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Sovereign Credit Default Swaps and the Currency Forward Bias

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ABSTRACT

We study the links between sovereign credit risk and the currency forward bias. In a setting of defaultable sovereign bonds, we show that the forward bias can be negatively linked to sovereign credit risk. We confirm empirically that the forward bias is negatively associated to sovereign CDS spreads and systematically across both developed and emerging countries but the effect is more pronounced for emerging countries. Furthermore, we show that the forward bias decreases after the inception of the sovereign CDS market. Overall, our results underscore the distinct role of the sovereign CDS market in enhancing price efficiency in currency forward and spot markets.

1. Introduction

There exists an established tradition of academic studies that have endeavored to shed light on the currency forward bias, defined as the expectation of the difference between the future spot price and the forward price, ever since Fama (1984) documented the presence of the forward bias in foreign exchange markets. Theoretically, if the interest parity holds, there is no currency forward bias or the forward bias should disappear over time (Bansal, 1997). However, empirically, the forward bias puzzle holds, implying that a high-yield (low-yield) currency appreciates (depreciates). The currency forward bias helps explain the popularity of carry trades, a trading strategy which involves borrowing a low-yield currency and simultaneously investing in high-yield currency bonds.[1] If the interest rate parity holds, the profit of such trading strategy will be offset by the loss due to the high-yield currency depreciation. However, the empirical evidence shows that a high-yield currency usually gives rise to profitable arbitrage opportunities (Fama, 1984). Given the persistence and the predictability of the currency forward bias, prior studies have demonstrated the abnormal gains from exploiting price deviations in the currency forward bias (Villanueva, 2007). By contrast, Bansal and Dahlquist (2000) find that the forward bias has been a typical feature of developed countries and somewhat a rare phenomenon in emerging market countries.

One of the cornerstones of modern financial economics is the efficient market hypothesis which states that if currency markets are efficient, then their prices should reflect all the available market information.[2] A large body of research documents that currency forward prices effectively predict future spot prices, leading to a random (or zero average of) bias between forward and future spot

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[1] Several papers such as Hodrick (1987), Lewis (1995) and Sarno (2005) have documented the execution of carry trades.

[2] The seminal papers of Fama (1970, 1991) and Malkiel (2003) provide more detailed discussion.

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prices. As mentioned above, academic researchers have puzzlingly found that the bias tends to be persistent and predictable, even though the foreign exchange market is the world's largest financial market. Thus, understanding the nature of the forward premium bias is of key importance from both an academic and a market practitioner perspective. Theoretically, the existence of the forward bias suggests a missing piece of information that is not reflected in foreign exchange markets prices; hence it is key for professional investors to understand the main determinants of the forward bias. Furthermore, it is important to note that recent research has documented that sovereign CDS spreads have strong predictive power for the cross-section of currency spot prices.

Understanding the relationship of sovereign CDS markets with the currency forward markets has thus been proven to be a key challenge in the economics and finance literature. Since the European sovereign debt crisis, sovereign CDS is still abundantly actively traded, with an outstanding notional of 1.2 billion USD at the end of 2019. By contrast, the size of the forward currency market is much smaller than that of the spot market. In terms of transactions, the forward currency turnover is in fact below half the spot transactions (BIS, 2019). In addition, currency forward markets are not accessible for all currencies, and are mostly being traded for developed countries (Eiteman et al., 2009). Hence, analyses of risks in currency forward markets are relatively underrepresented or absent in standard empirical works on currency markets. However, more recently there has been a rapid growth in currency forward transactions mostly due to the increasing need for multinational firms to hedge currency risk. Indeed, the trade in currency forwards contracts has doubled over the recent ten years, while spot transactions have increased by only 30%. Also, forward contracts have become common in more countries, especially in the Asia Pacific and Latin America regions. As noted above, academic research has been silent on the nature of the dynamic interaction between sovereign CDS premia and currency forward returns. Hence, we are motivated to understand whether and how sovereign CDS markets affect currency forward markets.

In this study, we analyze all the available currencies of 48 developed and developing countries over the 2003–2020 sample period. Typically, existing studies focus on the most popular currencies and on developed countries, because these currencies are widely traded, and the information is readily available.

Our study provides a new perspective on the economic sources of the forward premium puzzle. Based on our theoretical motivation, we propose that changes in sovereign credit risk affect the forward bias and we demonstrate that our theoretical prediction holds empirically.[3] In fact, the returns and predictability of foreign currency assets have been a long-standing challenging issue for both researchers and market practitioners. As mentioned, canonical interest rate parity states that the interest rate differential is the only driving factor of foreign currency pricing. However, empirically the theory is weak in explaining the innovations from foreign currency markets (Meese and Rogoff, 1983, 1988).

Our paper relates to the literature that seeks to understand the pricing of currencies. One strand of the literature explores the causes of the forward bias. Bansal (1997) provides an economic explanation on the negative relation between the forward bias and the interest rate differential (the difference between foreign and US interest rates). The author also documents the asymmetric effect of the interest rate differential on the forward bias. Based on the property of the asymmetry in forward bias, Backus et al. (2001), Sarno et al. (2006) and Coudert and Mignon (2013) further characterize the nonlinearity of the forward bias. On other hand, many studies have identified several attributes of the forward bias. Bansal and Dahlquist (2000) uncover country-specific macroeconomic variables, e.g. per capita GNP, inflation rate and inflation volatility, as important determinants of the magnitude of the forward bias. Other studies suggest that limit to arbitrage contributes to the persistence of the forward bias, e.g. order flow (Breedon et al, 2016), market liquidity shock (Kumar, 2020), and the large amount of credit-based transactions that prevent the access to bond instruments (Lee and Jung, 2019). Our paper empirically supports these previous analyses, by showing that sovereign credit risk is an additional determinant of the currency forward bias.

Our paper also relates to the recent literature that emphasizes the crucial role of sovereign credit risk for asset prices. In doing so, we go one step beyond by exploring how the pricing of currency forward bias is related to sovereign credit risk. Notably, recent studies provide evidence that a foreign currency is also exposed to sovereign credit risk, as a country's currency normally tends to depreciate with rising sovereign credit risk.

Prior studies show that sovereign credit risk can affect the currency exchange rate. Empirically, Della Corte et al. (2022) and Foroni et al. (2018) demonstrate that heightened sovereign credit risk may actually determine the devaluation of the country's currency. This is because the increased default probability on sovereign bond is accompanied by a contemporaneous reduction in the return from the interest rate differential, as well as a depreciation of the exchange rate.[4] In addition, Calice and Zeng (2019) show that sovereign CDS term premia, the difference between the 10- and 1-year CDS spreads, reveal valuable information about country-specific risks and matter for the cross-section of exchange rates and currency risk premia. These studies show that sovereign credit risk (as measured by sovereign CDS) is a prominent factor for predicting foreign currency rates (see, e.g. Foroni et al., 2018; Calice and Zeng, 2019; Della Corte et al., 2022), and the predictive power is higher than the interest rate differential (Foroni et al., 2018). Hence, our paper builds on the insights of these studies to examine the influence of sovereign CDS markets on currency forwards.

Nevertheless, the potential informativeness role of sovereign credit risk for the forward bias is controversial in theory and has not been addressed empirically yet. Consequently, our approach builds on the idea that the pricing of sovereign credit risk (as reflected in the sovereign CDS market) matters for asset prices through their global and idiosyncratic components. This paper however differs from existing studies by focusing on the relation and dynamic co-movement between currency forward and spot markets to explore whether the sovereign CDS premia also lead pricing discovery in the currency forward market. Notably, this is distinct from prior work that

[3] Note that, in this paper, we use sovereign credit risk and sovereign default risk interchangeably.

[4] The features of the theoretical framework are outlined in Section 2, where we detail the impact on the return in cross-country investment in the case of a sovereign bond default.

focuses on how spot currency returns can be predicted (priced) by sovereign CDS premia. In practice, a variety of currency trading strategies, e.g. carry trades, are constructed based on the existence of the forward premium bias. Therefore, our analysis is motivated by the insights of recent research that shows that currency returns can be also explained by sovereign CDS premia to shed light on these empirical patterns about international currency returns and the pricing of risk of carry trading strategies.

Our study contributes to a better understanding of the relationship between sovereign credit risk and foreign currency markets. We shed light on the dynamic co-movement between currency spot and forward prices due to the impact of sovereign CDS, as sovereign credit risk also affects currency markets. There exists an emerging strand of the literature on sovereign credit risk and foreign currency. However, these studies focus solely on a single currency market, mainly the currency spot market. Our findings on the currency price gap between spot and forward markets enrich the state of art and are in line with existing studies on sovereign credit risk and foreign currencies. Coudert and Mignon (2013), Della Corte et al. (2022) and Foroni et al. (2018) provide evidence that currency depreciates when sovereign risk increases, while Della Corte et al. (2022) further shows that global sovereign risk is a significant factor driving foreign currency depreciation. Global sovereign default risk is echoed also in prior studies, which provide evidence of a strong systematic (common set of global market factors) component in sovereign CDS spreads (Augustin and Tédongap, 2016; Longstaff et al., 2011). Calice and Zeng (2019) show that the sovereign CDS term structure cross-sectionally explain foreign currencies. Thus, these studies provide new insights on foreign currency determinants alongside traditional pricing factors, e.g., monetary policy (Du et al., 2018), commodity price (Ferraro et al., 2015; Sun et al, 2020), and VIX (Brunnermeier et al., 2008).

Our paper also contributes to the relatively small strand of the literature on price discovery of spot and forward prices. Interestingly, studies on this subject are primarily centered on commodity markets: e.g. Garbade and Silber (1983) link price discovery to the different speeds of new information in spot and futures markets; Figuerola-Ferrettia and Gonzalob (2010) propose a novel measurement methodology for price discovery in commodity markets; and Dolatabadi et al. (2015) and Yang et al. (2001) empirically explore commodity price discovery. On the other hand, there are only a few papers on price discovery in currency markets, partly because the currency forward bias has been well documented in the financial economics literature since the seminal work of Fama (1984). Yes, there have been a few theoretical and empirical contributions in this area. Andersen et al. (2007) study the cross-market price discovery among stock, bond, and currency markets and find that currency markets are relatively more responsive to the new flow of information. Osler et al. (2011) study the currency spot price discovery process between customers and the interdealer markets, and show that dealers do not always pass the new information to their customers; hence the pricing process relies on more fundamental information. However, these studies focus on cross-market dynamics (Andersen et al., 2007) or on market participants (Osler et al., 2011).

Our paper is closely related to Rosenberg and Traub (2009), who find evidence that intra-day currency futures prices lead the currency spot prices. However, a limitation of their study is that only five currency pairs are examined, subject to the available contracts in CME. Hence, our paper contributes to this strand of the literature by quantifying the price discovery mechanism of currency markets; in particular, we provide new evidence that the appearance of the sovereign CDS can reduce the currency forward bias, suggesting that the information in the sovereign CDS markets does impact on pricing dynamics within currency markets.

Our paper is similar in spirit to the recent work of Du and Schreger (2016) who introduce a new measure of country-specific sovereign credit risk: the local currency credit spread (i.e. the difference between the local currency bond over the synthetic local currency risk-free rate), and find that, unlike sovereign CDS (which contains both global and local risk), it is less responsive to global risk factors. Our theoretical analysis starts with a setting where we assume that the US Treasury bond is default risk free while the foreign treasury bond is defaultable. We propose a framework that links directly currency prices to sovereign credit risk which is in line with prior empirical studies on currency spot prices, but we further extend to shed further light on the joint dynamics of both currency spot and forward markets. In particular, the model that guides our empirical analysis predicts a loss in cross-country investment if the foreign bond defaults which suggests that sovereign credit risk leads to the devaluation of the foreign currency. We further show that the timing of the information is also important in explaining the bias between the forward price and the future spot price, because currency prices (both forward and spot) reflect two distinctive and dissimilar informational components of sovereign credit risk.

Our paper is related to the vast literature studying empirically the forward premium puzzle and is also closely related to a still relatively small theoretical literature which has put forward possible explanations of the forward premium puzzle (Backus et al., 2010; Burnside et al., 2009). Bansal (1997) develops a framework under rational expectations to explain the forward premium puzzle. Under the rational expectations, the interest rate parity still holds, and further assumes that the forward bias is correlated to the interest rate differential. Hence, the author shows that the forward bias is asymmetric and depends on the sign of the interest rate differential. In a subsequent study, Bansal and Dahlquist (2000) empirically test the economic model of Bansal (1997) and find that the forward bias is only confined to developed countries. Following the framework on modelling of the forward bias by Bansal (1997) and Bansal and Dahlquist (2000), several similar models have been developed. For example, Bacchetta and Van Wincoop (2010) model the forward bias based on the assumption of rational inattention, a less strong assumption than rational expectations; Lustig and Verdelhan (2007) and Lustig et al. (2011) ascribe the forward bias puzzle to the behavior of the risk premium. Burnside et al. (2009) and Corsetti et al. (2004) explore how market participants behavior affects the forward bias.

Other theoretical studies are built upon the mechanism of the monetary system. Introduced by Lucas (1982) and expanded subsequently by Lagos and Wright (2005), the representative economy has two countries with their own currency, financial, and goods trading systems. In this setting, frictions in goods trades and trading of financial assets are the primary drivers of the persistent violation of interest rate parity (Lee and Jung, 2019). This result is supported by the empirical findings of prior research (see, e.g., Krishnamurthy and Vissing-Jorgensen (2012), Nagel (2016), and Lee (2020)).

In this paper, to formalize the theoretical characterization of our approach, we follow Bansal (1997); thus, we assume that the rational expectations theory holds and that the sovereign bond is not default-free. Furthermore, our model and main empirical results

support and complement the analysis of the currency sport market provided in Coudert and Mignon (2013). Our results have also important implications for the literature which attempt to explain the forward currency bias. Sarno et al. (2006) explore the non-linearity of the forward. Bansal (1997) and Bansal and Dahlquist (2000) show that the interest rate differential is related to the sign of the forward bias. Several studies identify the driving factors of the forward bias from a risk-based perspective, e. g. currency systematic risk (Lustig et al., 2011) and currency market liquidity risk (Breedon et al., 2016; Kumar, 2020). It is important to stress that we further contribute to this strand of the literature by showing that this set of external factors are also primary pricing determinants of the currency forward bias. In particular, our novel contribution is to identify sovereign credit risk as a separate and more important contributor to the currency forward bias and we do so drawing from theory and empirically. Our analysis uncovers a negative association between sovereign credit risk and the currency forward bias. Moreover, when we compare the sovereign CDS variable with a latent variable (the UIRP) which is used to capture the unobserved effect of the uncovered interest rate parity, we find that the currency forward bias appears to be very highly correlated with sovereign CDS spreads.[5] In all, there seems to be a strong direct relation between the sovereign CDS market and the currency forward bias. Consequently, our evidence suggests that models that seek to explain the forward premium puzzle should also consider—as an additional relevant element—the sources of information embedded in the sovereign CDS markets, in addition to existing drivers such as currency systematic factors and market liquidity (Breedon et al., 2016; Kumar, 2020; Lustig et al., 2011). Furthermore, credible models should also account for some of the commonalities in developed and emerging market economies as our novel results show that the negative association between sovereign credit risk and currency forward bias is uniformly prevailing across both developed and emerging market countries. Thus, we build a bridge from the abovementioned risk-based models to new insights on how sovereign credit risk can be a modeling principle applicable also to the forward premium puzzle. Our results also complement extant theoretical models by simplifying their empirical testing.

We establish that there is a direct negative relation between sovereign credit risk premia and the currency forward bias, which in turn implies that an increase in sovereign credit risk premia translates into a decline in the magnitude of the forward bias. This finding is in line with Bansal and Dahlquist (2000), who show that the forward premium puzzle is confined to developed market economies which have typically lower sovereign credit spreads. We are therefore able to conclude from our analysis that the inclusion of the information on sovereign credit risk helps explain the forward bias. We introduce a setting of defaultable sovereign bonds, and model a negative relation between sovereign risk and the forward bias. Our approach differs from previous studies that address the persistence of the forward bias, and we show that the currency forward bias is negatively affected by shocks in the sovereign CDS market.

Overall, our empirical evidence supports our model's theoretical predictions. Furthermore, we find that the impact of sovereign risk is uniform across both developing and developed countries. Importantly, the results are also robust to different measures of the currency forward bias as well as after controlling for the uncovered interest rate parity. Moreover, a much broader set of foreign currencies and the consistency of the results ensures the robustness of our investigation.

The negative association between the country's sovereign CDS and the currency forward bias highlights the importance of the information embedded in sovereign CDS premia for pricing discovery in the foreign exchange markets, as the forward bias is inversely related with sovereign credit risk. To test our theoretical framework, we apply our empirical approach to a longer sample period which dates back to 1990, and find evidence that the forward bias is of relatively small magnitude following the inception of the CDS market. Therefore, taken together, our results demonstrate that sovereign CDS markets can actually increase the "covariability" between currency forward and spot markets.

