRD or even may cause a loss on the CT strategy. Thus, the profit and loss (hereinafter PNL) of a carry trade strategy depends on both current IRD and future exchange rate changes. A carry trade will be profitable for a given period as long as the target currency does not depreciate against the funding currency more than the IRD between them. Moreover, the PNL from the IRD is often gradual and stable over time while exchange rate movements of the currencies that are involved in the carry trade strategy are sharp and unexpected. This asymmetrical PNL is described as "picking up nickels in front of steamrollers: you have a long run of small gains but eventually get squashed" (Economist (2007)).

Carry trade may be of interest to not only sophisticated investors but also to policy makers of small open economies, since free capital flows and a floating exchange rate regime enable large international investors to significantly influence both exchange rates and local interest rates. In addition, carry traders usually behave as momentum investors in the sense that they exploit deviations from the UIP and act in the same direction as the deviation (positive feedback). As a result, they cause deviations from the UIP to grow larger and last longer (Baillie and Chang (2011)). Such influential activity may disturb both monetary policy and financial stability.

From a theoretical point of view, carry trading challenges the traditional theory regarding the relationship between IRD and future exchange rate changes. The relationships, called Uncovered Interest rate Parity (hereainfter UIP), state that in equilibrium the expected future exchange rate changes will fully offset the current IRD. Otherwise, many investors will implement carry trade strategies, which in turn will strengthen the target currency vis-a-vis the funding currency again and again in a 'vicious circle'.

Despite its attractiveness on a theoretical level, the evidence in general does not support the UIP hypothesis: On average the PNL from the (positive) IRD is not slashed by a depreciation of the target currency in the next period, making carry trade activity profitable, on average, though it has caused heavy losses in the past during short periods (Doskov and Swinkels (2015))[2]. This phenomenon is called the 'forward premium puzzle'.

However, Fama (1984), Chinn and Meredith (2005) and Frankel and Poonawala (2010), like many others, show that the PNL is period, country, and time span dependent. It was found that the UIP does not prevail for developed countries. For instance, taking a loan in Japanese

[1] An alternative CT strategy is buying a forward contract on the target currency (see a description of various CT strategies in Santealla et al. (2015)). This strategy, which yields the same profit as the basic one, is preferable if the forward market is more liquid than the bond markets.

[2] That profitability encouraged large financial institutions to construct ETFs and other indices that mimic carry trade strategies (see a list of indices in Gyntelberg and Remolona (2007)).

yen and depositing the proceeds in US Dollars was a profitable strategy. It was also found that the UIP does not hold for specific sub-periods and for shorter terms.
There are two versions of the UIP. In the first version the future realized exchange rate changes should equal the forward premium (hereinafter FP) which is defined as the log of the forward rate minus the log of the spot rate. In the second version the future realized exchange rate changes should equal the IRD. The two versions of the UIP are based on the underlying assumption that there are no 'arbitrage opportunities' in the local (efficient) market. This underlying assumption is called Covered Interest rate Parity (CIP) and provided that it prevails, it enables us to use both versions alternatively[3].
Many studies have tried to solve the puzzle by adding a time-varying risk premium to the tested UIP hypothesis (see: Engel (1996); Chinn and Meredith (2005); Sarantis (2006); Campbell et al. (2009)). Generally, risk premium is explained by 'risk averse' investors who demand a premium on their risky investment[4].
Our contribution to the literature is threefold. First, we examine several Carry To Risk (hereinafter CTR) measures for the attractiveness of carry trade strategies in the Israeli FX market. Second, we estimate the shekel/dollar (ILS/USD) risk premium implementing Kalman filter estimation and Gibbs sampling. Third, we introduce a measure for carry trade attractiveness (CTR3) based on the estimated shekel/dollar risk premium. By doing so we combine two strands of the literature: the risk-reward of carry trade strategies and the estimation of a currency risk premium. We also examine historically the capability of several attractiveness measures to predict[5] foreign investments in Israeli government bonds and *makam* (similar to US T-bills) using a proprietary data set of foreigners' positions (which will be described later).
The rest of the study is structured as follows. Section 2 presents the definitions of the CIP, UIP, and the carry trade measures we use in this study including currency risk premium, and surveys the literature. Section 3 describes the data sample, Section 4 depicts the estimated results, Section 5 discusses the results and conducts robustness checks for the proposed carry trade attractiveness measure (CTR3) and for the BOI's steps. Section 6 concludes.

2 Term Definitions and Survey of Literature

In this section we provide a short survey of the terms that are used in this study and the empirical evidence. We start by defining and describing the terms: Forward Premium (FP), Covered Interest rate Parity (CIP), Uncovered Interest rate Parity (UIP), excess return

[3] In some periods (even months) the CIP does not hold due to regulatory measures, tax differentiations, liquidity constraints etc. Thus, arbitrage opportunities are not immediately fulfilled (see Fong et al. (2010)).

[4] In this study, we use the commonly-used term 'risk premium', although there is neither a unique definition nor a practical way to assess it. In addition, there is no simple understanding regarding the extent or type of investors' risk aversion. For instance, if the risk premium is a compensation against return volatility then only 'risk averse' investors may demand it while both 'risk averse' and 'risk neutral' investors may demand a compensation against credit or liquidity risks.

[5] We use the term 'predict' or 'predictive capability' throughout the paper in order to improve the paper's readability. However, a comprehensive forecasting model is beyond the scope of the paper, mainly, due to the relatively short sample period.

(hereinafter ER), Currency Risk Premium (hereinafter P), and Carry To Risk (CTR). Then we survey the empirical evidence.

2.1 Term Definitions

2.1.1 Forward Premium (FP)

Currency forward contracts are agreements to buy or sell an amount of one currency for a certain price in terms of another currency at a certain future time. Currency forwards are usually traded in the Over the Counter (OTC) market between two financial institutions or between a financial institution and its customers. The definition of the forward premium at time t is: $FP_t = f_t - s_t$ where f_t is the log of the forward price (F_t) at time t and s_t is the log of the spot rate (S_t) at time t.

2.1.2 Covered Interest rate Parity (CIP)

Arbitrage constraints imply that the forward premium should approximately equal the IRD between any two currencies i.e. $FP_t = f_t - s_t \approx i_t - i_t^* = IRD_t$ where i and i^* are the domestic and foreign risk-free interest rates, respectively[6]. The latter are usually derived by a financial institution for a particular customer based on his lending conditions and are often higher than the risk-free short-term interest rates set by the central bank. This parity apparently holds in the FX market at least for longer periods. Deviations from this equality can indicate a lack of liquidity, credit risk, or other frictions in the markets (see Baba and Packer (2009) and Fong et al. (2010)).

2.1.3 Uncovered Interest rate Parity (UIP)

Following the above explanation, this parity states that the forward rate is an unbiased predictor of the future realized spot rates, and if the CIP holds, the IRD is also an unbiased predictor of future spot rate changes. The difference between UIP and CIP is that the former is a theoretical hypothesis which is based on several underlying assumptions (efficient markets, risk neutrality, and free flows) while the latter is the result of the 'no arbitrage opportunities' argument only. Later we examine whether the forward premium is indeed an unbiased conditional predictor of future exchange rate changes in Israel[7]. We test this hypothesis by regressing the forward premium on the future changes of the spot rate, as follows:

$$\Delta s_{t+1} = s_{t+1} - s_t = \alpha + \beta(f_t - s_t) + u_{t+1} \tag{1}$$

where $f_t - s_t$ is the FP at time t and is approximately equal to the IRD_t, and u_{t+1} is the error term. The UIP common test is for $\alpha = 0, \beta = 1$ and $E(u_{t+1}) = 0$ however, the findings of many studies exploring this issue are that $\hat{\beta} < 0$, which indicates that forward rates are not unbiased predictors of future spot rates (see Froot and Thaler (1990)). As mentioned

[6] The exact term is: $FP_t = \frac{F_t}{S_t} - 1 = \frac{1+i_t}{1+i_t^*} - 1 = \frac{i_t - i_t^*}{1+i_t^*}$.

[7] If the CIP hypothesis does not hold we will also examine the alternative using IRD rather than FP as a regressor.

before, these findings are referred to as the 'forward premium puzzle', and we elaborate on the meaning of $\hat{\beta}$ when there is a risk premium in Appendix A.

2.1.4 Excess Return (ER)

The evidence that there are deviations from the UIP, i.e., the 'forward premium puzzle', led researchers to calculate the extent of that deviation from the theoretical UIP equilibrium. ER can be used as a measure of profitability in the CT literature, and is defined as follows:

$$ER_{t+1} \equiv IRD_t - \Delta s_{t+1} = f_t - s_{t+1} \tag{2}$$

The deviation from the UIP i.e., $f_t - s_{t+1} \neq 0$ can be exploited by carry traders in both directions.

2.1.5 Currency Risk Premium (P)

As ER is found to be, on average, non-zero in many studies (see for example, Curcuru et al. (2010)), a time-varying currency risk premium (P) is perceived as a possible explanation to the forward premium puzzle (Engel (1996)). By adding and subtracting $E_t(s_{t+1})$ to ER_{t+1} in (2) where $E(.)$ is the expectation function, P is defined as the expected part of ER, as follows:

$$ER_{t+1} = (f_t - E_t(s_{t+1})) + (E_t(s_{t+1}) - s_{t+1}) \tag{3}$$

The excess return in the equation is decomposed into two parts: an expected component and an unexpected component. The expected component is called risk premium ($P = f_t - E_t(s_{t+1})$) and it can be perceived as the compensation an investor demands for the risk he incurs by holding a specific currency. That risk is the exposure to unanticipated changes in the exchange rates at which a currency trades against another currency. Under the UIP hypothesis, P should be zero. However, if for example P_t is temporarily negative, it means that the market demands a lower premium than the one implied by the UIP hypothesis ($f_t = E_t(s_{t+1})$). This can be a result of less 'risk aversion', more asymmetrical information, or more market frictions. The second component is the unexpected depreciation/appreciation of the domestic currency which, under the assumption of rational expectation, should equal zero, on average. The latter is expected to be 'white noise' and it appears in $t+1$ only, compared to the former.

As risk premium is an unobservable component, a natural approach for the modeling of risk premium is by state-space models. As in Wolff (1987) and Cheng (1993), a reasonable state-space model is

$$f_t - s_{t+1} = P_t + \epsilon_{t+1} \tag{4}$$

$$P_{t+1} = \phi P_t + u_t \tag{5}$$

where $\epsilon_{t+1} = E_t(s_{t+1}) - s_{t+1}$ is the expectation error, which is $i.i.d$ under the assumption of rational expectations, and we further assume that $\epsilon_{t+1} \sim N(0, \sigma^2)$ and $u_t \sim N(0, \sigma_u^2)$. The estimation of the unknown parameters, i.e, $\phi, \sigma^2, \sigma_u^2$, and of the risk premium P_t is constructed in this study by means of Kalman filter and Gibbs sampling as described in Kim and Nelson (1999). The Gibbs sampling has two main stages. Conditional upon the state variable P_t, we sample from the posterior distribution of the unknown parameters $\phi, \sigma^2, \sigma_u^2$

and conditional upon the unknown parameters, we sample from the posterior distribution of the state variable. The algorithm is described in more detail in Appendix B.

