The "Forward Premium Puzzle" and the Sovereign Default Risk

Virginie Coudert[1] and Valérie Mignon[23]

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ABSTRACT

Carry-trade strategies which consist of buying forward high-yield currencies tend to yield positive excess returns when global financial markets are booming, whereas they generate losses during crises. Firstly, we show that the sovereign default risk, which is taken on by investing in high-yield currencies, may increase the magnitude of the gains during the boom periods and the losses during crises. We empirically test for this hypothesis on a sample of 18 emerging currencies over the period from June 2005 to September 2010, the default risk being proxied by the sovereign credit default swap spread. Relying on smooth transition regression (STR) models, we show that default risk contributes to the carry-trade gains during booms, and worsens the losses during busts. Secondly, we turn to the "Fama regression" linking the exchange-rate depreciation to the interest-rate differential. We propose a nonlinear estimation of this equation, explaining the puzzling evolution of its coefficient by the change in the market volatility along the financial cycle. Then, we introduce the default risk into this equation and show that the "forward bias", usually evidenced by a coefficient smaller than unity in this regression, is somewhat alleviated, as the default risk is significant to explain the exchange-rate change.

JEL Classification: G15, G01, C3.

Keywords: Carry trades; UIP puzzle; Default risk; Smooth transition regression models.

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[1] Bank of France, Financial Stability Directorate, and EconomiX-CNRS, University of Paris Ouest, France. Email: virginie.coudert@banque-france.fr. Postal address: Bank of France, code 35-1537, 31 rue Croix des Petits Champs, 75001 Paris, France. Tel: 33142924292, Fax: 33142924867.

[2] EconomiX-CNRS, University of Paris Ouest, and CEPII, France. Email: valerie.mignon@u-paris10.fr.

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1. INTRODUCTION

The existence of excess returns on the exchange rate markets is one of the most famous puzzles in international economics. It is a breach in the uncovered interest parity (UIP), and as such referred to as the 'UIP puzzle'. It also contradicts the forward rate being the market rational expectation, justifying the denomination of "forward premium puzzle". Indeed, stylized facts have long shown that interest differentials do not compensate for currency changes observed in the next period. This puzzle has been widely documented in the economic literature (for a survey, see Engel, 1995; Obsteld and Rogoff, 2000) and still gives rise to extensive research (Chaboud and Wright, 2005; Sarno, 2005; Chinn, 2006; Clarida et al., 2009).

This is a logical consequence of the findings by Meese and Rogoff (1983) that exchange rates follow a random walk. If exchange rates are expected to stay at their current values, then it is rewarding to invest in high-yield currencies by borrowing low-yield currencies, as one can expect to earn the interest-rate differential without losing on the exchange rate. For the same reason, it is often profitable to buy forward a currency with a forward discount. Empirical evidence shows that currencies with high interest rates generally do not depreciate as much as the UIP would imply. On the contrary, they often appreciate slowly. Conversely, currencies with low interest rates do not appreciate as much as they should do if the UIP held; instead, they even tend to depreciate. This "forward bias" has been evidenced by numerous studies on various samples of currencies and periods since Froot and Thaler (1990) (Lustig and Verdelhan, 2007; Burnside et al., 2008). Only Chinn (2006) has shown that the UIP could be restored at least for some currencies when considering long-run horizon of 5 years.

Carry trades—which consist in taking long unhedged positions on high-yield currencies and short positions on low-interest rate currencies—have become an important activity on the

exchange markets (Fan and Lyons, 2003; Gagnon and Chaboud, 2007; Galati et al., 2007). These risky strategies are especially popular in periods of "booms" in global financial markets, when investors' risk appetite is high and volatility is low; in these times, the built-up of carry-trade positions helps to strengthen the high-yield currencies. (Brunnermeir et al., 2008; Clarida et al., 2009). The situation is reversed during "bust" periods, as carry-trades are abruptly unwound under adverse financial market conditions (Kohler, 2010; Coudert et al., 2011). As risk aversion increases during crises, investors sell the high-yield currencies, which sharply depreciate and turn to –the low-interest currencies, seen as safe-havens, which appreciate consequently (Ranaldo and Söderlind, 2007; McCauley and McGuire, 2009). In the long run, despite the severe losses occurred during the reversals, carry trades are found profitable.

The existence of these excess returns has been explained by different theories, going from market frictions, namely transaction costs (Huisman et al., 1998) to a "career risk hypothesis" (Liu and Sercu, 2009) in which the traders on the forex market are incited to shun any cash loss from going against the flow. Indeed, as soon as investors are risk-averse, the expected returns in different currencies have no reasons to equalize. The UIP does not hold, as investors require a risk premium to hold the most risky currencies. This idea was already present in the portfolio models of the seventies (Frankel, 1973), in which the risk premia depended on the investors' positions. It is now apprehended through stochastic discount factors (SDF) in the framework of asset pricing-based models (Cochrane, 2001), which have been used to address the "equity premium puzzle" as well (Mehra and Prescott, 1985, 2003). For example, Verdelhan (2010) uses a habit-based preferences model linking the excess returns on the forex market to the domestic consumption growth shocks under assumptions regarding the pro-cyclicality of real interest rates and time-varying risk aversion.

Another possible explanation for the puzzle goes through the "peso problem": the exchange rate risk premia would compensate agents for extremely negative returns, that they are exposed to with very low probabilities. Indeed, carry trades can be hedged against a sharp depreciation of their investment currency by buying a put option (or against the appreciation of the funding currency by a call). Holding a portfolio containing a carry trade and an option allows investors to hedge against negative events, while profiting from excess returns in the other states of nature. Using at-the-money option prices, Burnside et al. (2008) show that this portfolio also yields excess returns, not very different from those of the unhedged carry trades. This result is quite challenging, and the authors interpret it as a sign of a peso problem, coming from high SDF in the extremely unfavourable states of nature.

Here, we suggest a different interpretation of the forward premium puzzle based on the default risk. Indeed, an asset in a high-yield currency generally includes two types of risk: an exchange-rate risk and a default risk. Whereas the exchange-rate risk has been extensively studied, the default risk is generally neglected in the literature on carry trades. However, the recent global crisis has shown that this risk is not negligible as banks may collapse. A sovereign default can also spark a default of the whole banking system of the country and lead to the impossibility for investors to recover the totality of their funds. Our aim in this paper is thus to shed light on the default risk that investors implicitly take on when implementing carry trades. To this end, we focus on emerging countries as they are the most likely to present a significant default risk.

In this framework, after showing that the exchange-rate premium is positively linked to the default risk, we aim at investigating the hypothesis that the default risk of a country's banks tends to increase excess returns on its currency in boom periods, while it worsens the carry trade losses during busts. We test for this hypothesis empirically on a sample of daily data for

the gains of carry trade invested in 18 emerging currencies and funded in USD from June 2005 to September 2010. To avoid arbitrary dating boom and bust periods, we rely on ~~a~~ smooth-transition regression (STR) models which account for nonlinearities across the financial cycle by using the VIX as a transition variable. We show that default risk contributes to increase the carry-trade gains during low-volatility periods, whereas it deepens the losses, once volatility has exceeded a certain threshold.

To verify that the above results do not come from a correlation between interest rate and default risk, we rely on the "Fama regression" linking the exchange-rate depreciation to the interest-rate differential. In this context, one contribution of the paper is to provide a nonlinear estimation of the Fama regression along the financial cycle, using a STR model with the VIX as a transition variable. Another contribution is to introduce the default risk in it. The "forward bias", usually evidenced by a coefficient smaller than unity in this regression, is somewhat alleviated, as the default risk is significant to explain the exchange-rate change.

The rest of the paper is organized as follows. Section 2 shows that the exchange-rate premium increases with the default risk. Section 3 presents the data and some descriptive statistics about the gains of carry-trades invested in emerging currencies. Section 4 gives econometric estimations of the relationship between the carry-trade gains and the default risk. Section 5 provides nonlinear estimations of the Fama regression according to the volatility in financial markets and proposes an "augmented Fama regression" by introducing the default risk into the equation. Section 6 is devoted to some robustness checks regarding our results, and Section 7 concludes the paper.

2. INTERPRETING THE FORWARD BIAS

2.1. The exchange-risk premium

Let us take the point of view of a risk-averse US resident investing in a foreign currency from t to t+1 while funding in USD. There are no restrictions in capital movements between the two countries, no risk of default in both countries. Her foreign currency investment gives the following ex post excess return in USD, denoted r_{t+1} and expressed as follows:

$$(1+r_{t+1}) = \frac{(1+i_t)}{(1+i_t^{US})}\frac{S_t}{S_{t+1}} \tag{1}$$

where i_t is the interest rate in the foreign country, i_t^{US} the US interest rate and S_t is the exchange rate of the foreign currency against the USD, measured as the number of units of the foreign currency for one dollar.

This excess return is the yield of the carry trade invested in the foreign currency, while funded in USD. The exchange-rate risk premium ρ_t^S is defined as this expected excess return on the foreign currency:

$$(1+\rho_t^S) = E_t\left[\frac{1+i_t}{1+i_t^{US}}\frac{S_t}{S_{t+1}}\right] = E_t\left(1+r_{t+1}\right) \tag{2}$$

where $E_t\left[X_{t+1}\right]$ stands for the agents' expectation on X_{t+1} at time t.

As the covered interest rate parity holds, the exchange-rate premium is equivalent to the expected return of buying the foreign currency forward and selling it in the next period on the spot market. In other words, the exchange-rate premium is equal to the gap between the forward rate F_t and the expected spot rate for the next period:

$$(1+\rho_t^S) = E_t\left[\frac{F_t}{S_{t+1}}\right] \tag{2'}$$

As all risk premiums, in the framework of an asset pricing model (see Cochrane, 2001), the exchange-rate premium can be written as:

$$\rho_t^S = -Cov\left(r_{t+1}^S, m_{t+1}\right)\frac{1}{E_t\left[m_{t+1}\right]} \tag{3}$$

where m_t is the stochastic discount factor (SDF), which depends on the inter-temporal utility function of a representative agent.