Our results are relevant to both market participants and regulatory authorities and inform the debate about the efficiency and integrity of the over-the-counter (OTC) derivatives markets. Market participants, e.g. institutional investors, mutual funds, hedge funds, and banks are the major players in the currency market and they are active in both forward and spot markets for hedging FX risk or speculative (e.g. carry trade) purposes. Hence, our findings on sovereign CDS highlight a quantitatively important and previously neglected price discovery channel of the forward bias. Moreover, our results emphasize that the efficiency of the currency market also depends on the informativeness of the markets for sovereign CDS. To summarize, our evidence suggests that sovereign CDS markets enhance the overall transparency of currency market prices which in turn benefits professional global investors largely diversified in international financial assets such as foreign equity and sovereign debt. Similarly, market efficiency is a key goal of financial securities regulation. Our results imply that currency market efficiency also depends on the extent to which sovereign CDS markets are informative on sovereign credit risk. Therefore, sovereign CDS markets induce greater immediacy and transparency of currency market prices for internationally active investors.

The rest of the paper proceeds as follows. In Section 2, we lay out a basic theoretical framework on the links between sovereign credit risk and the currency forward bias. Section 3 describes the data. Section 4 presents the empirical analysis on the currency forward bias. Sections 5 and 6 discuss the core empirical results including sorting exercise and regression analysis to confirm the impact of sovereign CDS on the currency forward bias. Section 7 concludes.

2. Theoretical motivation

In this section, we start by developing a framework that will be utilized for the interpretation of the paper's main findings.
The interest rate parity without default risk:

[5] See Section 6.4 for detailed results and discussions. We use UIRP variable, shown in Equation (17) to capture the unobserved effect of uncovered interest parity.

$$E_t(s_{t+1}) = f_t = s_t + \left(i_t^{US} - i_t^{FX}\right), \tag{1}$$

where s_t is the logarithm of the FX spot rate, f_t is the logarithm of the FX forward rate, i^{FX} is the foreign interest rate and i^{US} is the US interest rate. The definition of the FX spot rate is USD per foreign currency.

2.1. Introducing sovereign credit risk

Considering a scenario of cross-country investment, an investor borrows \$1 with the borrowing rate of i_t^{US} at time t and converts it into the foreign country with FX rate S_t. FX rate is USD per foreign currency. Then, the investor invests the $1/S_t$ in the foreign defaultable fixed income security with yield i_t^{FX}.

Next moving the timeline from t to $t+1$, without default events, the investor will get $\frac{1}{S_t} \times e^{i_t^{FX}} \times S_{t+1}$ in the foreign country investment, but she needs to return $\$1 \times e^{i_t^{US}}$ at $t+1$. Under the assumption of interest rate parity, then the overall payoff is cancelled off:

$$\frac{S_{t+1}}{S_t} \times e^{i_t^{FX}} - e^{i_t^{US}} = 0. \tag{2}$$

It implies that the logarithm of the FX spot rate in $t+1$ can be denoted as $s_{t+1} = s_t + \left(i_t^{US} - i_t^{FX}\right)$, where $s = \log(S)$.

Note that the above analysis is in line with the standard definition of interest rate parity without default risk. Assuming default occurs only at $t+1$, both countries can default with default intensity of λ_{t+1}^{FX} and λ_{t+1}^{US}, with recovery rates R_{t+1}^{FX} and R_{t+1}^{US}. If the foreign country defaults, then the foreign component becomes $\frac{1}{S_t} \times e^{R_{t+1}^{FX} i_t^{FX}} \times S_{t+1}$, implying that the logarithm of the FX sport rate at $t+1$ is equal to:

$$s_{t+1} = s_t + \left(i_t^{US} - R_{t+1}^{FX} i_t^{FX}\right) \tag{3}$$

with default probability of $\lambda_{t+1}^{FX} \times \left(1 - \lambda_{t+1}^{US}\right)$. Notice that the inclusion of default risk is based on Coudert and Mignon (2013).

If two countries are allowed to default, then at $t+1$, there are four possibilities:

Scenario	Probability	s_{t+1}
Default of the foreign country	$\lambda_{t+1}^{FX} \times \left(1 - \lambda_{t+1}^{US}\right)$	$s_t + \left(i_t^{US} - R_{t+1}^{FX} i_t^{FX}\right)$
Default of US	$\lambda_{t+1}^{US} \times \left(1 - \lambda_{t+1}^{FX}\right)$	$s_t + \left(R_{t+1}^{US} i_t^{US} - i_t^{FX}\right)$
Both countries default	$\lambda_{t+1}^{FX} \times \lambda_{t+1}^{US}$	$s_t + \left(R_{t+1}^{US} i_t^{US} - R_{t+1}^{FX} i_t^{FX}\right)$
No default	$\left(1 - \lambda_{t+1}^{FX}\right) \times \left(1 - \lambda_{t+1}^{US}\right)$	$s_t + \left(i_t^{US} - i_t^{FX}\right)$

Hence,

$$\begin{aligned} s_{t+1} = E_{t+1}(s_{t+1}) = &\left[\lambda_{t+1}^{FX}\left(1 - \lambda_{t+1}^{US}\right)\right]\left(i_t^{US} - R_{t+1}^{FX} i_t^{FX}\right) + \left[\lambda_{t+1}^{US}\left(1 - \lambda_{t+1}^{FX}\right)\right]\left(R_{t+1}^{US} i_t^{US} - i_t^{FX}\right) \\ &+\left[\lambda_{t+1}^{FX}\lambda_{t+1}^{US}\right]\left(R_{t+1}^{US} i_{t+1}^{US} - R_{t+1}^{FX} i_t^{FX}\right) + \left[\left(1 - \lambda_{t+1}^{FX}\right)\left(1 - \lambda_{t+1}^{US}\right)\right]\left(i_t^{US} - i_t^{FX}\right) + s_t. \end{aligned} \tag{4}$$

If we assume the US bond is default free, then $\lambda^{US} = 0$. We can simply the Equation (4) by

$$\begin{aligned} s_{t+1} &= \lambda_{t+1}^{FX}\left(i_t^{US} - R_{t+1}^{FX} i_t^{FX}\right) + \left(1 - \lambda_{t+1}^{FX}\right)\left(i_t^{US} - i_t^{FX}\right) + s_t \\ &= \lambda_{t+1}^{FX}\widetilde{R}_{t+1}\left(i_t^{US} - i_t^{FX}\right) + \left(1 - \lambda_{t+1}^{FX}\right)\left(i_t^{US} - i_t^{FX}\right) + s_t \\ &= \left(1 - \left(1 - \widetilde{R}_{t+1}\right)\lambda_{t+1}^{FX}\right)\left(i_t^{US} - i_t^{FX}\right) + s_t \\ &= \left(1 - k_{t+1}^{FX}\right)\left(i_t^{US} - i_t^{FX}\right) + s_t, \end{aligned} \tag{5}$$

where k^{FX} is the sovereign CDS spread of the foreign country[6] and $\widetilde{R}$ is the adjusted recovery rate by $\left(i^{US} - R^{FX} i^{FX}\right)/\left(i^{US} - i^{FX}\right)$. Equation (5) suggests that the return of the foreign currency (i.e.$s_{t+1} - s_t$) is negatively related to the sovereign CDS spreads, given a positive interest rate differential $\left(i^{US} - i^{FX}\right)$. Recall that the extant empirical evidence also shows that sovereign credit risk is negatively related to the currency return (see, e.g., Della Corte et al., 2022).

Then, we turn our analysis to the forward rate. Recalling Equation (1), the forward price is the expectation of the spot price at $t+1$, $f_t = E_t(s_{t+1})$. From Equation (4), the $s_{t+1} = E_{t+1}(s_{t+1}) = E_{t+1}^{\lambda}(s_{t+1})$, where we calculate the expectation of the currency spot over the default probability. Comparing Equations (1), (4) and (5), the currency forward is equal to:

$$f_t = E_t(s_{t+1}) = E_t\left[\left(1 - k_{t+1}^{FX}\right)\left(i_t^{US} - i_t^{FX}\right) + s_t\right]. \tag{6}$$

If we assume that sovereign default risk and the interest rate differential are independent, that λ is a martingale process, i.e.

[6] CDS spreads, k, can be estimated by $(1 - R)\lambda$.

Table 1
Summary Statistics.

	Mean	STD	Max	Min	N
5-year Sovereign CDS (bp)	148.518	339.651	12268.203	1.083	24,840
FX Spot	0.772	0.640	2.666	0.000	24,840
1-month FX Forward	0.771	0.640	2.679	0.000	24,840
1-month Interest Rate (%)	2.967	3.371	34.204	−0.771	24,840

$E_t(\lambda_{t+1}) = \lambda_t$, and that default risk is estimated under the risk-neutral probability measure, then we can re-write Equation (6) as:

$$f_t = \left(1 - k_t^{FX}\right)\left(i_t^{US} - i_t^{FX}\right) + s_t. \tag{7}$$

This illustrates that the currency forward is also affected by concurrent shocks in sovereign credit risk. The currency forward bias is defined as the observed currency spot rate at time $t+1$ subtracted by the currency forward rate at time t:

$$fwdbias = s_{t+1} - f_t. \tag{8}$$

Relating the two terms at the right-hand side of Equation (8) with Equations (5) and (7), the expectation of the FX forward bias can be expressed as follows:

$$E(s_{t+1} - f_t) = -\left(k_{t+1}^{FX} - k_t^{FX}\right)\left(i_t^{US} - i_t^{FX}\right). \tag{9}$$

The model predicts that foreign sovereign credit risk and the interest rate difference (i.e. $i^{US} - i^{FX}$) are jointly negatively related to the currrency forward bias. In addition, in the conventional interest rate parity, the interest rate difference is positively related to the forward bias. This implies that a higher (lower) sovereign credit risk will have a negative (positive) impact on the forward bias.

3. Data

We obtain our data from Thomson Reuters Eikon and Markit. Our sample of sovereign CDS is a combination of Markit (2003–2017) and Thomson Reuters Eikon (2018–2020). Further data, e.g. FX spot and forward prices, and currency interbank interest rates, are obtained from Thomson Reuters Eikon. We use interbank offered rate (IBOR) as a proxy of the risk-free rate. We collect the LIBOR rate or other IBORs for different currencies. The IBOR rates are usually deemed risk free and not collateralized. In our analysis, we focus on sovereign CDS with 5-year tenor, as the 5-year tenor is the most liquid contract, and consider the 1-month FX forward bias, which is commonly studied in academic papers, e.g. Fama (1984).

Our daily observations cover 48 countries globally over the sample period from December 2003 to April 2020. It is important to note that we use the weekly averaged values to reduce noise in prices.

Table 1 reports the summary statistics for the variables. The number of the observations is 24,840. The sample average of the sovereign CDS spreads is 149 bps, with a maximum of 12,268 bps and a minimum of merely 1 bp. Normally, the maximum value of the CDS spread does not exceed 10,000 bps, but for the annualized data the CDS spread exceed 10,000. Hence, in our sample, the maximum value of sovereign CDS spreads is slightly above 10,000 bps, and this picks up the effects of the European sovereign debt crisis at the peak of the crisis in 2010–2011. The FX spot and 1-month forward track each other closely with the sample average of 0.77 and the standard deviation of 0.64. The sample average of the 1-month interest rate is 2.97% with a maximum of 34% and a minimum of −0.771%. Again, notice that the negative interest rate reflects the extreme high price of sovereign bonds during the crisis. We do not winsorize or remove those extreme values to include a more comprehensive collection of market events also in turbulent periods. Note that since the sovereign CDS spread data is different for each country, we further provide the data coverage information in the Appendix A.1.

4. Forward bias detection

We now test for the existence of the forward bias in our sample. The forward bias is defined as the difference between the spot rate at time $t+j$ (we denote it as S_{t+j}) and the corresponding FX forward rate for time $t+j$, at time t (we denote it as $F_{t,t+j}$). Under the assumption of continuous compounding, we use the logarithm values to calculate the FX forward bias that is given by:

$$fwdbias_{i,t+j} = s_{i,t+j} - f_{i,t,t+j} \tag{10}$$

where s_t is the logarithm of the FX spot rate and $f_{t,t+j}$ is the logarithm of the FX forward rate for country i. In this study, we focus on the 1-month interval (i.e. $j = 1$), and hence, to simplify our notation, we use $f_{i,t}$ to represent $f_{i,t,t+1}$.

We first use panel regression following the set-up in Fama (1984) to empirically test for the existence of the forward bias:

$$s_{i,t+1} - s_{i,t} = \beta_0 + \beta_1\left(f_{i,t} - s_{i,t}\right) + \varepsilon_{i,t}. \tag{11}$$

In the case of non-existence of the forward bias, i.e. $s_{i,t+1} = f_{i,t}$, in the Fama (1984) regression set-up the constant coefficient is zero and the slope coefficient is 1. Hence, when testing the slope coefficient, the null hypothesis is $\beta_1 = 1$. We also include country and time fixed effects in the panel regression. In addition, we also run the regression in each individual country setting.

Table 2
Currency Forward Bias.

Dependent Variable: $s(t+1) - s(t)$

Panel A: Panel Regression

	Slope		R2	Country FE	Time FE	N
All Countries	0.504	[−1.90]	0.002	Yes	Yes	24,840

Panel B: Country-specific Regressions

	Model 1			Model 2					
	Slope		R2	Intercept		Slope		R2	N
Austria									
Belgium									
Finland									
France									
Germany									
Greece	1.268	[0.24]	0.003	−0.001	[−0.50]	1.721	[0.51]	0.005	776
Ireland									
Netherlands									
Italy									
Spain									
Portugal									
Bulgaria	−0.052	[−0.45]	0.013	−0.005	[−1.54]	−1.874	[−1.22]	0.008	280
Bahrain	−0.006	[−80.82]	0.001	0.000	[0.09]	−0.006	[−77.19]	0.001	398
Brazil	14219.464	[1.73]	0.038	−0.016	[−1.82]	−11411.801	[−0.71]	0.007	194
Canada	0.868	[−0.06]	0.000	−0.001	[−0.26]	0.574	[−0.15]	0.000	454
China	0.818	[−0.48]	0.024	0.001	[1.69]	0.844	[−0.60]	0.051	402
H.K.	−0.097	[−6.20]	0.000	0.000	[2.03]	−0.312	[−6.87]	0.011	526
Cyprus	2.407	[0.98]	0.008	−0.001	[−0.53]	2.941	[1.14]	0.009	684
Czech	20.103	[1.22]	0.003	0.001	[0.35]	17.983	[0.96]	0.004	561
Denmark	0.892	[−0.04]	0.006	−0.008	[−1.66]	7.388	[1.49]	0.044	243
Estonia	2.533	[0.90]	0.006	0.000	[0.02]	2.517	[0.81]	0.006	531
Iceland	31.117	[0.57]	0.000	−0.002	[−0.41]	8.950	[0.17]	0.000	680
Indonesia	10097.032	[1.67]	0.019	−0.024	[−1.96]	−45960.120	[−1.54]	0.034	252
Israel	1.376	[0.31]	0.002	0.003	[1.43]	2.507	[1.19]	0.015	587
Japan	54.471	[0.58]	0.001	−0.002	[−0.87]	143.080	[1.06]	0.006	724
Jordan	−0.001	[−255.21]	0.000	−0.000	[−0.23]	−0.005	[−66.68]	0.001	143
Kazakhstan	49.052	[1.57]	0.006	−0.001	[−0.50]	41.118	[1.73]	0.008	413
S. Korea	−0.194	[−0.49]	0.000	0.000	[0.05]	−0.216	[−0.93]	0.000	582
Latvia	2.704	[1.00]	0.008	−0.001	[−0.26]	2.917	[1.02]	0.008	506
Lithuania	2.365	[0.85]	0.006	0.000	[0.01]	2.361	[0.82]	0.006	563
Malaysia	0.329	[−1.02]	0.000	0.001	[0.82]	0.692	[−0.40]	0.004	623
Nigeria	−32.723	[−4.14]	0.008	−0.000	[−0.05]	−34.327	[−1.30]	0.008	57
Norway	2.326	[0.57]	0.011	0.001	[0.64]	2.686	[0.76]	0.013	643
Poland	0.603	[−0.24]	0.001	0.006	[1.62]	2.435	[0.81]	0.009	568
Qatar	0.440	[−1.70]	0.015	−0.000	[−0.24]	0.429	[−1.45]	0.015	116
Russia	0.885	[−0.39]	0.044	0.002	[0.85]	1.018	[0.07]	0.047	585
Slovakia	2.527	[0.94]	0.007	0.000	[0.02]	2.516	[0.90]	0.007	561
Slovenia	2.334	[0.86]	0.006	0.000	[0.01]	2.329	[0.81]	0.006	565
S. Africa	0.703	[−0.41]	0.004	−0.011	[−1.03]	−1.338	[−0.97]	0.003	634
Sri Lanka	34.080	[1.25]	0.015	0.002	[0.89]	58.332	[1.46]	0.019	338
Sweden	3.005	[1.31]	0.014	−0.001	[−0.21]	3.032	[1.28]	0.014	676
Taiwan	1262.344	[0.09]	0.005	0.001	[0.42]	−2617.247	[−0.15]	0.001	109
Thailand	−4.158	[−7.83]	0.002	0.001	[0.61]	−3.848	[−10.75]	0.005	497
Turkey	0.413	[−1.38]	0.004	−0.006	[−0.78]	−0.146	[−1.17]	0.000	449
UAE	0.018	[−78.35]	0.010	0.000	[0.68]	0.017	[−81.07]	0.011	325
UK	6.464	[1.53]	0.027	−0.001	[−0.33]	6.201	[1.50]	0.027	595
Ukraine	0.952	[−0.09]	0.015	−0.010	[−1.32]	0.281	[−1.46]	0.002	313
Vietnam	7.598	[1.01]	0.087	−0.003	[−3.35]	−0.217	[−0.24]	0.000	453

This table provides the results for the currency forward bias. We follow Fama (1984) regression to test the existence of the forward bias analysis:

$$s_{i,t+1} - s_{i,t} = \beta_{i,0} + \beta_{i,1}\left(f_{i,t} - s_{i,t}\right) + \varepsilon_{i,t+1}$$

with the null hypotheses $H_0 : \beta_{i,0} = 0$ and $H_0 : \beta_{i,1} = 1$. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients.