2.1.6 Carry To Risk (CTR)

As mentioned above, a significant CT profitability can be problematic to policy makers, especially in small open economies, as it may disturb the conduct of monetary and exchange rate policies as well as market stability. One way to measure the attractiveness of CT strategies is called CTR. The CTR ratio aims to measure the ex-ante, risk-adjusted profitability of a CT position and is based on the IRD between two economies, adjusted for a depreciation risk. Thus, many CTR measures include both IRD and the inherent risk of CT strategies, namely a sharp depreciation of the target currency against the funding currency. Moreover, an asymmetric distribution of the currencies involved in the CT strategy are included in several CTR measures. Following Gyntelberg and Remolona (2007), Dobrynskaya (2014), and Santealla et al. (2015), we calculate several CTR measures, as follows. The first risk measure is based on the historical volatility:

$$CTR1_t = \frac{IRD_t}{Std(\Delta s)} \tag{6}$$

where $\Delta s_t = s_t - s_{t-1}$ and Std is the standard deviation of exchange rate changes for the last n periods ($n = 24$ months in our basic scenario). This measure assumes that CT investors are exposed to a symmetrical variability of the exchange rate changes and the historical variability will be the same in the future.
The second risk measure takes into account the expected future volatility:

$$CTR2_t = \frac{IRD_t}{IV_t} \tag{7}$$

where IV (implied volatility) is an *ex-ante* measure of future exchange rate variability. IV is derived from options written on the exchange rate (ILS/USD) with 'time to expiration' of one month[8]. This measure is an improvement to the first one as it looks ahead for variability rather than looking back to historical variability.
The third risk measure takes into consideration the currency risk premium using the Kalman filter and Gibbs sampling, as follows:

$$CTR3_t = \frac{\frac{1}{n}\sum_{j=0}^{n} P_{t-j}}{IV_t} \tag{8}$$

where P is the time varying currency risk premium defined above, n is the number of periods (24 months for the basic scenario and 2 to 6 months for the robustness checks - hereinafter) and IV is the implied volatility defined in CTR2.[9] This proposed CTR, which

[8] We use three-month ILS/USD options for the long period of twelve-month as there are no other liquid options available.

[9] We also examined a CTR that takes into account the downside risk (see Dobrynskaya (2014)). As the results were quite disappointing we decided to exclude that popular CTR.

is highly correlated with the ER (see Appendix B), is based on the assumption that one-sided gradual deviations from the UIP may signal an increasing carry trading attractivness. For example, if the average excess return for the last n periods (ER^n) is positive, $ER_t^n = \frac{1}{n}\Sigma_{i=1}^{n} ER_{t+1-i} = \frac{1}{n}\sum_{i=1}^{n}(IRD_{t-i} - \Delta s_{t+1-i}) = \frac{1}{n}\sum_{i=1}^{n} IRD_{t-i} - \frac{1}{n}(s_t - s_{t-n+1}) > 0$, we might see more and more carry traders jumping in and taking advantage of such a situation just like momentum traders. However, we use $CTR3_t$ to sum up only the expected component of the ER so that in the above case the average premium of the last n periods is $P_t^n = \frac{1}{n}\Sigma_{i=1}^{n} ER_{t+1-i} - \frac{1}{n}\Sigma_{i=1}^{n} v_{t+1-i}$, while the latter term should equal zero over a long period (24 months). Thus, in comparison with ER, $CTR3_t$ is an *ex-ante* measure, the signal without the noise, and it has an economic meaning. $CTR3_t$ is consistent with Lewis (1989) and with Chakraborty and Evans (2008), who explain deviations from the UIP hypothesis ($ER \neq 0$) by models of learning even under both risk neutrality and rational expectation assumptions. It is also in line with Harris and Yilmaz (2009), who show that momentum trading based on the low frequency component of the exchange rate can be very profitable.[10] Yet, the optimal number of periods, n, is an empirical question, as on the one hand there is much noise in FX markets while on the other hand hesitating can cause a loss of profitable opportunities. In this respect, the dynamics of the behavior of carry traders is quite similar—a gradual and continuous build up of positions in the FX market and a sudden downside or crash (see Brunnermeier et al. (2008)).

2.2 The empirical evidence

The evidence regarding CT's profitability and risk has been growing in the last few years. The crash in the global target currencies such as the New Zealand Dollar (NZD) and the Australian Dollar (AUD) during the sub-prime crisis and the mirror phenomena in the global funding currencies—the Japanese Yen (Y) and the Swiss Franc (SFC), raised interest in this topic among policy makers, researchers, and investors. The latter two examine the outstanding yields such strategies achieved compared to alternative investments, and the inconsistency with the UIP hypothesis (see: Brunnermeier et al. (2008); Gyntelberg and Remolona (2007); Lustig et al. (2013); Dobrynskaya (2014); Doskov and Swinkels (2015)). As reflected by the various CTRs presented above and by the evidence, CT profitability is positively influenced by interest rate differentials while negatively influenced by the *ex-ante* risk (Clarida et al. (2009)). However neither the CIP nor the UIP hypotheses explicitly include risk measures, while most studies that examine the deviations from both the CIP and the UIP hypotheses (i.e., arbitrage opportunities and the forward premium puzzle, respectively) usually focus on risks embedded in CT activity.
It is widely accepted that in efficient markets there are no 'arbitrage opportunities' at least for longer horizons (see Akram et al. (2008)). Thus, CIP should prevail up to the bid-ask spreads, regardless of investors' preferences. However, Akram et al. (2008), Baba and Packer (2009), Fong et al. (2010), and Santealla et al. (2015) find temporarily large deviations as a result of, *inter alia*, liquidity risk, credit risk, or regulatory interventions. These short-lived arbitrage opportunities are inconsistent with the CIP condition. Moreover, the size of CIP arbitrage opportunities can be economically significant.

[10] Using an HP filter suitable for monthly data yields similar results.

Another, and even more important, source of profitability related to the CT activity is the deviation from the UIP hypothesis. Following the seminal paper of Fama (1984), Froot and Thaler (1990), among others, report on a few studies where $\hat{\beta}$ of equation (1) was found to be positive, and even in these few instances it was less than the UIP hypothesis, i.e., $\hat{\beta} = 1$. The average $\hat{\beta}$ of the studies Froot and Thaler (1990) surveyed was -0.88. Yet, Baillie (2011), like many others, shows that the bias is both period and country dependent. The evidence is that for developed countries $\hat{\beta} < 0$ significantly, for emerging markets $0 < \hat{\beta} < 1$, but for all countries the explanatory power of the UIP's OLS regressions is very poor, i.e., $Adj.R^2 \approx 0$, and the residuals are usually not a 'white noise' series (see Frankel and Poonawala 2010). Thus, the UIP hypothesis holds relatively more in emerging markets and for longer horizons. (See Alper et al. (2009) for a survey on the UIP hypothesis in emerging markets)
In order to explain the 'forward premium puzzle', several studies add a time-varying risk premium to the standard OLS regression (see a survey in Engel (1996)). The deviation from the UIP hypothesis is explained by relaxing the assumptions of: (1) 'risk neutrality' of the investors and 'rational expectations', (2) 'efficient markets' where there are no 'arbitrage opportunities', no transaction costs, and liquid markets, (3) symmetrical information and no 'learning process', (4) Free capital flows and a central bank that does not intervene in the markets either directly or indirectly through its monetary policy, and (5) the linearity or the symmetry of the ER. This phenomenon is sometimes called 'downside risk' or 'crash risk' and it might be a result of the Peso problem[11]. The relaxation of any one of these assumptions limits the potential of profitable CT strategies. This is found by Dobrynskaya (2014) to be a significant explanatory variable in explaining both high profitability for carry trade in developed countries and the stock market profitability. In contrast, Jurek (2014) estimates the crash risk as being as low as one-third of the total CT returns. In any case the interpretation of the risk premium and its capability to predict CT activities are not simple (see Doskov and Swinkels (2015)).
In this regard, Kumar and Trück (2014) and Inci and Lu (2007) find that three-month FPs convey more information regarding future spot rates than one-month FPs due to the relative importance of a time-varying risk premium. This is because risk is not materially embedded in short-term FPs. As a result, short-term FPs are consistent with the UIP hypothesis while the former are not.
The explanations of Kumar and Trück (2014) and Inci and Lu (2007) for the negligible influence of the time-varying risk premium on short-term FPs versus its substantial influence on long-term FPs are: (1) Hedgers who have long-term exposure in the spot market (for example, exporters/importers) take an opposite exposure in the forward market. In addition, the longer the maturity of the forward contract the less roll-over costs may be until the end of the exposure period. In contrast, speculators which are usually not exposed in the spot market prefer short-term forward contracts. (2) Short-term interest rates, which are supposed to equal the FP according to the CIP hypothesis, are mostly influenced by the central bank, and are thus less affected by risk premiums. This means that the UIP hypothesis is likely to hold for short-term contracts but unlikely to hold for longer-term contracts. Furthermore, when there are many hedgers compared to speculators in the market, the risk premium is likely to be an important factor of the FP, and thus the UIP is unlikely to hold, while the

[11] A situation when low frequency expected events may lead to long periods of ex-post excess returns.

opposite is true if there are many speculators in the market (all other things being equal, particularly time to maturity).
Finally, in a similar study, Santealla et al. (2015) tried to find relationships between foreign investors' flows and carry trade attractiveness measures (mainly CTR2) in Latin American countries. They inded found strong contemporaneous relationships between CTR2 and foreign investors' positions, but not a predictive capability. They also report on deviations from the CIP that were exploited by foreigners in some countries.
As the time-varying risk premium is a latent variable, it is estimated in the literature using various methodologies (see Engel (1996)). We follow Cheng (1993) and Wolff (1987), and assess that premium using the Kalman filter and Gibbs sampling (see Appendix B). By doing so we combine two strands of the literature: the carry trade attractiveness using interest rate differentials adjusted for various measures of risk and the estimation of a (latent) time-varying risk premium.

3 Data and descriptive statistics

Our data consist of daily data of the shekel/dollar spot rate (the 'representative exchange rate' calculated and reported by the Bank of Israel), one- and twelve-month forward rates, and one- and twelve-month local and global interest rates (the *makam* and the LIBOR, respectively). The forward rates are the interbank quotes reported by Bloomberg information system. The data sample spans the period from January 1, 2003 to December 31, 2014 . For each day, we carefully adjust the appropriate interest rate and the forward 'time to expiration' so that forward premiums and interest rate differentials are consistently determined. Then, we aggregate the daily data into monthly averages as common CT strategies may last several months. Following Curcuru et al. (2010), who report on a sharp decrease of CT attractiveness and unwinding of CT positions following the sub-prime crisis in 2008, we test for different behavior of the CIP and UIP hypotheses in various sub periods. (see Alper et al. (2009) recommendations for emerging markets and Baillie (2011).) In particular, we carry out the Bai and Perron (1998) procedure, which tests for possible breakpoints in the UIP regression parameters, and then estimate several regression models for each sub-period accordingly. The results from the Bai and Perron test are depicted in Figure 1. The test identifies three breakpoints, all located in the years 2008-2009. Therefore, we define three periods: first (1/2003-7/2008), second (4/2009-3/2014), and all (1/2003-12/2014). The latter is the entire sample including both the sub-prime sub period and the short period 4/2014-12/2014 where, by the Bai and Perron test, a new sub period started.