A positive risk premium, $\rho_t > 0$, is the standard situation for a risky asset as its returns are negatively correlated with the SDF. Indeed, risky assets normally yield positive returns in the favourable states of nature with low SDF—typically during periods of booming financial markets—and, conversely, provide low or negative returns in the unfavourable states of nature, which have high SDF—typically during financial crises when all asset prices plummet. As the exchange rate is a relative price between two countries' assets, the risk premium is symmetrical. This means that if some currencies have positive risk premiums, others necessarily have negative ones.

Equation (3) states that currencies bear positive exchange risk premiums, if their returns are negatively correlated to the SDF. A simple example of SDF is linked to market returns, like in the capital asset pricing model (CAPM). In this case, the risk premium of an asset is proportional to the correlation of its returns (r^S in our case) with market returns (r^M):

$$\rho_t^S = \mathrm{cov}\left(r_{t+1}^S, r_{t+1}^M\right)\frac{1}{E_t\left[r_{t+1}^M\right]} \tag{4}$$

Consequently, currencies which yield positive returns when global markets are bullish and depreciate during bear markets, should offer positive risk premiums to compensate investors for their risks. These positive correlations with market returns characterize most emerging countries' currencies, as this will be illustrated in Section 3.

Conversely, currencies which are able to offer returns negatively correlated to market returns have negative risk premia. These latter currencies can be viewed as safe havens, as they tend to appreciate in the unfavourable states of nature (such as the Japanese yen or the Swiss franc for example).

2.2. Introducing default risk

We now consider a one-period investment in the foreign country when a default is possible. The default will occur during this period with a given probability p, $0< p <1$. In this case, the ex post return of this investment is:

$$\begin{cases} 1+r_{t+1} = \dfrac{1+i_t}{1+i_t^{US}} \dfrac{S_t}{S_{t+1}} & \text{if no default} \\ 1+r_{t+1} = \mu \dfrac{1+i_t}{1+i_t^{US}} \dfrac{S_t}{S_{t+1}} & \text{if default} \end{cases} \tag{5}$$

where μ is the recovery rate, $0<\mu<1$.

The total risk premium on this investment, denoted ρ_t^T, now stems from two factors: the exchange-rate risk premium ρ_t^S, as in the previous section, and the default risk premium. It can be expressed as:

$$(1+\rho_t^T) = E_t\left[\frac{1+i_t}{1+i_t^{US}} \frac{S_t}{S_{t+1}} \left[1 - I_{t+1}(1-\mu)\right]\right] \tag{6}$$

where I_t is the indicator function of default (equal to 1 in case of default, 0 otherwise). Replacing the exchange-rate premium ($1+\rho_t^S$) by its value as in Equation (2), we can re-write Equation (6) as the following:

$$(1+\rho_t^T) = (1+\rho_t^S)\left[1 - p(1-\mu)\right] + (1-\mu)Cov_t\left[(s_{t+1} - s_t), I_{t+1}\right] \tag{7}$$

Let us remark that the quantity $p(1-\mu)$ stands for the expected loss on the investment due to default, at constant exchange rates, which can be thought of as the default risk premium, which we denote *d*:

$$d = p(1-\mu) \tag{8}$$

The covariance term can be expressed as:

$$Cov_t\left[(s_{t+1}-s_t), I_{t+1}\right] = p\left[E_t\left[(s_{t+1}-s_t)/I_{t+1}=1\right] - E_t[s_{t+1}-s_t]\right]$$

where $E_t\left[(s_{t+1}-s_t)/I_{t+1}=1\right]$ stands for the expected depreciation given default.

The exchange-rate premium can therefore be written as:

$$(1+\rho_t^S) = \frac{(1+\rho_t^T) - d\left[E_t\left[(s_{t+1}-s_t)/I_{t+1}=1\right] - E_t(s_{t+1}-s_t)\right]}{1-d} \tag{9}$$

Equation (9) shows that the exchange risk premium is a positive function of the default risk premium *d* under reasonable assumptions on the parameters. (i) If the investment in the foreign currency is considered as risky, theoretically, the total risk premium ρ_t^T should be positive. (ii) The covariance term is also positive as a default generally sparks a depreciation in the currency, however it is likely to be much smaller than 1.[4] Hence, ρ_t^S is a positive function of *d*. In other words, the exchange-risk premium positively depends on the default premium.

[4] It even stays smaller than 1 under unrealistically pessimistic assumptions. For instance if *p*=30%, the depreciation in case of default at 80% (above the unconditional depreciation) and μ=30%, the covariance term (0.84) is still smaller than 1.

2.3. Risk-neutral framework

To better clarify the link between the exchange-rate premium and the default risk, let us consider a risk-neutral approach. In this framework, the total risk premium ρ_t^T on the foreign investment is null. As investors disregard the risk, there are no expected excess returns. Therefore, Equation (6) would be expressed as:

$$1 = E_t^{RN}\left[\frac{1+i_t}{1+i_t^{US}}\frac{S_t}{S_{t+1}}\left[1 - I_{t+1}(1-\mu)\right]\right] \tag{10}$$

where $E_t^{RN}\left[X_{t+1}\right]$ is the expected value of X_{t+1} under the risk-neutral distribution.

In the risk-neutral approach, the foreign exchange premium still exists, but only to compensate risk-neutral investors for the expected loss due to a possible default:

$$(1+\rho_t^{RN,S}) = \frac{1}{1-d}\left[1 - d\left[E_t^{RN}\left[(s_{t+1} - s_t)/I_{t+1} = 1\right] - E_t^{RN}\left[(s_{t+1} - s_t)\right]\right]\right] \tag{11}$$

where $\rho_t^{RN,S}$ stands for the risk-neutral exchange-rate premium. In this framework, the exchange-rate premium only stems from the possibility of default.

3. DATA AND DESCRIPTIVE STATISTICS

3.1. Data to calculate excess returns

We consider the realised gains in investing in emerging currencies, while funding in USD.[5] The sample is made of the currencies of the main emerging countries for which there is (i) capital mobility, hence the possibility of carry trades, which excludes countries like China for example, and (ii) an active sovereign CDS market, hence daily data for their premia. It includes 18 currencies in Latin America, those of Argentina (ARS), Brazil (BRL), Chile

[5] Other funding currencies have also been considered as robustness checks (see Section 6).

(CLP), Colombia (COP), Mexico (MXN), Peru (PEN); in Asia, Indonesia (IDR), Korea (KRW), Malaysia (MYR), Philippines (PHP), Thailand (THB); in emerging Europe, the Czech Republic (CZK), Hungary (HUF), Poland (PLN), Romania (RON), as well as the Russian rouble (RUB), the Israel shekel (ILS) and the South African rand (ZAR).

Some currencies in the sample are traded on small foreign exchange markets, a characteristic that may question the relevance of carry trades in these countries. However, these currencies may still be attractive destinations for carry trades, due to high interest rate differentials. For example, Hoffmann (2011) showed that carry trades were lucrative in Romania before the recent world crisis, due to both high interest rates and stable exchange-rate volatility.

For all these currencies, we consider the one-year change in their exchange rate versus dollar (in logarithms), and the one-year interest rate taken in difference with the one-year interest rate in the US, all series being extracted from Bloomberg. The choice of a one-year horizon is justified in Section 3.2 below.

3.2. Gauging the situation on financial markets and the default risk premium

To gauge the situation on the global financial markets, we use the VIX, extracted from Bloomberg. This indicator measures the implied volatility of the S&P500 index options for the next 30 days, calculated by the Chicago Board Options Exchange (CBOE). Strong volatility spillovers in stock markets all around the world make the situation on the S&P500 representative of the global financial markets. More specifically, the VIX is an indicator for the investors' attitude towards risk not only on equity markets, but also on credit markets (Collin-Dufresne et al., 2001), and more generally a good gauge for risk aversion (Coudert and Gex, 2008). As highlighted by Brunnermeier et al. (2009), it is also linked to various

other markets, such as the changes in the risk premia in sovereign CDS[6], and many of the financial crises all around the world were accompanied by strong increases in the VIX. This variable is thus a good proxy for identifying booms and busts on global financial markets (see Becker et al., 2009). Low volatility on the stock market matches boom periods, whereas high volatility typically goes with crises.

Concerning the default risk on an investment in a given currency, we proxy it by the sovereign credit default swap (CDS) premium of the country. The sovereign default is able to represent the default of the country's banks, as both credit events are highly correlated because of the close financial interactions between the two types of agents. As the CDS is aimed at hedging the default risk, its premium is a good proxy for the default risk premium; it is roughly equal to the expected loss due to default $d = p(1-\mu)$ in Equation (8).

We extract the one-year sovereign CDS premia from Datastream. We choose a one-year horizon which is the smallest maturity for the CDS data. Of course, to compare this default risk to the exchange-rate risk, we have to consider the same one-year maturity for interest rates and exchange-rate changes.

3.3. Periods

The sample spans from June 2005 to September 2010 on a daily periodicity. This period includes three different phases of the financial cycle:[7]

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[6] See also Acharya et al. (2011) who show that financial sector bailouts and sovereign credit risk are closely linked.

[7] This decomposition of periods may seem somewhat arbitrary. Note that it will only be used to present descriptive statistics and linear regressions on the carry-trade gains in the following sections. Both the stylized facts and the linear regressions only aim at giving an insight of the reversals occurring along the financial cycle. Our main results will not depend on this time decomposition, as they will be obtained through nonlinear estimations over the whole period.