Table 3
Sorting Forward Bias Average.

Panel A: Pooled Forward Bias Average by Sovereign CDS and Interest Rate Differential

			ΔLNSCDS				
		All	G1 (Most Negative)	G2	G3	G4 (Most Positive)	G4 − G1
	All	−1.483	83.682	20.241	−7.610	−102.244	−185.926
		[−0.86]	[26.59]	[7.21]	[−2.66]	[−23.68]	[−34.80]
IRDiff	G1 (Most Negative)	−6.393	141.174	47.717	6.216	−218.670	−359.844
		[−1.48]	[18.72]	[7.97]	[0.97]	[−20.31]	[−27.37]
	G2	−13.935	84.282	−2.769	−27.337	−87.714	−171.997
		[−3.88]	[12.27]	[−0.43]	[−4.79]	[−10.84]	[−15.68]
	G3	2.778	71.122	32.362	−8.600	−107.105	−178.227
		[0.92]	[13.81]	[5.91]	[−1.53]	[−14.99]	[−20.77]
	G4 (Most Positive)	11.308	33.367	3.239	−0.934	14.605	−18.762
		[4.34]	[6.67]	[0.69]	[−0.18]	[2.41]	[−2.38]
	G4 − G1	17.701	−107.807	−44.478	−7.150	233.275	−126.569
		[3.51]	[−11.56]	[−5.91]	[−0.88]	[18.28]	[−12.89]

Panel B: Pooled Forward Average by Region:

	ΔLNSCDS				
	G1 (Most Negative)	G2	G3	G4 (Most Positive)	G4 − G1
E. Europe	120.682	42.079	−14.497	−132.603	−253.285
	[17.02]	[6.31]	[−2.19]	[−12.17]	[−19.49]
W. Europe	61.591	9.050	−11.050	−89.058	−150.650
	[13.24]	[2.07]	[−2.41]	[−14.44]	[−19.50]
America	83.158	−20.480	−27.801	−153.480	−236.638
	[4.27]	[−1.56]	[−1.42]	[−6.31]	[−7.62]
Asia	47.153	3.706	−0.819	−63.738	−110.891
	[8.38]	[0.85]	[−0.19]	[−8.65]	[−11.96]
Middle East	115.155	56.320	31.525	−76.743	−191.899
	[10.98]	[7.63]	[4.45]	[−5.73]	[−11.30]
Africa	285.386	135.006	−48.807	−359.163	−644.548
	[11.15]	[6.30]	[−2.47]	[−9.97]	[−14.63]

This table provides the pooled average of the forward bias for different sorts. Panel A reports the results for the double sorts on the weekly changes of sovereign CDS and the interest rate differential. Panel B reports the regional results for the single sort on the weekly change of sovereign CDS.

Table 2 reports the results of the regression. Panel A reports the panel regression results. The slope coefficient is 0.504, statistically significantly different from $\beta_1 = 1$ at the 10% level. This result indicates that the monthly change of the spot rate is smaller than the forward rate.

Panel B reports the regression results for the individual countries. We perform the regressions with and without intercept. We can see that the forward bias does not exist but only in some countries, e.g. mid-east countries (Bahrain, Jordan and UAE) and some Asian countries (Thailand and Hong Kong). Moreover, the forward bias results are not as strong as documented in the literature (Fama, 1984). There are several possible explanations. First of all, we consider weekly averaged values and hence the smoothing may mitigate the observed forward bias. In addition, for each country-specific regression, the statistical significance may suffer from the lower number of observations. However, the result highlights an important implication. Recall that the start of our sample period is set at the onset of the sovereign CDS market and that, as shown in our theoretical framework described in the previous section, the currency forward bias is negatively related to sovereign credit risk. Consequently, the "appearance" of the sovereign CDS market may actually reduce the absolute values of the forward bias. We will test this conjecture in Section 6.7.

5. Currency forward bias and sovereign CDS: Sorting exercise

The theoretical model derived in the previous section illustrates that both the sovereign CDS spread and the interest rate differential are negatively related to the currency forward bias. For our empirical analysis, we first identify the possible linkages of the sovereign CDS spread and interest rate differential with the forward bias. We conjecture a negative relation between the forward bias and sovereign CDS spreads and the interest rate differential. We also acknowledge that the sort for sovereign CDS and interest rate differential is quantitatively constrained by the number of observations available on a weekly basis. Hence, we employ pooled sorting for sovereign CDS spreads and the interest rate differential, instead of traditional cross-sectional sorting methods.

Table 3 reports the double sorting estimated results. The left-upper value in Panel A is the pooled average of the forward bias in our sample. The pooled average is −1.483 bps, but it is statistically insignificant.

The first row (column) presents the results of the sorting exercises of the forward bias by sovereign CDS spreads and by interest rate differential, while the last number of the row or column shows the difference of the forward bias between the highest and the lowest of the groups. We can see that the higher (lower) the sovereign CDS in absolute terms, the lower the forward bias. By contrast and consistent with the extant literature, we document a positive relation between the forward bias and the interest rate differential.

Table 4
Baseline Regression.

	Dependent Variable: Forward Bias							
	Intercept		$\Delta LNSCDS$		*IRDiff*		R-sqr	N
Austria	−0.000	[−0.05]	−0.036	[−3.14]	0.294	[1.44]	0.083	724
Belgium	−0.001	[−0.29]	−0.038	[−2.87]	0.237	[1.15]	0.071	751
Bulgaria	−0.005	[−1.65]	−0.088	[−8.65]	−0.264	[−1.51]	0.299	280
Bahrain	−0.003	[−3.99]	−0.001	[−1.07]	−0.589	[−6.13]	0.632	398
Brazil	−0.010	[−1.55]	−0.131	[−9.58]	−0.005	[−0.04]	0.513	194
Canada	−0.001	[−0.25]	−0.014	[−4.17]	0.069	[0.32]	0.017	454
China	0.000	[0.49]	−0.002	[−0.49]	−0.035	[−1.42]	0.015	402
H.K.	0.000	[0.37]	−0.001	[−1.24]	−0.099	[−6.05]	0.180	526
Cyprus	0.000	[0.02]	−0.018	[−1.78]	0.418	[1.98]	0.031	684
Czech	0.002	[0.67]	−0.061	[−2.10]	0.260	[1.74]	0.117	561
Germany	−0.001	[−0.41]	−0.033	[−2.76]	0.184	[0.90]	0.057	750
Denmark	−0.008	[−1.76]	0.009	[0.31]	0.807	[1.66]	0.044	243
Estonia	0.001	[0.29]	−0.016	[−0.98]	0.381	[1.23]	0.025	531
Finland	−0.000	[−0.07]	−0.027	[−2.14]	0.350	[1.65]	0.043	724
France	−0.000	[−0.14]	−0.039	[−3.23]	0.290	[1.40]	0.080	731
Greece	0.001	[0.44]	−0.035	[−3.68]	0.288	[1.39]	0.088	650
Iceland	0.009	[1.29]	−0.073	[−3.99]	0.149	[1.33]	0.171	680
Indonesia	−0.011	[−1.89]	−0.070	[−5.51]	−0.109	[−0.86]	0.379	252
Ireland	−0.000	[−0.25]	−0.030	[−1.94]	0.221	[1.00]	0.050	746
Israel	0.003	[1.39]	−0.038	[−1.67]	0.120	[1.13]	0.058	587
Italy	−0.000	[−0.25]	−0.031	[−2.40]	0.196	[0.94]	0.051	776
Japan	−0.002	[−0.67]	0.022	[1.53]	0.081	[0.85]	0.026	724
Jordan	0.025	[5.82]	0.002	[0.74]	0.053	[0.44]	0.005	143
Kazakhstan	0.001	[0.44]	−0.011	[−1.81]	0.096	[1.61]	0.026	413
S. Korea	0.001	[0.56]	−0.094	[−4.84]	0.086	[0.79]	0.357	582
Latvia	0.001	[0.24]	−0.034	[−1.61]	0.427	[1.42]	0.057	506
Lithuania	0.001	[0.30]	−0.031	[−1.69]	0.374	[1.36]	0.050	563
Malaysia	0.002	[1.55]	−0.037	[−5.56]	0.034	[0.60]	0.245	623
Netherlands	0.000	[0.19]	−0.033	[−2.76]	0.446	[2.27]	0.076	638
Nigeria	−0.011	[−1.25]	−0.163	[−4.64]	−0.082	[−1.39]	0.282	57
Norway	0.001	[0.61]	−0.064	[−3.31]	0.068	[0.45]	0.123	643
Poland	0.009	[1.82]	−0.115	[−3.76]	0.234	[1.46]	0.311	568
Portugal	−0.000	[−0.25]	−0.028	[−2.33]	0.209	[0.95]	0.047	757
Qatar	−0.000	[−3.60]	−0.002	[−3.34]	−0.074	[−4.31]	0.329	116
Russia	0.002	[0.76]	−0.083	[−5.28]	−0.008	[−0.17]	0.410	585
Slovakia	0.001	[0.29]	−0.037	[−2.06]	0.273	[1.03]	0.084	561
Slovenia	0.001	[0.37]	−0.037	[−2.37]	0.207	[0.81]	0.081	565
S. Africa	−0.012	[−1.72]	−0.165	[−6.96]	−0.247	[−2.05]	0.446	634
Spain	−0.000	[−0.22]	−0.039	[−2.43]	0.228	[1.04]	0.066	763
Sri Lanka	−0.000	[−0.05]	−0.021	[−2.30]	0.014	[0.39]	0.064	338
Sweden	0.000	[0.05]	−0.058	[−3.36]	0.156	[1.45]	0.142	676
Taiwan	0.001	[0.52]	−0.021	[−4.09]	0.071	[0.69]	0.048	109
Thailand	0.003	[1.47]	−0.036	[−5.20]	0.156	[1.65]	0.131	497
Turkey	−0.008	[−1.37]	−0.178	[−6.92]	−0.120	[−1.82]	0.474	449
UAE	−0.000	[−3.12]	0.000	[0.90]	−0.050	[−4.11]	0.295	325
UK	0.000	[0.21]	−0.034	[−2.98]	0.398	[1.48]	0.096	595
Ukraine	0.000	[0.11]	−0.078	[−2.30]	0.022	[0.28]	0.131	313
Vietnam	0.000	[0.18]	−0.003	[−0.88]	0.040	[2.23]	0.044	453

This table reports the baseline regression results for the sovereign CDS on the currency forward bias. The specification of the regression is: $fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\,\Delta LNSCDS_{i,t+1} + \beta_{i,2}\,IRDiff_{i,t} + \varepsilon_{i,t+1}$.
Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

Moreover, we observe that the forward bias is steady monotonic decreasing for sorted sovereign CDS spreads, and the largest positive changes in sovereign CDS spreads contribute to the forward bias by 186 bps on average, compared to highest negative changes in sovereign CDS spreads.

Panel A reports the pooled forward bias estimates sorted by both the sovereign CDS and the interest rate differential. Note that the averaged forward bias for the most positive sovereign CDS and interest rate differential is 127 bps smaller than that for the most negative sovereign CDS and interest rate differential.

Overall, this set of results suggests that sovereign CDS is negatively associated with the forward bias, conditional on different magnitudes of interest rate differential (see the last column of Panel A). On the other hand, the conditional (on sovereign CDS) association between the interest rate differential and the forward bias is in general negative, although non-monotonic decreasing (see the last column of Panel A). Overall, the sorting results empirically support our model's predictions, that is, the forward bias has a negative

Table 5
Sub-sample Baseline Regression.

	Dependent Variable: Forward Bias							
	Sample Period: 2003–2009				Sample Period: 2010–2020			
	Δ*LNSCDS*		Δ*IRDiff*		Δ*LNSCDS*		Δ*IRDiff*	
Austria	−0.036	[−2.44]	0.099	[0.34]	−0.045	[−3.05]	0.244	[0.68]
Belgium	−0.027	[−1.82]	0.236	[0.71]	−0.059	[−4.74]	0.082	[0.33]
Bulgaria	−0.090	[−7.13]	−0.805	[−1.19]	−0.078	[−6.10]	0.803	[1.22]
Bahrain	−0.001	[−0.62]	−0.541	[−6.97]	0.000	[0.07]	−0.881	[−3.63]
Brazil					−0.131	[−9.58]	−0.005	[−0.04]
Canada	−0.010	[−3.18]	0.043	[0.15]	−0.035	[−2.51]	0.567	[0.74]
China	0.003	[1.14]	−0.002	[−0.05]	−0.009	[−1.95]	−0.109	[−2.34]
H.K.	−0.000	[−0.50]	−0.127	[−6.62]	−0.004	[−3.60]	−0.275	[−2.34]
Cyprus	−0.016	[−1.37]	0.287	[0.75]	−0.023	[−1.35]	0.404	[1.13]
Czech	−0.045	[−1.36]	0.177	[0.98]	−0.150	[−6.13]	−0.345	[−0.33]
Germany	−0.031	[−1.98]	0.159	[0.54]	−0.040	[−2.63]	0.064	[0.24]
Denmark					0.009	[0.31]	0.807	[1.66]
Estonia	−0.008	[−0.49]	0.304	[0.71]	−0.098	[−3.19]	−0.435	[−0.51]
Finland	−0.020	[−1.33]	0.222	[0.68]	−0.050	[−3.50]	0.293	[0.81]
France	−0.036	[−2.00]	0.162	[0.54]	−0.046	[−3.72]	0.246	[0.81]
Greece	−0.033	[−2.15]	0.177	[0.52]	−0.037	[−3.74]	0.231	[0.72]
Iceland	−0.077	[−4.04]	0.216	[1.08]	−0.013	[−0.50]	−0.190	[−1.13]
Indonesia					−0.070	[−5.51]	−0.109	[−0.86]
Ireland	−0.036	[−1.66]	0.124	[0.37]	−0.031	[−2.21]	0.110	[0.38]
Israel	−0.025	[−1.03]	0.149	[1.16]	−0.095	[−3.78]	0.029	[0.25]
Italy	−0.028	[−1.77]	0.230	[0.68]	−0.036	[−2.07]	0.079	[0.31]
Japan	0.023	[1.08]	−0.027	[−0.22]	0.019	[1.24]	0.453	[0.88]
Jordan					0.002	[0.74]	0.053	[0.44]
Kazakhstan	−0.011	[−1.63]	0.092	[1.30]	−0.018	[−1.54]	0.511	[1.41]
S. Korea	−0.092	[−3.61]	0.378	[1.72]	−0.097	[−5.44]	−0.188	[−1.85]
Latvia	−0.026	[−1.12]	0.330	[0.83]	−0.087	[−4.13]	−0.109	[−0.15]
Lithuania	−0.020	[−1.04]	0.310	[0.88]	−0.087	[−3.92]	0.111	[0.18]
Malaysia	−0.027	[−3.03]	0.123	[1.51]	−0.049	[−4.23]	−0.184	[−1.04]
Netherlands	−0.028	[−2.15]	0.366	[1.12]	−0.053	[−2.53]	0.236	[0.65]
Nigeria					−0.163	[−4.64]	−0.082	[−1.39]
Norway	−0.067	[−2.74]	0.054	[0.28]	−0.064	[−2.73]	−0.228	[−0.67]
Poland	−0.098	[−2.73]	0.189	[0.91]	−0.177	[−7.25]	0.246	[1.20]
Portugal	−0.041	[−1.89]	0.228	[0.68]	−0.023	[−2.03]	0.080	[0.29]
Qatar					−0.002	[−3.34]	−0.074	[−4.31]
Russia	−0.040	[−3.75]	−0.064	[−1.09]	−0.138	[−15.97]	0.061	[0.49]
Slovakia	−0.028	[−1.38]	0.264	[0.83]	−0.080	[−4.49]	−1.045	[−1.25]
Slovenia	−0.032	[−1.63]	0.231	[0.73]	−0.053	[−3.11]	−0.459	[−0.80]
S. Africa	−0.150	[−4.76]	−0.205	[−1.52]	−0.202	[−11.88]	−0.568	[−1.75]
Spain	−0.037	[−1.58]	0.254	[0.75]	−0.048	[−3.15]	0.058	[0.22]
Sri Lanka	−0.015	[−1.40]	0.074	[1.54]	−0.030	[−2.11]	−0.014	[−0.15]
Sweden	−0.058	[−2.93]	0.312	[1.87]	−0.066	[−4.05]	−0.135	[−0.76]
Taiwan					−0.019	[−3.94]	0.818	[1.47]
Thailand	−0.030	[−3.76]	0.184	[0.98]	−0.052	[−4.52]	0.143	[1.06]
Turkey	−0.194	[−7.35]	−0.005	[−0.03]	−0.159	[−4.09]	−0.122	[−1.46]
UAE	0.000	[0.81]	−0.060	[−4.54]	0.000	[1.34]	−0.006	[−2.64]
UK	−0.024	[−2.02]	0.840	[2.59]	−0.052	[−2.96]	0.158	[0.58]
Ukraine	−0.107	[−3.36]	−0.077	[−0.43]	−0.014	[−0.62]	0.019	[0.21]
Vietnam	−0.003	[−0.61]	0.052	[1.78]	−0.004	[−0.55]	0.040	[1.25]

This table reports the sub-sample results for sovereign CDS on the currency forward bias. The regression specification is:
$fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\,\Delta LNSCDS_{i,t+1} + \beta_{i,2}\,IRDiff_{i,t} + \varepsilon_{i,t+1}$.
Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses. To save space, the intercept term is included in the regression but not reported.

relationship with sovereign CDS spreads and the interest rate differential.

