# A Credit-Based Theory of the Currency Risk Premium*

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ABSTRACT

This paper extends the work of Kremens and Martin (2019) and uncovers a novel component for exchange rate predictability. Our theory shows that currency returns compensate investors for the expected currency depreciation in the case of a severe but rare credit event. We compute this risk compensation – the credit-implied risk premium (CRP) – by exploiting the price difference between sovereign credit default swaps denominated in different currencies. Using data for 17 Eurozone countries over the period 2010-19, we find that CRP positively forecasts the euro-dollar exchange rate return between one-week and six-month horizon, both in-sample and out-of-sample. We also show that currency trading strategies that exploit the informative content of CRP generate substantial out-of-sample economic value.

Keywords: Exchange rate, predictability, risk premium, credit risk, sovereign default
JEL codes: F31, F37, F47, G12, G15.

# 1 Introduction

Investors holding foreign government bonds face two major sources of risk. First, they bear the risk of a potential depreciation of the foreign currency and, second, there is a risk of bond value erosion caused by a deterioration in sovereign creditworthiness. These sources of risk are highly intertwined as sovereign defaults are commonly accompanied by large currency depreciation (e.g., Reinhart, 2002; Na et al., 2018).[1] This interaction has fundamental asset pricing implications for investors exposed to such currencies. Augustin et al. (2019) show that the risk-adjusted probability of a currency depreciation conditional on a default is substantially higher than the true probability, which implies that such events tend to occur in bad economic times for investors. Hence, currencies that are expected to depreciate severely in times of default are particularly risky and should deliver higher excess returns. Despite abundant empirical evidence pointing to a tight relation between depreciation and default, we lack a theoretically-motivated measure of currency depreciation conditional on sovereign default as well as empirical evidence on its implications for exchange rate predictability.

This paper makes an attempt to fill this gap in the literature in two respects. First, we develop a simple yet intuitive theory that uncovers a novel component for exchange rate predictability, which we label the *credit-implied risk premium*. This component adds to the *quanto-implied risk premium* of Kremens and Martin (2019), upon which our theory builds. While the quanto-implied risk premium captures the expected covariation between currency returns and frequent but small changes in US market conditions, the risk premium that we derive reflects investors' expectations about excess currency movements in times of rare but severe events abroad, such as a sovereign default. This risk premium compensates investors for a currency depreciation conditional on default *in excess* of what is unconditionally predicted by the uncovererd interest rate parity. Second, we empirically assess the exchange rate predictive ability of this novel risk premium and find robust evidence using statistical and economic criteria both in-sample and out-of-sample. We thus overturn the general wisdom that exchange rates are well approximated by a random walk model (e.g., Meese and Rogoff, 1983; Engel and West, 2005).

We exploit a unique feature of the sovereign CDS market to measure the risk-neutral expected

[1] Specifically, Reinhart (2002) shows that the probability of a severe currency depreciation around a sovereign default is about 84%, based on a sample of 58 countries between 1970 and 2002. Herz and Tong (2008) show that debt crises Granger cause currency crises in a sample of 108 emerging countries over the 1975-2005 period, while Mano (2013) shows, over the 1873-2008 period, that currencies fall on average by 17.6% during the default year and by 29.2% compared to five years earlier. Na et al. (2018) use data for 70 countries for the period 1975-2013 and report that the median exchange rate depreciates by 45% in a three-year window around a default event. A related literature finds that countries with increased sovereign credit risk, measured using sovereign credit default swaps (CDS), also experience a significant currency depreciation (e.g., Della Corte et al., 2018).

currency depreciation conditional on default, which is the primary determinant of our risk premium. Sovereign CDS for the same entity, maturity, and restructuring clauses are quoted in various currencies, but prices can be different because a CDS provides insurance against different risks depending on the currency denomination. A long position in a country's CDS quoted in local currency protects a US investor against a default in that country, whereas a long position in the same CDS but quoted in USD provides an additional hedge against the risk of a local currency depreciation upon default. Since the probability of default underlying a CDS quoted in different currencies is the same, we show that, under no arbitrage, the difference in prices reflects investors' expectations about a depreciation of the local currency relative to the USD upon a sovereign default.

The Eurozone is a perfect setting for our study because most member states have a liquid market for CDS instruments denominated both in USD and EUR.[2] In particular, the price difference between CDS in USD and EUR, which is commonly referred to as the "quanto spread", has been fairly large for a number of Eurozone countries with the unfolding European sovereign debt crisis. For example, Spain's 5-year CDS spread quoted on January 2, 2012 was worth 378 basis points in USD but only 290 basis points in EUR. Figure 1 shows that quanto spread has been always positive since 2010.

Armed with daily CDS spreads on 17 Eurozone member states quoted both in USD and EUR, we first compute a measure of expected currency depreciation conditional on default for each member state. We then obtain an aggregate measure of credit-implied risk premium for the Eurozone by weighting the country-specific components by their outstanding debt. We find that the risk-adjusted expected depreciation of the EUR relative to USD conditional on a sovereign default is 25%, on average, which is economically large. In addition, the expected (excess) depreciation level displays a substantial amount of volatility, thus suggesting that the associated risk premium strongly varies over time.

We focus on the return predictability of the USD/EUR exchange rate, which is the most liquid currency pair according to the latest Triennial Survey of the Bank for International Settlements (2016) with a daily turnover that exceeds half a trillion dollars. The risk premium related to expected depreciation in default positively predicts future USD/EUR returns, at any horizon between 1 week and 1 year. This finding is robust to controlling for existing currency predictors, such as the interest rate differential (Londono and Zhou, 2017), the quanto-implied risk premium (Kremens and Martin, 2019), the volatility risk premium (Della Corte et al., 2016a; Londono and Zhou, 2017), the liquidity premium (Karnaukh

[2] Alternatively, it is possible to estimate the expected currency depreciation from sovereign bond data, but this approach is less straightforward. It requires extrapolating credit risk from debt financial instruments written on the same entity with the same maturity and in at least two currencies. Government bonds respecting these constraints may be hard to find, especially for industrialized countries that tend to issue debt in their own currency.

et al., 2015; Mancini et al., 2013), and portfolio-based currency factors (Lustig et al., 2011; Menkhoff et al., 2012). At the 3-month horizon, the $t$-statistic of the slope coefficient associated with our risk premium is 3.16 and a one-standard-deviation increase in the risk premium induces an annualized excess currency return of 5.39%.[3] The risk premium associated with the tight relation between depreciation and default therefore contains valuable information for exchange rate predictability.