As in Fama (1984), we calculate basic statistics of exchange rate changes ($s_{t+1} - s_t$), FP($f_t - s_t$), ER($f_t - s_{t+1}$), and IRD($i_t - i_t^*$). The results are presented in Table I.

Table I contains one-month shekel-dollar exchange rate changes, ER, FP, and interest rate differentials (IRD) between one-month *makam* and LIBOR. It is divided into three periods: all (1/2003-12/2014), first (1/2003-7/2008), and second (4/2009-3/2014). The standard deviation of the one-month FP ($f_t - s_t$) is much smaller than the standard deviation of the other variables, which may indicate that the FP conveys relatively little information regarding future returns (see Inci and Lu (2007)). The autocorrelation of ER_{t+1} is larger in the first lag than the autocorrelation of $s_{t+1} - s_t$, perhaps indicating, a more substantial auto-

Figure 1. Breakpoints in the OLS regression of the UIP (Equation 1)

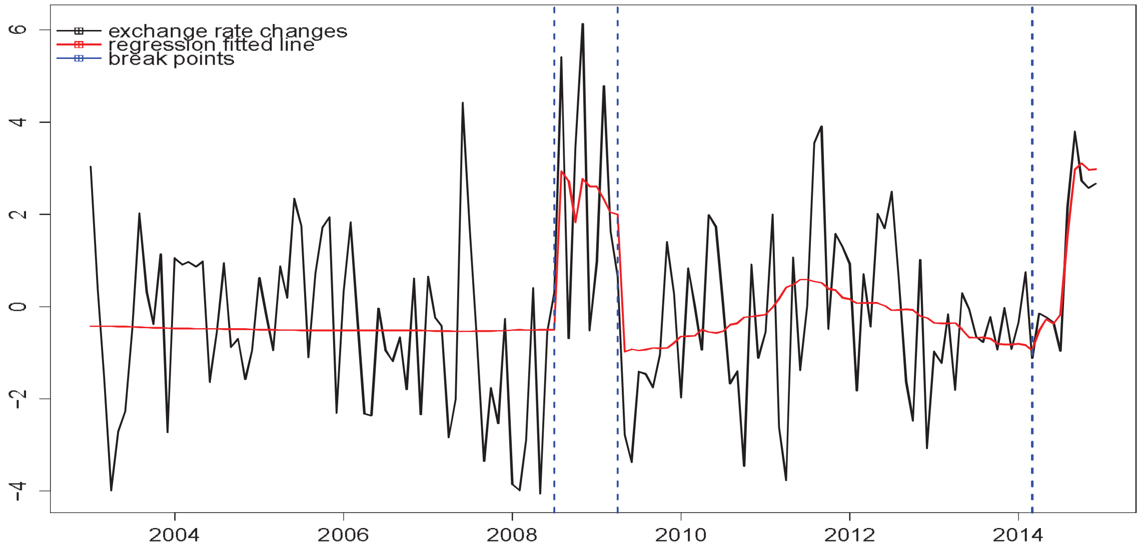


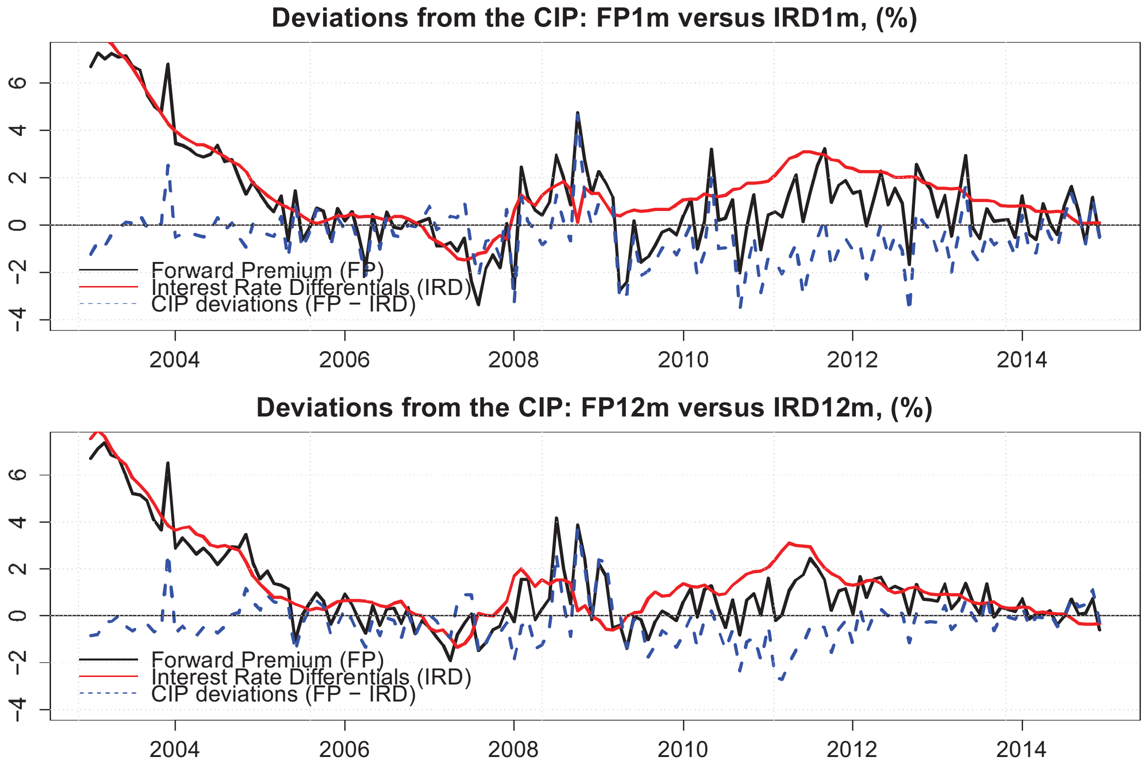
Table I

Descriptive statistics of the main variables

VARIABLE	PERIOD	ρ_1	ρ_2	ρ_3	ρ_4	MEAN	SD	MAX	MIN
$\Delta s_{t+1} = s_{t+1} - s_t$									
dusd	all	0.307	0.116	0.134	0.083	-0.124	1.952	6.138	-4.06
dusd	first	0.255	0.002	-0.038	0.13	-0.497	1.798	4.419	-4.06
dusd	second	0.226	0.082	0.167	-0.024	-0.317	1.679	3.908	-3.77
$ER_{t+1} = IRD_t - \Delta s_{t+1} = f_t - s_{t+1}$									
ER1m	all	0.76	0.704	0.713	0.62	1.087	2.167	7.274	-3.325
ER1m	first	0.873	0.848	0.792	0.742	1.592	2.765	7.274	-3.325
ER1m	second	0.201	0.117	0.406	0.136	0.552	1.247	3.23	-2.362
$FP_t = f_t - s_t$									
FP1m	all	0.758	0.701	0.711	0.619	0.09	0.181	0.605	-0.28
FP1m	first	0.873	0.847	0.791	0.742	0.132	0.231	0.605	-0.28
FP1m	second	0.203	0.119	0.405	0.135	0.046	0.105	0.269	-0.2
$IRD_t = i_t - i_t^*$									
IRD1m	all	0.947	0.889	0.828	0.765	1.594	1.89	7.957	-1.472
IRD1m	first	0.952	0.897	0.837	0.774	1.824	2.619	7.957	-1.472
IRD1m	second	0.95	0.893	0.825	0.75	1.613	0.76	3.087	0.532

The variable 'dusd' is the exchange rate changes, 'ER1m' refers to Excess Return, 'FP1m' is the Forward Premium, and 'IRD' is the Interest Rate Differential between Israel (Telbor) and the US (Libor). All variables are for one month. ρ_j {j = 1..4} are the autocorrelation coefficients of the examined series.

Figure 2. The development of the deviations from the CIP hypothesis



correlation in the risk premium. The autocorrelation of $f_t - s_t$ is relatively large compared to both $f_t - s_{t+1}$ and $s_{t+1} - s_t$. This may be explained by the fact that the two latter include unexpected components with high variability. These results are quite similar to those found in Fama (1984). Finally, the autocorrelations during the second sub-period are smaller than the respective figures during the first period and the autocorrelations of the IRD are much larger than the respective FPs. This phenomenon of the second sub-period is inconsistent with the CIP, and thus calls for an examination of the CIP hypothesis within the various sub-periods.

3.1 Testing the CIP hypothesis

As mentioned above, the CIP states that the forward premium should equal the IRD. The differences between forward premium and IRD are depicted in Figure 2. It is easy to see that the CIP does not perfectly hold in the entire sample data. This is especially true for the second sub-period where the CIP deviations (the dotted line) are scattered below the zero horizontal line. Additionally, the deviations from the CIP are more prominent with regard to the one-month contract than with regard to the twelve-month horizon. Such a relatively long sub-period in which the CIP does not hold is uncommon by the CIP empirical evidence, and may reflect market frictions as in some Latin American countries (see Santealla et al. (2015)) and during the sub-prime crisis (see Baba and Packer (2009)).[12] A more rigorous way to assess the CIP hypothesis is to run the following OLS regression:

$$FP_t = f_t - s_t = \alpha + \beta(i_t - i_t^*) + u_t \tag{9}$$

where the null is: $\beta = 1$ or $\alpha = 0; \beta = 1$. The results for the entire sample are depicted in Table II. For both maturities (one and twelve months) in Table II, the difference between the

Table II : CIP hypothesis test

H_0: (FP = IRD)	Period	α	β	Adj.R^2	Wald ($\beta = 1$)	Wald ($\alpha = 0; \beta = 1$)
1m	all	-0.475†	0.979‡	0.73	0.59	0
1m	first	-0.249†	1.007‡	0.91	0.85	0.07
1m	second	-0.902‡	0.899‡	0.3	0.55	0
12m	all	-0.138	0.894‡	0.75	0	0
12m	first	-0.195	0.971‡	0.89	0.34	0.02
12m	second	0.018	0.486†	0.23	0	0

‡, †, and * denote 0.01, 0.05, and 0.1 significance level, respectively. The OLS regressions employ the Newey and West (1987) heteroskedasticity and autocorrelation consistent standard errors.

second sub-period and other periods (all and first) is quite prominent. For example, β of the second period for twelve months is less than 0.5 while that of the first period is higher than 0.97. In addition, the second period total goodness of fit for all maturities is less than or equal to 0.3 while in the other periods is higher or equal to 0.73. Moreover, the CIP hypothesis that $\beta = 1$ is rejected by the Wald test in the second period for twelve months but not for one month. However, the null $\alpha = 0; \beta = 1$ is rejected by the Wald test for all cases (at the 10 percent significance level). The regression results corroborate Figure 2 regarding the partial similarity between the forward premium and the IRD, especially during the second sub-period when the β coefficient is as low as 0.899 for one month and 0.486 for twelve months. The deviations from the CIP, however, are inconsistent with the evidence that in efficient markets the CIP holds—i.e., there are no arbitrage opportunities (Akram et al., 2008). This surprising result can be partially explained by the massive foreign investments in Israeli short-term bonds (mainly *makam*) during the second half of the sample period which apparently caused a liquidity shortage for domestic financial companies and domestic banks. Such a shortage was reflected by large swap spreads offered by domestic banks compared to foreign banks. This is also consistent with Fong et al. (2010), who state that deviations from the CIP can reflect liquidity and credit risks and with Baba and Packer (2009) who document a similar shortage to European banks of US dollars during the sub-prime crisis.