Next, we further explore the robustness of our main findings with respect to our sample of countries. As we can see from Panel B, our results confirm the evidence of a robust statistically negative association between the forward bias and sovereign CDS spreads for all of the regions.

6. Currency forward bias and sovereign CDS: Regression analysis

6.1. Baseline regression

In this section, we present our baseline regression analyses on the determinants of the currency forward bias. Recall that in our model the forward bias is estimated as the negative sovereign CDS multiplied by the interest rate differential. Guided by our theory and previous analysis of our model, we now run the following baseline regression:

$$fwdbias_{i,t+1} = s_{i,t+1} - f_{i,t} = \beta_{i,0} + \beta_{i,1}\Delta LNSCDS_{i,t+1} + \beta_{i,2} IRDiff_{i,t} + \varepsilon_{i,t+1}, \tag{12}$$

where *LNSCDS* is the logarithm of the sovereign CDS spread, *IRDiff* is the interest rate differential (i.e. $i^{US} - i^{FX}$), and Δ is the operator between months $t+1$ and t. A prediction of the model is that β_1 and β_2 have negative coefficients. As noted above, we estimate the baseline regression framework for each individual country.

Table 4 reports the results for the 48 countries and shows that both the coefficients for the Δ*LNSCDS* and *IRDiff* are negative for most of the countries.

Furthermore, 38 countries exhibit negative coefficients for Δ*LNSCDS* at 10% level. These regression results therefore accord with our previous findings further suggesting that sovereign credit risk premia are an underlying important force driving the forward bias. However, we find weak evidence of the positive association with *IRDiff* for 7 countries[7] at 10% level. We can see that the negative impact of sovereign credit risk on forward bias is unambiguous, while the interest rate differential impacts primarily in developed countries (only two countries—Tailand and Vietnam—are developing countries). Overall, these results indicate that the interest rate differential has a small effect on the forward bias, and thus confirm our main results that the sovereign CDS premia is a much stronger determinant of the forward bias.

One aspect that is worth noting is that in general we do not find significant intercept coefficients in the baseline regressions, suggesting that no abnormal premium in the forward bias derive from the sovereign CDS premia and interest rate differential variables.[8]

We further perform our main empirical analysis on the sub-sample periods 2003–2009 and 2010–2020. The results are reported in Table 5. Overall, our empirical evidence aligns with our full sample period main findings. Interestingly, we also find that the statistical significance for the interest rate differential is stronger over the second-half sub-sample period 2010–2020.

6.2. Alternative currency forward bias measure

In our previous regression, we define the FX forward bias by $s_{i,t+1} - f_{i,t}$. Here, we also consider a different measure of the forward bias as a robustness check of our results. Formally, the forward bias is given by:

$$fwdbias_{i,t+1} = \left(s_{t+1} - s_{i,t}\right) - \widehat{\beta}_{i,1}\left(f_{i,t} - s_{i,t}\right), \tag{13}$$

where $\widehat{\beta}_{i,1}$ is the slope coefficient in Fama (1984) regression fitted in the previous section. The equation is a generalized forward bias. If $\widehat{\beta}_{i,1}$ is equal to 1, then the forward bias is the original definition. The generalized forward bias assumes that there is a persistent underlying driving factor for the existence of the forward bias, e.g. transaction costs, or other illiquidity problems (e.g. Breedon et al, 2016; Lee and Jung, 2019). Hence, the slope coefficient of the Fama (1984) regression is used to identify the existence of the forward bias. As such, the generalized forward bias takes the persistent bias into consideration.

Next, we employ the proposed alternative forward bias measure as a new regressor in our baseline model. The results are reported in Table 6, which shows that our results remain unchanged with this alternative forward bias measure.

6.3. Forward bias and local sovereign credit risk

Research on sovereign credit risk documents a strong co-movement of sovereign CDS spreads among countries (e.g. Longstaff et al., 2011), indicating that sovereign CDS spreads may also contain information related to global sovereign default risk. As such, sovereign CDS spreads may also track non-local sovereign credit risk. To investigate whether local sovereign credit risk actually contributes to the forward bias, we decompose the sovereign CDS into the two separate components of global and local sovereign risk. We can then express the relationship of interest as:

$$\Delta LNSCDS_{i,t+1} = \alpha_{i,0} + \alpha_{i,1}\Delta LNSCDS^{US}_{t+1} + e_{i,t+1}, \tag{14}$$

[7] The 7 countries are: Cyprus, Czeck, Denmark, Finland, Netherlands, Tailand, and Vietnam.

[8] It is a challenge to adopt a right proxy for the short-term risk-free rate in cross-countries studies. To confirm that our results are not affected by the choice of risk-free rates, we also use the overnight REPO rate as another proxy of the risk-free rate and repeat the main regression analyses. We find that the results are qualitatively the same. To save space, we do not report the results, but they are available upon request. However, we acknowledge that all these proxies may not actually be risk-free, as they may still likely contain counter-party risk. Du and Schreger (2016) and Borri and Shakhnov (2021) also provide alternative estimations of the risk-free rate across different countries.

Table 6
Alternative Forward Bias.

	Dependent Variable: Alternative Forward Bias							
	Intercept		$\Delta LNSCDS$		*IRDiff*		R-sqr	N
Austria	−0.001	[−0.44]	−0.036	[−3.11]	0.127	[0.62]	0.076	724
Belgium	−0.001	[−0.50]	−0.037	[−2.84]	0.145	[0.70]	0.067	751
Bulgaria	−0.005	[−1.62]	−0.086	[−8.81]	0.028	[0.16]	0.295	280
Bahrain	−0.000	[−0.45]	0.000	[0.46]	−0.004	[−0.50]	0.010	398
Brazil	−0.008	[−1.27]	−0.132	[−9.67]	0.103	[0.98]	0.508	194
Canada	−0.001	[−0.24]	−0.014	[−4.16]	0.107	[0.50]	0.017	454
China	0.001	[0.63]	−0.002	[−0.50]	−0.027	[−1.07]	0.009	402
H.K.	0.000	[2.09]	−0.001	[−1.16]	−0.002	[−0.10]	0.022	526
Cyprus	−0.001	[−0.48]	−0.017	[−1.70]	0.232	[1.10]	0.020	684
Czech	0.001	[0.53]	−0.062	[−2.09]	0.130	[0.86]	0.111	561
Germany	−0.001	[−0.59]	−0.033	[−2.75]	0.107	[0.52]	0.054	750
Denmark	−0.009	[−1.97]	0.008	[0.26]	0.139	[0.28]	0.002	243
Estonia	0.000	[0.12]	−0.017	[−0.98]	0.184	[0.59]	0.019	531
Finland	−0.001	[−0.46]	−0.026	[−2.10]	0.184	[0.87]	0.034	724
France	−0.001	[−0.46]	−0.039	[−3.19]	0.155	[0.74]	0.073	731
Greece	0.000	[0.03]	−0.034	[−3.57]	0.128	[0.61]	0.078	650
Iceland	0.009	[1.32]	−0.073	[−3.98]	0.146	[1.30]	0.171	680
Indonesia	−0.017	[−2.75]	−0.068	[−5.30]	0.138	[1.04]	0.349	252
Ireland	−0.001	[−0.46]	−0.030	[−1.92]	0.131	[0.59]	0.045	746
Israel	0.003	[1.25]	−0.038	[−1.67]	0.008	[0.07]	0.054	587
Italy	−0.001	[−0.45]	−0.031	[−2.39]	0.117	[0.56]	0.048	776
Japan	−0.002	[−0.83]	0.022	[1.52]	−0.033	[−0.34]	0.022	724
Jordan	−0.000	[−0.09]	−0.000	[−0.20]	0.001	[0.13]	0.000	143
Kazakhstan	0.001	[0.49]	−0.011	[−1.82]	0.068	[1.15]	0.019	413
S. Korea	0.001	[0.58]	−0.094	[−4.84]	0.088	[0.81]	0.357	582
Latvia	0.000	[0.02]	−0.034	[−1.60]	0.179	[0.59]	0.050	506
Lithuania	0.000	[0.16]	−0.031	[−1.69]	0.193	[0.70]	0.046	563
Malaysia	0.002	[1.61]	−0.037	[−5.57]	0.058	[1.05]	0.246	623
Netherlands	−0.001	[−0.50]	−0.033	[−2.70]	0.189	[0.95]	0.060	638
Nigeria	−0.010	[−1.17]	−0.156	[−4.18]	−0.075	[−1.26]	0.258	57
Norway	0.001	[0.60]	−0.064	[−3.28]	−0.069	[−0.46]	0.116	643
Poland	0.009	[1.82]	−0.115	[−3.75]	0.119	[0.74]	0.308	568
Portugal	−0.001	[−0.47]	−0.028	[−2.31]	0.116	[0.52]	0.043	757
Qatar	−0.000	[−2.64]	−0.002	[−3.61]	−0.040	[−2.54]	0.325	116
Russia	0.002	[0.75]	−0.083	[−5.28]	−0.010	[−0.23]	0.410	585
Slovakia	0.000	[0.12]	−0.037	[−2.06]	0.071	[0.27]	0.079	561
Slovenia	0.000	[0.22]	−0.037	[−2.37]	0.028	[0.11]	0.077	565
S. Africa	−0.012	[−1.78]	−0.165	[−6.98]	−0.047	[−0.39]	0.440	634
Spain	−0.001	[−0.41]	−0.039	[−2.41]	0.146	[0.66]	0.062	763
Sri Lanka	−0.002	[−0.35]	−0.020	[−2.14]	−0.029	[−0.79]	0.056	338
Sweden	−0.000	[−0.11]	−0.059	[−3.35]	−0.016	[−0.15]	0.134	676
Taiwan	0.001	[0.67]	−0.021	[−4.15]	0.088	[0.84]	0.057	109
Thailand	0.003	[1.37]	−0.036	[−5.21]	0.149	[1.60]	0.129	497
Turkey	−0.007	[−1.13]	−0.179	[−7.06]	−0.013	[−0.19]	0.468	449
UAE	−0.000	[−0.05]	0.000	[0.69]	−0.000	[−0.43]	0.004	325
UK	−0.000	[−0.16]	−0.034	[−2.94]	−0.018	[−0.07]	0.061	595
Ukraine	−0.001	[−0.28]	−0.075	[−2.14]	0.077	[0.98]	0.128	313
Vietnam	0.000	[0.18]	−0.003	[−0.89]	0.040	[2.23]	0.044	453

This table reports the sovereign CDS and the currency forward bias, where the currency forward bias is constructed with alternative methods. The regression specification is:

$$fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\,\Delta LNSCDS_{i,t+1} + \beta_{i,2}\,IRDiff_{i,t} + \varepsilon_{i,t+1}$$

with $fwdbias_{i,t+1} = (s_{t+1} - s_{i,t}) - \widehat{\beta}_{i,1}\left(f_{i,t} - s_{i,t}\right)$, where $\widehat{\beta}_{i,1}$ is the slope coefficient in Fama (1984) regression fitted in the previous section. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

where $LNSCDS^{US}_{t+1}$ is the logarithm of the US sovereign CDS spread. In our setting, we use US sovereign CDS as a proxy for global sovereign credit risk. It is important to note that in the literature, there is no consensus on a specific measure of global sovereign risk. For instance, Longstaff et al. (2011) use the first principal component method while Della Corte et al. (2022) use cross-sectional average to analyze commonalities in global sovereign risk. We choose the US sovereign CDS as a proxy for two reasons. First, the US is the most important financial market. Moreover, US sovereign risk also affects other countries' sovereign risk (Longstaff et al., 2011). The second reason is that since the FX rates are all expressed against the US dollar, we aim also to control for the sovereign risk of the US economy. From the equation, $\alpha_{i,1}\Delta LNSCDS^{US}_{t+1}$ denotes global sovereign risk and the local sovereign risk is equal to

Table 7
Local Sovereign Risk.

	Dependent Variable: Forward Bias							
	Intercept		$\Delta LNSCDS^{Local}$		*IRDiff*		R-sqr	N
Austria	−0.000	[−0.16]	−0.048	[−4.26]	0.291	[1.45]	0.107	724
Belgium	−0.001	[−0.36]	−0.048	[−3.69]	0.261	[1.26]	0.089	751
Bulgaria	−0.004	[−1.35]	−0.095	[−6.42]	−0.177	[−0.90]	0.285	280
Bahrain	−0.003	[−3.92]	−0.001	[−1.12]	−0.590	[−6.03]	0.633	398
Brazil	−0.010	[−1.54]	−0.132	[−9.58]	−0.006	[−0.06]	0.514	194
Canada	−0.001	[−0.25]	−0.009	[−1.95]	0.059	[0.27]	0.007	454
China	0.000	[0.46]	−0.002	[−0.76]	−0.035	[−1.42]	0.016	402
H.K.	0.000	[0.37]	−0.001	[−1.47]	−0.100	[−6.27]	0.181	526
Cyprus	−0.000	[−0.01]	−0.020	[−2.09]	0.424	[1.98]	0.032	684
Czech	0.001	[0.51]	−0.073	[−2.62]	0.271	[1.73]	0.128	561
Germany	−0.001	[−0.51]	−0.042	[−3.03]	0.189	[0.93]	0.071	750
Denmark	−0.008	[−1.76]	0.007	[0.25]	0.805	[1.65]	0.044	243
Estonia	0.001	[0.24]	−0.018	[−1.00]	0.402	[1.23]	0.026	531
Finland	−0.000	[−0.14]	−0.036	[−2.64]	0.349	[1.64]	0.052	724
France	−0.000	[−0.23]	−0.049	[−3.56]	0.303	[1.46]	0.098	731
Greece	0.001	[0.41]	−0.039	[−4.28]	0.294	[1.40]	0.098	650
Iceland	0.010	[1.41]	−0.080	[−4.24]	0.174	[1.43]	0.172	680
Indonesia	−0.011	[−1.79]	−0.069	[−5.08]	−0.107	[−0.81]	0.364	252
Ireland	−0.001	[−0.30]	−0.036	[−2.07]	0.230	[1.04]	0.057	746
Israel	0.003	[1.32]	−0.043	[−1.92]	0.117	[1.16]	0.062	587
Italy	−0.001	[−0.29]	−0.036	[−2.67]	0.206	[0.97]	0.058	776
Japan	−0.002	[−0.69]	0.016	[1.14]	0.092	[0.92]	0.015	724
Jordan	0.025	[5.82]	0.002	[0.83]	0.053	[0.44]	0.005	143
Kazakhstan	0.001	[0.46]	−0.013	[−1.99]	0.099	[1.67]	0.028	413
S. Korea	0.001	[0.44]	−0.091	[−4.96]	0.098	[0.83]	0.297	582
Latvia	0.000	[0.18]	−0.038	[−1.71]	0.466	[1.48]	0.061	506
Lithuania	0.000	[0.22]	−0.038	[−2.07]	0.412	[1.43]	0.060	563
Malaysia	0.002	[1.39]	−0.039	[−6.02]	0.019	[0.38]	0.243	623
Netherlands	0.000	[0.09]	−0.043	[−3.43]	0.449	[2.31]	0.090	638
Nigeria	−0.013	[−1.56]	−0.164	[−4.49]	−0.097	[−1.73]	0.286	57
Norway	0.001	[0.54]	−0.064	[−2.72]	0.095	[0.58]	0.098	643
Poland	0.008	[1.93]	−0.130	[−4.79]	0.231	[1.55]	0.313	568
Portugal	−0.001	[−0.29]	−0.032	[−2.52]	0.217	[0.98]	0.052	757
Qatar	−0.000	[−3.61]	−0.002	[−3.19]	−0.075	[−4.41]	0.317	116
Russia	0.001	[0.60]	−0.085	[−5.71]	−0.011	[−0.23]	0.403	585
Slovakia	0.000	[0.17]	−0.045	[−2.53]	0.295	[1.08]	0.100	561
Slovenia	0.001	[0.29]	−0.043	[−2.61]	0.222	[0.85]	0.092	565
S. Africa	−0.009	[−1.09]	−0.165	[−6.53]	−0.185	[−1.15]	0.404	634
Spain	−0.000	[−0.26]	−0.047	[−2.80]	0.247	[1.10]	0.077	763
Sri Lanka	−0.000	[−0.03]	−0.022	[−2.33]	0.015	[0.42]	0.070	338
Sweden	−0.000	[−0.09]	−0.060	[−3.04]	0.179	[1.53]	0.110	676
Taiwan	0.001	[0.51]	−0.019	[−3.45]	0.071	[0.68]	0.044	109
Thailand	0.003	[1.30]	−0.035	[−4.45]	0.144	[1.48]	0.115	497
Turkey	−0.007	[−1.03]	−0.182	[−6.77]	−0.101	[−1.42]	0.468	449
UAE	−0.000	[−3.09]	0.000	[1.76]	−0.050	[−4.10]	0.295	325
UK	0.001	[0.26]	−0.041	[−3.01]	0.456	[1.65]	0.108	595
Ukraine	−0.000	[−0.05]	−0.070	[−2.28]	0.018	[0.22]	0.089	313
Vietnam	0.000	[0.16]	−0.005	[−1.56]	0.040	[2.27]	0.047	453

This table reports the local sovereign risk impact on the currency forward bias. The model specification is:
$fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\,\Delta LNSCDS^{Local}_{i,t+1} + \beta_{i,2}\,IRDiff_{i,t} + \varepsilon_{i,t+1}$
where $\Delta LNSCDS^{Local}_{i,t+1}$ is the unexplained part of the $\Delta LNSCDS_{i,t+1} = \alpha_{i,0} + \alpha_{i,1}\,\Delta LNSCDS^{US}_{t+1} + e_{i,t+1}$. Mathematically, $\Delta LNSCDS^{Local}_{i,t+1} = \widehat{\alpha}_{i,0} + e_{i,t+1}$. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

$(\alpha_{i,0} + e_{i,t+1})$. Hence, we can rewrite our baseline regression specification as:

$$fwdbias_{i,t+1} = s_{i,t+1} - f_{i,t} = \beta_{i,0} + \beta_{i,1}\Delta LNSCDS^{Local}_{i,t+1} + \beta_{i,2} IRDiff_{i,t} + \varepsilon_{i,t+1}. \quad (15)$$

We expect, again, both $\beta_{i,1}$ and $\beta_{i,2}$ coefficients to be negative.