We also find superior out-of-sample exchange rate predictability relative to the traditional benchmark random walk model (or historical average) of Meese and Rogoff (1983). We assess this predictive power using the out-of-sample $R^2$ of Campbell and Thompson (2008) as well as the tests of equal accuracy for nested models developed by McCracken (2007) and Clark and West (2007), while bootstrapping the $p$-values using the algorithm of Mark (1995) and Kilian (1999). Our results lead to the conclusion that the USD/EUR exchange rate returns remain predictable out-of-sample over a horizon that varies between 1 week and 1 year. In addition to statistical evidence of exchange rate predictive ability, we find that the information content of CRP generates tangible out-of-sample economic gains to an investor using exchange rate forecasts in active portfolio management. Following Fleming et al. (2001) and Della Corte et al. (2009), among many others, we design an international asset allocation strategy whereby a US investor allocates her wealth between a dollar-denominated bond and a euro-denominated cash account while using CRP to predict the exchange rate return. We evaluate the performance of a dynamically rebalanced portfolio using a mean-variance analysis. This approach allows us to measure how much a risk-averse investor is willing to pay for switching from a portfolio strategy based on the random walk model to a competing one that exploits information in CRP. We find that the CRP-based strategy generates a substantial amount of out-of-sample economic value that outperforms the performance of the benchmark model using a sample of weekly non-overlapping observations. The profitability of our strategy survives to reasonably high transaction costs.

Currencies are typically exposed to crash risk, which relates to rare but severe adverse aggregate shocks (Farhi et al., 2009; Chernov et al., 2018). Currencies can crash for many reasons and a sovereign default is one of them. Yet we provide evidence that the credit-implied risk premium is only weakly related to crash risk, as measured by the currency option risk-reversal on the USD/EUR (with correlation of -0.44). Moreover, we find that the economic and monetary news that help explain fluctuations in this

[3] The results are robust to various specifications. First, we aggregate each country's CRP using either a country's outstanding sovereign debt or its GDP to compute weights. Second, we derive a daily synthetic version of the quanto-implied risk premium based on option prices and use it as a control. Third, we control for dealers' counterparty risk and, fourth, we investigate the return predictability in the case of non-overlapping data based on weekly observations. In all these cases, we find no substantial difference in our results.

currency risk premium substantially differ from those that drive crash risk, implied volatility, and quanto-implied risk premium. Hence, the risk premium associated with the interaction between default and depreciation complements existing types of risk prevalent in currency markets.

Our work relates to a growing literature on the currency denomination of sovereign CDS. Mano (2013) is the first to exploit the difference between sovereign CDS denominated in USD and in local currency. He concludes that a segmented markets model can generate predictions consistent with his empirical findings on the currency depreciation during sovereign defaults. Du and Schreger (2016) quantify the expected currency depreciation in emerging markets from the credit spread differential between sovereign bonds denominated in USD and in local currency.[4] De Santis (2015) uses the quanto spread to analyze the risk of currency redenomination in the Eurozone, while Corradin and Rodriguez-Moreno (2014) and Buraschi et al. (2015) exploit quanto spreads to explain pricing anomalies between bond yields denominated in different currencies.[5]

This paper also complements two contemporaneous studies. Lando and Bang Nielsen (2018) show that quanto spreads reflect the risk that a currency depreciates not only at the time of default but also as default risk increases. Their contribution is to decompose theoretically and empirically these two effects. Augustin et al. (2019) use the term structure of quanto spreads to offer an asset-pricing perspective on the relation between sovereign defaults and currency depreciation in the Eurozone and on the possibility of credit contagion. They address the debate on whether a default has an immediate or gradual impact on the exchange rate, thereby contributing to a better of understanding of the "Twin Ds" (depreciation and default). Their findings provide strong support in favor of the first channel. They also show that the currency risk premium associated with depreciation in default has an upward term structure (i.e., increases with the horizon), while we exploit the conditional properties of the currency risk premium at a given horizon to study exchange rate predictability. Taken together, these papers offer a broad picture on the asset pricing implications of the interaction between currency depreciation and default for the credit derivative and currency markets.

Overall, we contribute to the existing literature by identifying a credit-based currency risk premium that has strong implications for exchange rate return predictability. Since the path-breaking contribution of Meese and Rogoff (1983), a vast body of empirical studies finds that economically meaningful variables fail to empirically predict exchange rate returns. While there is some evidence that exchange

[4] The authors compute the credit risk components of sovereign yields in local and foreign currencies by creating an artificial local risk-free rate, based on the US treasury bonds, the US LIBOR rates, local LIBOR rates, and currency swaps.

[5] In a related study, Ehlers and Schoenbucher (2004) use Japanese corporate CDS data denominated in USD and JPY and analyze the expected exchange rate.

rates and economic fundamentals move together over long horizons (Mark, 1995), the general view is that exchange rates are not predictable, especially at short horizons. The central contribution of our paper is that we find robust empirical evidence that the USD/EUR exchange rate is predictable at short horizons, both in-sample and out-of-sample, using a novel theoretically-motivated risk premium measure. Our findings confirm the view that a risk premium capturing investors' expectations about a currency depreciation upon default is informative about future currency excess returns.

The remainder of the paper is organized as follows. Section 2 develops a theory that identifies a credit-implied currency risk premium. Section 3 quantifies this risk premium using the price difference of CDS denominated in USD and EUR and provides a descriptive analysis. Sections 4 and 5 conduct an analysis of exchange rate predictability in-sample and out-of-sample, respectively. We discuss the economic value of such predictability in Section 6, while Section 7 concludes. The Appendix contains technical details and presents additional results not included in the main body of the paper.

# 2 Theory

In this section, we extend the theory of Kremens and Martin (2019) and identify a novel source of currency risk premium.

## 2.1 Environment

Today is time $t$, and we focus on assets with payoffs at time $t+1$. We write $\mathbb{E}_t$ for the expectation operator (under the physical measure) conditional on all information available at time $t$, and $M_{t+1}$ for a stochastic discount factor (SDF) that prices assets denominated in dollars. An asterisk on the expectation (or covariance) term indicates that it is computed using the risk-neutral measure.

Let $S_t$ be the price in dollars at time $t$ of a unit of the foreign currency such that an increase in $S_t$ reflects an appreciation of the foreign currency relative to the dollar. The expected gross exchange rate return is then given by

$$\mathbb{E}_t\left[\frac{S_{t+1}}{S_t}\right] = \underbrace{\frac{R^{\$}_{f,t}}{R^{i}_{f,t}}}_{IRD_t} + RP_t, \tag{1}$$

where $R^{\$}_{f,t}$ is the domestic one-period gross interest rate, $R^{i}_{f,t}$ is the corresponding foreign currency gross interest rate, and $RP_t$ is a time-varying risk premium. Under risk-neutrality, Equation (1) is equivalent to the Uncovered Interest Parity (UIP) condition:

$$\mathbb{E}_t^*\left[\frac{S_{t+1}}{S_t}\right] = \frac{R_{f,t}^{\$}}{R_{f,t}^{i}}, \tag{2}$$

which implies that the higher interest-rate currency is expected to depreciate. UIP fails empirically, a stylized fact known as the forward premium puzzle, and the resulting excess return can be interpreted as compensation for some time-varying currency risk premium (e.g., Fama, 1984; Lustig et al., 2011). We now identify a potential driver of this risk premium that relates to expected currency movements in the case of a severe but rare credit event.