- a pre-crisis period of booming financial markets. We date this period from the beginning of the sample (1st June 2005) up to the Friday 12 September 2008, the working day before the failure of Lehman Brothers;
- the global financial crisis, from 15 September 2008 to 30 September 2009. We consider that global financial crisis started the day after the Lehman Brothers bankruptcy, on the 15 September 2008. Indeed, the banking crisis had already been brewing in the US and European countries for a year when Lehman Brothers collapsed. However neither the stock market, nor the forex market of emerging countries had been under fire before this event. This is confirmed by the behaviour of the VIX, which jumped from about 25% up to 80% in a few days after Monday 15 September 2008;
- a recovery period. We date the recovery from 1rst October 2009, as attested by capital flowing back to emerging countries. At that time, Brazil implemented a new tax on capital inflows to counter the appreciation in its exchange rate; most other emerging currencies also appreciated strongly. This date is consistent with the evolution of the VIX that went back to its pre-crisis level precisely at that time.

3.4. Descriptive statistics on carry-trade gains

We now take a look at the returns of the carry trades invested in the different currencies of the sample and funded in USD. As no default occurred over the sample period, we calculate the carry-trade gains by using Equation (1). Table 1 displays these one-year carry-trade excess returns for each of the currencies in the sample, along with their average interest-rate differentials with the US, their one-year depreciation against USD and volatility.

During the pre-crisis period when global markets were booming, the excess returns are positive for all currencies (except South Africa). Over this period, carry trades yielded 8.1% per year on average; that came from both positive interest-rate differentials (2.1% on average) and currency appreciation (5.8% on average). The highest excess returns on this period are obtained by the Brazilian real, which yielded 26.2% per year (11.6% from the interest differential and 14.7% from the currency appreciation). On the top of being so profitable, carry-trade gains also exhibited a lower volatility in this boom period.

Strikingly, all the excess returns suddenly fell to very negative values during the crisis (12.9% on average). Interest-rate differentials stayed positive and even slightly increased on average, but currencies depreciated dramatically, losing 16.3% on average. Meanwhile, their volatility soared in most cases. The recovery period is characterized by a renewal of carry-trade gains. Returns sharply rallied for all currencies, reaching 11.8% in the year following the crisis, due to both very positive interest-rate differentials (5.1% on average) and currency appreciation (6.6% on average).

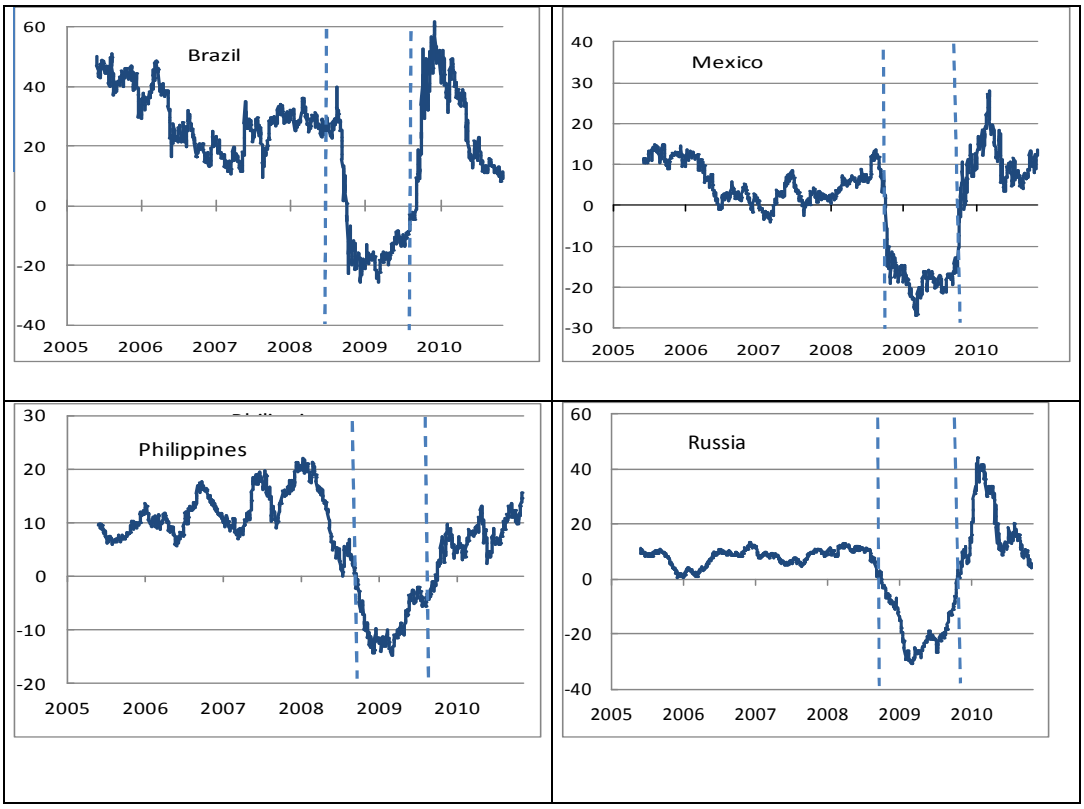
Table 1. Average interest-rate differentials with the US, one-year change in the exchange rates vs. USD, and one-year carry-trade excess returns, in %

	Pre-crisis period				Crisis period				Recovery period			
	01/06/2005-12/09/2008				15/09/2008-30/09/2009				1/10/2009-30/09/2010			
	Interest rate differential	Exchange rate change	Excess returns	Volatility	Interest rate differential	Exchange rate change	Excess returns	Volatility	Interest rate differential	Exchange rate change	Excess returns	Volatility
Argentina	5.5	1.7	3.9	3,0	11.4	13.8	-2.5	7,5	19.2	7.1	12.1	6,6
Brazil	11.6	-14.7	26.2	7,3	8.2	20.7	-12.5	9,9	9.2	-18.2	27.4	16,7
Chile	-4.0	-6.3	2.3	6,1	-2.9	18.2	-21.1	10,0	-1.7	-12.5	10.9	12,9
Colombia	3.0	-8.1	11.1	8,7	5.6	15.6	-10.0	10,2	5.9	-14.7	20.7	13,5
Mexico	3.7	-1.9	5.6	3,8	4.4	22.9	-18.5	7,2	4.4	-5.4	9.8	11,4
Peru	0.0	-4.1	4.1	4,7	2.1	5.0	-2.9	4,8	0.2	-7.4	7.6	6,0
Korea	-0.1	-3.0	2.9	5,0	2.4	28.6	-26.2	4,5	2.2	-12.6	14.8	6,1
Indonesia	5.9	-0.1	6.0	5,8	5.6	14.4	-8.9	7,6	8.3	-15.4	23.7	11,2
Malaysia	-1.1	-4.4	3.4	2,7	0.2	7.8	-7.6	4,1	0.6	-7.9	8.5	6,1
Philippines	3.3	-7.7	11.0	5,1	2.5	10.4	-8.0	4,6	2.5	-4.5	7.1	8,4
Thailand	-0.4	-5.7	5.4	6,5	0.2	4.5	-4.2	11,9	0.3	-6.4	6.8	16,2
Czech Rep.	-1.8	-10.8	9.1	3,4	0.6	13.1	-12.5	11,2	1.1	-2.7	3.8	13,2
Hungary	3.4	-4.8	8.2	7,5	4.7	20.3	-15.6	7,1	7.5	-1.5	9.0	14,2
Poland	0.7	-10.8	11.5	9,7	2.7	27.6	-24.9	11,3	3.1	-4.2	7.3	13,7
Romania	4.8	-7.5	12.3	11,0	6.5	23.3	-16.8	10,3	10.8	2.0	8.8	15,0
Russia	3.0	-4.9	7.9	9,0	4.0	24.0	-19.9	15,8	13.7	-3.5	17.2	18,5
Israel	0.6	-5.9	6.6	7,2	1.2	7.7	-6.5	16,1	0.1	-4.4	4.4	15,5
South Africa	4.0	5.5	-1.5	8,2	8.9	18.3	-9.4	10,0	7.2	-18.3	25.5	9,6
Average	2.2	-5.8	8.1	6,4	3.5	16.3	-12.9	9,1	5.1	-6.6	11.8	11,9

Source: authors' calculations. Excess returns = interest rate differential – exchange-rate change. "Average" is the unweighted average.

Figure 1 depicts the pattern in the carry-trade gains for four countries: Brazil, Mexico, Philippines and Russia. The four charts show how abrupt the reversal of carry-trade gains was during the crisis. Returns became negative only a few days after the Lehman Brothers' bankruptcy. They turned positive again in late 2009, as capital flowed back to emerging countries.