The results are reported in Table 7. We can clearly see that local sovereign credit risk contributes to the forward bias. Furthermore, we observe a negative $\beta_{i,1}$ in most of the countries. Moreover, consistent with our prior findings, $\beta_{i,2}$ are statistically significant only for a few countries.

Table 8
Controlling for UIRP.

	Dependent Variable: Forward Bias									
	Intercept		$\Delta LNSCDS$		$\Delta UIRP$		*IRDiff*		R-sqr	N
Austria	0.000	[0.01]	−0.036	[−3.26]	−6.472	[−0.91]	0.284	[1.42]	0.089	724
Belgium	−0.000	[−0.24]	−0.037	[−2.83]	−4.044	[−0.51]	0.236	[1.17]	0.074	751
Bulgaria	−0.005	[−1.65]	−0.087	[−8.84]	4.664	[1.15]	−0.264	[−1.48]	0.303	280
Bahrain	−0.003	[−3.98]	−0.001	[−0.86]	0.318	[2.68]	−0.584	[−6.12]	0.654	398
Brazil	−0.010	[−1.56]	−0.131	[−9.55]	21530.661	[0.65]	−0.006	[−0.05]	0.515	194
Canada	−0.000	[−0.15]	−0.011	[−3.27]	55.221	[3.71]	0.092	[0.47]	0.100	454
China	0.000	[0.47]	−0.003	[−0.91]	0.683	[1.58]	−0.036	[−1.50]	0.023	402
H.K.	0.000	[0.37]	−0.001	[−1.27]	−0.533	[−0.44]	−0.099	[−5.99]	0.182	526
Cyprus	0.000	[0.07]	−0.018	[−1.82]	−6.085	[−0.80]	0.408	[1.94]	0.036	684
Czech	0.002	[0.66]	−0.061	[−2.10]	15.838	[0.64]	0.261	[1.74]	0.117	561
Germany	−0.001	[−0.36]	−0.032	[−2.66]	−4.109	[−0.53]	0.182	[0.91]	0.059	750
Denmark	−0.008	[−1.80]	0.005	[0.18]	−7.692	[−1.61]	0.822	[1.73]	0.052	243
Estonia	0.001	[0.29]	−0.016	[−0.98]	−1.306	[−0.13]	0.378	[1.19]	0.025	531
Finland	−0.000	[−0.02]	−0.027	[−2.23]	−5.517	[−0.65]	0.343	[1.64]	0.047	724
France	−0.000	[−0.09]	−0.039	[−3.22]	−4.792	[−0.59]	0.288	[1.42]	0.083	731
Greece	0.001	[0.51]	−0.035	[−3.88]	−6.367	[−0.83]	0.278	[1.35]	0.094	650
Iceland	0.009	[1.31]	−0.073	[−4.01]	−53.506	[−1.22]	0.152	[1.35]	0.174	680
Indonesia	−0.010	[−1.75]	−0.070	[−5.71]	38992.845	[0.99]	−0.100	[−0.77]	0.385	252
Ireland	−0.000	[−0.19]	−0.030	[−2.06]	−5.815	[−0.74]	0.214	[0.99]	0.054	746
Israel	0.003	[1.42]	−0.039	[−1.76]	−7.638	[−1.62]	0.120	[1.14]	0.067	587
Italy	−0.000	[−0.20]	−0.031	[−2.42]	−4.374	[−0.69]	0.191	[0.93]	0.054	776
Japan	−0.002	[−0.66]	0.022	[1.50]	−271.635	[−0.46]	0.080	[0.84]	0.027	724
Jordan	0.023	[5.56]	−0.001	[−0.40]	0.477	[8.65]	0.009	[0.07]	0.183	143
Kazakhstan	0.001	[0.40]	−0.011	[−1.83]	44.858	[1.28]	0.093	[1.63]	0.035	413
S. Korea	0.001	[0.56]	−0.094	[−4.83]	4.627	[2.93]	0.086	[0.79]	0.358	582
Latvia	0.001	[0.24]	−0.034	[−1.63]	−2.097	[−0.23]	0.423	[1.38]	0.057	506
Lithuania	0.001	[0.30]	−0.031	[−1.70]	−1.195	[−0.12]	0.372	[1.32]	0.050	563
Malaysia	0.002	[1.55]	−0.035	[−4.78]	−5.509	[−1.13]	0.032	[0.57]	0.255	623
Netherlands	0.000	[0.20]	−0.033	[−2.71]	−4.186	[−0.50]	0.440	[2.24]	0.079	638
Nigeria	−0.011	[−1.27]	−0.164	[−4.56]	−22.607	[−0.81]	−0.083	[−1.42]	0.284	57
Norway	0.001	[0.63]	−0.062	[−2.99]	−7.562	[−0.49]	0.069	[0.47]	0.125	643
Poland	0.009	[1.82]	−0.115	[−3.75]	−4.282	[−0.72]	0.233	[1.46]	0.312	568
Portugal	−0.000	[−0.19]	−0.029	[−2.58]	−6.316	[−0.81]	0.201	[0.93]	0.053	757
Qatar	−0.000	[−3.38]	−0.002	[−3.29]	−0.013	[−0.03]	−0.074	[−4.48]	0.329	116
Russia	0.002	[0.68]	−0.083	[−5.39]	−0.662	[−1.62]	−0.011	[−0.26]	0.415	585
Slovakia	0.001	[0.29]	−0.037	[−2.08]	−1.627	[−0.18]	0.270	[1.00]	0.084	561
Slovenia	0.001	[0.37]	−0.037	[−2.40]	−1.465	[−0.16]	0.204	[0.79]	0.081	565
S. Africa	−0.012	[−1.72]	−0.166	[−7.10]	−0.782	[−0.24]	−0.248	[−2.04]	0.446	634
Spain	−0.000	[−0.17]	−0.039	[−2.51]	−5.216	[−0.68]	0.222	[1.03]	0.070	763
Sri Lanka	0.000	[0.09]	−0.021	[−2.41]	92.782	[1.04]	0.020	[0.62]	0.073	338
Sweden	0.000	[0.06]	−0.058	[−3.30]	−0.508	[−0.06]	0.157	[1.48]	0.142	676
Taiwan	0.001	[0.49]	−0.021	[−4.44]	6098.165	[0.22]	0.072	[0.68]	0.049	109
Thailand	0.003	[1.47]	−0.036	[−5.20]	0.851	[0.82]	0.155	[1.65]	0.132	497
Turkey	−0.008	[−1.37]	−0.177	[−6.97]	0.915	[0.45]	−0.119	[−1.83]	0.476	449
UAE	−0.000	[−3.29]	0.000	[0.33]	0.895	[5.82]	−0.045	[−4.67]	0.577	325
UK	0.000	[0.19]	−0.034	[−3.01]	7.234	[1.11]	0.399	[1.47]	0.099	595
Ukraine	0.001	[0.12]	−0.079	[−2.30]	−0.090	[−0.10]	0.023	[0.29]	0.131	313
Vietnam	0.000	[0.18]	−0.003	[−0.84]	−8.784	[−1.38]	0.040	[2.23]	0.044	453

This table reports the robustness check for the sovereign CDS on the currency forward bias. The model specification is:
$fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\,\Delta LNSCDS_{i,t+1} + \beta_{i,2}\,IRDiff_{i,t} + \beta_{i,3}\,\Delta UIRP_{i,t+1} + \varepsilon_{i,t+1}$.
Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

6.4. Controlling for uncovered interest rate parity

If the forward rate deviates markedly from the interest rate parity, then the forward bias is more likely to materialize. In addition, the literature documents that sovereign credit risk is also an important determinant of the uncovered interest rate parity. However, for the purposes of our study, in our setting, we do not include any of these UIRP drivers, but we simply rely on the whole UIRP as a control variable. Specifically, we estimate the regression model based on the interest rate parity by applying this revised specification:

$$\Delta\left(f_{i,t+1} - s_{i,t+1}\right) = \alpha_{i,0} + \alpha_{i,1}\Delta IRDiff_{i,t+1} + e_{i,t+1}. \tag{16}$$

Note that the changes in the UIRP are derived from the unexplained part of the regression, i.e. $\Delta UIRP_{i,t+1} = \alpha_{i,0} + e_{i,t+1}$. Since the variable *UIRP* is based on the residual of the interest rate parity, Equation (16), we treat the factors which drive the currency forward

Table 9
Controlling for CDS and FX Liquidity.

	Dependent Variable: Forward Bias									
	$\Delta LNSCDS$		$IRDiff$		ΔBAS^{FX}		ΔBAS^{CDS}		R-sqr	N
Austria	−0.036	[−3.16]	0.296	[1.46]	1.432	[0.14]	1.421	[0.49]	0.084	724
Belgium	−0.038	[−2.87]	0.238	[1.16]	2.061	[0.20]	0.745	[0.25]	0.072	751
Bulgaria	−0.090	[−9.17]	−0.282	[−1.61]	−15.069	[−0.75]	4.400	[2.20]	0.314	280
Bahrain	−0.001	[−1.98]	−0.589	[−6.57]	0.746	[4.29]	0.029	[0.58]	0.667	398
Brazil	−0.127	[−10.99]	−0.006	[−0.06]	36.998	[1.86]	−5.235	[−1.10]	0.528	194
Canada	−0.013	[−4.31]	0.062	[0.34]	44.035	[2.09]	−11.813	[−2.84]	0.083	454
China	−0.002	[−0.48]	−0.035	[−1.44]	−21.218	[−1.17]	0.398	[1.57]	0.020	402
H.K.	−0.001	[−1.28]	−0.099	[−6.15]	17.609	[1.75]	−0.071	[−0.85]	0.187	526
Cyprus	−0.018	[−1.78]	0.421	[2.00]	2.824	[0.25]	0.900	[0.24]	0.031	684
Czech	−0.060	[−2.11]	0.241	[1.66]	106.763	[1.17]	3.496	[0.97]	0.134	561
Germany	−0.033	[−2.76]	0.185	[0.90]	0.780	[0.08]	0.458	[0.14]	0.057	750
Denmark	0.005	[0.18]	0.772	[1.61]	106.854	[1.88]	8.738	[0.79]	0.062	243
Estonia	−0.017	[−1.00]	0.383	[1.24]	11.867	[0.87]	1.507	[0.41]	0.027	531
Finland	−0.027	[−2.15]	0.353	[1.67]	3.694	[0.35]	1.130	[0.35]	0.043	724
France	−0.039	[−3.24]	0.292	[1.42]	2.056	[0.20]	0.715	[0.23]	0.080	731
Greece	−0.035	[−3.64]	0.290	[1.40]	11.976	[0.98]	1.631	[0.47]	0.091	650
Iceland	−0.069	[−3.74]	0.187	[1.66]	1262.532	[2.64]	−7.979	[−1.35]	0.214	680
Indonesia	−0.067	[−5.39]	−0.109	[−0.85]	−25597.558	[−1.44]	0.723	[0.30]	0.383	252
Ireland	−0.030	[−1.95]	0.223	[1.01]	3.115	[0.30]	0.620	[0.20]	0.050	746
Israel	−0.038	[−1.82]	0.127	[1.30]	67.147	[4.12]	0.416	[0.16]	0.148	587
Italy	−0.031	[−2.41]	0.196	[0.94]	3.006	[0.29]	0.736	[0.29]	0.051	776
Japan	0.021	[1.47]	0.079	[0.86]	3433.288	[3.03]	0.770	[0.36]	0.048	724
Jordan	−0.000	[−0.18]	0.066	[0.55]	1.350	[2.01]	−1.016	[−0.76]	0.055	143
Kazakhstan	−0.012	[−1.79]	0.096	[1.62]	38.400	[0.20]	0.372	[0.40]	0.026	413
S. Korea	−0.097	[−4.62]	0.087	[0.80]	5179.926	[1.17]	5.173	[1.31]	0.375	582
Latvia	−0.034	[−1.62]	0.424	[1.41]	20.307	[1.55]	−1.071	[−0.27]	0.062	506
Lithuania	−0.031	[−1.69]	0.373	[1.36]	11.159	[0.89]	0.207	[0.06]	0.052	563
Malaysia	−0.036	[−5.41]	0.033	[0.59]	7.640	[0.63]	−0.898	[−0.98]	0.247	623
Netherlands	−0.033	[−2.75]	0.448	[2.27]	6.001	[0.52]	−0.174	[−0.05]	0.077	638
Nigeria	−0.158	[−4.33]	−0.080	[−1.38]	508.356	[1.11]	−1.925	[−0.45]	0.292	57
Norway	−0.064	[−3.25]	0.059	[0.40]	36.653	[2.05]	−6.827	[−2.63]	0.139	643
Poland	−0.113	[−3.81]	0.232	[1.50]	7.511	[1.09]	4.898	[1.40]	0.319	568
Portugal	−0.029	[−2.35]	0.212	[0.97]	6.963	[0.65]	0.769	[0.24]	0.047	757
Qatar	−0.001	[−3.55]	−0.064	[−4.98]	−4.067	[−5.03]	−0.070	[−0.50]	0.622	116
Russia	−0.082	[−5.59]	−0.007	[−0.16]	−244.103	[−0.89]	−0.964	[−0.35]	0.413	585
Slovakia	−0.036	[−2.05]	0.275	[1.05]	6.967	[0.55]	0.784	[0.25]	0.084	561
Slovenia	−0.037	[−2.37]	0.206	[0.81]	10.505	[0.86]	−0.413	[−0.34]	0.082	565
S. Africa	−0.164	[−6.84]	−0.244	[−1.97]	11.574	[1.81]	2.759	[1.09]	0.453	634
Spain	−0.039	[−2.44]	0.229	[1.05]	3.233	[0.33]	0.333	[0.10]	0.066	763
Sri Lanka	−0.018	[−2.03]	0.021	[0.58]	−309.566	[−1.44]	−2.680	[−3.43]	0.107	338
Sweden	−0.058	[−3.31]	0.153	[1.43]	58.866	[2.30]	1.268	[0.27]	0.152	676
Taiwan	−0.021	[−3.93]	0.081	[0.79]	68.919	[1.55]	−0.112	[−0.06]	0.064	109
Thailand	−0.040	[−6.13]	0.168	[1.97]	114.563	[4.66]	−0.330	[−0.31]	0.225	497
Turkey	−0.178	[−6.95]	−0.119	[−1.77]	−1.430	[−0.49]	0.487	[0.09]	0.475	449
UAE	0.000	[0.80]	−0.050	[−4.06]	−0.009	[−0.01]	0.058	[0.70]	0.297	325
UK	−0.033	[−2.92]	0.405	[1.54]	2.313	[0.27]	−5.354	[−1.80]	0.103	595
Ukraine	−0.074	[−2.09]	0.024	[0.29]	−6.619	[−1.24]	−2.375	[−1.03]	0.151	313
Vietnam	−0.003	[−0.78]	0.039	[2.19]	−4849.965	[−1.32]	−0.447	[−1.62]	0.060	453

This table displays the robustness check for the sovereign CDS on the currency forward bias. The model specification is:
$fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\Delta LNSCDS_{i,t+1} + \beta_{i,2}IRDiff_{i,t} + \beta_{i,3}\Delta BAS^{FX}_{i,t+1} + \beta_{i,4}\Delta BAS^{CDS}_{i,t+1} + \varepsilon_{i,t+1}$.
The intercept coefficients are not reported to save space. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

rate to deviate from the theory of interest rate parity as latent or unobservable.[9] We include this variable in our baseline regression; therefore, the revised regression specification for controlling for uncovered interest parity is as follows:

$$fwdbias_{i,t+1} = \beta_{i,0} + \beta_{i,1}\Delta LNSCDS_{i,t+1} + \beta_{i,2}IRDiff_{i,t} + \beta_{i,3}\Delta UIRP_{i,t+1} + \varepsilon_{i,t+1}. \quad (17)$$

The purpose of this regression is twofold. First, we expect $\beta_{i,1}$ coefficients still to be negative when we add the additional control variable *UIRP*. Hence, this indicates that our central prediction is economically significant and robust. Interestingly, it also provides a benchmark to evaluate the relative importance of sovereign credit risk and other factors across currency markets.