### 2.2 Global portfolio

Consider an investor holding a global portfolio measured in dollars, whose gross return $X_{t+1}$ satisfies the fundamental equation of asset pricing:

$$1 = \mathbb{E}_t\left[M_{t+1}X_{t+1}\right]. \tag{3}$$

We assume that the gross return $X_{t+1}$ consists of two risky parts:

$$X_{t+1} = 1 + r_{t+1} + d_{t+1}. \tag{4}$$

The first component, $r_{t+1}$, captures the dollar-return of a diversified portfolio of stocks (e.g., S&P 500), as in Kremens and Martin (2019). The second one, $d_{t+1} = a - \mathbf{1}_D b$, is the payoff of a dollar-denominated contingent claim. This claim pays a constant return $a$ and incurs a loss of $b$ when $\mathbf{1}_D = 1$, where $\mathbf{1}_D$ is an indicator function that takes the value 1 under a specific event and 0 otherwise. Denote by $\mathbb{Q}$ the risk-neutral probability that this event occurs between $t$ and $t+1$. The contingent claim paying $d_{t+1}$ is akin to an investment or an insurance contract whose payoff is conditional to the realization of a rare but severe event in the foreign country.[6] This environment nests Kremens and Martin (2019) when neglecting the contingent part of the portfolio return, i.e., $a = b = 0$.

We focus on a contingent claim related to sovereign credit events, which are rare but with severe financial consequences for investors. To fix ideas, think of a foreign government bond that returns a periodic coupon $a$, while losing $b$ (the unrecovered bond value) in the case of a default. Alternatively,

[6] The contingent claim can be any bond, derivative, or structured product whose payoff depends on a specific event, such as a recession, a natural disaster, a war, a political crisis, or a default, among others.

the investor can write a CDS on such bond, which entails receiving a periodic premium $a$ and delivering $b$ (the unrecovered bond value) to the protection buyer in the case of a default. Both cases induce a severe loss to the investor in the case of a sovereign default.

The global portfolio does not expose the investor to currency risk, as both portfolio return components ($r_{t+1}$ and $d_{t+1}$) are measured in dollars. However, a foreign currency becomes risky for the investor if it covaries with the performance of her global portfolio. The real-world expected return from holding the foreign currency between $t$ and $t+1$ is thus given by the following identity (see Appendix A.1 for the derivation):

$$\begin{aligned}\mathbb{E}_t\left[\frac{S_{t+1}}{S_t}\right] &= \frac{R^{\$}_{f,t}}{R^{i}_{f,t}} + \underbrace{\frac{1}{R^{\$}_{f,t}} cov^*_t\left(\frac{S_{t+1}}{S_t}, r_{t+1}\right)}_{\text{Quanto-implied risk premium}} \\ &+ \underbrace{\frac{\mathbb{Q}b}{R^{\$}_{f,t}}\left(\mathbb{E}^*_t\left[\frac{S_t - S_{t+1}}{S_t}\ |_{\mathbf{1}_D=1}\right] - \mathbb{E}^*_t\left[\frac{S_t - S_{t+1}}{S_t}\right]\right)}_{\text{Credit-implied risk premium}} + A_t, \qquad (5)\end{aligned}$$

where $A_t$ is a residual term that we assume equal to zero akin to Kremens and Martin (2019).

The identity displayed in Equation (5) indicates that currency excess returns compensate investors for two distinct sources of risk. The first term captures the conditional risk-neutral covariance between the exchange rate and the stock portfolio return $r_{t+1}$. Investors demand a positive currency risk premium if the foreign currency is expected to depreciate (under the risk-neutral measure) when the return on the domestic stock portfolio is low. Kremens and Martin (2019) refer to it as the *quanto-implied risk premium* (QRP).

The new risk premium component – the second term – reflects risk-neutral expectations about currency variations in the case of a sovereign credit event in the foreign country ($\mathbf{1}_D = 1$). We label this term the *credit-implied risk premium* (CRP). It is mostly determined by the risk-neutral expected currency variation conditional on this event, the risk-neutral probability of occurrence $\mathbb{Q}$, and by the magnitude of the contingent loss $b$.[7] If the foreign currency is expected to depreciate in the case of the default, this risk premium is positive. By contrast, a currency that is expected to appreciate in the case of the default provides an hedge to investors, and the risk premium becomes negative. We expect CRP to be positive, given that sovereign defaults are typically associated with a depreciation of the domestic

[7] This risk premium vanishes if either the exchange rate is independent on the event, the corresponding probability is nill ($\mathbb{Q} = 0$), or if the portfolio has no exposure to the event ($b = 0$).

currency (Na et al., 2018).

CRP consists of the difference between two expectations, as shown in Equation (5). The first one captures the expected currency depreciation conditional on a default, while the second one reflects the unconditional expected currency depreciation, based on the uncovered interest rate parity. Hence, investors receive a premium for the risk associated with the expected *excess* currency depreciation upon default, as given by the difference between the two expectations, rather than for the risk associated with the expected *total* currency depreciation.

This new source of risk premium adds to the expected covariance between the exchange rate return and the stock portfolio return $r_{t+1}$, which arises from small but frequent shocks in the domestic country. By contrast, CRP reflects the currency depreciation during a rare but severe foreign shock that does not necessarily impact the investor's portfolio return. For example, the sovereign default of Greece in 2012 has induced large losses to the holders of the Greek government bonds (and to the sellers of the corresponding CDS contracts), without significantly affecting U.S. stock market investors. Our generalization of Kremens and Martin (2019)'s framework thus highlights the presence of two separate and complementary currency risk premia. This additional risk premium shall improve our comprehension of currency excess returns, which is what we aim to verify empirically. Appendix A.1 provides the analytical derivation of Equation (5).

# 3 Computation of the credit-implied risk premium

In this section, we first exploit sovereign CDS data to determine the expected currency depreciation conditional on sovereign default under the risk-neutral measure. We then use this expectation to compute CRP and present some descriptive statistics.

## 3.1 Sovereign CDS and currency denomination

We first discuss a specific feature of the sovereign CDS market, which is the multiple currency denomination of CDS contracts. We explain how the price difference of CDS denominated in various currencies allows extracting expectations about future currency movements conditional on default.