[12] Despite their volatile behavior, FPs perceived as more reliable than IRDs so, many studies use FP rather than IRD as regressors in the UIP hypothesis tests (see Kim and Song (2014)).

3.2 Testing the UIP hypothesis

The deviation from the CIP calls for a test of the UIP, which is even more relevant to the CT strategy than the CIP. As the CIP held only partially during the sample period, we shall examine the UIP with both the forward premium and the IRD as regressors. The UIP is basically tested by regressing the future exchange rate changes on the forward premium (Equation 1) or alternatively (assuming the CIP) on the IRD, as follows:

$$\Delta s_{t+1} = \alpha + \beta(i_t - i_t^*) + u_{t+1} \tag{10}$$

In both cases if the UIP holds we expect that: $\alpha = 0; \beta = 1$, and $E(u_{t+1}) = 0$.

Table III : OLS regression results of two alternative UIP tests

Δs_{t+1}	FP/IRD	Period	α	β	Adj.R^2	Wald ($\beta = 1$)	Wald ($\alpha = 0; \beta = 1$)
1m	IRD	all	-0.08	-0.04	0	0	0
1m	IRD	first	-0.43	-0.017	0	0	0
1m	IRD	second	-1.37‡	0.68†	0.1	0.29	0
1m	FP	all	-0.336†	0.177*	0.04	0	0
1m	FP	first	-0.571†	0.069	0.01	0	0
1m	FP	second	-0.613‡	0.62‡	0.22	0.01	0
12m	IRD	all	-3.783‡	1.28‡	0.08	0.56	0
12m	IRD	first	-4.989‡	1.162†	0.09	0.73	0
12m	IRD	second	-5.1†	3.594†	0.14	0.06	0.06
12m	FP	all	-3.112‡	1.058‡	0.06	0.88	0
12m	FP	first	-4.792‡	1.219‡	0.1	0.62	0
12m	FP	second	-1.776	1.782	0.04	0.56	0.25

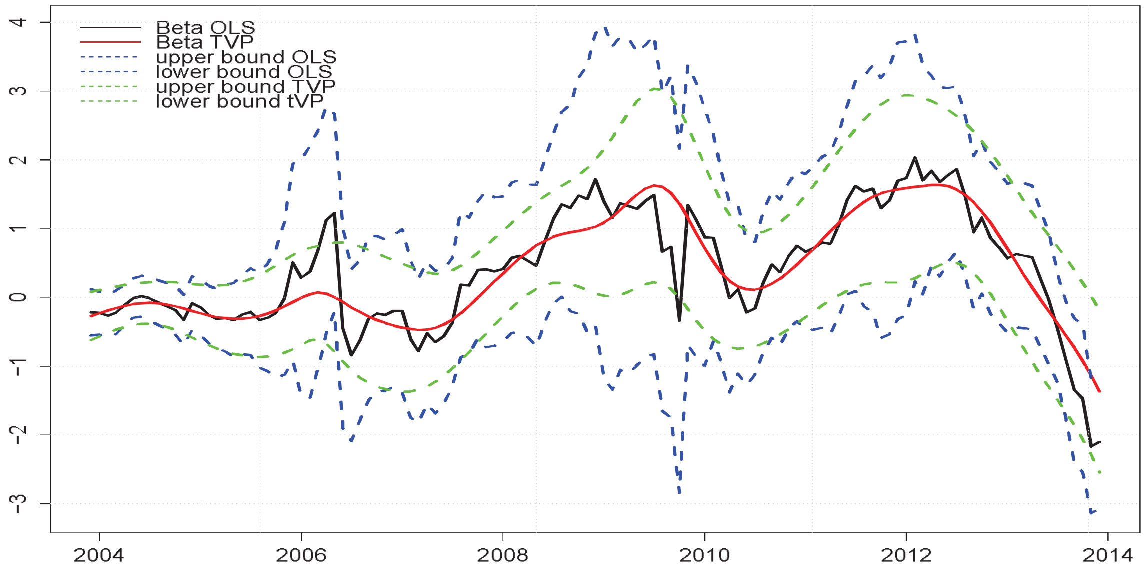
‡, †, and * denote 0.01, 0.05, and 0.1 significance level, respectively. The OLS regressions employ the Newey and West (1987) heteroskedasticity and autocorrelation consistent standard errors.

The results, which are depicted in Table III, go hand in hand with the evidence that the total goodness of fit ($Adj.R^2$) of the models is quite poor. In all cases $\alpha < 0$, which is against the UIP hypothesis. Generally, the coefficients of FP and IRD (β in Equations 1 and 13, respectively) become large and more significant, and the UIP hypothesis is less rejected by the Wald test, with maturity. For example, for the longer maturity of twelve months $\hat{\beta} > 1$, the UIP hypothesis that $\beta = 1$ cannot be rejected by a one percent significance level, and the null hypothesis that $\beta = 1$ or alternatively $\alpha = 0; \beta = 1$ cannot be rejected by the Wald tests for the second period. These results are inconsistent with Kumar and Trück (2014) and Inci and Lu (2007), especially for the one-month horizon, as they could not reject the null for one-month horizon but did reject it for three-month horizon. In contrast, our results show the opposite, namely the UIP hypothesis is rejected by the one-month horizon but not by the twelve-month horizon.

In general, our results regarding the twelve-month horizon are consistent with Chinn and Meredith (2005), who find that the UIP holds in longer time horizons (at least quarterly data and more than a year).

The results of Table III and the Baillie and Cho (2014) argument regarding the β instability

Figure 3. The UIP's $\hat{\beta}$: OLS rolling regression versus TVP (Time Varying Parameter)



call for a model of the time varying β. Figure 3 displays the $\hat{\beta}$ from a rolling OLS regression with a 24-month (centered) window using Equation 1. Following Baillie and Cho (2014) we additionally draw a Time Varying Parameter (TVP), which is based on a kernel estimation. (For a description of the TVP method see Appendix C.) In addition we depict upper and lower bounds, one for the OLS $\hat{\beta}$ (5 percent standard errors) and the other for the TVP $\hat{\beta}$. According to both methods the $\hat{\beta}$ shows a significant change over time, fluctuating between -2 and 2. Obviously, the TVP is smoother than the rolling regression $\hat{\beta}$ since the TVP is a weighted average of all observations that assigns decaying weights to far observations while the rolling regression assigns equal weights to observations within a moving window. This causes to a jump in $\hat{\beta}$ whenever a large observation enters or exits the moving window. It is also evident that the bounds of the TVP are beneath those of the rolling regression in the first period until 2012 and are relatively close in the subsequent period. The unstable β is consistent with the empirical evidence, and is the result of either the breaks during the sample period or the non-normal distribution of the ER (i.e., downside risk). These risks usually reflect the Peso problem (see FN 11) and the potential of significant negative returns.

4 Estimation results

This section describes episodes of carry trade activity in Israel and compares the capabilities of the CTRs described above to predict the nonresidents' position in government bonds and *makam*.

4.1 Carry trade activity in Israel

Usually, carry trade activity is more intense in periods of high IRD and low exchange rate volatility (Clarida et al. (2009)), especially where the exchange rate regime is not freely floating, e.g., managed in a band by the central bank or any other regulatory body. Despite the existence of attractive conditions for carry trade in the past in Israel, substantial activity began to take place only in 2006. This is partially explained by the prominent activity of foreign banks' in the domestic market since that time. The CT's attractiveness is measured in this study by three measures (CTRs) including the proposed time-varying risk premium. In what follows, we examine the relationships between the various CTRs and nonresidents' position in government bonds and in *makam* using a proprietary dataset of data reported to the Bank of Israel (BOI). This dataset consists of all foreign investors' holdings in any Israeli bonds from January 2003 to December 2014.

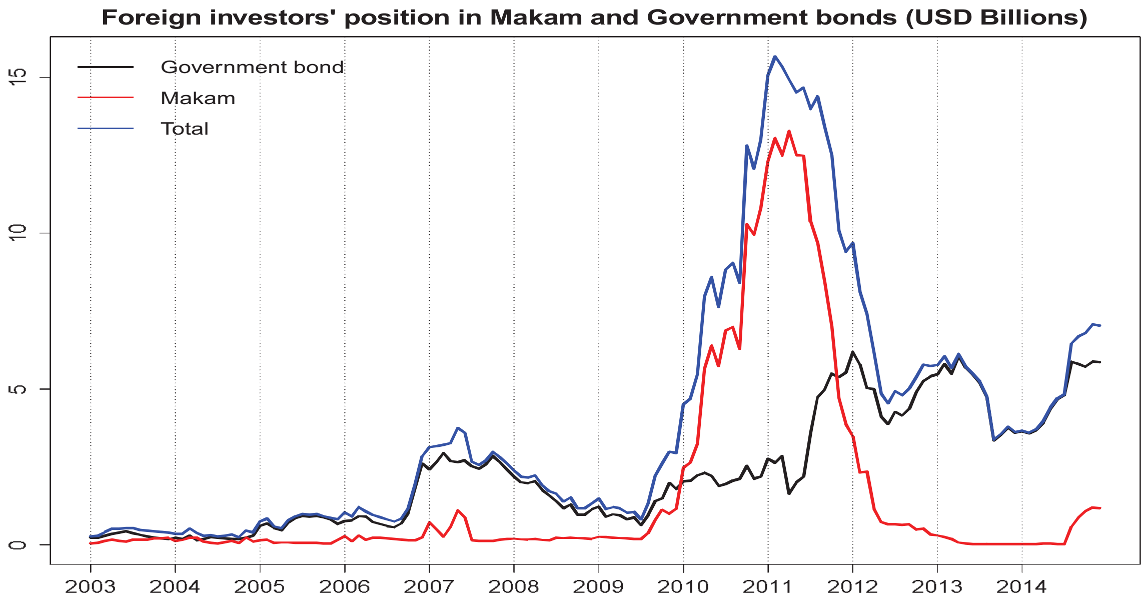
4.2 Historical examination of nonresident positions and CTRs

The carry trade measures introduced above supposedly measure the potential profitability of carry trade activity. By no means do these measures serve as leading indicators to carry trade activity based on rigorous forecasting models. However, assuming these measures indeed reflect the potential profitability of this type of activity, we wish to examine historically to what extent these investment opportunities were exploited by various players in the FX market. Since carry trade activity is usually an uncommon knowledge, we measure that activity by the nonresidents' position as in Santealla et al. (2015)). The position of foreign investors in *makam*, in government bonds, and total position (in both *makam* and bonds) is depicted in Figure 4.