[9] For example, Breedon et al. (2016) argue that currency order flow is one of the drivers to the forward bias, due to limit to arbitrage.

The regression results controlling for UIRP are reported in Table 8. Overall, we find that sovereign CDS spreads are an important determinant of the forward bias, because the sovereign CDS coefficient is still significant and consistent with our previous results. In addition, we find that the UIRP results are mainly insignificant.[10] As discussed, this means that when comparing sovereign CDS and the latent variable for UIRP, the impact from sovereign CDS is stronger than that from the *UIRP* variable. Therefore, sovereign credit risk appears as the primary driver of the currency forward bias. Interestingly, we observe that in five countries (i.e. Bahrain, Canada, Jordan, South Korea, and UAE), the impact of the UIRP variable on the forward bias is of positive sign, and for two countries (i.e. Bahrain and UAE) the coefficient on sovereign CDS is insignificant. Our evidence suggests that the forward bias in Jordan and the UAE are more likely to be driven by other factors than sovereign credit risk. However, we also find that both sovereign CDS and UIRP statistically explain the forward bias in Bahrain, Canada and South Korea, indicating that they are both important factors.

While Equation (16) may contain a measurement error on the uncovered interest rate parity, we further examine whether our results are still robust by regressing the UIRP (defined as $fwdbias_{t+1} - IRDiff_t$) on the change of the sovereign CDS spread. The results, reported in Table A2 in the Appendix, confirm that the coefficient for $\Delta LNSCDS$ is significantly negative.

6.5. Controlling for CDS liquidity and FX liquidity

There is a well-established empirical evidence showing that currency prices are affected by market liquidity (see, e.g., Glassman, 1987; Ding and Hiltrop, 2010; Karnaukh et al., 2015). In addition, CDS are traded over the counter; hence, the CDS spread is also highly impacted by its own liquidity (Badaoui et al, 2013). To test if our findings on the association between sovereign CDS and the currency forward bias are also affected by liquidity risk in both CDS and currency markets, we repeat our main regression analyses with additional control variables for the CDS liquidity and foreign exchange liquidity.

To proxy for liquidity risk, we adopt the commonly used CDS spread and FX bid-ask spread, respectively. Note that CDS spreads obtained from Markit are composite prices; therefore, there is no bid-ask spread information available (see, Lin et al., 2021; Markit, 2012 for a discussion). Therefore, the CDS bid-ask spreads are sourced from Reuters Eikon. However, since the scope of CDS in Eikon is comparatively smaller than in Markit, the CDS bid-ask spreads are not available for all the countries included in the complete sample: hence, for these countries we use the cross-sectional average of the bid-ask spreads.[11]

Table 9 shows that our core finding on the association between the sovereign CDS spread and the currency forward bias is not altered after controlling for the changes in CDS liquidity and FX market liquidity. Hence, this suggests that the effect of sovereign credit risk on the currency forward bias is not fully captured by liquidity risk in both markets.

With respect to liquidity risk, we find a positive association between the currency forward bias and the FX bid-ask spread for a few currencies, e.g., Bahrain, Canada, Demark, etc. Notably, our results support prior studies, e.g., McGroarty et al. (2009), that FX market illiquidity causes uncertainty in spot prices, thereby the forward bias is enlarged.[12] On the other hand, we do not find unambiguous evidence on the impact of the sovereign CDS bid-ask spread on the forward bias as the loadings on ΔBAS^{CDS} are mostly statistically insignificant.

6.6. Panel regression results

In this section, we conduct all our previous regression specifications in the form of panel regressions. In this analysis, we further control for time and country fixed effects. The panel regression results are presented in Table 10.

Models 1, 4, 6, and 8 provide the panel regression results for local sovereign credit risk, controlling for uncovered interest rate parity, and the alternative forward bias measure. The results show that sovereign CDS premia have a negative impact on the forward bias. Nevertheless, we find no statistical significance for the interest rate differential in all of the panel regression models. Recall that while we observe some significance for the country-specific results, overall the impact of the interest rate differential is negligible. As illustrated in the previous sections, our proposed model focuses on the effect of sovereign credit risk on the forward bias. Therefore, it is not surprising that the interest rate differential does not seem to contribute to the forward bias.

We also conduct additional robustness tests, including controlling for the AR(1) effect (as reported in Models 2, 5, 7, and 9). All in all, the results are similar. Moreover, for the lagged variables (reported in Model 3), the effect is only marginally significant at the 10%

[10] Also note that some countries have rather large coefficient of ΔUIRP. Recall that ΔUIRP is the regression residuals of Eq. (16). If the residual on average is small, it means that the interest rate parity explains well the currency changes while ΔUIRP resembles a random error. Therefore, those large coefficients are less likely to exhibit statistical significance.

[11] The sovereign CDS bid-ask spreads are only available for the following 30 countries: Belgium, Bulgaria, Bahrain, Brazil, China, Czech, Germany, France, Indonesia, Ireland, Israel, Italy, Kazakhstan, S. Korea, Latvia, Lithuania, Malaysia, Poland, Portugal, Qatar, Russia, Slovakia, Slovenia, S. Africa, Spain, Thailand, Turkey, UK, Ukraine, and Vietnam.

[12] Note that some currencies have pretty small changes in the FX bid-ask spread in our sample and hence we observe some large regression coefficients. The small change in the bid-ask spread implies that the FX liquidity does not dramatically change over time. However, given the statistical significance, the effect should not be ignored. The reported regression coefficient provides the magnitude of the impact given 1 unit movement of the independent variable, which may not happen to all countries. One could use the economic impact (i.e., the impact given 1 STD movement in the independent variable) for further interpretation. Taking Icelandic Krona as an example, the forward bias is expected to increase by 0.0074 in terms of the economic impact given one STD movement of $\Delta BAS^{FX}_{Iceland}$ ($0.0074 = 5.83E{-}6 \times 1262.532$, where 5.83E-6 is the STD of $\Delta BAS^{FX}_{Iceland}$ and 1262.532 is the corresponding coefficient).

Table 10
Panel Regression Results.

	Original FwdBias							Alternative FwdBias	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9
Δ*LNSCDS*	−0.033	−0.032				−0.033	−0.032	−0.033	−0.032
	[−6.10]	[−5.70]				[−6.10]	[−5.70]	[−6.09]	[−5.70]
$\Delta LNSCDS^{Local}$				−0.034	−0.033				
				[−6.59]	[−6.17]				
Δ*URIP*						−0.060	−0.105		
						[−0.22]	[−0.39]		
IRDiff	0.004	0.007		0.005	0.008	0.004	0.007	0.029	0.032
	[0.13]	[0.26]		[0.16]	[0.28]	[0.13]	[0.26]	[1.00]	[1.17]
fwdbias (lag1)		0.050			0.052		0.050		0.048
		[1.28]			[1.30]		[1.28]		[1.22]
Δ*LNSCDS* (lag1)			−0.008						
			[−1.72]						
IRDiff (lag1)			0.003						
			[0.10]						
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R-sqr	0.048	0.049	0.003	0.047	0.048	0.048	0.049	0.049	0.050
N	24,840	23,939	23,939	24,840	23,939	24,840	23,939	24,840	23,939

This table reports the panel regression results for the previous analysis. Heteroscedasticity and autoregressive robust standard errors are applied to test statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

level, indicating the weak predictive power of sovereign CDS on the forward bias.

We further explore whether the effect of sovereign default risk varies in terms of economic development. Prior studies document (e.g. Borri and Shakhnov, 2021) that the UIRP holds more strongly for emerging market currencies even though they are more exposed to sovereign risk. Here, we hypothesize that since emerging market currencies are more sensitive to sovereign default risk, sovereign default risk is a more pronounced driver of the forward bias in emerging market countries than in developed countries.

To validate our hypothesis, we repeat our main panel regression analysis for emerging and developed countries, separately. Therefore, we test whether the magnitude of the effect is related to the divergent paths of economic development.

We use the Human Development Index (HDI) to proxy for the level of economic development of a country. At the time of data collection, the latest HDI values are more complete as at the end of 2019. Therefore, we use the 2019 values and assume the HDI values to be constant throughout the whole sample period. Such assumption is reasonable as the HDI values are stable for individual economies over time. Moreover, since there are not conventional criteria to classify a country as a developed or emerging market economy, we use 0.86, the sample average of the HDI values in our sample, and we can split our sample into developed economies (HDI higher than 0.86) and emerging market economies (HDI lower than 0.86).[13]

As shown in Table 11, we find that the coefficients of the Δ*LNSCDS* are both significantly negative at the 1% level for both developed and emerging market countries, but the magnitude of the coefficient is larger for the emerging market countries. Therefore, this evidence lends support to our conjecture that emerging market currencies are more sensitive to sovereign default risk and as such sovereign default risk is a stronger driving factor of the forward bias in emerging market countries. In the last column, we report the results for the whole sample and add a dummy variable (*D_Emer*) for the group of emerging market economies (i.e. taking the value of 1 if the country is an emerging country and zero otherwise). We find that the coefficient of the Δ*LNSCDS* is still negatively significant and the interactive term $\Delta LNSCDS \times D\_Emer$ is also negatively significant. The result again confirms our hypothesis.

6.7. Sovereign CDS price discovery

In the previous sections, we have demonstrated that sovereign CDS spreads are negatively associated with the forward bias, both from a theoretical and an empirical perspective. The results suggest that the information embedded in the sovereign CDS market translates into declines in the forward bias. Hence, this result underscores the prominent role of the sovereign CDS market in improving price discovery between the currency spot and forward markets.

Ever since the identification of a persistent and predictable forward bias (Bansal and Dahlquist, 2000; Fama, 1984; Villanueva, 2007), researchers have long established the existence of market inefficiency in FX markets, e.g., due to limit to arbitrage on asset pricing (Breedon et al, 2016; Kumar, 2020). However, we hypothesize that markets for sovereign CDS provide additional information that improves market efficiency, beyond the existing factors documented in the literature. As a consequence, the magnitude of the forward bias should be smaller in presence of sovereign CDS markets. Therefore, our empirical research strategy is to compare the forward bias before and after the inception of the sovereign CDS market. If we observe a reduced value of the forward bias in absolute

[13] There are 18 emerging countries, including Bulgaria, Bahrain, Brazil, China, Indonesia, Jordan, Kazakhstan, Malaysia, Nigeria, Qatar, Russia, Slovakia, S. Africa, Sri Lanka, Thailand, Turkey, Ukraine, and Vietnam, while the remaining 30 countries are developed economies.

Table 11
Sovereign Default Risk and Economic Development.

Dependent Variable: Forward Bias			
	Emerging Ct	Developed Ct	All Ct
Δ*LNSCDS*	−0.053	−0.026	−0.030
	[−8.97]	[−4.55]	[−4.55]
Δ*LNSCDS* × *D_Emer*		−0.012	
			[−1.79]
IRDiff	−0.022	0.113	0.004
	[−0.79]	[1.54]	[0.15]
Country FE	Yes	Yes	Yes
Time FE	Yes	Yes	Yes
R-sqr	0.088	0.042	0.050
N	6708	18,132	24,840

This table reports the panel regression results for different levels of economic developments. Heteroscedasticity and autoregressive robust standard errors are applied to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

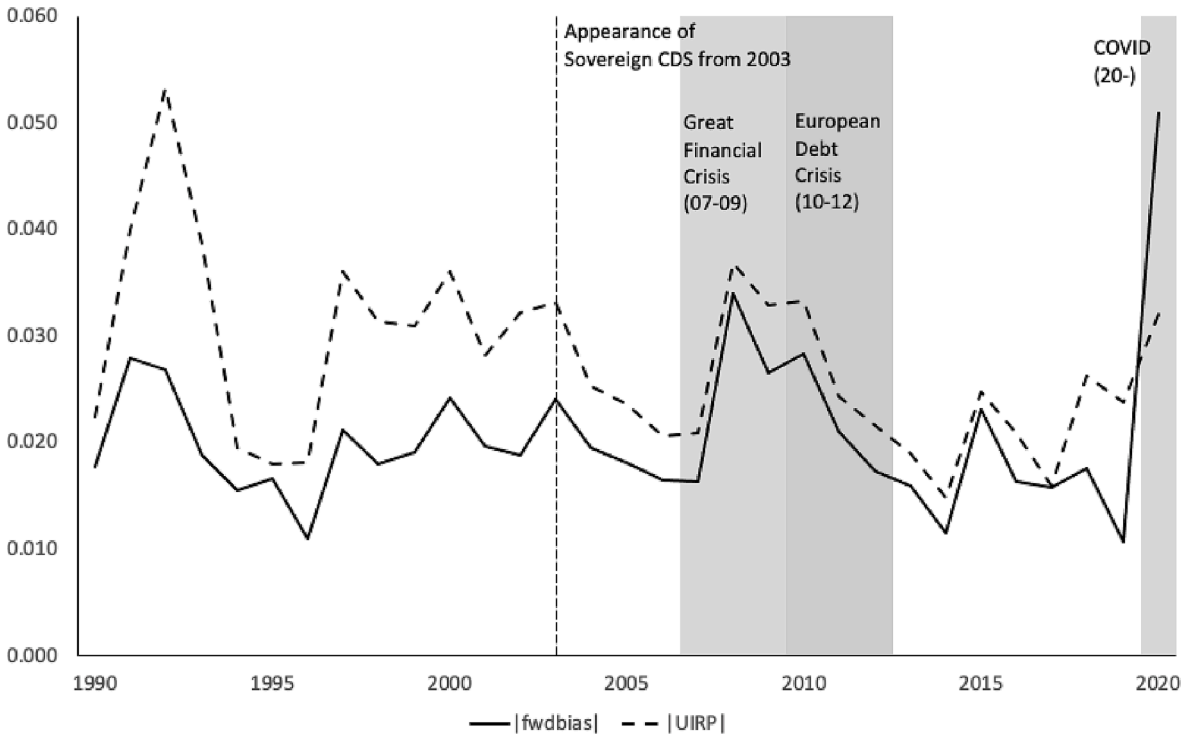

Fig. 1. Yearly Pooled Average Forward Bias and Uncovered Interest Rate Parity.

term, we can conclude that the sovereign CDS market also affects the price discovery process.

Because our original sample covers only the 2003–2020 sample period, we extend our sample backward to 1990. Hence, we can capture the individual country's forward bias over the periods before-and after-the inception of the sovereign CDS market. Most of the currencies data is available since 1990. However, for some of the countries, data availability is insufficient.[14] As such, we are left with 26 countries. The Appendix provides the information about the data period availability for each country.

As a preliminary analysis, we simply pool the average absolute values of the forward bias for each year to shed light on the yearly trend of the forward bias. As shown in Fig. 1, interestingly, we find that the magnitude of the forward bias exhibits a declining trend over time after the appearance of the sovereign CDS market. However, we also observe that the magnitude of the forward bias widens multiple times during the crisis periods (illustrated by the shaded areas).

[14] For this extended sample, we only analyze 26 countries; in particular, they are: Austria, Belgium, Canada, H.K., Cyprus, Czech, Germany, Estonia, Finland, France, Greece, Iceland, Ireland, Italy, Japan, Latvia, Lithuania, Malaysia, Netherlands, Poland, Portugal, Slovakia, Slovenia, S. Africa, Spain, and UK.

Table 12
Sovereign CDS Inception.

	$\|fwdbias\|$			$\|UIRP\|$		
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
SCDS Inception	−0.003	−0.004	−0.006	−0.011	−0.014	−0.017
	[−2.84]	[−3.79]	[−4.64]	[−6.81]	[−8.69]	[−10.38]
GFC	0.009	0.009	0.009	0.010	0.009	0.009
	[2.64]	[2.70]	[2.75]	[3.24]	[3.17]	[3.08]
EDC	0.005	0.006	0.006	0.006	0.005	0.005
	[2.68]	[2.84]	[2.98]	[2.76]	[2.69]	[2.56]
COVID	0.034	0.033	0.033	0.011	0.001	−0.009
	[15.40]	[14.72]	[11.23]	[5.78]	[0.56]	[−0.79]
Intercept	0.020			0.032		
	[23.14]			[22.49]		
Country FE	No	Yes	Yes	No	Yes	Yes
Time FE	No	No	Yes	No	No	Yes
rsqr	0.022	0.026	0.013	0.028	0.057	0.030
N	32,997	32,997	32,997	32,997	32,997	32,997

This table reports the panel regression results for the inception of the sovereign CDS market. Heteroscedasticity and autoregressive robust standard errors are applied to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

In addition, since there has been a secular downward trend in the path of interest rates globally after 2000, with narrower interest rate differential, the forward bias is weaker. To test this conjecture, we plot the value based on the uncovered interest rate parity by $|UIRP_t| = |fwdbias_t - IRDiff_{t-1}|$. We find a similar trend in $|UIRP|$, indicating that the reduced magnitude of the forward bias is indeed related to the emergence of the sovereign CDS market.