A sovereign CDS is a credit derivative that protects its buyer from a sovereign credit event.[8] The

[8]The specified sovereign government, the bonds of which are protected, is called the reference entity. A credit event is commonly called a "default." There are four main types of credit events, as defined by the International Swaps and Derivatives Association (ISDA, 2003): Obligation acceleration, failure to pay the interest or principal, restructuring of debt, and repudiation or moratorium of debt.

default-protection buyer pays the seller a premium on a periodic basis. The annualized premium determines the CDS spread, which is quoted in fraction of the notional specified at the inception of the contract. In the case of a credit event, the protection seller compensates the holder with a contingent payment that reflects the unrecovered value of the underlying government bond. Identical sovereign CDS contracts can be denominated in various currencies, although the protection relates to the same credit event. While USD-denominated CDS are generally the most common and traded contracts, CDS denominated in local currency remain important for asset-liability management of banks' balance sheets and risk management purposes (Barclays, 2011). For example, European banks and investment funds largely use sovereign CDS (of European countries) denominated in EUR to offset a credit valuation adjustment.

The denomination of sovereign CDS in local currency exposes a US protection buyer to exchange rate fluctuations. Such exposure is risky because local currencies tend to depreciate in the event of a default, precisely when the protection buyer receives a positive payoff in that currency. A CDS quoted in USD protects its holders from such currency depreciation, which would erode the value of their claim in default. Consequently, investors price sovereign CDS in USD and local currency differently: They typically pay a higher premium to be hedged in USD, although the contracts are otherwise identical. For example, Spain's 5-year CDS spreads were quoted 378 basis points in USD and 290 basis points in EUR on January 2nd, 2012, for the same maturity, recovery rate, and restructuring clauses. Figure 1 shows that quanto spread has been always positive since 2010. The relatively higher price of USD-denominated CDS appears to be salient feature of the data. The difference between the CDS quotes in distinct currencies is commonly called the "quanto spread" or the "CDS quanto" by reporting dealers (J.P. Morgan, 2010).

FIGURE 1 ABOUT HERE

### 3.2 Implied currency depreciation

This section exploits the multiple currency denomination of sovereign CDS to quantify the expected currency depreciation conditional on a sovereign default, i.e., $\mathbb{E}_t^*\left[\frac{S_t - S_{t+1}}{S_t} \,|_{\mathbf{1}_D=1}\right]$ in Equation (5). This component, which we label as *implied currency depreciation* or simply ICD, is a key ingredient of CRP.

We compute ICD by implementing an arbitrage-free strategy that uses sovereign CDS denominated in two different currencies, namely EUR and USD. For ease of exposition, we consider a 2-period case and relegate the comprehensive derivation to Appendix B. Current time is $t$ and we consider a potential default at time $t+1$. Let $CS_t^{€}$ and $CS_t^{\$}$ be, respectively, the date-$t$ CDS spread denominated in EUR and USD, and $N^{€}$ and $N^{\$}$ their respective notional. The exchange rate $S_t$ denotes the date-$t$ USD price a unit of EUR.

A US investor implements the following long-short strategy: She is long a CDS denominated in EUR with nominal $N^{€}$ that pays $CS_t^{€}N^{€}$ each period, and short a CDS denominated in USD with nominal $N^{\$}$ that pays $CS_t^{\$}N^{\$}$ each period. The long-short position in two identical CDS contracts, except for the currency denomination, offsets the credit risk dimension. The agent hedges the currency risk associated with the EUR-denominated premium payments by entering a fixed-for-fixed currency swap in which EUR are received and USD are delivered at a swap rate. We set the swap rate equal to spot rate $S_t$ at the inception of the strategy, as in Du and Schreger (2016).

At inception, the date-$t$ cash flow of the long-short position, denoted by $CF_t$, is given by

$$CF_t \quad = \quad CS_t^{\$}N^{\$} - CS_t^{€}N^{€}S_t, \tag{6}$$

which implies that, for a self-financing strategy ($CF_t = 0$), the notional of the CDS denominated in USD is equal to $N^{\$} = \frac{CS_t^{€}N^{€}}{CS_t^{\$}}S_t$. At default, a CDS protection buyer receives the fraction of the unrecovered bond value $(1-R)$ times the exposure of the CDS contract, where $R \in (0,1)$ is the recovery rate of the underlying sovereign bond. The investor's long position in the EUR-denominated CDS implies receiving the amount $(1-R)N^{€}$ converted in USD at the exchange rate $S_{t+1}$, while the short position in the USD-denominated CDS implies a delivery of $(1-R)N^{\$}$.

In absence of arbitrage, the risk-neutral expectation at time $t$ of next period cash flow $CF_{t+1}$, conditional on default $1_{D=1}$, must then satisfy

$$\mathbb{E}_t^*\left[CF_{t+1|\mathbf{1}_{D=1}}\right] \quad = \quad (1-R)N^{€}\mathbb{E}_t^*\left[S_{t+1}\mid_{\mathbf{1}_{D=1}}\right] - (1-R)\frac{CS_t^{€}N^{€}}{CS_t^{\$}}S_t = 0, \tag{7}$$

where $\mathbb{E}_t^*\left[S_{t+1}\mid_{\mathbf{1}_{D=1}}\right]$ is the risk-neutral expected USD/EUR exchange rate conditional on a default at time $t+1$. It is then easy to show that the date-$t$ risk-neutral measure of the implied depreciation of the EUR relative to the USD conditional on default (denoted by $ICD_t$), corresponds to

$$\underbrace{\mathbb{E}_t^*\left[\frac{S_t - S_{t+1}}{S_t}\mid_{\mathbf{1}_{D=1}}\right]}_{ICD_t} = \frac{CS_t^{\$} - CS_t^{€}}{CS_t^{\$}}, \tag{8}$$

which shows that we can directly exploit the relative difference in CDS spreads in both currencies to quantify the implied currency depreciation at time $t$.[9]

There are two additional aspects that may be important in practice. The first one relates to the case in which a default occurs in between two premium payments. We must then consider the accrued premium, which is the remaining part of the CDS premium the protection buyer (seller) has to pay (receive). The cash flow at default time then equals $CF_{t+1} = [(1-R)N^{€} - \chi^{€}]S_{t+1} - [(1-R)N^{\$} - \chi^{\$}]$, where $\chi^{€}$ and $\chi^{\$}$ denote the accrued premiums of the CDS positions in EUR and USD, respectively. Figure 2 illustrates the cash flows before and at default with and without these accrued premiums. Appendix B compares summary statistics pertaining to our 1-year CRP and ICD measures computed in both cases (Table I). We verify that the results of this paper remain similar whether or not we account for this practical consideration.

Second, hedging the EUR flows into USD involves entering a currency swap that must be closed in the event of a default. For CDS contracts with maturity $T$ longer than one period, a default can occur before the maturity of the currency swap. In that case, one must determine the residual value of the currency swap, denoted by $V_{t_D}$ with $t < t_D \leq T$. The cash flow at default time $t_D$ becomes $CF_{t_D} = S_{t_D}[(1-R)N^{€}] - [(1-R)N^{\$} + V_{t_D}]$, which varies with default time $t_D$ because $V_{t_D}$ depends on the remaining time to maturity. We must, therefore, determine the cash flow $CF_{t_D}$ for each possible default time $t_D$ and weigh it with the corresponding risk-neutral default probability, which can be extracted from the term structure of CDS spreads. Appendix A provides the derivation of this general case. Closed-form solutions for ICD and CRP no longer exist in that case, and we must rely on an iterative procedure. The consideration of short maturity CDS contracts helps to conveniently overcome this issue. Specifically, with one-period maturity CDS contracts, the currency swap has no residual value in default, as it exactly coincides with the swap's maturity. This is our benchmark case throughout the paper.