Figure 1. Returns on a one-year carry trade funded in USD and invested in local currency, in %

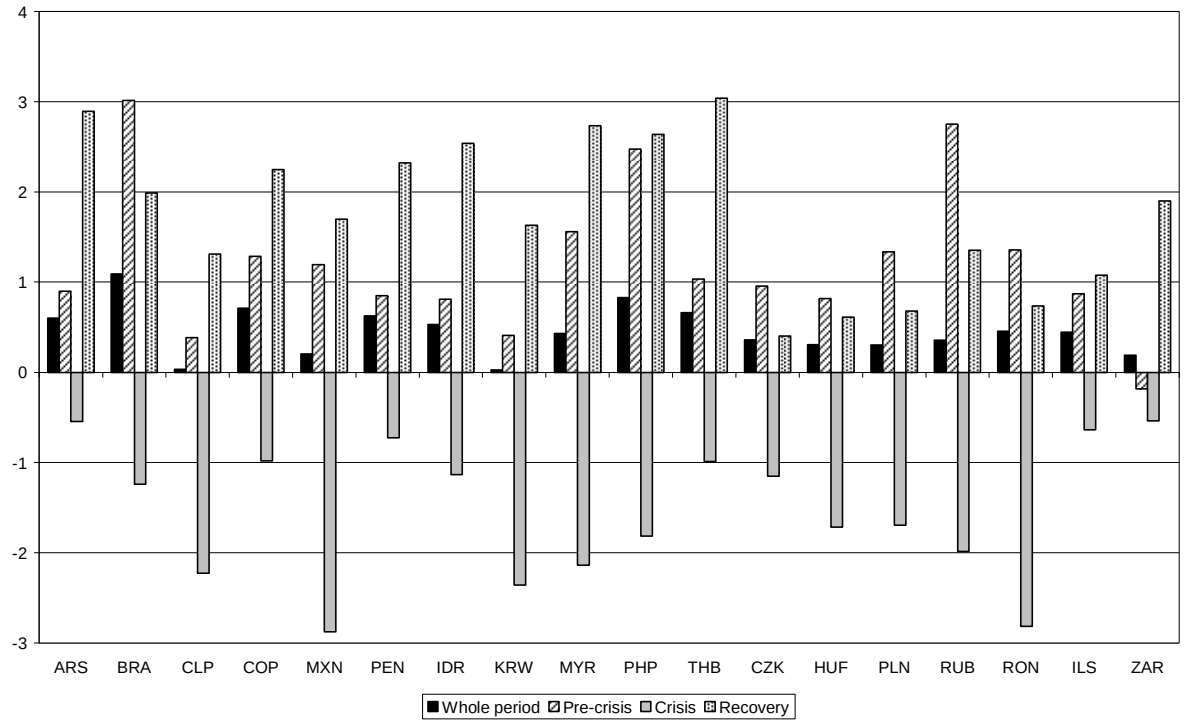


Note: Dotted vertical lines indicate the start and the end of the crisis period. Source: authors' calculations.

The relevance of the carry-trade strategies can also be observed through the Sharpe ratios, that compare the excess returns to their volatility (Jordà and Taylor, 2009). As shown in Figure 2, Sharpe ratios are positive over the whole period as well as during the pre-crisis and recovery periods, whereas negative during the crisis. Latin American currencies were very attractive during the pre-crisis period considering these ratios; the Brazilian real is still one of most rewarding currency along with the Thailand's bath on this criterion. Strikingly, the Sharpe ratios are even higher during the recovery period than they were before the crisis for 10 countries in 11 in Latin America and Asia. Sharpe ratios are boosted by the sharp appreciation of these currencies in the immediate aftermath of the crisis, which more than cancelled the adverse effect of a higher volatility. Only for Central and Eastern European markets Sharpe

ratios tend to be lower on the last period, due to persistent depreciation of their currencies and important volatility.

Figure 2. Sharpe ratios



Source: authors' calculations.

4. ESTIMATING THE RELATIONSHIP BETWEEN CARRY-TRADE GAINS AND DEFAULT RISK

The stylized facts highlighted above, as well as the theoretical relationship in Equation (9), suggest a positive correlation between the carry-trade gains invested in emerging currencies and global market returns. Indeed, ex post carry-trade gains tend to be positive on average and to co-move with market returns, as evidenced by former studies (Clarida et al., 2009). This implies a positive exchange-risk premium for these currencies as stated by Equation (4).

Now, we want to test empirically for the hypothesis that the exchange-rate premium increases with the default risk, as established in Equation (9), i.e. that the correlation of realized carry-

trade gains with market returns increases with default risk. To check for this hypothesis, we aim at empirically verifying the two following propositions:

(P1) "The carry-trade gains observed during periods of booming financial markets are higher when invested in countries with higher default risk".

(P2) "Conversely, the losses on carry trades during crises are deeper for high default risk currencies".

Hence the excess returns of carry trades depend on default risk in a nonlinear way. The relation between both variables is positive as far as financial markets are booming, i.e. that volatility is low, and turns negative when volatility increases above a certain threshold.

4.1. Linear estimations across periods

We first assess the link between the ex post returns on carry trades and the default risk, proxied by the sovereign CDS premia, by estimating the following regression:

$$r_{it+1} = \lambda_i d_{it} + \alpha_i + \varepsilon_{i\,t+1} \tag{12}$$

where r_{it+1} is the ex post excess return in the one-year carry trade invested at time t in currency i (i = 1, …, 18), d_{it} is country i's default risk proxied by its sovereign CDS premium. The estimations are run successively for the 18 currencies in the sample and over the three sub-periods under study: pre-crisis period, crisis, and recovery (see Section 3.3). As carry-trade gains are measured over one year but are observed at a daily frequency, this overlapping structure introduces a moving average component in the errors (Hansen and Hodrick, 1980). Various methods exist for dealing with the overlapping observations' issue (see Harri and Wade Brorsen, 2009), among which: (i) reducing the sample to obtain no overlapping data, (ii) averaging the data, (iii) using heteroskedasticity and autocovariance consistent estimators, such as Hansen and Hodrick (1980) or Newey and West (1987), (iv) using a bootstrap

method. The first two methods are generally not recommended since they are inefficient, strongly reduce the number of degrees of freedom and do not eliminate the moving average component in the errors. The Newey-West's covariance estimate performs quite well on large sample, although it can be biased downward in small samples (Nelson and Kim, 1993). The parametric bootstrap used by Mark (1995) among others has the drawback of being based on the inefficient OLS estimator. Hence, given the large size of our sample (more than 1300 observations), we rely on the Newey and West (1987)'s autocorrelation and heteroskedasticity consistent covariance estimator (Mark, 2001; Britten-Jones et al., 2011).

To establish proposition (P1), we have to find the coefficients λ_i significantly positive when estimated over the booming market period. Conversely for proposition (P2) to hold, we should find the λ_i significantly negative over the crisis.

The results obtained by running regression (12) on each of the 18 currencies lend support to the two propositions. As expected, all coefficients λ_i are significantly positive in the first period (except only for 2 currencies) and also for the recovery period (Table A.1 in the Appendix). Conversely, all coefficients λ_i are significantly negative when estimated over the crisis episode. Consequently, the default risk contributes to amplify the carry-trade gains in the boom period and to worsen the losses during the crisis. This situation prevails even if there was no sovereign default observed over the sample. The carry-trade losses only occurred through the sharp currency depreciation observed during the crises.

4.2. Nonlinearities depending on the volatility on financial markets

As the relation between carry-trade gains and default risk is reversed during crises, it is interesting to study how it evolves along the financial cycle. To do so, we use a smooth transition regression (STR) model, which links the evolution of the coefficient between the

two variables to the volatility of global financial markets measured by the VIX. According to this specification, the carry-trade gains depend nonlinearly on the default risk, their relationship being dependent on the level of the VIX. More specifically, our STR model is given by:

$$r_{it+1} = \lambda_{0i} d_{it} + \alpha_{0i} + [\lambda_{1i} d_{it} + \alpha_{1i}] g(v_t; \gamma_i, c_i) + \varepsilon_{it+1} \tag{13}$$

where ε_{it} is $iid(0, \sigma^2_{i\varepsilon})$, v_t is the VIX which acts as the transition variable and $g(v_t; \gamma_i, c_i)$ is the transition function which by convention is bounded by zero and one. $\gamma_i > 0$ denotes the slope parameter that determines the smoothness of the transition from one regime to the other (*i.e.* the abruptness of the transition dynamics at c_i), and c_i is the threshold parameter.

In this model, two regimes—linear and nonlinear—characterize the dynamics of the carry-trade gains, depending on market volatility. The transition from one regime to the other is smooth, implying that there exists a continuum of states between extreme regimes. Two transition functions are commonly considered (Teräsvirta and Anderson, 1992):

$$g(v_t; \gamma_i, c_i) = (1 + \exp(-\gamma_i (v_t - c_i)))^{-1} \text{: logistic STR model (LSTR)} \tag{14}$$

$$g(v_t; \gamma_i, c_i) = 1 - \exp(-\gamma_i (v_t - c_i)^2) \text{: exponential STR model (ESTR)} \tag{15}$$

The LSTR specification accounts for asymmetric realizations, in the sense that the two regimes are associated with small and large values of the transition variable relative to the threshold value. On the other hand, in the ESTR model, increases and reductions in the transition variable have similar effects, but the middle grounds are characterized by different dynamics. In both cases—LSTR and ESTR—when γ_i goes to zero, the STR process reduces to a linear model. When γ_i tends to infinity, the LSTR model becomes a two-regime threshold model with abrupt transition (Tong, 1990). To specify the STR model, we follow the methodology proposed by Teräsvirta (1994). We first test for linearity and, if the null

hypothesis is rejected, choose between the LSTR and ESTR specifications using the sequential strategy developed by Teräsvirta (1994). Once this choice has been made, we estimate the STR model and apply various misspecification tests: test of no residual autocorrelation (Teräsvirta, 1998), LM-test of no remaining nonlinearity (Eitrheim and Teräsvirta, 1996), and ARCH-LM test (Engle, 1982).[8]

The nonlinear form of the relationship is confirmed as the null hypothesis of linearity is rejected, for all our 18 currencies. Hence, the response of the carry-trade gains to the CDS spread does differ according to the level of financial volatility. Moreover, the LSTR specification is retained for all the currencies. This means that there are two regimes involved. The first one prevails when the volatility is low on financial markets, the VIX being under the estimated threshold c_i. The relationship between carry-trade gains and default risk then only involves the linear parameter λ_{0i}, as the second term in the right-hand side of Equation (13) is close to 0 (g being close to 0). The second, high-volatility regime occurs when the VIX exceeds the threshold, the relationship then also includes the second term of Equation (13) and coefficient λ_{1i} is involved (as well as the sum of λ_{0i} and λ_{1i}).