Based on our discussion, we estimate the following baseline regression:

$$|fwdbias_{i,t}| = \beta_0 + \beta_1 SCDSInception_{i,t} + \beta_2 GFC_t + \beta_3 EDC_t + \beta_4 COVID_t + \varepsilon_{i,t}, \quad (18)$$

where *SCDSInception* is a binary variable, taking the value of 1 at the start date of the sovereign CDS quotes of the country in the Markit database and 0 otherwise.[15] A negative coefficient for *SCDSInception* indicates a reduced forward bias after inception, and hence supports our hypothesis that sovereign CDS contributes to price discovery. We also construct three dummy variables—*EDC* (European Sovereign Debt Crisis), *GFC* (Global Financial Crisis), and *COVID*—and include these dummy variables as control variables. The *EDC* dummy variable takes the value of 1 during the years from 2010 and 2012 and 0 otherwise; the *GFC* dummy variable takes the value of 1 during the years from 2007 to 2009 and 0 otherwise; the *COVID* dummy variable takes the value of 1 in the year 2020 and 0 otherwise.

Table 12 presents the results. Model 1 shows the baseline regression results. We observe that the coefficient is of negative sign and statistically significant at the 1% level. This result suggests that the magnitude of the forward bias has fallen after the inception of the sovereign CDS market. Notice that since we carry out our baseline regression using panel data, the unobserved time and country effect can affect the statistical significance. Thus, to control for the unobserved effects, we further repeat our baseline regression with fixed effects. Importantly, with this added control, the results from Models 2 (Country FE) and 3 (Country and Time FE) remain unchanged. Overall, the results in Table 10 confirm our main hypothesis that sovereign CDS markets do improve price discovery in spot and forward currency markets.

There are some noteworthy findings: although the magnitude of the forward bias decreases after the inception of the sovereign CDS market, it becomes larger during the 2008–2009 global financial crisis, because we observe significantly positive coefficients for the crisis dummies, and the coefficients for these dummies are larger than that of *SCDSInception*. This suggests that during the period of the global financial crisis, the magnitude of the forward bias is mostly attributable to macroeconomic uncertainty. Nevertheless, our empirical evidence strongly validates our conjecture that the sovereign CDS market improves price discovery.

We assess the robustness of our findings in several ways. First, we re-estimate Equation (18) with $|UIRP|$ as dependent variable, to address our concern that our baseline result is not driven by the narrowed interest rate differential. In Table 12 (Models 4 to 6), we confirm that our results are still robust.

One may cast doubt on our sample representativeness as it only includes countries with traded sovereign CDS and that may give rise to potential sample bias concerns. However, our main argument about the sovereign CDS market inception concerns the time-series impact, i.e. comparing the forward bias values (before- and after-inception) of each individual country. Hence, it would not reinforce our approach to include also countries that never have traded sovereign CDS. Furthermore, to corroborate our evidence, we further run time-series regressions for the individual countries. We report the main results in Table A3, in the Appendix. Overall, the findings are consistent with the panel regression results, indicating that indeed the currency forward bias is lower in the sub-sample period subsequent to the inception of the sovereign CDS market. Specifically, most of the *SCDSInception* variables are significantly negative at the

[15] The first sovereign CDS contracts begun trading at around 1997 (Packer and Suthiphongchai, 2003), but the trading volume and liquidity of the market was relatively thin. Hence, we rely only on the Markit quote data as an indicator of the trading initiation period of the sovereign CDS.

Table 13
ADF and MDH Tests for Sovereign Default Risk Process.

Country	MDH Test *t*-statistic	ADF Test *t*-statistic	N	Country	MDH Test *t*-statistic	ADF Test *t*-statistic	N
Austria	0.772	−7.276	724	S. Korea	−0.026	−8.370	582
Belgium	0.620	−7.571	751	Latvia	2.602	−5.809	506
Bulgaria	0.472	−5.239	280	Lithuania	1.136	−6.580	563
Bahrain	2.430	−5.460	398	Malaysia	0.854	−9.579	623
Brazil	0.486	−6.190	194	Netherlands	0.898	−7.113	638
Canada	−1.549	−12.593	454	Nigeria	−1.090	−3.387	57
China	0.879	−8.623	402	Norway	1.070	−7.522	643
H.K.	0.579	−8.087	526	Poland	0.337	−7.250	568
Cyprus	1.335	−8.163	684	Portugal	1.138	−8.801	757
Czech	0.749	−7.153	561	Qatar	−0.492	−5.007	116
Germany	0.014	−8.096	750	Russia	0.420	−8.878	585
Denmark	−4.419	−5.127	243	Slovakia	0.597	−7.073	561
Estonia	1.457	−7.783	531	Slovenia	1.382	−7.355	565
Finland	0.681	−8.210	724	S. Africa	1.026	−8.007	634
France	0.687	−7.865	731	Spain	1.539	−8.953	763
Greece	4.230	−7.746	650	Sri Lanka	−1.398	−8.387	338
Iceland	1.625	−6.924	680	Sweden	0.734	−6.799	676
Indonesia	0.352	−6.997	252	Taiwan	1.480	−4.547	109
Ireland	0.997	−8.015	746	Thailand	0.657	−8.921	497
Israel	0.277	−7.446	587	Turkey	0.400	−8.252	449
Italy	1.698	−9.624	776	UAE	2.809	−7.742	325
Japan	0.790	−7.996	724	UK	1.869	−6.965	595
Jordan	0.948	−5.347	143	Ukraine	1.480	−5.267	313
Kazakhstan	−0.366	−8.037	413	Vietnam	0.297	−6.762	453

This table reports the results of the ADF and MDH tests for the sovereign default risk process. We follow Greene (2012) to test whether $\Delta LNSCDS$ is a martingale difference process (with null hypothesis that the sovereign default risk is a martingale process) as well as the Augmented Dickey–Fuller test (with the null hypothesis that sovereign default risk is a unit root process) for $\Delta LNSCDS$. The tests are conducted at the country level, and the corresponding *t*-statistics are reported in the table.

1% level. Taken together, the results for individual countries provide unambiguous and strong confirmation of our central argument that the sovereign CDS enhances price discovery between currency spot and forward markets as these empirical patterns can be directly connected to developments in the sovereign CDS market.

6.8. *Model verification and non-linearity relationship*

In this section, we further explore whether our assumptions discussed in Section 2 are reasonable or not. Recall the first assumption is that sovereign default risk is a martingale process. Here we provide further empirical evidence. A martingale process is a sequence of a random variable for which, at a particular time, the conditional expectation of the next value in the sequence is equal to the present value, regardless of all prior values. One important characteristic of a martingale process is the random walk. We use Greene (2012) to test if sovereign default risk is a martingale process. Specifically, we test if the change of the logarithm of sovereign CDS spreads, which is the main variable of interest, follows a martingale difference sequence, i.e. $E(\Delta LNSCDS_t|\Delta LNSCDS_{t-1},\cdots) = 0$.

Empirically, to test the martingale difference hypothesis for a random variable, we examine the zero average difference. If the historically averaged difference is statistically significantly different from zero, the process is not a martingale process; otherwise, it is a martingale process. We follow Greene (2012) to construct the test statistic, $t = \frac{\sqrt{T}\mu_{\Delta LNSCDS,i}}{\sigma_{\Delta LNSCDS,i}}$, where $\mu_{\Delta LNSCDS,i}$ (or $\sigma_{\Delta LNSCDS,i}$) is the sample mean (or sample standard deviation) of the $\Delta LNSCDS$ for country *i*. Under the null hypothesis, $\Delta LNSCDS$ is a martingale difference process.

We test the martingale difference hypothesis for each country and the results are reported in Column MDH Test in Table 13. We can see that for the majority of countries sovereign default risk exhibits a martingale process. Specifically, only 7 (or 5) countries reject the null hypothesis that sovereign default risk is a martingale process at the 10% (or 5 %) level.[16] Thus, the evidence confirms our prediction.

Furthermore, since our main regression is based on time-series analysis, we also test whether $\Delta LNSCDS$ is stationary. We conduct the Augmented Dickey–Fuller test under which the null hypothesis is that the time-series variable is non-stationary. Column ADF Test in Table 13 shows that $\Delta LNSCDS$ is stationary for all countries at the 1 % level. It indicates that the results are not driven by autoregressive effects.

Next, we discuss our second hypothesis that sovereign default risk and the interest rate differentials are independent. If this assumption does not hold, then there will be additional terms on Equation (9), that is,

[16] Specifically, Bahrain, Denmark, Greece, Latvia, and UAE reject the null hypothesis that the default risk process is a martingale process at 5% level; Italy and UK further reject the null hypothesis at 10% level.

Table 14
Linear Dependence between Sovereign Default Risk and IR differential.

	Dependent Variable: Δ*LNSCDS*			
	Model 1	Model 2	Model 3	Model 4
Intercept	0.007			
	[0.61]			
IRDiff	−0.026	0.005		−0.512
	[−0.24]	[0.05]		[−1.31]
$IRDiff(t-1)$			0.032	0.523
			[0.31]	[1.27]
$\Delta LNSCDS(t-1)$				0.044
				[1.56]
Country FE	No	Yes	Yes	Yes
Time FE	No	Yes	Yes	Yes
R-sqr	0.000	0.000	0.000	0.002
N	24,840	24,840	23,939	23,939

This table reports the results for the linear dependence between sovereign default risk and the interest rate differential. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

$E(s_{t+1}-f_t) = -\left(k_{t+1}^{FX}-k_t^{FX}\right)\left(i_t^{US}-i_t^{FX}\right)+COV\left(-\left(k_{t+1}^{FX}-k_t^{FX}\right),\left(i_t^{US}-i_t^{FX}\right)\right)$, where the second term is the covariance term between sovereign default risk and the interest rate differential.[17] To test whether our second assumption holds, we regress Δ*LNSCDS* on *IRDiff*. Recall that the interest rates we use in our analysis is the interbank interest rate, and these interest rates are often used as the risk-free rate. Although this proxy for the risk-free rate may not be optimal because it may still contain counterparty risk, it is not supposed to reflect sovereign credit risk. Unsurprisingly, as shown in Table 14, we do not find evidence of statistical significance on the interest rate differential.

Given the previous result of zero correlation between sovereign default risk and the interest rate differential, we also further test if our results are affected by possible covariance terms. We use, as a proxy of the covariance, the interactive term of changes in sovereign CDS and the interest rate differential. The result is reported in Column 1 of Table 15, and as expected our main results are not altered when the interactive term is controlled for. In fact, we find that the coefficient on the interactive term is insignificant.

Finally, we test if our results are affected by possible non-linearities of the regressors, as in practice a linear association is less likely to hold. In the main regression setting, we examine the linear association between sovereign risk and the currency forward bias; here we control for non-linearity of sovereign risk by adding the quadratic term of Δ*LNSCDS*.

As we can see from Column 2 of Table 15, the coefficient of Δ*LNSCDS* is still negatively significant at the 1% level. The quadratic term is also negatively significant, indicating a non-linear and concave association between sovereign risk and the forward bias. The forward bias decreases when sovereign default risk increases, and the effect becomes stronger with higher sovereign credit risk. The last column reports the regression results when both non-linearity and the covariance terms are controlled for. Again, we find the results are qualitatively the same as the main results.

7. Conclusion

Interest on sovereign credit risk has been growing rapidly in recent years since the European sovereign debt crisis. Recent research on sovereign credit risk mostly focuses on its connection with other financial markets. We build on the insights of previous studies (e.g. Foroni et al., 2018; Calice and Zeng, 2019; Della Corte et al., 2022) to explore the relationship between sovereign credit risk and the currency forward bias. In a setting of defaultable sovereign bonds, we model a negative relation between country's sovereign credit risk and the forward bias, suggesting that the information from the sovereign CDS market is transmitted to the forward bias.

Empirically, using sovereign CDS spreads for 48 countries over the sample period 2003 to 2020, we provide evidence that sovereign CDS spreads negatively explains the currency forward bias. Specifically, we find significant negative coefficients of the weekly changes in sovereign CDS spreads for the weekly changes in the currency forward bias, defined as the difference between the 1-month forward price and the spot rate. Hence, the empirical results support our main hypotheses. Furthermore, we show that the impact of sovereign credit risk on the forward bias is uniform across both developed and emerging countries, with the effect larger for emerging countries. Notably, the results are also robust to a different measure of the forward bias as well as after controlling for uncovered interest rate parity and market liquidity.

Our study provides insights on how sovereign credit risk affects the currency forward bias, and contributes to a better understanding of the linkages between sovereign debt and foreign exchange markets. We also contribute to prior studies identifying several driving factors of the currency forward bias by providing novel evidence on the pricing discovery process of the currency forward bias. The implications of our results can be beneficial to both market participants and regulators. Market participants, e.g. institutional investors, mutual funds, hedge funds, and banks, are the leading currency market players and they enter both forward and spot markets regularly for hedging FX risk or speculation (e.g. carry trade) purposes. Therefore, the magnitude of the forward bias can guide the

[17] If X and Y are random variables, then $E(XY) = E(X)\,E(Y) + COV(X,Y)$.

Table 15
Control for Covariance Term and Non-linearity.

	Dependent Variable: Forward Bias		
	Model 1	Model 2	Model 3
$\Delta LNSCDS$	−0.031	−0.029	−0.027
	[−5.63]	[−6.91]	[−5.80]
$\Delta LNSCDS^2$		−0.016	−0.016
		[−2.00]	[−2.18]
IRDiff	0.002	0.003	0.001
	[0.07]	[0.11]	[0.05]
$\Delta LNSCDS \times IRDiff$	1.992		2.054
	[1.28]		[1.34]
Country FE	Yes	Yes	Yes
Time FE	Yes	Yes	Yes
R-sqr	0.053	0.052	0.057
N	24,840	24,840	24,840

This table reports the results of controlling for the covariance term between sovereign default risk and the interest rate differential and the non-linearity of sovereign default risk. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses.

investment decision process of professional investors. For example, if the forward bias is small, then the currency hedge is expected to be effective; on the other hand, carry trade return strategies likely underperform in the same context. Hence, our empirical evidence clearly shows that sovereign credit risk as measured by sovereign CDS is key to gauge the forward bias. Similarly, for regulators, market efficiency is an important goal. Our results suggest that currency market efficiency also depends on the information about sovereign risk. Therefore, sovereign credit risk induces greater immediacy and transparency of currency market prices for internationally active investors.

In addition, shocks in sovereign CDS markets are informative to investors who need to hedge the risk of holding foreign currencies, e.g., multinational firms. If the interest rate parity tends to hold (i.e. smaller currency forward bias for countries under severe default risk), then firms should hedge currency risk for currencies heavily exposed to sovereign credit markets. Hence, our findings highlight a quantitatively important and previously neglected price discovery channel of the forward bias. Furthermore, market efficiency is a primary objective for regulators to maintain transparency and protect investors. For example, the stated mission of the US CFTC (Commodity Futures Trading Commission) is to "[…] foster open, transparent, fair, competitive, and secure markets through oversight of derivatives platforms […]". Our results suggest that the efficiency of the currency market also depends on the economic information aggregated in the markets for sovereign CDS. Hence, from both a regulatory and a market design perspective, regulators might consider to use sovereign CDS (e.g. the cross-sectional average of sovereign CDS spreads) as a useful structural indicator to gauge the overall efficiency of currency markets. Taken together, our evidence illustrates that sovereign CDS benefits market participants as they enhance the overall transparency of currency market prices.

Although we present new evidence, our analysis leaves a number of questions open. Future research should further explore the linkages between sovereign CDS and currency markets. Since our results indicate that sovereign credit risk affects both currency forward and spot markets, the implementation of currency trading strategies, many of which, e.g. currency carry trade, are built upon the arbitrage of spot and forward prices, might be affected by the information embedded in the sovereign CDS market. A fruitful avenue for future research would be extending our work to alternative econometric specifications on other CDS maturities (e.g. 10 year) and on the CDS term premium and applying them to shed light on other cross-sectional and possibly time series patterns in currency markets prices. In addition, recent academic studies (Longstaff et al., 2011; Augustin, 2018; Calice and Zeng, 2019) provide evidence on how the sovereign CDS markets reflect information on two distinctive components, a global and a local risk component, and on their predictive power on foreign exchange markets. Hence, it would be desirable for future research to investigate whether the global and local components of sovereign credit risk have different effects on currency markets. It is also an interesting subject for future research to explore the market informativeness of sovereign CDS for other financial markets. In this paper, we limit our research to currency markets. Recently, Lettau et al. (2014) show that downside risk premia exist in many asset classes. Thus, investigating the pricing of sovereign CDS markets and their effects on the price efficiency of other financial markets is an important area of future work.