FIGURE 2 ABOUT HERE

## 3.3 Data and construction

We now quantify CRP for the Eurozone. Our analysis uses mid-quotes CDS spreads in EUR and USD with the complete restructuring clause, as provided by Markit. Although USD-denominated CDS are

[9] The conditional depreciation of the EUR only relates to a default in Europe, thus implicitly ignoring the possibility of a default in the United States. That is, we do not analyze the case in which the EUR appreciates against the USD if a credit event is triggered in the United States.

the most traded contracts, sovereign CDS quoted in Euro are also reasonably liquid for each member of the Eurozone (Barclays, 2011). We consider the 17 countries that are part of the Eurozone, which are Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Ireland, Italy, Latvia, Lithuania, the Netherlands, Portugal, Slovenia, Slovakia, and Spain.[10] The sample spans the period between August 20, 2010, and April 26, 2019. The dataset contains a total of 38,046 quotes per maturity. We focus on 1-year CDS contracts because short-maturity CDS spreads tend to be particularly informative compared to longer-maturity CDS (Augustin, 2018). Moreover, the consideration of the 1-year maturity addresses the aforementioned issue related to the currency swap value at default.

Following our theoretically-motivated measure of the credit-implied risk premium (Equation 5), we first compute the empirical counterpart for each Eurozone country $i$ as follows:

$$CRP_{i,t} = \frac{b\mathbb{Q}_i}{R^{\$}_{f,t}} \left( \underbrace{\frac{CS^{\$}_{i,t} - CS^{€}_{i,t}}{CS^{\$}_{i,t}}}_{ICD_{i,t}} - \underbrace{\frac{R^{€}_{f,t} - R^{\$}_{f,t}}{R^{€}_{f,t}}}_{ECD_t} \right), \tag{9}$$

where $ICD_{i,t}$ employs 1-year CDS spreads on country $i$ denominated both in EUR and USD, and $ECD_t$ uses 1-year zero-coupon interest rates in gross terms derived from Bloomberg data. The difference between $ICD_{i,t}$ and $ECD_t$ determines country $i$'s expected currency depreciation conditional upon default in excess of the expected depreciation dictated by the uncovered interest rate parity condition. For the construction of country-specific measures of CRP, we also use the unconditional risk-neutral probability of default $\mathbb{Q}_i$ extracted from the first year of CDS spreads in USD (see Appendix C) and the loss given default $b = 0.6$ based on the ISDA convention.

We then construct the aggregate measure $CRP_t$ for the Eurozone as the weighted average of each country's $CRP_{i,t}$,

$$CRP_t = \sum_i \omega_i CRP_{i,t}, \tag{10}$$

where the weight $\omega_i$ reflects the relative size of the country $i$'s total sovereign debt outstanding at the start of the sample period (i.e., 2010), which are obtained from Bloomberg. Countries with larger sovereign debt are expected to have a greater impact on the EUR in the case of a default, and thus naturally contribute more to CRP.

[10] We only include Greece in the robustness analysis presented in Section 4.4 because of infrequent quotes.

### 3.4 Analysis of the credit-implied risk premium

There is strong variation in ICD over time and across countries, as illustrated by Figure 3. Table 1 indicates that the safest but economically large countries (e.g., Austria, France, Germany) exhibit the greatest risk-neutral expectation of the EUR depreciation in default. The level of creditworthiness is also highly heterogenous among Eurozone countries, which implies that CRP varies substantially in the cross section. As expected, the five countries with the highest level of CRP are (by decreasing order) Portugal, Ireland, Cyprus, Spain, and Italy, which are relatively high credit-risky countries. The countries with the lowest CRP are either countries playing a negligible financial role in the Eurozone such as Estonia, Finland, and Latvia, or high creditworthy countries such as Germany.

FIGURE 3 AND TABLE 1 ABOUT HERE

At the Eurozone level, the EUR is expected to depreciate by 25% conditional on a sovereign default within a 1-year horizon. Such depreciation is sizable, by any standard. The corresponding risk premium (CRP) is 0.11%, on average.[11] This risk premium for the USD/EUR appears to be small, but this is due to the probability of a sovereign default being negligible at the 1-year horizon.[12] Figure 4 suggests substantial time variation in ICD and CRP at the aggregate level. Over the sample period, the expected EUR conditional on default is between 0.8% and 46.2% lower than the current spot exchange rate, while CRP fluctuates between 0.03% and 0.20%.

FIGURE 4 ABOUT HERE

## 4 Exchange rate return predictability

In this section, we examine the role of CRP for exchange rate predictability. In particular, we investigate the predictive power of CRP beyond that of the interest rate differential, the quanto-implied risk premium, FX liquidity, FX volatility, and traditional currency factors. We start by describing the set of control variables and then discuss the results.

[11] The behavior of CRP remain very similar if we consider the accrued premium. As shown in Appendix D, the difference in CRP with and without the accrued premium is negligible and the sample correlation between them is above $99\%$. To keep our analysis simple, we will ignore the accrued premium throughout the paper.

[12] Consistent with this finding, Augustin et al. (2019) estimate the risk premium associated with the EUR depreciation in default with an affine non-arbitrage model and find that the risk premium has an upward-sloping term structure. They obtain a risk premium that is close to zero over the 1-year horizon.

## 4.1 Control variables

The identity of Equation (5) indicates that the interest rate differential (IRD), which reflects the UIP forecast, and the quanto-implied risk premium are two theoretically-founded predictors of excess currency returns. We thus control for IRD, as in Londono and Zhou (2017) and Kremens and Martin (2019), which we define as the difference between the zero-coupon 1-year rates in EUR and in USD. We also control for Kremens and Martin (2019)'s quanto-implied risk premium (QRP) on the USD/EUR exchange rate, which we obtain from the authors. The correlation between CRP and QRP is 0.38, confirming the view that these two predictors capture distinct components of the currency risk premium.

Empirical evidence suggests that liquidity and volatility in the FX market also predict future exchange rate returns (Mancini et al., 2013; Menkhoff et al., 2012). We thus employ the updated illiquidity measure of Karnaukh et al. (2015) for the USD/EUR, as kindly shared by Angelo Ranaldo, and the realized volatility akin to Menkhoff et al. (2012) on the USD/EUR. We additionally consider the volatility risk premium (VRP), which has been shown to drive exchange rate returns (Della Corte et al., 2016a; Londono and Zhou, 2017). We construct the VRP for the USD/EUR as the difference between the lagged 1-year realized volatility and the 1-year model-free implied volatility. Figure 5 illustrates the time series of the main variables.