Figure 4 shows that the total position of foreign investors in the local bond market peaked around 2011 due to a relatively large increase in their holdings of *makam*. Since that peak, and as a result of the BOI's compulsory reporting and liquidity demands, nonresidents' position reverted back to the long-run levels. In contrast, nonresidents' position in government bonds are less volatile and are characterized by a positive trend. This unique episode, which is local rather than global, demonstrates the potential impact of nonresidents' CT strategies (see Santealla et al., 2015).

In order to learn about the relationships between the various CTRs (see equations 6-8 above) and nonresidents' position in Israel's government bonds, we examine the cross-correlation functions up to 18 months. We are interested especially in CTRs that precede or contemporaneously correlate with nonresidents' positions by finding the largest cross-correlation coefficient (hereinafter ccc) between each one of the three CTRs and the three positions (bond, *makam*, and total). Particularly, we calculate the cross-correlation function (hereinafter ccf) for each combination of CTR and position (3*3). The ccf displaces one series versus the second one and calculates the contemporaneous correlation coefficient between the two. We are looking for abnormal (larger than 0.2 in absolute terms) increasing cross-correlation coefficients i.e., significant statistical causation from any CTR to the nonresidents' position in bond, in *makam*, or in total holdings. Moreover, the maximal ccc should occur immediately before or at lag 0 and should decay afterwards. Concerning our proposed CTR measure (CTR3), recall that $P_t = f_t - E_t(s_{t+1})$. Thus, a positive P means that the

Figure 4. Nonresidentss' position in *makam* and goverment bonds



risk-adjusted market expectations of the market for a given depreciation are smaller than the realized FP. In other words, 'risk averse' investors or hedgers are willing to pay higher forward rates compared to market expectations (all other things being equal). Therefore, the sign of a particular ccc is important, as significant positive ccc immediately before lag=0 reflects a situation when a positive P predicts or statistically causes larger positions in bonds or *makam* by carry traders and vice-versa. The PNL of CT is positively influenced by IRD, which is equal to the FP by the CIP hypothesis, and negatively affected by the realized future exchange rate depreciation. In contrast, a significant negative ccc before or at lag=0 is less intuitive for CT, as larger FP or IRD compared to the expected depreciation statistically causes smaller positions in *makam* or bonds. Finally, an insignificant ccc can reflect a situation where the risk premium is negligible compared to FP or IRD.
Based on the evidence we present in Figures 5-8, the various CTRs using one-month horizon during the second period, when nonresident investors were active in domestic bonds.

The results depicted in Figures 5-7 show that CTR1, CTR2, and CTR3 fulfill the requirements for a precedence CTR namely, they positively peak around lag=0 and decays thereafter for foreign positions in government bond, only. However, there is a substantial difference among these three CTRs: the first two CTRs peak several months before lag=0 while CTR3 peaks at that point.[13] In order to further examine the differences between the

[13] For the sake of brevity we do not present the first period. However, the examination of the two subperiods show a different behavior of the ccf. During the first period all ccc's negatively peak before lag=0 and decay afterwards. The ccf differences between the first and the second periods are probably the result of both the high IRD and the absence of foreign investors from the local market during the first period.

Figure 5. Cross-correlation between nonresidents' positions and CTR1

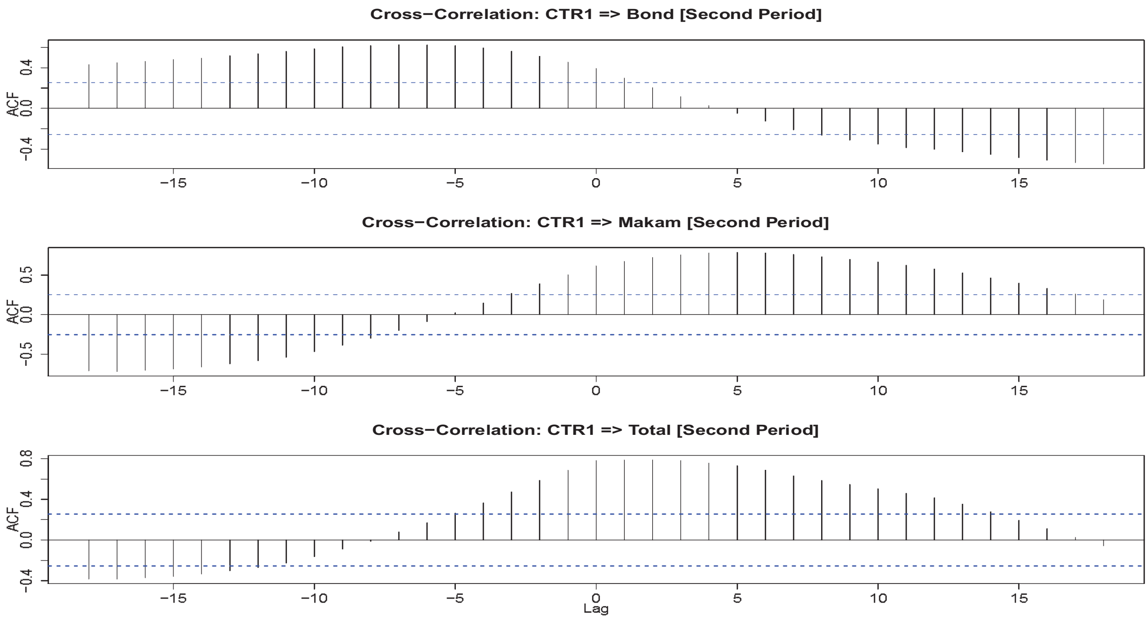


Figure 6. Cross-correlation between nonresidents' positions and CTR2

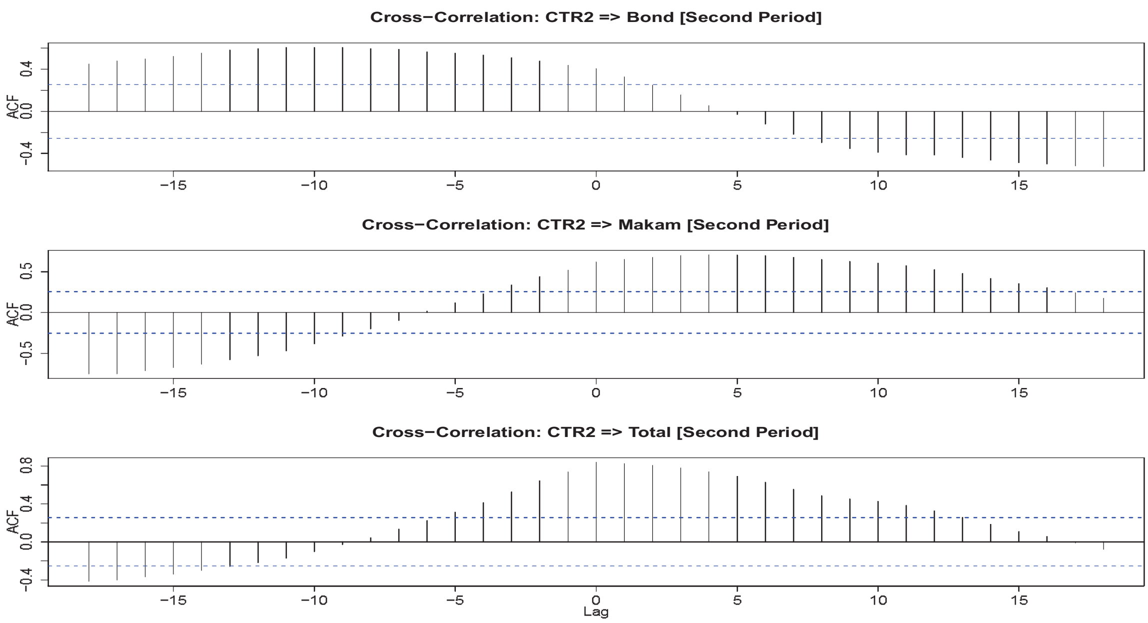
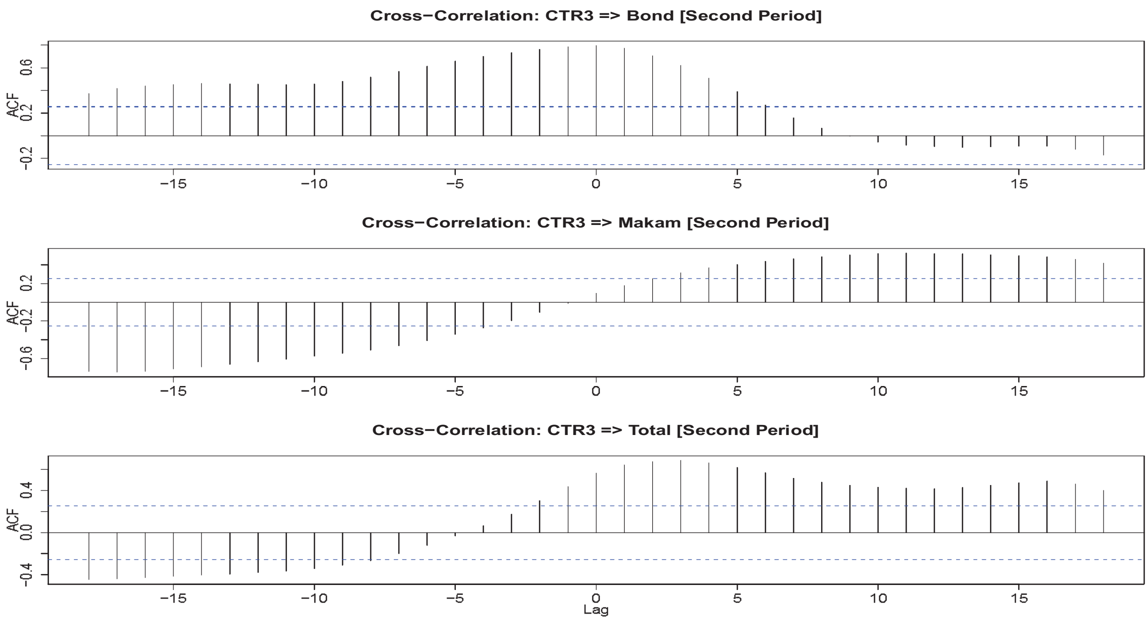


Figure 7. Cross-correlation between nonresidents' positions and a currency risk premium (CTR3)



various CTRs and their capability to predict nonresidents' positions we regress the foreign investors' position in government bonds and in *makam* on the lags of the various CTRs during the second sub-period, as follows:

$$Pos_{j,t} = \alpha + \beta CTR_{t-k} + u_t \tag{11}$$

where j $\in$ (government bond, *makam*, total holdings), and k $\in (0, ..., 6)$. The results of the OLS regressions are presented in Table IV and corroborate those of the ccfs of Figures 5-7. For the positions in *makam* most βs are insignificant and the goodness of fit estimates ($Adj.R^2$) are poor. The main difference between both CTR1 and CTR2 versus CTR3 is the robustness of the lag coefficients. For example, CTR3—our proposed measure—is positively significant from lag=0 to lag=6 while CTR1 and CTR2 are positively significant from lag=3 only. Additionally, the goodness of fit ($Adj.R^2$) of CTR3 is larger than those of CTR1 and CTR2. To accomplish the examination of the cross-correlations between the various CTRs and the nonresidents' positions we estimate Granger's causality tests in Table V. A standard Granger's causality test addresses the question (in the case of two time-series variables, X and Y), of wether Y can be better predicted using the histories of both X and Y than it can by using the history of Y alone. If the answer is affirmative one can say that, X Granger-causes Y. The formal test with p lags is given by,

$$Y_t = a_0 + a_1 Y_{t-1} + ... + a_p Y_{t-p} + b_1 X_{t-1} + ... + b_p X_{t-p} + u_t \tag{12}$$

$$X_t = c_0 + c_1 X_{t-1} + ... + c_p Y_{t-p} + d_1 Y_{t-1} + ... + d_p Y_{t-p} + v_t \tag{13}$$