Several features evidenced by the results reported in Table 2 confirm our former hypothesis. First, when financial volatility is low, the relationship between carry-trade gains and default risk is significantly positive in most cases. This is evidenced by $\lambda_{0i} > 0$ for 13 currencies out of 18. Second, when financial markets get more volatile, the relationship turns negative (for 14 cases in 18), and the default risk then contributes to deepen the losses. Third, the level of the VIX that determines the two regimes varies across countries. As an example, it is around 23%

[8] To save space, results of both linearity and misspecification tests are not reported here but are available upon request to the authors.

for most Latin America countries, meaning that the relation between carry-trade gains and CDS premia starts to be reversed when the VIX crosses the value of 23%. This figure roughly fits our definition of periods given in Section 3, as the crisis is identified when the VIX suddenly jumps above its former level of 25% and ends when it gets back to this value. During the 2008-2009 crisis, as the VIX has suddenly jumped from a value of 25% to nearly 80% in a few days, the transition between the two regimes has been very brutal.

Table 2. Coefficient λ_i in the nonlinear estimations

	Linear coefficient		Nonlinear coefficient		Transition parameters	
Country *i*	λ_{0i}		λ_{1i}		γ_i	c_i
Argentina	1.009	(16.99)	-1.194	(-19.61)	159.67	22.97
Brazil	34.177	(16.08)	-51.700	(-23.59)	12.57	22.51
Chile	42.927	(13.75)	-61.084	(-19.38)	52.66	23.13
Colombia	37.161	(16.47)	-48.603	(-21.63)	8.94	22.84
Mexico	22.039	(10.99)	-32.321	(-15.96)	15.30	21.72
Peru	-5.214	- (-5.30)	-0.435	(-0.44)	39.88	16.38
Indonesia	-27.461	(-15.32)	21.678	(12.03)	210.35	15.55
Korea	144.184	(3.79)	-155.207	(-4.08)	3.30	13.71
Malaysia	19.871	1 (7.34)	-29.757	(-11.09)	6.98	18.72
Philippines	-8.964	(-30.06)	6.941	(13.27)	5.61	36.58
Thailand	-120.091	(-10.02)	112.169	(9.35)	25.84	13.05
Czech Republic	47.255	(4.66)	-71.079	(-7.06)	9.16	17.18
Hungary	24.903	(7.52)	-33.422	(-10.14)	7.37	17.73
Poland	36.296	(5.67)	-56.967	(-8.87)	10.47	19.40
Romania	2.211	(2.33)	-7.054	(-7.14)	8.41	21.89
Russia	-11.445	(-19.74)	9.910	(16.22)	4.30	37.65
Israel	15.469	(5.13)	-30.679	(-10.08)	17.98	18.39
South Africa	0.368	(3.43)	-0.609	(-1.97)	1.81	55.90

Note: This table reports the results from the estimation of Equation (13) $r_{it+1} = \lambda_{0i} d_{it} + \alpha_{0i} + [\lambda_{1i} d_{it} + \alpha_{1i}] g(v_t; \gamma_i, c_i) + \varepsilon_{it+1}$, where r_{it+1} is the ex post excess return in the one-year carry trade, d_{it} the sovereign CDS premium, v_t the VIX (transition variable), g the transition function, γ_i the slope parameter that determines the smoothness of the transition and c_i the threshold parameter. *t* statistics are given in parentheses.

5. REVISITING THE FAMA REGRESSION

As interest rates are highly correlated with default risk, we have to verify whether our results persist if the regression also includes the interest-rate differential. To do so, we first run the "Fama regression" (Fama, 1984), which links the exchange-rate change to the interest differential, and then introduce the default risk in it.

5.1. The standard Fama regression

The forward bias is generally evidenced by estimating a regression between the realized exchange-rate change and the interest-rate differential (Chaboud and Wright, 2005; Chinn, 2006; Clarida et al., 2009). The regression between the two variables is referred to as the "Fama regression" and expressed as followed:

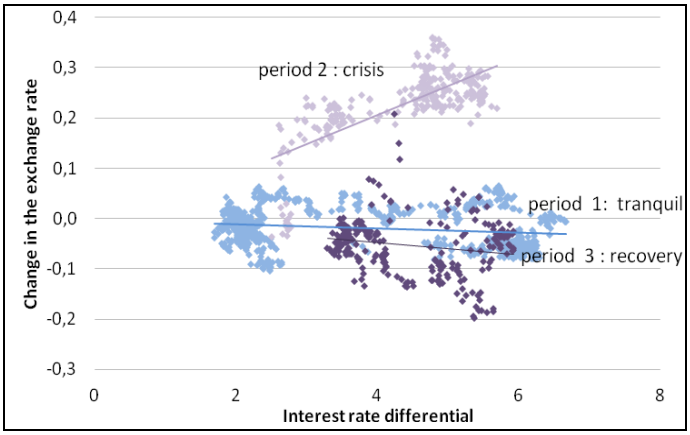
$$s_{i_{t+1}} - s_{i_t} = \beta_i (i_{i_t} - i_t^{US}) + \alpha_i + \varepsilon_{i_{t+1}} \qquad (16)$$

where $s_{i_{t+1}} - s_{i_t}$ stands for the one-year change in country's *i* exchange rate versus dollar (in logarithms), $(i_{i_t} - i_t^{US})$ the one-year interest-rate differential between country *i* and the US; β_i and α_i being parameters to estimate.

If the UIP held, the estimation of Equation (16) would give $\beta_i = 1$ and $\alpha_i = 0$. However, as previous literature has shown,[9] β_i is often found smaller than 1, or even negative. More precisely, Brunnermeier and Pedersen (2009) have shown that the β_i are smaller than 1 during periods of low volatility, and switch to strong positive values during crises. As shown in Table A.2 in the Appendix, the estimations made on our sample of 18 currencies confirm those previous results. Figure 3 illustrates these changes in the value of β across periods for Mexico. The regression line between the interest-rate differential and the currency depreciation has a negative slope in the pre-crisis period as well as during the recovery, whereas the slope turned positive during the crisis.

[9] See Engel (1995) for a survey.

Figure 3. Interest-rate differential and change in the exchange rate, regression lines across the three sub-periods, Mexico



5.2. Estimating the Fama regression with nonlinearities

The previous findings clearly highlight a reversal in the relationship between the exchange-rate change and the interest-rate differential across periods, as the coefficients β_i are suddenly shifted upwards during the crises. This leads us to test for nonlinearities in the Fama regression, by using a STR specification. The exchange-rate change now depends nonlinearly on the interest differential, their relationship fluctuating according to the evolution of market volatility. More specifically, the STR model is given by:

$$s_{it+1} - s_{it} = \beta_{0i}(i_{it} - i_t^{US}) + \alpha_{0i} + \left[\beta_{1i}(i_{it} - i_t^{US}) + \alpha_{1i}\right] g(v_t; \gamma_i, c_i) + \varepsilon_{it+1} \qquad (17)$$

As the null hypothesis of linearity is rejected in favour of the LSTR specification, the exchange rate responds differently to the interest differential according to market volatility. More specifically, as for the estimation of Equation (13), there are two regimes involved: the first one prevails when the volatility is low on financial markets, and the second one intervenes under high volatility conditions—the VIX exceeding a certain threshold.

Several findings stand out from the results (Table 3). First, the coefficient β_{0i} of the interest-rate differential is lower than one for the majority of countries, and even negative for 10

countries in the low-volatility regime. This illustrates the forward bias under low-volatility environment, as high-yield currencies tend to appreciate. Second, the coefficient sharply rises with financial market volatility, i.e. when the VIX exceeds a certain threshold *c;* this is evidenced by β_{1i} being positive and greater than 1 in the high-volatility regime for all currencies (except for the Russian rouble). In other words, the forward bias disappears in high-volatility periods. Third, the threshold value of the VIX that triggers the nonlinear regime is above 25% for all currencies, matching the previous conclusions.

Table 3. Results of the nonlinear estimations of the Fama regression

	Linear	Nonlinear	Transition parameters	
Country *i*	β_{0i}	β_{1i}	γ_i	c_i
Argentina	-7.958 (-3.423)	8.524 (4.831)	2.64	39.57
Brazil	0.512 (4.070)	3.783 (4.026)	3.09	32.14
Chile	-2.520 (-13.183)	6.056 (7.328)	3.46	33.69
Colombia	0.025 (0.170)	5.578 (11.585)	6.93	29.58
Mexico	-0.618 (-4.036)	11.639 (22.504)	5.01	26.18
Peru	-6.588 (-9.573)	9.894 (12.543)	1.17	95.34
Indonesia	-2.032 (-20.180)	5.325 (5.203)	3.91	38.08
Korea	-9.388 (-7.172)	21.970 (11.475)	1.36	27.27
Malaysia	0.969 (7.485)	3.018 (7.253)	4.67	33.40
Philippines	1.240 (8.056)	6.923 (7.760)	1.32	34.79
Thailand	-1.788 (-6.173)	8.329 (12.309)	3.64	30.91
Czech Republic	0.919 (2.901)	12.195 (17.649)	3.83	25.65
Hungary	1.327 (9.180)	3.805 (6.257)	4.49	31.06
Poland	1.242 (5.929)	12.847 (21.612)	4.20	25.98
Romania	0.128 (1.763)	2.995 (9.997)	2.65	28.52
Russia	-0.039 (-0.571)	-0.678 (-2.465)	4.56	28.70
Israel	3.401 (10.545)	9.587 (18.842)	5.10	21.16
South Africa	-6.860 (-9.820)	20.152 (7.797)	1.03	41.28

Note: This table reports the results from the estimation of Equation (17): $s_{it+1} - s_{it} = \beta_{0i}(i_{it} - i_t^{US}) + \alpha_{0i} + \left[\beta_{1i}(i_{it} - i_t^{US}) + \alpha_{1i}\right] g(v_t; \gamma_i, c_i) + \varepsilon_{i\,t+1}$, β_{0i} and β_{1i} denote the interest-rate differential coefficients in the linear and nonlinear regimes, respectively. γ_i is the slope parameter and c_i is the threshold value of the VIX (in %). *t* statistics are given in parentheses.