FIGURE 5 ABOUT HERE

A recent stream of the literature also highlights the role of global factors for exchange rate returns based on currency portfolios (e.g., Lustig et al., 2011). We thus construct currency factors based on the 20 most liquid currency pairs extracted from Datastream. The strategies are based on six criteria resulting in six factors that we consider in our set of controls. We have Lustig et al. (2011)'s dollar and carry-trade factors, as well as the global imbalance, momentum, value, and risk-reversal currency factors. Appendix E provides details on the computation of all control variables and presents descriptive statistics.

### 4.2 Baseline Spedicfication

We evaluate the in-sample predictive content of CRP by running regressions based on the following specification at the daily frequency:

$$\frac{1}{k}\Delta s_{t+k} = a_k + b_k CRP_t + c_k IRD_t + d_k X_t + \varepsilon_{t+k}, \tag{11}$$

where $\Delta s_{t+k}$ is the log USD/EUR exchange rate between days $t$ and $t+k$ in annual terms, $CRP_t$ is the percentage credit-implied risk premium for the Eurozone on day $t$, $IRD_t$ is the 1-year interest rate differential, and $X_t$ captures a set of control variables. Table 2 reports the least-squares estimates of $b_k$, for an horizon $k$ running between 1-week and 1-year, with and without control variables.

TABLE 2 ABOUT HERE

Consistent with our theory, we find that CRP positively predicts future USD/EUR exchange rate returns. In Panel A, we do not include any control variables and report statistically significant estimates of $b_k$ ranging between $2.229$ at the 1-week horizon and $1.009$ at the 1-year horizon. This evidence is robust to including various control variables such as USD/EUR realized liquidity and volatility (Panel B), and global factors obtained from currency strategies based on carry, momentum, value, external imbalances, and option risk reversals. In Panel D, we further include all controls but CRP remains a strong predictor of future USD/EUR exchange rate returns. Our findings are also important in economic terms. The coefficient estimate of 1.685 at 3-month horizon in Panel D, for example, suggests that a one-standard-deviation increase in CRP predicts a future exchange rate appreciation of about $5.39\%$ per annum. Overall, our analysis indicates that an increase in CRP helps predict the future USD/EUR exchange rate return. The information we exploit, moreover, is not spanned by FX volatility, FX liquidity, and traditional global currency factors.

### 4.3 Controlling for QRP

We now account for the quanto-implied risk premium in the analysis. Specifically, we first predict USD/EUR exchange rate returns using the QRP measure of Kremens and Martin (2019) at various horizons $k$ and then check whether the regression residuals $\eta_{t+k}$ are related to our CRP measure using the following specification

$$\frac{1}{k}\Delta s_{t+k} = \alpha_k + \beta_k QRP_t + \gamma_k IRD_t + \eta_{t+k} \tag{12}$$

$$\eta_{t+k} = \phi_k + \omega_k CRP_t + \delta_k X_t + \psi_{t+k}. \tag{13}$$

Since QRP is only available at the monthly frequency until November 2015, we first retrieve the daily missing observations by a forward filling procedure and then restrict our analysis between August 2010 and November 2015. We report the estimates associated with the first equation in Panel A of Table 3 and confirms that QRP predicts future USD/EUR exchange rate returns between 1-month and 2-year ahead.[13] CRP drives the unexplained component $\eta_{t+k}$ at any horizon $k$ without controls (Panel B) and with controls (Panel C), respectively. The magnitude of the coefficient ($\omega_k$) and its statistical significance are similar across both cases. In sum, this analysis provides evidence that CRP and QRP contains complementary information for exchange rate predictability, as suggested by our theory.

TABLE 3 ABOUT HERE

We also augment our baseline specification in Equation (11) and include QRP as an additional explanatory variable. We present the results in Panel A of Table 4 and find that CRP predicts the USD/EUR exchange rate returns at all horizons, while controlling for QRP, IRD, and all other additional variables. Recall that QRP is only available at the monthly frequency and we obtain daily observations by forward filling, i.e., we keep the last observation constant until a new observation becomes available. This procedure, however, may underestimate the information content of QRP. To overcome this concern, we propose a theoretically-motivated proxy for QRP based on the implied variance of the USD/EUR exchange rate return. In particular, if the investor holds a foreign risk-free bond as opposed to a diversified portfolio of domestic stocks, QRP is equal to $IVAR_t \times R^{€}_{f,t}/R^{\$}_{f,t}$, where $IVAR_t$ is the risk-neutral variance of the USD/EUR exchange rate return. We report the details of this derivation in Appendix A.2. For this exercise, we employ 1-year over-the-counter currency options and quantify $IVAR_t$ using the model-free approach of Britten-Jones and Neuberger (2000). As a robustness, we also implement the method recently proposed by Martin (2017) as well as use the squared implied volatility from ATM options.[14] Armed with our synthetic versions of QRP, we revisit the predictability of CRP,

[13] We extend the analysis to a 2-year horizon to match the exercise reported in Kremens and Martin (2019).

[14] Appendix F.1 shows empirically that our synthetic version of QRP, labeled as QRP$^{S}$, is tightly related to QRP using a panel of 11 exchange rates analyzed in Kremens and Martin (2019).

report the results in Table 4, and show that CRP remains a strong predictor of the future USD/EUR exchange rate returns.

TABLE 4 ABOUT HERE

### 4.4 Robustness and extensions

In this section, we provide several extensions and additional analyses for robustness.

#### 4.4.1 Alternative specifications

We first verify that our empirical findings are robust to the aggregation method of CRP. Table 5 reproduces the analysis of Table 2 when we aggregate each country's CRP using GDP weights (at the start of the sample) rather than weighting by a country's outstanding public debt level. The results remain very similar.

TABLE 5 ABOUT HERE

Although our focus is on the aggregate measure of CRP, it is important to determine which countries drive the predictive power of CRP for USD/EUR returns. In Appendix F.2, we consider country-level CRP and find that the predictability of CRP arises from the economically impactful countries and not by a few small European countries with negligible consequences for the Eurozone.[15]

One may argue that the difference in EUR- and USD-denominated CDS spreads on the same underlying entity could be attributed to dealers' credit risk, as opposed to interaction between default and depreciation. This would be the case if, for example, CDS in EUR are largely quoted by European banks, while CDS in USD are mostly quoted by US banks. We address this concern and provide evidence in Appendix F.3 that CRP is robust to controlling for dealers' counterparty risk.

Finally, we verify in Appendix F.4 that the predictability of USD/EUR excess returns comes from the time variation in implied currency depreciation and not from the default probability.[16] This test

[15]The countries exhibiting the highest predictability of USD/EUR excess returns are the risky countries that contributed largely to the European debt crisis, such as Ireland, Italy, Portugal, and Spain, and the two economically large creditworthy entities, which are France or Germany.