Table IV : OLS regression results: nonresidents' holdings as a function of CTR lags during the second sub-period

CTR	Lag	α_{mak}	β_{mak}	$Adj.R^2_{mak}$	α_{bond}	β_{bond}	$Adj.R^2_{bond}$
CTR1	0	-1.361	6.079‡	0.4	2.305	1.447	0.1
CTR1	1	-0.468	5.035†	0.2	2.139	1.688	0.2
CTR1	2	0.513	3.917*	0.1	1.98*	1.924	0.3
CTR1	3	1.654	2.626	0.0	1.86‡	2.118†	0.4
CTR1	4	2.864	1.264	-0.0	1.767‡	2.273‡	0.5
CTR1	5	4.083	-0.098	-0.0	1.692‡	2.396‡	0.6
CTR1	6	5.249	-1.388	0.0	1.645‡	2.484‡	0.6
CTR2	0	-2.538	437.5	0.4	1.946	108.364	0.1
CTR2	1	-1.703	379.2	0.3	1.869	117.124	0.2
CTR2	2	-0.887	324.455	0.2	1.731†	129.075*	0.2
CTR2	3	0.211	251.112	0.1	1.6†	140.703‡	0.3
CTR2	4	1.404	172.28	0.0	1.488†	150.442‡	0.4
CTR2	5	2.726	85.803	-0.0	1.432†	155.964‡	0.4
CTR2	6	4.027	1.759	-0.0	1.385*	160.725‡	0.5
CTR3	0	2.701	18.538	0.0	2.491‡	19.056‡	0.5
CTR3	1	3.211	9.941	-0.0	2.512‡	19.235‡	0.5
CTR3	2	3.691	2.2	-0.0	2.501‡	19.998‡	0.6
CTR3	3	4.172	-5.318	-0.0	2.547‡	19.937‡	0.6
CTR3	4	4.714	-13.719	0.0	2.607‡	19.434‡	0.6
CTR3	5	5.13	-19.823	0.0	2.657‡	19.065‡	0.6
CTR3	6	5.56	-26.1	0.1	2.719‡	18.502‡	0.6

The subscripts 'mak' and 'bond' refer to foreign investor positions in *makam* and government bonds, respectively. *, †, and ‡, denote 0.1, 0.05, and 0.01 significance level, respectively. The OLS regressions employ Newey and West (1987) heteroskedasticity and autocorrelation consistent standard errors.

Then, an F test is conducted to test whether $H_0 : b_1 = ... = b_p = 0$ against H_A: at least one of $b_i \neq 0$. If one cannot reject the H_0, X does not Granger-cause Y. Similarly, testing $H_0 : d_1 = ... = d_p = 0$ against H_A tests wether Y does not Granger-cause X.
Using standard Granger's tests is inappropriate in our case, as the underlying assumption of the test is normal distribution (by using the F test) while all of the unit root tests regarding the various CTRs and the nonresidents' positions indicate an integration level of at least one (I(1) at the five percent significance level). This can be seen in the columns UR_{CTR} and UR_{POS} of Table V, where the number of lags are automatically selected by the Schwartz Information Criteria - SIC (not shown due to space consideration). These results mean that almost all the examined series are non-stationary. We also test for co-integration between each CTR and nonresident positions by running the Johansen co-integration test (maximum eigenvalue) with constant and no trend and with a maximum lag of twelve months. In six out of 12 combinations there are co-integration relations; three of them are CTR3 (risk premium for the entire sample period). While the existence of co-integration relations yields an indication regarding the co-movement of two non-stationary series, it is still not enough to discover how the examined CTRs predicted the nonresidents' positions. As we examine in this study non-stationary data, we hereinafter use the Toda and Yamamoto (1995) non-

stationary Granger's causality test. The detailed procedure is described in Appendix D.

Table V : Non-Stationary Granger causality tests: various CTRs and nonresidents' holdings during the second sub-period

CTR	Position	UR_{CTR}	UR_{POS}	coint.	$select_{lag}$	$optimal_{lag}$	Grang	$best_{CTR}$	$best_{pos}$
CTR1	bond	-4	-4	1	1	1	0.011	0.006	0.384
CTR1	mak	-4	-3	1	1	1	0	0	0
CTR1	tot	-4	-4	1	1	3	0.03	0.059	0.01
CTR2	bond	-5	-4	2	1	2	0.143	0.059	0.14
CTR2	mak	-5	-3	0	1	1	0.002	0.003	0.009
CTR2	tot	-5	-4	0	1	4	0.462	0.125	0.007
CTR3	bond	-6	-4	2	1	1	0.103	0.004	0.649
CTR3	mak	-6	-3	0	1	1	0	0	0.104
CTR3	tot	-6	-4	1	1	2	0.011	0.002	0.116

'UR' refers to unit root tests (ADF), 'coint.' refers to the cointegration level, 'Grag' to the common granger causality tests (for comparison purposes only), and 'best' refers to the Wald test of causality coefficient by Toda and Yamamoto (1995) procedure.

Table V presents the results of the Toda and Yamamoto (1995) non-stationary Granger causality test. By this procedure, we automatically select the number of lags by Schwartz Information Criterion (SIC) and in all cases it is p = 1 month (the column $select_{lag}$). Then for asymptotic purposes we add more and more lags as exogenous variables, until we have no more auto-correlations in the residuals of the VAR model. The maximum $optimal_{lag}$ we find is m = 4 months i.e., p + m = 5 while the minimum is p + m = 2. The next column presents the (inappropriate) standard Granger's causality test for comparison purposes. The two columns farthest to the right show the results of Wald tests of excluding either a CTR variable or a position variable. These columns are related to the "Optimal lag" i.e., p + m lags. A small figure of the Wald test (0.01 or less) means a rejection of the Granger no-causality hypothesis (H_0) in favor of the H_A. As can be seen, the currency risk premium (CTR3) significantly causes the foreign investors' positions (*makam*, government bonds, and total at the one percent significance level) but not the other way around while other CTRs do not.

5 A discussion of the results

The goal of this study is to find a successful CTR that can predict CT activity that is reflected by nonresidents' positions in *makam* and domestic government bonds. Few studies have tried to find such relationships and even fewer report successful results. While Curcuru et al. (2010) did not find convincing evidence of such relationships, Santealla et al. (2015) found contemporaneous positive relationships, between foreign investors' flows and carry trade attractiveness measures (mainly CTR2) in Latin American countries. However, they did not find a predictive capability on the part of any examined CTR. Possible explanations for the unsuccessful tries regarding both CT and the UIP hypothesis are (see Alper et al. (2009)): (1) the lack of public data on foreign investors' positions and CT activity, (2)

inappropriate CTRs e.g., symmetrical ones such as CTR2, (3) discarding breakpoints in the sample period, (4) assuming that the CIP hypothesis holds, (5) examining short-term CT strategies.
By controlling for each of these points we can obtain some significant results in this paper. Our proposed CTR3, has the capability to forecast foreign investor positions in government bonds during the second period, as it is asymmetric and gradually developed, similar to CT (Baillie and Chang (2011)). Additionally, by testing for different behavior of the CIP and the UIP hypotheses in various sub-periods using the Bai and Perron (1998) procedure, we divide the sample period into two different sub-periods. The first period is characterized by relatively large IRD but an absence of nonresidents in the local market (Table I and Figure 4). In contrast, during the second period the presence of foreign investors was quite prominent (e.g., their investment patterns in *makam*), despite the smaller IRD. Moreover, their significant CT activity partially caused market frictions that were reflected in deviations from the CIP hypothesis[14]. The latter was more significant for the twelve-month horizon than the one-month horizon (Table II and Figure 2). The dramatic and unique nonresident investment patterns in *makam* and the (successful) reaction of the Bank of Israel is probably the reason why none of the examined CTRs could predict foreigners' positions in *makam* during the second period (Figures 5-7). In contrast with the *makam*, several CTRs did predict nonresidents' positions in government bonds during the second period (CTR2 in Figure 6 and CTR2/CTR3 in Table IV). However, CTR3 was the only one that gradually increased before the build up of the nonresidents' positions in those bonds and peaked with them (Figure 7 and Table 4). The advantageous performance of CTR3 can also be seen in the non-stationary Granger's causality test (Table V) introduced by Toda and Yamamoto (1995). CTR3 was the only one that Granger-caused nonresidents' positions but not the other way around.
Our results are inconsistent with Kumar and Trück (2014) and Inci and Lu (2007), who find that three-month FPs convey more information regarding future spot rates than one-month FPs due to the relative importance of a time-varying risk premium. Thus, the UIP hypothesis is likely to hold for short-term contracts but unlikely to hold for longer-term contracts. With Table III, one cannot reject the UIP hypothesis, namely that $\beta = 1$ only for FP and IRD for twelve months while this hypothesis is easily rejected for a one-month horizon. This result is also seen in Appendix B where P, which was estimated by twelve-month IRD, is less negative than a respective P that was estimated using one-month IRD. In contrast with Kumar and Trück (2014) and Inci and Lu (2007), our results regarding the twelve-month horizon are consistent with Chinn and Meredith (2005), who find that the UIP holds in longer time horizons (at least quarterly data and more than a year). According to Kumar and Trück (2014) and Inci and Lu (2007), if there are many hedgers compared to speculators in the market, the risk premium is likely to be an important factor of the FP. Thus, the UIP is unlikely to hold, while the opposite is true if there are many speculators in the market. Their conjecture, which has not been directly tested in the literature yet, calls

[14] The refutation of the CIP hypothesis can still be consistent with the 'no arbitrage' argument as arbitrage opportunities must be examined with regard to a particular investor and a point in time rather than the mean of many investors along the month. Such an examination should also take into account both the bid-ask spreads and commissions with regard to the particular investor.

for examination using micro structure data and is left for future research.

5.1 Robustness checks

In this section we present robustness checks regarding CTR3 - the proposed predictive measure for foreign investor positions. First, we examine the influence of CTR3's window size on its predictive capability. Second, we control for the unique episode of Makam's massive purchasing by foreign investors during 2009-2010 and the BOI's steps during 2011 (will be described later). For both checks we add another predictive measure, CTR4, which is similar to CTR2 but with FP rather than IRD in the nominator, as follows:

$$CTR4_t = \frac{FP_t}{IV_t} \tag{14}$$

This measure is necessary when the CIP does not hold i.e., $FP \neq IRD$. Additionally, in order to assess the importance of IRD themselves, we change in the following tables CTR1 such that it is IRD, only (namely, not divided by historical volatility).