5.3. The Fama regression augmented with default risk

To verify the hypothesis that the default risk explains the forward bias at least partly, we now introduce the default risk proxy into the Fama regression. This "augmented Fama regression" is written as follows:

$$s_{it+1} - s_{it} = \beta_i'(i_{it} - i_t^{US}) + \alpha'_i + \lambda_i d_{it} + \varepsilon_{it+1} \quad (18)$$

Note that the term captured in the Fama regression if the coefficient β_i was unity, $s_{it+1} - s_{it} - (i_{it} - i_t^{US})$, is the opposite of the ex post excess return in the carry trade. To be consistent with propositions (P1) and (P2) stated above, we therefore expect to find λ_i negative during boom periods, and positive during the crisis.

To comfort our hypothesis, we would also require that the bias found in the coefficient β_i estimated in the former section is reduced. We therefore expect the β_i' estimated during the low volatility periods to be greater than the corresponding coefficient β_i estimated over the same period in Equation (16), that does not take into account the default risk.

The estimation results meet our expectations, lending support to our hypothesis. Results for the three considered sub-periods are reported in the appendix (Table A.3). Firstly, the coefficient on the default risk λ_i is negative as expected in the pre-crisis period for most currencies (15 in 18). This means that default risk reduces the depreciation in the currency during the booming market period, contributing to the forward bias. Secondly, the coefficient turns positive during the crisis for 16 currencies out of 18, as the default risk accentuates the depreciation of the currency during the turmoil. Thirdly, the coefficient λ_i remains positive for 11 in 18 countries in the recovery period. This can be interpreted by a more cautious behaviour of investors in the immediate aftermath of the crisis, as they discriminate more their assets according to the default risk. Fourthly, the coefficient on the interest-rate differential β_i' is found higher than its counterpart in Equation (16) for the majority of countries (12 out of 18) in the boom period, as expected.

5.4. Nonlinearities in the "augmented" Fama regression

Our "augmented Fama regression" including the default risk (Equation (18)) is also likely to exhibit strong nonlinearities for two reasons: (i) the coefficient β varies across periods, on the same grounds as in the Fama regression, and (ii) default risk tends to boost carry-trade gains when markets are booming, and to worsen the losses in financial turmoil as shown in Section 4.2; therefore the coefficient λ also fluctuates according to the volatility on financial markets. To check that, we estimate the following STR model, where the transition from one state to the other is governed by the level of the VIX:

$$s_{it+1} - s_{it} = \beta_{0i}(i_{it} - i_t^{US}) + \lambda_{0i} d_{it} + \alpha_{0i} + \left[\beta_{1i}(i_{it} - i_t^{US}) + \lambda_{1i} d_{it} + \alpha_{1i}\right] g(v_t; \gamma_i, c_i) + \varepsilon_{it+1} \tag{19}$$

Results of the tests confirm both the existence of nonlinearities and the logistic form of the transition function *g*. Estimations reported in Table 4 give support to our hypothesis about the role of the default risk in the forward bias. First, when volatility is low, the default risk is associated with less exchange-rate depreciation in the majority of cases (as λ_{0i} is significantly negative for 9 currencies in 17). Second, when market volatility rises above a given threshold, the depreciation in the exchange rate is accentuated by the default risk (as λ_{1i} is significantly positive in 11 cases out of 17).

Table 4. Results of the nonlinear estimations of the "augmented Fama regression"

	Linear		Nonlinear	Nonlinear	Transition parameters	
Country *i*	β_{0i}	λ_{0i}	β_{1i}	λ_{1i}	γ_i	c_i
Brazil	0.069 (0.612)	-34.661 (-17.126)	2.202 (5.429)	52.809 (25.320)	18.682	23.034
Chile	-8.600 (-4.651)	-83.374 (-2.022)	7.192 (3.793)	104.140 (2.529)	3.933	13.154
Colombia	1.293 (7.675)	-38.390 (-15.334)	1.465 (4.235)	47.792 (18.855)	7.773	23.492
Mexico	-0.563 (-4.435)	-18.966 (-10.809)	5.264 (16.428)	26.886 (15.313)	12.714	22.354
Peru	-6.507 (-8.999)	6.642 (7.389)	9.614 (13.027)	-2.559 (-2.785)	15.864	17.128
Korea	-6.559 (-15.667)	9.574 (4.276)	5.898 (9.905)	3.046 (1.340)	10.087	22.651
Indonesia	-0.349 (-2.091)	19.438 (11.764)	-2.048 (-10.461)	-14.675 (-8.852)	212.49	15.551
Malaysia	1.998 (7.625)	-28.946 (-6.547)	-0.839 (-2.672)	38.690 (8.809)	7.203	17.807
Philippines	1.398 (20.237)	9.599 (31.501)	-0.871 (-1.914)	-7.393 (-14.963)	6.143	36.502
Thailand	-12.657 (-15.037)	63.537 (7.574)	12.901 (14.954)	-55.217 (-6.579)	20.364	13.415
Czech Republic	2.092 (6.899)	-59.567 (-8.864)	4.933 (9.301)	71.645 (10.670)	5.283	19.707
Hungary	1.251 (8.440)	-23.662 (-7.820)	-0.697 (-2.801)	32.404 (10.792)	8.137	17.800
Poland	0.899 (4.176)	-44.577 (-8.565)	7.220 (16.279)	55.927 (10.793)	5.648	19.679
Romania	-0.142 (-2.890)	7.667 (29.070)	2.045 (7.105)	-6.532 (-12.924)	5.070	33.855
Russia	-0.509 (-10.648)	15.374 (31.536)	3.264 (7.564)	-13.082 (-24.374)	6.251	36.481
Israel	3.129 (10.015)	3.237 (0.979)	6.309 (12.784)	2.624 (0.782)	8.452	17.817
South Africa	-4.805 (-19.154)	-8.572 (-5.129)	2.256 (6.814)	20.548 (12.034)	76.831	23.075

Note: This table reports the results from the estimation of Equation (19) $s_{it+1} - s_{it} = \beta_{0i}(i_{it} - i_t^{US}) + \lambda_{0i} d_{it} + \alpha_{0i} + \left[\beta_{1i}(i_{it} - i_t^{US}) + \lambda_{1i} d_{it} + \alpha_{1i}\right] g(v_t; \gamma_i, c_i) + \varepsilon_{i_{t+1}}$, β_{0i} and β_{1i} denote the interest-rate differential coefficients in the linear and nonlinear regimes, λ_{0i} and λ_{1i} the CDS premia coefficients in the linear and nonlinear regimes, respectively. γ_i is the slope parameter and c_i is the threshold value of the VIX (in %). The model does not converge for Argentina. *t* statistics are given in parentheses.

6. ROBUSTNESS CHECKS

6.1. Choice of the funding currency

The former results concern carry trades funded by USD. However carry-trade can be funded by any low interest rate currencies or a basket of them. The Japanese yen (JPY) and the Swiss Franc (CHF) have been particularly popular to fund carry trades in the last decade, because of their very low interest rates. To investigate the robustness of our results to the choice of the funding currency, we replicate all our empirical analysis using the Japanese yen and the Swiss franc as alternative funding currencies. Results corresponding to the estimation of Equation (12) are reported in Table A.4 in the Appendix. Comparing results in Table A.4 with those in Table A.1 clearly show that our findings are robust to the choice of the funding currency: for

the three considered funding currencies, coefficients switch from positive (except for only two countries) to negative values from the first to the second period, and become again positive in the last, recovery period. All the other results obtained using the Japanese yen and the Swiss franc as funding currencies are also very similar to those obtained with the US dollar.[10]

6.2. Estimations by panels of countries

To further check our findings and provide more synthetic results, we also run all the estimations using panel data. To this end, we classify our countries into four groups: the complete panel includes all the 18 currencies in the sample, whereas the others group together the countries belonging to the same geographic areas (Latin America, Asia, and emerging Europe). All the results obtained using panel data confirm those issued from the country-by-country analysis. We report some of them below as an illustration.

To estimate nonlinear equations on a panel basis, we use the panel smooth transition regression (PSTR) model introduced by González et al. (2005), which involves three steps. In the first, specification step, we test for homogeneity against the PSTR alternative, using the LM-test statistic provided by González et al. (2005). In the second, estimation step, nonlinear least squares are used to obtain the parameter estimates, once the data have been demeaned. Finally, in the evaluation step, we apply misspecification tests in order to check the validity of the estimated PSTR model.

Results obtained for Equation (13) through the PSTR model (Table 5) comfort our previous findings. First, when volatility is low, the default risk contributes to swell the carry-trade

[10] To avoid too many tables, we do not report the complete results, but they are available upon request to the authors.

gains. This is attested by the significantly coefficient λ_{0i}, for the whole sample, as well as for Latin America and emerging Europe, Asia being an exception. Second, when the volatility overcomes a certain threshold, the coefficient sign changes, and the default risk contributes to deepen the losses. This is shown by the negative value of coefficient λ_{1i}, with a higher absolute value than λ_{0i} for all panels. Third, the level of the VIX which determines the transition between the two regimes is estimated around 23% for all the panels, which, as previously mentioned, roughly fits our definition of periods given in Section 3.