[16]As a counterfactual, we construct an alternative measure of CRP, where implied currency depreciation is constant but the default probability varies over time, and show that this alternative CRP measure is unable to predict USD/EUR excess returns at any horizon.

confirms that the time variation in markets' expectations on the currency depreciation upon default contains novel information for understanding exchange rate returns.

#### 4.4.2 Counterfactual analysis

We now consider a counterfactual exercise to verify that CRP reflects a risk premium specifically related to the EUR depreciation in the case of a sovereign default. If CRP were a mere proxy for a global currency risk premium, as opposed to what our paper suggests, variations in CRP should also predict non-EUR exchange rate returns. Similarly, one may be worried that the predictability of the USD/EUR with CRP is driven by variations in the USD and not in the EUR.

We now address these concerns. For this exercise, we replace the USD/EUR by the USD/JPY in our baseline predictability analysis and test the hypothesis that CRP should not help explain the evolution of the USD/JPY. The JPY is a natural candidate because it is a highly liquid currency that is only weakly related to the EUR, as opposed to the CHF or the GBP. Confirming our hypothesis, Table 6 shows that the predictability of the USD/JPY returns with CRP is never statistically significant at any horizon. Therefore, a higher CRP predicts an appreciation of the EUR relative to the USD through the increased risk embedded in the EUR, and not through a global risk premium driving the USD.

TABLE 6 ABOUT HERE

#### 4.4.3 Non-overlapping analysis

One may be concerned by the persistence in the cumulative exchange rate returns, especially at a relatively long horizon. To show that this issue is not driving our findings, we investigate the return predictability in the case of non-overlapping data. Table 7 reports the results based on weekly observations for a horizon between 1 and 4 weeks, using our benchmark specification.[17] We use middle of the week's observations to reduce the microstructure noise associated with the beginning and the end of the week trading activity.[18] Table 7 shows the results of the regressions based on Tuesday's, Wednesday's, or Thursday's observations. The coefficients are almost always statistically significant and are similar across the different days considered to sample the data. The predictability of CRP thus continues to hold when forecasting exchange rate returns with non-overlapping observations.

[17] We control for the 12-month EUR-USD interest rate differential but not for QRP, given the frequency of this analysis.

[18] Using weekly data also overcomes the concern of a time zone differential arising because CDS in EUR (USD) appear to be traded mostly in London (New York), according to informal conversations with practitioners.

TABLE 7 ABOUT HERE

#### 4.4.4 Residual covariance term

The identity in Equation (5) contains a residual covariance term, which is considered to be constant over time as in Kremens and Martin (2019). We now relax this assumption and measure this residual covariance term as the regression residuals $\varepsilon_{t+k}$ of the following specification:

$$\frac{1}{k}\Delta s_{t+k} = a_k + b_k CRP_t + c_k IRD_t + d_k QRP_t + \varepsilon_{t+k}. \tag{14}$$

We study the correlations between the regression residuals, based on the 3-month horizon, and our main variables to have a better understanding of what this term captures. Figure 6 reports the results. The residual covariance term is strongly related to VRP, realized volatility, and to the level of FX illiquidity. This result confirms the importance of using these variables as controls for currency excess return predictability, consistent with the existing literature. The residual covariance term is however almost independent of QRP. Importantly, we find the residual covariance term is positively related to CRP, thereby providing further support to our theory. That is, higher CRP is associated with higher future USD/EUR excess returns. These findings remain very similar when we estimate the residuals from the predictive regression (Panel A) or when we compute their theoretical counterparts (i.e., with $a_k = 0$ and $b_k = c_k = d_k = 1$, as reported in Panel B), which demonstrates the stability in the results.

FIGURE 6 ABOUT HERE

#### 4.4.5 Comparison with alternative sources of risk

Different types of risk affect the currency market and they are potentially related to one another. We address this concern and show that CRP reflects fundamentally different information from that embedded in other kinds of risk. Farhi et al. (2009) and Chernov et al. (2018) find that most currencies are exposed to crash risk, which relates to rare but severe adverse aggregate shocks such as a recession or a sovereign default. It is thus particularly important to verify that CRP is distinct from crash risk and to the level of credit risk in the Eurozone.

Following the literature, we measure crash risk for the USD/EUR as the difference in implied volatility between USD/EUR puts and calls with 25 deltas. We measure credit risk as the debt-weighted average of each country's 1-year CDS spreads denominated in USD. Panels A and B of Figure 7 show that a high level of CRP does not proxy for elevated levels of crash risk or credit risk. In fact, the correlation between CRP and crash risk is -0.44 in our sample, while the correlation is -0.22 with credit risk.

FIGURE 7 ABOUT HERE

To better understand how CRP differs from crash risk, credit risk, and QRP, we provide a comprehensive analysis of their determinants based on a regression analysis. We present and discuss the results in Appendix F.5. In sum, CRP is pro-cyclical, whereas the other sources of risk are countercyclical. Also, we find that real, financial, and nominal conditions drive CRP differently than they drive crash risk, credit risk, and QRP. This analysis confirms that CRP is a novel source of currency premium.

# 5 Forecast evaluation: A statistical perspective

In this section, we provide statistical tests of out-of-sample predictive ability, which are critical for evaluating the success or failure of empirical exchange rate models.

## 5.1 Benchmark model

Since the seminal contribution of Meese and Rogoff (1983), a large body of research finds that empirical specifications based on economically meaningful variables fail to provide accurate out-of-sample exchange rate forecasts, thus leading to the prevailing view that exchange rates are not predictable, especially at short horizons (e.g., Mark, 1995). As a result, the random walk (RW) model has become *de facto* the benchmark model to evaluate the predictive ability for exchange rate returns. We now test the null hypothesis of equal predictive ability between the RW-based model and a CRP-based specification. The latter can be seen as an unrestricted model that nests the former, restricted one.

## 5.2 Out-of-sample tests of forecast accuracy

We describe a set of statistical criteria based on out-of-sample forecasts. First, we follow Campbell and Thompson (2008) and compute, for each forecast horizon $k$, the out-of-sample $R^2$ statistic as

$$R^2_{OOS,k} = 1 - \frac{\sum_M \left(r_{t+k} - \widehat{r}_{t+k|t}\right)^2}{\sum_M \left(r_{t+k} - \overline{r}_{t+k|t}\right)^2} = 1 - \frac{MSE_{CRP,k}}{MSE_{RW,k}},$$

where $r_{t+k}$ is the USD/EUR exchange rate return between times $t$ and $t+k$, $\overline{r}_{t+k|t}$ is the $k$-period ahead forecast of the benchmark model made at time $t$, $\widehat{r}_{t+k|t}$ is the corresponding forecast of our alternative model, and $M$ is the total number of out-of-sample forecast observations. This statistics is effectively comparing the forecast errors of the RW model to the forecast errors of a CRP-based model, and a positive value implies that the alternative model outperforms the benchmark model. This statistic can be also rewritten in terms of out-of-sample mean-square errors ($MSE$) with the subscript $CRP$ ($RW$) referring to the CRP-based (random walk) model.