5.1.1 Sensitivity of CTR3 to the window size

Table VI presents the cross correlation coefficients (ccc) of the four CTR measures: (1) five months before (-5M), (2) at the same month (0), and (3) five months after (+5M) versus the time of both foreigner positions in government bonds, Makam, and total. Recall that we are interested in CTRs that precede or contemporaneously correlate with foreigners' positions and have an inverted U shape i.e., a maximal (positive) ccc that occurs just before or at lag=0 and decays afterwards. Such a pattern may indicate an appropriate candidate for foreigner position predictors. Panel A of Table VI compares the four CTRs where the window size which is relevant only for CTR3 is set to 2 months. A prominent result for all examined CTRs is that positive ccc can be seen in the second period, only. CTR3 with regard to government bonds during that period yields the highest ccc at lag = 0 (bnd0) and decays afterward (0.74 at lag = 0 and 0.26 at lag +5M) while CTR2 obtains the highest ccc at lag 0 for foreigner total positions (0.84 at lag 0 versus 0.31 and 0.69 at lag -5M and +5M, respectively). In that respect, CTR4 shows similarity with CTR2 which means that FP rather than IRD does not improve the predicting capability of CTR2. This is also the case with CTR1 which is inferior to CTR2 namely, adding risk to the IRD (CTR1) improves its predicting capability. Panel B examines the sensitivity of CTR3 to the window size starting from 2 month up to 10 month. It can be seen that an inverted U shape appears from 2 months up to 10 month with the highest (positive) ccc at 5-6 month window size only for government bonds. This evidence reflects an insensitivity of CTR3 to the selected window size.

5.1.2 Controlling for the BOI's steps

As monetary policy became less and less expansionary during the period 2009-2011, the interest rate gap between Israel and the developed countries widened, and the inflow of capital increased. From August 2009 until April 2011, the value of the asset portfolio

Table VI : CTR3's window size versus foreigner positions: Cross correlations

CTR	period	window	bnd-5M	bnd0	bnd+5M	mak-5M	mak0	mak+5M	tot-5M	tot0	tot+5M
PANEL A: COMPARISON OF CTRS											
CTR1	all	NA	-0.19	-0.25	-0.10	0.00	0.12	0.24	-0.35	-0.32	-0.08
CTR2	all	NA	-0.30	-0.35	-0.22	-0.04	0.04	0.12	-0.46	-0.46	-0.26
CTR3	all	2	-0.40	-0.39	-0.26	-0.19	-0.12	-0.03	-0.60	-0.57	-0.37
CTR4	all	NA	-0.36	-0.35	-0.25	-0.17	-0.10	-0.02	-0.55	-0.52	-0.35
CTR1	first	NA	-0.70	-0.68	-0.39	-0.20	-0.32	-0.25	-0.74	-0.74	-0.46
CTR2	first	NA	-0.75	-0.77	-0.53	-0.25	-0.34	-0.25	-0.80	-0.85	-0.59
CTR3	first	2	-0.73	-0.73	-0.53	-0.23	-0.27	-0.25	-0.81	-0.82	-0.58
CTR4	first	NA	-0.70	-0.71	-0.52	-0.19	-0.25	-0.25	-0.76	-0.80	-0.57
CTR1	second	NA	0.62	0.39	-0.05	0.02	0.61	0.79	0.26	0.78	0.73
CTR2	second	NA	0.55	0.40	-0.03	0.12	0.62	0.70	0.31	0.84	0.69
CTR3	second	2	0.67	0.74	0.26	-0.27	0.20	0.43	0.05	0.65	0.56
CTR4	second	NA	0.39	0.43	0.06	-0.12	0.17	0.24	0.06	0.42	0.26
PANEL B: CTR3'S WINDOW SIZE											
CTR3	second	3	0.67	0.79	0.35	-0.32	0.13	0.41	-0.00	0.60	0.60
CTR3	second	4	0.66	0.80	0.40	-0.35	0.09	0.40	-0.04	0.56	0.62
CTR3	second	5	0.64	0.81	0.46	-0.38	0.05	0.39	-0.08	0.51	0.65
CTR3	second	6	0.63	0.81	0.54	-0.43	-0.02	0.35	-0.13	0.43	0.66
CTR3	second	7	0.60	0.80	0.59	-0.46	-0.07	0.33	-0.18	0.36	0.68
CTR3	second	8	0.58	0.78	0.66	-0.50	-0.13	0.29	-0.23	0.28	0.68
CTR3	second	9	0.58	0.79	0.72	-0.56	-0.20	0.23	-0.28	0.20	0.66
CTR3	second	10	0.57	0.76	0.75	-0.60	-0.24	0.22	-0.32	0.12	0.64

The table depicts the cross correlations: t-5M, t, and t+5M where, 5M is 5 month lag/lead with regard to month t (the cross correlation of t is simply the contemporaneous correlation coefficient). The subscripts 'mak' and 'bnd' refer to foreign investor positions in Makam and Government bonds, respectively. Window size is relevant for CTR3 only and is in months.

of foreigners grew by 700 percent. Most of the activity of foreigners was concentrated in Makam during that period. [15] During 2011 the bank of Israel (BOI) took steps to restore the equilibrium short-term rates by imposing on local banks a reserve requirement of 10 percent against foreigners' transactions in foreign currency derivatives, and a reporting requirement on various transactions. These steps actually brought down foreigners' positions in Makam to almost zero by 2013 (see Figure 4). In order to control for the period of the BOI's steps i.e., from 2011 on we exhibit in Table VII the results of the last table, during the second period but until December 2010.

Table VII : CTR3's window size versus foreigner positions before the BOI's steps

CTR	window	bnd-5M	bnd0	bnd+5M	mak-5M	mak0	mak+5M	tot-5M	tot0	tot+5M
PANEL A: COMPARISON OF CTRs										
CTR1	NA	-0.09	0.80	0.46	0.28	0.97	0.26	0.23	0.98	0.30
CTR2	NA	-0.08	0.82	0.42	0.32	0.92	0.19	0.28	0.94	0.23
CTR3	2	0.12	0.70	0.27	0.69	0.63	-0.12	0.63	0.66	-0.07
CTR4	NA	0.18	0.37	-0.01	0.40	0.30	-0.19	0.38	0.32	-0.18
PANEL B: CTR3'S WINDOW SIZE										
CTR3	3	0.05	0.68	0.39	0.62	0.68	-0.04	0.57	0.70	0.02
CTR3	4	0.00	0.59	0.52	0.53	0.68	0.04	0.47	0.69	0.11
CTR3	5	-0.06	0.49	0.59	0.42	0.67	0.13	0.36	0.67	0.19
CTR3	6	-0.12	0.33	0.68	0.27	0.66	0.24	0.22	0.64	0.31
CTR3	7	-0.17	0.17	0.74	0.11	0.64	0.36	0.07	0.59	0.42
CTR3	8	-0.21	-0.06	0.75	-0.08	0.52	0.43	-0.10	0.45	0.49
CTR3	9	-0.22	-0.19	0.75	-0.20	0.41	0.46	-0.21	0.34	0.52
CTR3	10	-0.22	-0.39	0.64	-0.34	0.24	0.48	-0.33	0.16	0.51

The table depicts the cross correlations: t-5M, t, and t+5M where, 5M is 5 month lag/lead with regard to month t (t is the contemporaneous correlation coefficient) during the period 4/2009 - 12/2010. The subscripts 'mak' and 'bnd' refer to foreign investor positions in Makam and Government bonds, respectively. Window size is relevant for CTR3 only and is in months.

Focusing on the period before the BOI's steps yields completely different results for all CTRs and particularly to CTR3. In comparison to Table VI the maximal ccc of foreigner positions in Makam, are higher especially around lag 0. Moreover, for most CTR3s in Panel B of Makam an inverted U shape can be seen while those of government bonds are less robust compared to Table VI. This can indicate a substitution between positions in Makam and in Government bonds during that period. The latter inference, however, is not unshakable as the examined period is both non-stationary and quite short thus, cross-correlation as well as contemporaneous correlation which are presented in tables VI and VII can be a misleading statistic for goodness of fit.

[15] The share of foreigners in the makam and short-term government bond markets grew from 1.8 and 0 percent in August 2008 to 34.5 and 17.2 percent in May 2011, respectively.

6 Conclusions

This study investigates the attractiveness of carry trade strategies in the Israeli FX market during the period from January, 2003 to December, 2014. We examine several Carry To Risk (CTR) popular measures including measures that take into account *ex ante* risk or asymmetric downside risk. In particular, we examine the predictive capability of various CTRs in forecasting CT activity, which is represented by nonresidents' investment in *makam* and government bonds. We also propose a new CTR measure which is based on a (latent) time-varying currency risk premium. This CTR is assessed using the Kalman filter and Gibbs sampling. By doing so we combine two strands of the literature: the carry trade attractiveness using interest rate differentials adjusted for various measures of risk, and the estimation of a time-varying currency risk premium. In order to estimate the currency risk premium we also assess the Covered Interest rate Parity (CIP) and the Uncovered Interest rate Parity (UIP) hypotheses using a proprietary dataset. Using the Bai and Perron (1998) procedure, which tests for possible breakpoints in the UIP regression parameters, we find two different sub-periods. We find that the CIP hypothesis does not hold during the second half of the sample period while the UIP hypothesis holds for the twelve-month horizon only. Although the latter is usually consistent with the empirical evidence, the former is less common and is explained by uncommon investments of nonresidents in *makam*—a phenomenon that created market frictions such as a shortage in financial instrument supplies by local banks. In cross-correlation functions, OLS regressions with lags, and causality tests of non-stationary series, we find that our proposed CTR predicted nonresidents' positions in government bonds during the second sub-period but not the other way around. This finding is both more robust and lasts for more months than the other examined CTRs.

Appendix A The Meaning of $\hat{\beta}$ while assuming a risk premium

The estimated coefficient $\hat{\beta}$ is defined as,

$$\hat{\beta} = \frac{cov(s_{t+1} - s_t, f_t - s_t)}{var(f_t - s_t)}$$

Adding and subtracting $E_t(s_{t+1})$, yields:

$$\hat{\beta} = \frac{cov(E_t(s_{t+1}) - s_t, f_t - s_t) + cov(s_{t+1} - E_t(s_{t+1}), f_t - s_t)}{var(f_t - s_t)}$$

If we further write according to Equation (2) $E_t(s_{t+1} - s_t) = f_t - s_t - P_t$ then, it is easy to check that,

$$\hat{\beta} = \frac{var(f_t - s_t) + cov(P_t, f_t - s_t) - cov(f_t - s_t, E_t(s_{t+1}) - s_t)}{var(f_t - s_t)}$$

$$\hat{\beta} = 1 + \frac{cov(P_t, f_t - s_t)}{var(f_t - s_t)} - \frac{cov(f_t - s_t, E_t(s_{t+1}) - s_t)}{var(f_t - s_t)}$$

The assumption of rational expectation means that there are no systematic prediction errors in this market implying that $cov(f_t - s_t, E_t(s_{t+1}) - s_t) = 0$ and

$$\hat{\beta} = 1 + \frac{cov(P_t, f_t - s_t)}{var(f_t - s_t)}$$

It is clear now that $\hat{\beta} = 1$ when $cov(P_t, f_t - s_t) = 0$ probably implying that $f_t - s_t$ does not include a risk premium. As mentioned before many studies including Fama (1984) find that $\hat{\beta} < 0$ indicating that $cov(f_t - s_t, E_t(s_{t+1}) - s_t) < 0$ since,

$$\begin{aligned} 0 > cov(f_t - s_t, s_{t+1} - s_t) &= var(E_t(s_{t+1} - s_t)) + cov(P_t, E_t(s_{t+1} - s_t)) \\ &\geq cov(P_t, E_t(s_{t+1} - s_t)) \end{aligned}$$

and we get that,

$$var(P_t) > var(E_t(s_{t+1}) - s_t)$$

since,

$$var(f_t - s_t) = var(E_t(s_{t+1} - s_t)) + 2cov(P_t, E_t(s_{t+1} - s_t)) + var(P_t) > 0$$

Thus, the meaning that $\hat{\beta} < 0$ should be consistent with the finding that the risk premium accounts for more variability of $f_t - S_{t+1}$ than that of the prediction error.