Table 5. Coefficient λ_i in the nonlinear estimations (panel)

	Linear	Nonlinear	Transition parameters	
Panel *i*	λ_{0i}	λ_{1i}	γ_i	c_i
Whole sample	1.926 (15.802)	-2.673 (-22.677)	6.784	23.009
Latin America	1.464 (20.627)	-1.707 (-24.564)	5.358	23.022
Asia	-3.297 (-6.964)	-3.811 (-9.324)	2.545	23.778
Emerging Europe	7.651 (15.257)	-12.857 (-26.638)	1.219	21.499

Note: see Table 2.

Nonlinearities in the Fama regression are also confirmed in the panel framework (Table 6). The coefficient on the interest-rate differential is lower than 1—and even negative for two groups of countries—in the low-volatility regime, while it is always greater than unity when the VIX exceeds its threshold.

Table 6. Results of the nonlinear estimations of the Fama regression

	Linear	Nonlinear	Transition parameters	
Panel i	β_{0i}	β_{1i}	γ_i	c_i
Whole sample	0.034 (1.284)	4.085 (49.212)	0.246	32.353
Latin America	-0.167 (-4.291)	2.637 (27.269)	0.241	29.782
Asia	-0.142 (-1.910)	7.979 (31.311)	0.274	36.593
Emerging Europe	0.156 (3.427)	7.227 (35.569)	0.202	33.097

Note: This table reports the results from the panel estimation of Equation (17): $s_{it+1} - s_{it} = \beta_{0i}(i_{it} - i_t^{US}) + \alpha_{0i} + [\beta_{1i}(i_{it} - i_t^{US}) + \alpha_{1i}]g(v_t;\gamma_i, c_i) + \varepsilon_{it+1}$, β_{0i} and β_{1i} denote the interest-rate differential

coefficients in the linear and nonlinear regimes, respectively. γ_i is the slope parameter and c_i is the threshold value of the VIX (in %). *t* statistics are given in parentheses.

Finally, estimating our "augmented" Fama regression (Table 7) on a panel-basis gives very similar results to those obtained on an individual currency basis. First, the coefficient on the risk default proxy λ_i is negative before the crisis as expected for all the considered areas. Second, its value increases during the crisis for all panels, being close to zero for the whole sample and reaching the positive territories for Asia and emerging Europe. Third, the recovery period appears different than the former booming market period, as the coefficient λ_i is positive for the whole sample, as well as for Latin America and emerging Europe. On the whole, the default risk tends to mitigate high-yield currency depreciation in booming markets, this effect being reversed during crises and in their immediate aftermath. Fourth, the coefficient on the interest-rate differential β_i' in the pre-crisis period is higher than the same coefficient estimated in Equation (16) for all the considered panels. Moreover, the coefficients β_i' are now positive for the pre-crisis period for two areas: Asia and emerging Europe, whereas they were negative for all panels in the standard Fama regression. Consequently, the bias on this coefficient estimated over the pre-crisis period is mitigated by taking into account the default risk.

Table 7. Results of the "augmented Fama regression" (panel estimation)

	Pre-crisis period		Crisis period		Recovery period	
Panel *i*	Interest differential β_i'	Default risk λ_i	Interest differential β_i'	Default risk λ_i	Interest differential β_i'	Default risk λ_i
Whole sample	-0.348	-2.127	2.681	-0.085	-1.264	3.159
	(-21.490)	(-16.792)	(32.776)	(-3.702)	(-37.001)	(38.909)
Latin America	-0.723	-0.231	2.049	-0.161	-1.685	4.126
	(-36.182)	(-4.525)	(33.588)	(-11.214)	(-33.730)	(29.391)
Asia	0.189	-5.281	0.106	5.137	-1.046	-7.087
	4.472)	(-17.488)	(1.059)	(31.170)	(-26.995)	(-24.477)
Emerging Europe	0.343	-22.650	3.689	1.540	-0.829	4.547
	(6.227)	(-39.283)	(29.181)	(11.901)	(-19.485)	(14.772)

Note: This table reports the results from the estimation of Equation (18), $s_{it+1} - s_{it} = \beta_i'(i_{it} - i_t^{US}) + \alpha'_i + \lambda_i d_{it} + \varepsilon_{it+1}$, where $s_{it+1} - s_{it}$ is the one-year change in exchange rate versus dollar, $(i_{it} - i_t^{US})$ the one-year interest-rate with the US, d_{it} is the sovereign CDS premium proxing the default risk. *t* statistics are given in parentheses.

7. CONCLUSION

The positive excess returns on carry trades invested in high-yield currencies has long constituted a famous puzzle for economists and paved the way for an extensive literature. Here, we state the hypothesis that carry-trade gains observed in booming market periods can be interpreted as a risk premium stemming both from exchange-rate change and default risk. If this hypothesis holds, two propositions should follow: (i) currencies issued by a country with a high default risk should yield higher excess returns in boom periods; (ii) conversely, these currencies should depreciate more severely during crises—the riskier the country in terms of default, the sharper the reversal of carry-trade gains during crises.

We have verified these propositions empirically on a sample of 18 emerging currencies by two means. First, we run a linear regression between carry-trade gains and a proxy of the default risk. We obtain the expected signs on the coefficients when splitting the estimations into different sub-periods corresponding to booms and crisis episodes. Then we run a nonlinear regression where the impact of the default risk on carry-trade gains changes according to market volatility. This allows us to confirm the two propositions above: (i) when

market volatility is low, carry-trade gains are higher when invested in currencies issued by countries with high default risk; (ii) conversely, over a certain threshold, the more volatile the financial markets become, the deeper the losses on investment in high default risk countries.

As carry-trade gains are also correlated with interest-rate differentials, we had to verify that the former results are still valid when accounting for interest differentials. To do that, we use the so-called "Fama regression", linking the exchange-rate change to the interest differential. First, we give a nonlinear estimation of this equation, which allows us to explain the evolution of its coefficient by the change in market volatility along the financial cycle. Second, we introduce the default risk proxy in this regression. The default risk is a significant variable with its sign switching across periods, as expected. Moreover, the coefficient on the interest-rate differential gets closer to unity once the default risk is included in the regression, which shows that the "forward bias" is somewhat mitigated by taking into account the default risk. On the whole, our findings give support to the hypothesis that default risk contributes to the excess returns in the exchange market during boom periods. Nevertheless, we are aware that this factor only partially explains the forward bias, as a large part of the carry trades are invested in advanced country currencies, such as the Australian dollar or the New Zealand dollar, with practically no sovereign default risk. The present results only show that the default risk contributes to increase the forward bias in the case of high-yield emerging countries.

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APPENDIX

Table A.1. Coefficient λ_i in the linear estimations

Country *i*	Pre-crisis period	Crisis period	Recovery period
Argentina	1.634 (16.261)	-0.0463 (-2.463)	1.332 (21.794)
Brazil	65.473 (37.156)	-7.509 (-15.577)	39.011 (15.317)
Chile	20.661 (4.619)	-15.514 (-24.715)	27.470 (6.884)
Colombia	25.131 (24.044)	-5.150 (-10.769)	25.947 (20.453)
Mexico	23.846 (13.649)	-8.284 (-23.102)	11.750 (9.531)
Peru	8.995 (7.533)	-1.833 (-8.614)	10.904 (18.721)
Indonesia	4.950 (6.245)	-2.602 (-14.682)	23.817 (24.776)
Korea	-4.217 (-1.838)	-10.376 (-29.924)	19.348 (10.414)
Malaysia	10.068 (7.858)	-5.293 (-15.783)	20.823 (10.393)
Philippines	12.131 (14.648)	-3.066 (-22.464)	7.141 (17.079)
Thailand	17.348 (8.009)	-3.180 (-9.044)	9.356 (18.117)
Czech Republic	142.568 (18.014)	-12.496 (-13.578)	5.301 (1.761)
Hungary	37.295 (11.372)	-4.451 (-19.259)	1.801 (1.942)
Poland	112.518 (21.255)	-13.819 (-16.480)	8.984 (4.214)
Romania	21.327 (6.721)	-3.661 (-20.488)	2.773 (3.874)
Russia	17.601 (22.011)	-2.599 (-8.607)	18.538 (11.058)
Israel	55.019 (17.147)	-6.076 (-7.841)	7.462 (8.049)
South Africa	-3.753 (-3.315)	-5.177 (-9.249)	30.072 (12.325)

Note: This table reports the results from the estimation of Equation (12): $r_{it+1} = \lambda_i d_{it} + \alpha_i + \varepsilon_{it+1}$ where r_{it+1} is the ex post excess return in the one-year carry trade invested at time *t* in currency *i* and d_{it} is country *i*'s sovereign CDS premium. *t* statistics are given in parentheses.