8. Authorship statement

All the persons who meet the authorship criteria are listed as authors, and all the authors certify that they have participated sufficiently in the work to take public responsibility for the content, including participation in the concept, design, analysis, writing, or revision of the manuscript. All the work has been equally split between the authors.

Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Table A1
Sovereign CDS Data Coverage by Countries.

Sovereign CDS Data Range			
Austria	2003-Dec to 2017-Dec	S. Korea	2004-Aug to 2020-Apr
Belgium	2003-Dec to 2019-Oct	Latvia	2004-Sep to 2014-Sep
Bulgaria	2008-Apr to 2014-Sep	Lithuania	2003-Dec to 2018-Jun
Bahrain	2006-Dec to 2018-Sep	Malaysia	2003-Dec to 2020-Apr
Brazil	2012-Jan to 2020-Apr	Netherlands	2003-Dec to 2017-Dec
Canada	2003-Dec to 2014-Sep	Nigeria	2012-Aug to 2014-Aug
China	2006-Oct to 2020-Apr	Norway	2003-Dec to 2017-Dec
H.K.	2004-Aug to 2014-Sep	Poland	2003-Dec to 2019-May
Cyprus	2003-Dec to 2017-Dec	Portugal	2003-Dec to 2019-Dec
Czech	2003-Dec to 2014-Sep	Qatar	2013-Jun to 2020-Apr
Germany	2003-Dec to 2019-Oct	Russia	2003-Dec to 2020-Apr
Denmark	2013-Feb to 2017-Dec	Slovakia	2003-Dec to 2014-Sep
Estonia	2004-Jun to 2014-Sep	Slovenia	2003-Dec to 2019-May
Finland	2003-Dec to 2017-Dec	S. Africa	2003-Dec to 2020-Apr
France	2003-Dec to 2018-Dec	Spain	2003-Dec to 2019-Dec
Greece	2003-Dec to 2017-Dec	Sri Lanka	2007-Dec to 2014-Sep
Iceland	2004-Mar to 2017-Dec	Sweden	2003-Dec to 2017-Dec
Indonesia	2010-Nov to 2020-Apr	Taiwan	2006-Aug to 2014-Apr
Ireland	2003-Dec to 2019-Oct	Thailand	2005-Jan to 2020-Apr
Israel	2003-Dec to 2019-Dec	Turkey	2006-Jul to 2020-Apr
Italy	2003-Dec to 2020-Apr	UAE	2007-Feb to 2014-Sep
Japan	2003-Dec to 2017-Dec	UK	2006-Jun to 2018-Nov
Jordan	2011-Mar to 2014-Sep	Ukraine	2008-Jun to 2019-Oct
Kazakhstan	2004-Jun to 2014-Sep	Vietnam	2005-Nov to 2019-May

This table shows the sovereign CDS data coverage by countries.

Table A2
Uncovered Interest Rate Parity and Sovereign CDS.

Country	$\beta_{i,1}$		Country	$\beta_{i,1}$	
Austria	−0.032	[−3.12]	S. Korea	−0.094	[−5.11]
Belgium	−0.035	[−3.02]	Latvia	−0.034	[−1.78]
Bulgaria	−0.079	[−6.54]	Lithuania	−0.031	[−1.88]
Bahrain	−0.007	[−2.60]	Malaysia	−0.041	[−5.80]
Brazil	−0.152	[−11.19]	Netherlands	−0.030	[−2.69]
Canada	−0.016	[−4.71]	Nigeria	−0.307	[−5.46]
China	−0.003	[−0.30]	Norway	−0.050	[−2.95]
H.K.	−0.002	[−1.23]	Poland	−0.121	[−4.24]
Cyprus	−0.015	[−1.53]	Portugal	−0.024	[−2.10]
Czech	−0.064	[−2.33]	Qatar	−0.002	[−3.39]
Germany	−0.028	[−2.72]	Russia	−0.098	[−4.31]
Denmark	0.010	[0.39]	Slovakia	−0.034	[−2.11]
Estonia	−0.016	[−0.99]	Slovenia	−0.034	[−2.35]
Finland	−0.023	[−2.02]	S. Africa	−0.162	[−6.31]
France	−0.036	[−3.32]	Spain	−0.037	[−2.49]
Greece	−0.031	[−3.66]	Sri Lanka	0.009	[0.38]
Iceland	−0.058	[−3.41]	Sweden	−0.053	[−3.35]
Indonesia	−0.080	[−4.97]	Taiwan	−0.010	[−1.15]
Ireland	−0.026	[−1.95]	Thailand	−0.040	[−6.05]
Israel	−0.041	[−1.93]	Turkey	−0.175	[−6.12]
Italy	−0.027	[−2.26]	UAE	−0.004	[−1.60]
Japan	0.015	[1.13]	UK	−0.025	[−2.15]
Jordan	0.004	[0.48]	Ukraine	−0.107	[−3.56]
Kazakhstan	−0.022	[−1.08]	Vietnam	0.022	[1.53]

This table presents the results for sovereign CDS on the uncovered interest rate parity. The regression specification is:

$UIRP_{t+1} = \left(fwdbias_{i,t+1} - IRDiff_{i,t}\right) = \beta_{i,0} + \beta_{i,1}\,\Delta LNSCDS_{i,t+1} + \varepsilon_{i,t+1}$.

Heteroscedasticity and autoregressive robust standard errors are used to test statistical significance of the coefficients. The *t*-statistic is reported in parentheses. To save space, we only report the $\beta_{i,1}$.

Table A3
Sovereign CDS Inception by Country.

	Dependent Variable				Sample Coverage
	|fwdbias|		*|UIRP|*		
Austria	−0.005	[−3.07]	−0.012	[−3.94]	1990–2017
Belgium	−0.006	[−3.34]	−0.012	[−3.89]	1990–2019
Canada	0.003	[1.91]	−0.006	[−2.38]	1990–2014
H.K.	−0.000	[−0.15]	−0.000	[−0.07]	1990–2014
Cyprus	−0.005	[−3.10]	−0.012	[−4.01]	1990–2017
Czech	−0.007	[−2.66]	−0.028	[−2.78]	1997–2014
Germany	−0.006	[−3.32]	−0.012	[−3.85]	1990–2019
Estonia	−0.006	[−3.19]	−0.013	[−3.78]	1990–2014
Finland	−0.005	[−3.07]	−0.012	[−3.94]	1990–2017
France	−0.005	[−3.13]	−0.012	[−3.92]	1990–2018
Greece	−0.005	[−2.95]	−0.012	[−3.88]	1990–2017
Iceland	−0.002	[−0.52]	−0.002	[−0.19]	1998–2017
Ireland	−0.006	[−3.29]	−0.012	[−3.87]	1990–2019
Italy	−0.006	[−3.54]	−0.012	[−3.96]	1990–2020
Japan	−0.004	[−1.70]	−0.014	[−2.54]	1990–2017
Latvia	−0.007	[−3.86]	−0.013	[−4.08]	1990–2014
Lithuania	−0.006	[−3.47]	−0.013	[−3.91]	1990–2018
Malaysia	−0.003	[−0.67]	−0.009	[−1.85]	1990–2020
Netherlands	−0.006	[−3.22]	−0.013	[−4.21]	1990–2017
Poland	0.001	[0.32]	−0.085	[−5.64]	1996–2019
Portugal	−0.006	[−3.39]	−0.012	[−3.94]	1990–2019
Slovakia	−0.006	[−3.44]	−0.013	[−3.92]	1990–2014
Slovenia	−0.006	[−3.55]	−0.013	[−3.92]	1990–2019
S. Africa	0.001	[0.23]	−0.032	[−1.93]	1999–2020
Spain	−0.006	[−3.45]	−0.012	[−3.94]	1990–2019
UK	0.001	[0.78]	−0.003	[−1.52]	1994–2018

This table shows the time-series regression for the inception of the sovereign CDS market. Heteroscedasticity and autoregressive robust standard errors are used to test the statistical significance of the coefficients. The *t*-statistic is reported in parentheses. To save space, we only report the coefficients for the *SCDSInception*.

Data availability

Data will be made available on request with the permission of the data provider.

Appendix

See Tables A1–A3.

References

Andersen, T.G., Bollerslev, T., Diebold, F.X., Vega, C., 2007. Real-time price discovery in global stock, bond and foreign exchange markets. J. Int. Econ. 73 (2), 251–277.

Augustin, P., 2018. The term structure of CDS spreads and sovereign credit risk. J. Monet. Econ. 96, 53–76.

Augustin, P., Tédongap, R., 2016. Real economic shocks and sovereign credit risk. J. Financ. Quant. Anal. 541–587.

Bacchetta, P., Van Wincoop, E., 2010. Infrequent portfolio decisions: A solution to the forward discount puzzle. Am. Econ. Rev. 100 (3), 870–904.

Backus, D.K., Gavazzoni, F., Telmer, C., Zin, S.E., 2010. Monetary policy and the uncovered interest parity puzzle. NBER working paper.

Backus, D.K., Foresi, S., Telmer, C.I., 2001. Affine term structure models and the forward premium anomaly. J. Financ. 56 (1), 279–304.

Badaoui, S., Cathcart, L., El-Jahel, L., 2013. Do sovereign credit default swaps represent a clean measure of sovereign default risk? A factor model approach. J. Bank. Financ. 37 (7), 2392–2407.

Bansal, R., 1997. An exploration of the forward premium puzzle in currency markets. Rev. Financ. Stud. 10 (2), 369–403.

Bansal, R., Dahlquist, M., 2000. The forward premium puzzle: Different tales from developed and emerging economies. J. Int. Econ. 51 (1), 115–144.

BIS Report, 2019. Triennial Central Bank Survey: Foreign exchange turnover in April 2019. URL: https://www.bis.org/statistics/rpfx19_fx.pdf.

Borri, N., Shakhnov, K., 2021. Global risk in long-term sovereign debt. Rev. Asset. Pric. Stud. 11 (3), 654–693.

Breedon, F., Rime, D., Vitale, P., 2016. Carry trades, order flow, and the forward bias puzzle. J. Money Credit Bank. 48 (6), 1113–1134.

Brunnermeier, M.K., Nagel, S., Pedersen, L.H., 2008. Carry trades and currency crashes. NBER Macroecon. Annu. 23 (1), 313–348.

Burnside, C., Eichenbaum, M., Rebelo, S., 2009. Understanding the forward premium puzzle: A microstructure approach. Am. Econ. J. Macroecon. 1 (2), 127–154.

Calice, G., Zeng, M., 2019. The term structure of sovereign CDS and the cross-section exchange rate predictability. Int. J. Financ. Econ.

Corsetti, G., Dasgupta, A., Morris, S., Shin, H.S., 2004. Does one Soros make a difference? A theory of currency crises with large and small traders. Rev. Econ. Stud. 71 (1), 87–113.

Coudert, V., Mignon, V., 2013. The forward premium puzzle and the sovereign default risk. J. Int. Money Financ. 32, 491–511.

Della Corte, P., Sarno, L., Schmeling, M., Wagner, C., 2022. Exchange rates and sovereign risk. Manag. Sci. 68 (8), 5591–5617.

Ding, L., Hiltrop, J., 2010. The electronic trading systems and bid-ask spreads in the foreign exchange market. J. Int. Finan. Markets. Inst. Money 20 (4), 323–345.

Dolatabadi, S., Nielsen, M.Ø., Xu, K., 2015. A fractionally cointegrated VAR analysis of price discovery in commodity futures markets. J. Futur. Mark. 35 (4), 339–356.
Du, W., Schreger, J., 2016. Local currency sovereign risk. J. Financ. 71 (3), 1027–1070.
Du, W., Tepper, A., Verdelhan, A., 2018. Deviations from covered interest rate parity. J. Financ. 73 (5), 915–957.
Eiteman, D.K., Stonehill, A.I., Moffett, M.H., 2009. Multinational Business Finance, 12th ed. Addison-Wesley, Reading, MA.
Fama, E.F., 1970. Efficient capital markets: A review of theory and empirical work. J. Financ. 25 (2), 383–417.
Fama, E.F., 1984. Forward and spot exchange rates. J. Monet. Econ. 14 (3), 319–338.
Fama, E.F., 1991. Efficient capital markets: II. J. Financ. 46 (5), 1575–1617.
Ferraro, D., Rogoff, K., Rossi, B., 2015. Can oil prices forecast exchange rates? An empirical analysis of the relationship between commodity prices and exchange rates. J. Int. Money Financ. 54, 116–141.
Figuerola-Ferretti, I., Gonzalo, J., 2010. Modelling and measuring price discovery in commodity markets. J. Econ. 158 (1), 95–107.
Foroni, C., Ravazzolo, F., Sadaba, B., 2018. Assessing the predictive ability of sovereign default risk on exchange rate returns. J. Int. Money Financ. 81, 242–264.
Garbade, K.D., Silber, W.L., 1983. Price movements and price discovery in futures and cash markets. Rev. Econ. Stat. 289–297.
Glassman, D., 1987. Exchange rate risk and transactions costs: Evidence from bid-ask spreads. J. Int. Money Financ. 6 (4), 479–490.
Greene, W.H., 2012. Econometric Analysis, 7th edition. Pearson.
Hodrick, R., 1987. The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets. Routledge.
Karnaukh, N., Ranaldo, A., Söderlind, P., 2015. Understanding FX liquidity. Rev. Financ. Stud. 28 (11), 3073–3108.
Krishnamurthy, A., Vissing-Jorgensen, A., 2012. The aggregate demand for treasury debt. J. Polit. Econ. 120 (2), 233–267.
Kumar, V., 2020. Liquidity shocks: A new solution to the forward premium puzzle. Econ. Model. 91, 445–454.
Lagos, R., Wright, R., 2005. A unified framework for monetary theory and policy analysis. J. Polit. Econ. 113 (3), 463–484.
Lee, S., 2020. Money, asset prices, and the liquidity premium. J. Money Credit Bank. 52 (6), 1435–1466.
Lee, S., Jung, K.M., 2019. A liquidity-based resolution of the uncovered interest parity puzzle. J. Money Credit Bank.
Lettau, M., Maggiori, M., Weber, M., 2014. Conditional risk premia in currency markets and other asset classes. J. Financ. Econ. 114 (2), 197–225.
Lewis, K.K., 1995. Puzzles in international financial markets. Handb. Int. Econ. 3, 1913–1971.
Lin, M.-T., Kolokolova, O., Poon, S.-H., 2021. Slow-and fast-moving information content of CDS spreads: New endogenous systematic factors. Eur. J. Financ. 27 (1–2), 136–157.
Longstaff, F.A., Pan, J., Pedersen, L.H., Singleton, K.J., 2011. How sovereign is sovereign credit risk? Am. Econ. J. Macroecon. 3 (2), 75–103.
Lucas Jr, R.E., 1982. Interest rates and currency prices in a two-country world. J. Monet. Econ. 10 (3), 335–359.
Lustig, H., Verdelhan, A., 2007. The cross section of foreign currency risk premia and consumption growth risk. Am. Econ. Rev. 97 (1), 89–117.
Lustig, H., Roussanov, N., Verdelhan, A., 2011. Common risk factors in currency markets. Rev. Financ. Stud. 24 (11), 3731–3777.
Malkiel, B.G., 2003. The efficient market hypothesis and its critics. J. Econ. Perspect. 17 (1), 59–82.
Markit. User Guide CDS & Bonds, 2012 June.
McGroarty, F., ap Gwilym, O., Thomas, S., 2009. The role of private information in return volatility, bid–ask spreads and price levels in the foreign exchange market. J. Int. Financ. Markets, Inst. Money 19(2), 387–401.
Meese, R.A., Rogoff, K., 1983. Empirical exchange rate models of the seventies: Do they fit out of sample? J. Int. Econ. 14 (1–2), 3–24.
Meese, R., Rogofp, K., 1988. Was it real? The exchange rate-interest differential relation over the modern floating-rate period. J. Financ. 43 (4), 933–948.
Nagel, S., 2016. The liquidity premium of near-money assets. Q. J. Econ. 131 (4), 1927–1971.
Osler, C.L., Mende, A., Menkhoff, L., 2011. Price discovery in currency markets. J. Int. Money Financ. 30 (8), 1696–1718.
Packer, F., Suthiphongchai, C., 2003. Sovereign credit default swaps. BIS Q. Rev.
Rosenberg, J.V., Traub, L.G., 2009. Price discovery in the foreign currency futures and spot market. J. Deriv. 17 (2), 7–25.
Sarno, L., 2005. Towards a solution to the puzzles in exchange rate economics: Where do we stand? Can. J. Econ. 38 (3), 673–708.
Sarno, L., Valente, G., Leon, H., 2006. Nonlinearity in deviations from uncovered interest parity: An explanation of the forward bias puzzle. Eur. Finan. Rev. 10 (3), 443–482.
Sun, X., Wang, J., Yao, Y., Li, J., Li, J., 2020. Spillovers among sovereign CDS, stock and commodity markets: A correlation network perspective. Int. Rev. Financ. Anal. 68, 101271.
Villanueva, O.M., 2007. Forecasting currency excess returns: Can the forward bias be exploited? J. Financ. Quant. Anal. 963–990.
Yang, J., Bessler, D.A., Leatham, D.J., 2001. Asset storability and price discovery in commodity futures markets: A new look. J. Fut. Markets: Fut. Opt. Other Deriv. Prod. 21 (3), 279–300.