Appendix B The Gibbs sampler algorithm

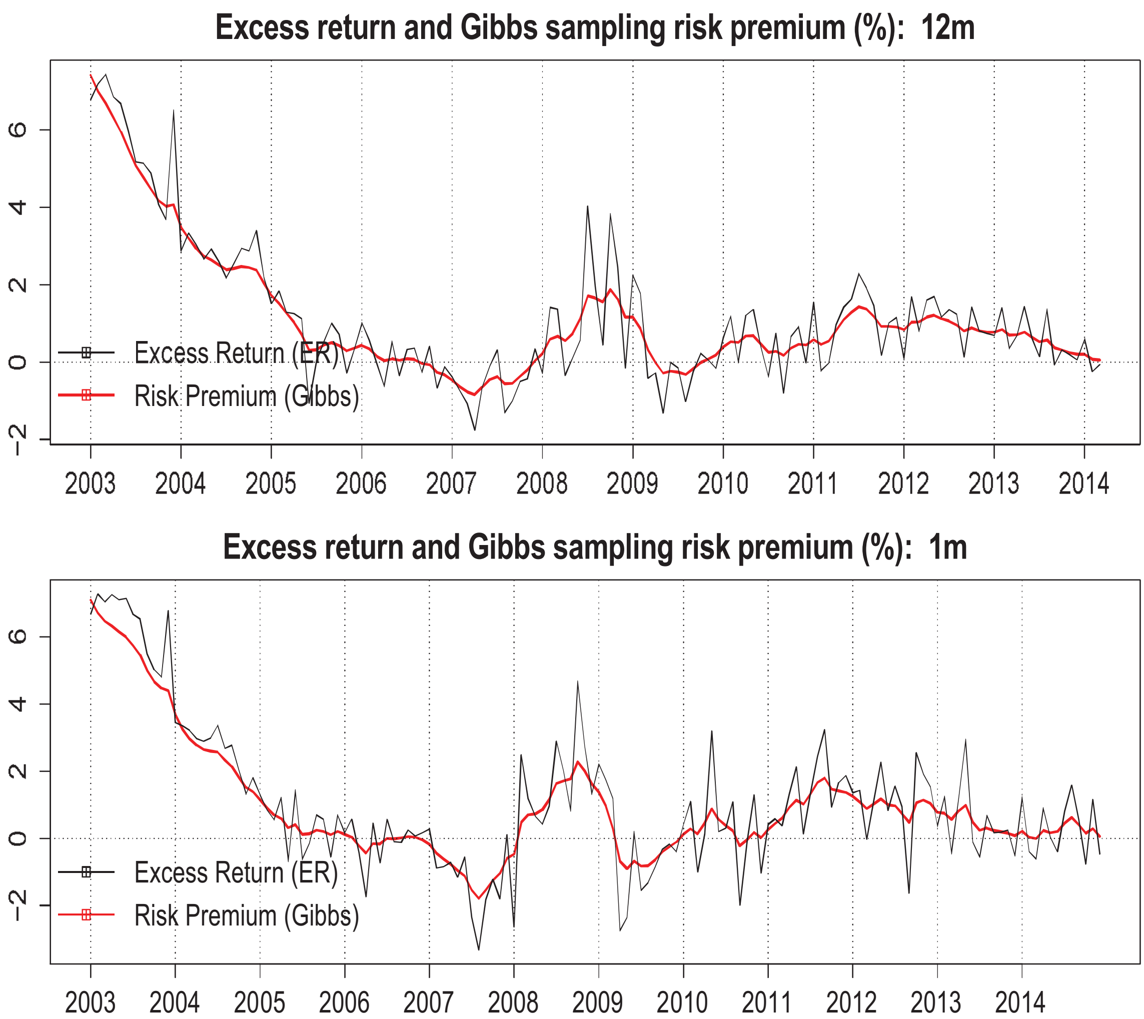
In this appendix we assume the following general state-space model:

$$y_t = \beta_t + \epsilon_t \quad t = 1, \ldots, T \tag{B1}$$

$$\beta_{t+1} = \phi\beta_t + u_t \quad t = 1, \ldots, T \tag{B2}$$

where $\epsilon_t \sim N(0, \sigma^2)$ and $u_t \sim N(0, \sigma_u^2)$.
Stage 0 (Initialization) - Run the Kalman filter with the maximum likelihood estimates of the unknown parameters and attain the updated $(\hat{\beta}_{t/t})$ and smoothed $(\hat{\beta}_{t/T})$ estimates of state variable β_t and its covariance matrices $P_{t/t}$. Estimate $\sigma_2(\hat{\sigma_2})$ and $\sigma_u^2(\hat{\sigma_u^2})$ with the sample variance of $y_t - \hat{\beta}_{t/T}$ and by regressing $\hat{\beta}_{t+1/T}$ on $\hat{\beta}_{t/T}$ and retaining the variance maximum likelihood estimates , respectively.
Stage 1 (Kalman filter step) - Draw from the Normal distribution with mean $\hat{\beta}_{T/T}$ and with variance $(P_{T/T})$. Denote this draw by $\hat{\beta}_T$. At times $t = T-1, \ldots, 1$ draw $\hat{\beta}_t$ from the normal distribution with mean $\hat{\beta}_{t/t} + P_{t/t}\hat{\phi}(\hat{\phi}^2 + \hat{\sigma}_u^2)^{-1}(\hat{\beta}_{t+1} - \hat{\phi}\beta_{t/t})$ and variance $P_{t/t} - P_{t/t}\hat{\phi}(\hat{\phi}^2 + \hat{\sigma}_u^2)^{-1}\phi P_{t/t}$.
Stage 2 (Linear Regression Step) - Conditional upon $\hat{\sigma_u^2}$ and $\hat{\beta}_1, \ldots, \hat{\beta}_T$ we first sample $\hat{\phi}$ from a normal distribution with mean $(1+\frac{1}{\sigma_u^2}\hat{\beta}^{lag'}\hat{\beta}^{lag})^{-1}(\frac{1}{\sigma_u^2}\hat{\beta}^{lag'}\hat{\beta})$ and variance $(1+\frac{1}{\sigma_u^2}\hat{\beta}^{lag'}\hat{\beta}^{lag})^{-1}$. By this we assume that the prior distribution of $\hat{\phi}$ is $N(0, 1)$. Next, we sample σ_u^2 from $\Gamma(1, \frac{T}{2}, D1)$ where T is the number of observations in the sample and $D1 = 0.1+(y-\hat{\beta})'(y-\hat{\beta})$ and σ_u^2 from $\Gamma(1, \frac{T}{2}, D2)$ where $D2 = 0.1 + (\hat{\beta} - \hat{\phi}\hat{\beta}^{lag})'(\hat{\beta} - \hat{\phi}\hat{\beta}^{lag})$.
We iterate on these two steps 500 times, burn the first 200 and average upon 300 samples of states and parameters. The estimation results are displayed in Figure 16.

Figure 8. Risk premium based on Kalman filter versus Excess return



Appendix C Estimation of a time varying parameter (TVP) in the UIP regression

The β from the UIP Fama regression was found in our study as in many other studies to be insignificant giving rise to the "forward premium puzzle". This finding and other tests rejecting the hypotheses of a stable β give rise to estimation of a time varying β. A very naive approach of estimation is by means of a rolling regression. This method of estimation yields as in Baillie and Cho (2014) non smooth estimates of the slope coefficient (β_t) and correspondingly large confidence intervals. In this paper we use an alternative method suggested by Baillie and Cho (2014) which yields smoother estimates of the slope coefficient and also relatively tighter bounds. This method is based on a kernel regression which assigns higher weights to close observations and decaying weights to remote ones. By this approach we estimate the slope coefficient by the following kernel estimator,

$$\hat{\beta}_t = (\sum_{k=1}^{T} w_{kt} x_k x_k')^{-1} (\sum_{k=1}^{T} w_{kt} x_k y_k)$$

where $w_{kt} = K(\frac{t-k}{H})$, y_t is the spot return at time t i.e. $s_{t+1} - s_t$ and $x_t' = [1, (f_t - s_t)]$ $t = 1, \ldots, T$. We take the Gausian kernel with infinite support and $H = 6.5$. The estimator of the variance of the TVP is given by,

$$Var(\hat{\beta}_t) = \hat{\sigma}_u^2 (\sum_{k=1}^{T} w_{kt} x_k x_k')^{-1} (\sum_{k=1}^{T} w_{kt}^2 x_k x_k') (\sum_{k=1}^{T} w_{kt} x_k x_k')^{-1}$$

where $\hat{\sigma}_u^2 = \frac{1}{T} \sum_{t=1}^{T} (y_t - x_t' \hat{\beta})^2$. As mentioned above, the TVP is smoother than the rolling regression β since the TVP is a weighted average of all observations which assigns decaying weights to remote observations while the rolling regression assigns equal weights to observations within a moving window. This causes a jump whenever a large figure enters or exits the moving window.

Appendix D Non-stationary Granger causality test of Toda and Yamamoto (1995)

The procedure suggested by Toda and Yamamoto (1995) for a Granger causality test when at least one of the two tested series is non- stationary, includes the following steps:
1. Let the maximum order of integration for X and Y be m.
2. Set up a VAR model in the levels of the data (X and Y).
3. Determine the appropriate maximum lag length for the variables in the VAR, say p, using the usual information criteria, such as Akaike Information Criterion (AIC).
4. Make sure that the VAR is well-specified e.g., ensure that there is no serial correlation in the residuals. If needed, increase p until there is no serial correlation in the residuals.
5. Add to the preferred VAR model m additional lags of each of the variables into each of the equations so that the model includes p + m lags.
6. Test the hypothesis that the coefficients of (only) the first p lagged values of X are zero in the Y equation, using a standard Wald test. Then do the same thing for the coefficients of the lagged values of Y in the X equation. The coefficients for the 'extra' m lags should not be included while performing the Wald tests as they are intended to fix up the test asymptotic in the VAR model.
The Wald test statistics will be asymptotically chi-square distributed with p d.o.f., under the null. Rejection of the null implies a rejection of Granger non-causality (H_0). That is, a rejection supports the presence of Granger causality. If the X and Y are co-integrated, then there must be Granger causality between them - either one-way or in both directions. However, the converse is not true.

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