Table A.2. Coefficient β_i in the Fama regression

Country *i*	Pre-crisis period	Crisis period	Recovery period
Argentina	0.209 (3.879) [0.000]	1.275 (19.826) [0.000]	0.381 (14.665) [0.000]
Brazil	-1.144 (-17.610) [0.000]	2.433 (12.047) [0.000]	-2.065 (-21.715) [0.000]
Chile	1.327 (9.491) [0.019]	-5.766 (-11.698) [0.000]	7.342 (10.729) [0.000]
Colombia	-1.698 (-9.860) [0.000]	2.681 (9.404) [0.000]	-2.414 (-15.393) [0.000]
Mexico	-0.479 (-5.556) [0.000]	5.216 (38.815) [0.000]	-1.209 (-7.176) [0.000]
Peru	-9.082 (-6.245) [0.000]	2.487 (12.402) [0.000]	-3.650 (-3.418) [0.000]
Indonesia	-0.148 (-1.759) [0.000]	2.262 (9.295) [0.000]	-1.863 (-20.094) [0.000]
Korea	-2.927 (-2.684) [0.000]	10.038 (11.133) [0.000]	-4.689 (-12.396) [0.000]
Malaysia	3.656 (27.748) [0.000]	4.602 (3.031) [0.018]	-8.381 (-7.504) [0.000]
Philippines	-1.207 (-8.269) [0.000]	3.714 (12.600) [0.000]	-1.776 (-16.466) [0.000]
Thailand	2.321 (3.619) [0.039]	3.136 (3.241) [0.027]	-5.593 (-7.973) 0.000]
Czech Republic	5.339 (14.381) [0.000]	14.461 (16.559) [0.000]	-2.497 (-2.306) [0.001]
Hungary	-0.320 (-1.505) [0.000]	4.369 (23.462) [0.000]	-0.237 (-0.992) [0.000]
Poland	0.380 (0.915) [0.135]	10.500 (30.450) [0.000]	-1.102 (-2.650) [0.000]
Romania	-0.751 (-6.884) [0.000]	3.286 (18.515) [0.000]	0.055 (0.525) [0.000]
Russia	-0.917 (-7.398)[0.000]	6.069 (20.964) [0.000]	-0.327 (-3.962) [0.000]
Israel	1.730 (3.538) [0.136]	7.686 (11.293) [0.000]	5.142 (3.228) [0.009]
South Africa	1.109 (7.817) [0.441]	1.824 (6.711)[0.002]	-2.615 (-13.987) [0.000]

Note: This table reports the results from the estimation of Equation (16) $s_{i_{t+1}} - s_{i_t} = \beta_i (i_{i_t} - i_t^{US}) + \alpha_i + \varepsilon_{i_{t+1}}$, where $s_{i_{t+1}} - s_{i_t}$ is the one-year change in exchange rate versus dollar, $(i_{i_t} - i_t^{US})$ the one-year interest-rate with the US. *t* statistics are given in parentheses. p-values of the Wald test of the null hypothesis $\beta_i = 1$ are reported in squared brackets.

Table A.3. Results of the estimation of the Fama regression with the default risk

	Pre-crisis period		Crisis period		Recovery period	
Country *i*	β_i'	λ_i	β_i'	λ_i	β_i'	λ_i
Argentina	0.350	-0.482	1.537	-0.101	0.485	-0.231
	(4.163)	(-2.813)	(13.959)	(-3.609)	(5.127)	(-1.222)
Brazil	-0.138	-35.261	0.897	7.880	-4.710	37.209
	(-1.479)	(-15.007)	(2.970)	(6.349)	(-18.402)	(10.645)
Chile	0.416	-38.999	-1.285	11.684	9.912	14.946
	(2.051)	(-5.357)	(-3.470)	(17.349)	(4.538)	(1.559)
Colombia	0.896	-24.530	0.855	5.477	-1.423	-7.929
	(4.004)	(-14.834)	(1.776)	(4.914)	(-2.829)	(-2.201)
Mexico	-0.117	-7.860	3.824	3.195	-3.917	16.313
	(-0.910)	(-4.227)	(16.950)	(7.698)	(-5.760)	(4.205)
Peru	-7.013	-8.236	1.716	1.189	-0.542	-10.524
	(-5.844)	(-7.876)	(7.219)	(5.746)	(-1.037)	(-17.020)
Indonesia	-0.659	6.015	0.244	3.428	-3.287	12.143
	(-7.184)	(7.228)	(1.462)	(10.386)	(-4.956)	(2.204)
Korea	-2.599	4.739	0.967	10.400	-2.083	-9.766
	(-2.077)	(1.986)	(1.567)	(14.741)	(-1.973)	(-2.264)
Malaysia	3.390	-1.814	3.396	5.117	-2.685	-15.679
	(17.498)	(-1.672)	(9.691)	(24.137)	(-2.157)	(-5.734)
Philippines	0.289	-9.344	0.555	3.409	-2.276	1.424
	(1.231)	(-7.285)	(1.909)	(9.939)	(-8.220)	(1.815)
Thailand	-6.569	-32.291	1.694	3.104	0.527	-9.158
	(-6.863)	(-13.339)	(2.828)	(9.822)	(0.696)	(-14.187)
Czech Rep.	2.488	-115.443	8.367	7.528	-8.175	14.635
	(9.895)	(-13.918)	(8.114)	(7.031)	(-3.243)	(2.039)
Hungary	0.622	-33.580	2.458	2.724	-3.249	11.904
	(3.073)	(-10.585)	(5.972)	(5.065)	(-13.515)	(21.556)
Poland	0.640	-112.201	8.582	3.535	-1.287	0.841
	(2.143)	(-21.644)	(12.886)	(3.963)	(-0.698)	(0.091)
Romania	-0.623	-6.345	1.849	2.589	-1.236	6.830
	(-5.427)	(-5.073)	(7.840)	(8.329)	(-4.398)	(6.423)
Russia	0.229	-12.941	5.705	0.302	-1.044	12.132
	(2.109)	(-14.786)	(13.448)	(1.059)	(-5.818)	(4.944)
Israel	2.858	-57.841	8.015	-0.389	6.159	-7.974
	(9.520)	(-22.545)	(5.619)	(-0.271)	(9.374)	(-11.117)
South Africa	0.666	6.954	-0.901	10.242	-4.469	17.408
	(3.139)	(4.147)	(-5.047)	(14.262)	(-8.509)	(3.247)

Note: This table reports the results from the estimation of Equation (18), $s_{it+1} - s_{it} = \beta_i'(i_{it} - i_t^{US}) + \alpha'_i + \lambda_i d_{it} + \varepsilon_{it+1}$, where $s_{it+1} - s_{it}$ is the one-year change in exchange rate versus dollar, $(i_{it} - i_t^{US})$ the one-year interest-rate with the US, d_{it} is the sovereign CDS premium proxing the default risk. *t* statistics are given in parentheses.

Table A.4. Coefficient λ_i in the linear estimations, alternative funding currencies (JPY and CHF)

Country *i*	Pre-crisis period	Crisis period	Recovery period
Argentina	1.277 (3.922)	-0.294 (-11.901)	0.828 (11.261)
	-0.093 (-0.300)	0.090 (4.273)	0.801 (6.138)
Brazil	72.880 (22.806)	-112.311 (-18.297)	31.938 (11.165)
	63.264 (21.106)	-5.294 (-13.260)	32.297 (19.206)
Chile	33.112 (3.765)	-21.215 (-18.997)	12.942 (3.077)
	0.331 (0.044)	-12.011 (-15.538)	16.199 (7.035)
Colombia	28.273 (11.840)	-9.133 (-21.895)	20.057 (13.811)
	19.868 (8.983)	-2.984 (-8.127)	19.692 (18.876)
Mexico	33.171 (7.033)	-12.271 (-18.037)	5.509 (3.683)
	15.330 (3.467)	-6.048 (-12.622)	4.957 (2.948)
Peru	16.901 (16.106)	-6.116 (-14.947)	3.922 (3.718)
	8.189 (8.700)	0.449 (1.547)	3.817 (3.489)
Indonesia	7.432 (4.672)	-4.609 (-16.827)	19.535 (17.065)
	2.090 (1.523)	-1.422 (-8.485)	18.331 (16.765)
Korea	-10.633 (-3.089)	-13.597 (-22.676)	12.327 (5.867)
	-22.653 (-7.556)	-8.655 (-21.387)	12.894 (11.607)
Malaysia	5.722 (2.141)	-11.268 (-17.272)	7.908 (4.143)
	5.703 (3.286)	-2.110 (-5.996)	8.769 (2.974)
Philippines	15.007 (11.102)	-6.153 (-15.505)	2.339 (4.417)
	10.443 (8.937)	-1.418 (-5.604)	1.511 (1.507)
Thailand	18.944 (4.404)	-8.897 (-21.467)	2.102 (2.905)
	3.631 (1.103)	-0.169 (-0.591)	2.909 (2.152)
Czech Republic	121.853 (10.829)	-18.808 (-9.049)	4.545 (1.317)
	82.844 (10.992)	-7.809 (-8.420)	3.376 (1.902)
Hungary	34.689 (6.653)	-6.697 (-24.254)	0.564 (0.594)
	19.003 (5.310)	-3.067 (-16.643)	0.282 (0.427)
Poland	109.203 (11.496)	-18.343 (-20.252)	2.198 (0.889)
	76.381 (11.413)	-11.019 (-15.070)	2.607 (1.431)
Romania	19.118 (3.946)	-5.604 (-21.886)	0.631 (0.809)
	9.112 (2.235)	-2.593 (-15.865)	0.919 (1.960)
Russia	20.885 (9.486)	-3.967 (-14.016)	13.128 (6.934)
	11.600 (6.407)	-1.939 (-7.742)	13.682 (7.743)
Israel	62.019 (15.046)	-12.590 (-13.769)	-0.575 (-0.425)
	33.571 (13.061)	-1.964 (-2.904)	-0.662 (-0.421)
South Africa	-6.983 (-3.219)	-8.234 (-11.161)	24.239 (9.358)
	-16.799 (-9.723)	-3.456 (-4.746)	24.637 (16.460)

Note: This table reports the results from the estimation of Equation (12): $r_{it+1} = \lambda_i d_{it} + \alpha_i + \varepsilon_{it+1}$ where r_{it+1} is the ex post excess return in the one-year carry trade invested at time *t* in currency *i* and d_{it} is country *i*'s sovereign CDS premium. *t* statistics are given in parentheses. For each country, the first line corresponds to the Japanese yen as the funding currency, while the second line reports the results corresponding to the use of the Swiss franc as the funding currency.