The second statistic is the out-of-sample root mean-squared error difference of Welch and Goyal (2008). This statistics is computed as

$$\Delta RMSE_k = \sqrt{MSE_{RW,k}} - \sqrt{MSE_{CRP,k}},$$

and a positive value means that a model exploiting the informational content of CRP outperforms the benchmark RW model.

The most popular method for testing whether an alternative model has a lower MSE than the benchmark is using the Diebold and Mariano (1995) and West (1996) statistic, which has an asymptotic standard normal distribution when comparing forecasts from non-nested models. However, as shown by McCracken (2007), this statistic has a non-standard distribution when comparing forecasts from nested models and is severely undersized when using standard normal critical values. To this end, we employ the $MSE$-$F$ statistic of McCracken (2007), a variant of the Diebold and Mariano (1995) and West (1996) statistic designed to test the equal predictive ability of nested models, defined as

$$MSE\text{-}F = (M - k + 1) \times \frac{MSE_{RW,k} - MSE_{CRP,k}}{MSE_{CRP,k}},$$

where $M$ is the number of out-sample forecasts, and $k$ is the degree of overlap.[19]

Finally, we apply the procedure developed by Clark and West (2007) to test the null of equal predictive ability of two nested models. This approach acknowledges the fact that under the null of no predictability the MSE of an alternative model is expected to be greater than the MSE of the benchmark because the former specification introduces noise into the forecasting process by estimating an additional parameter that is not helpful for the prediction. Finding that the benchmark model has a smaller MSE does not

[19]When the models are correctly specified, the forecast errors are serially uncorrelated and exhibit conditional homoskedasticity. In this case, McCracken (2007) numerically generates the asymptotic critical values for the $MSE$-$F$ test. When the above conditions are not satisfied, we use a bootstrap procedure to compute valid critical values.

necessarily provides clear evidence against the alternative model. Clark and West (2007) propose a correction for the MSE of the alternative model. A computationally convenient way of testing for equal MSE is to define

$$CW_{t+k|t} = (r_{t+k} - \overline{r}_{t+k|t})^2 - [(r_{t+k} - \widehat{r}_{t+k|t})^2 - (\overline{r}_{t+k|t} - \widehat{r}_{t+k|t})^2] \tag{15}$$

and then regress $CW_{t+k|t}$ on a constant while using the $t$-statistic for a zero coefficient.The $MSE\text{-}F$ statistic of McCracken (2007) and the $CW$ statistic of Clark and West (2007) test the null hypothesis that both the benchmark and the competing model have equal MSE against the alternative that the competing model has lower MSE.

For all these test statistics, we obtain bootstrapped critical values for a one-sided test by estimating the model and generating 1,000 bootstrapped time series under the null of no predictability, as in Mark (1995) and Kilian (1999). The procedure preserves the autocorrelation structure of the predictive variable and maintains the cross-correlation structure of the residual. The bootstrap algorithm is summarized in Appendix G.

## 5.3 Empirical evidence

We assess the statistical performance of CRP by reporting the out-of-sample tests of predictability against the benchmark RW model. On each day $t$, we regress the log USD/EUR exchange rate return on the lagged CRP measure and produce forecasts between 1-week and 1-year horizon using a 1-year rolling window of data. Following Campbell and Thompson (2008), we impose an economic sign restriction on the slope coefficient of CRP, i.e., we set the slope coefficient equal to zero when its estimate is negative. Such restriction is consistent with the theoretical prediction of our model. Moreover, it mitigates the parameter instability arising from using a short window of data and help improve the OOS performance. The role of these restrictions becomes negligible when we increase the rolling window between 18-months and 24-months.

We report the test statistics discussed above along with the bootstrapped $p$-values in Table 8. Panel A (column I) displays the results for the benchmark CRP measure and find evidence of superior predictive ability against the RW model for $k$ ranging between 1 month and 6 months. When $k = 3$ months, for instance, the $R^2_{OOS}$ is above 4% and is statistically different from zero at the 5% confidence level. This result is further corroborated by the $\Delta RMSE$, $MSE\text{-}F$, and $CW$ which point in the same direction. In Panel B, we use an alternative (GDP-based) weighting scheme for our CRP measure and find qualitatively similar results. We then implement the economic sign restrictions akin Campbell and Thompson (2008) in the column II of Panels A and B and find better out-of-sample results: we uncover evidence of superior predictive ability of CRP against RW for $k$ ranging between 1 week and 1 year.

TABLE 8 ABOUT HERE

## 6 Economic value of FX predictability

In this section, we assess the performance of an asset allocation strategy that exploits the predictability in exchange rate returns. We first describe the framework and then present the out-of-sample empirical assessment of dynamically rebalanced portfolios using 1-week ahead USD/EUR forecasts based on CRP.

### 6.1 Maximum expected return rule

We design a simple asset allocation strategy that involves trading the USD/EUR exchange rate in the spirit of West et al. (1993), Fleming et al. (2001), and Della Corte et al. (2009). Specifically, we consider a US investor who builds a portfolio by allocating her wealth between a short-term bond in USD and a cash account in EUR. While the domestic bond yields a riskless return $r^{\$}_{f,t}$, the cash account delivers a zero interest rate in local currency but a risky return in USD terms at time $t+1$. Its expected return at time $t$, then, amounts to the expected USD/EUR exchange rate return such that any positive excess return is purely driven by exchange rate predictability.[20]

Our investor rebalances her mean-variance portfolio every period using one-step ahead conditional mean return and conditional volatility forecasts. The former is given by $\mathbb{E}_t[r_{t+1}] = \mathbb{E}_t[\Delta s_{t+1}]$ and we consider two distinct models to form exchange rate return expectations, namely the random walk model of Meese and Rogoff (1983) and a predictive regression that exploits information in CRP. For the latter component, we simply employ the sample standard deviation of the regression residuals at the time we produce the exchange rate forecast, i.e., $\mathbb{V}_t[r_{t+1}] = \sigma_\varepsilon$. In other words, we do not model the dynamics of FX return volatility and the optimal weights will vary over time only to the extent that the predictive regressions produce better forecasts of the exchange rate returns. To sum up, conditional on the forecasts generated by a given model, the investor dynamically rebalances her portfolio every

[20]Results remain qualitatively identical if one considers a short-term bond in EUR since interest rate in the Eurozone are approximately zero in our sample. Our setting, moreover, avoids any concerns stemming from negative interest rates.

period by computing the new optimal weights. This setup is designed to assess the economic value of exchange rate predictability by informing us which empirical exchange rate model leads to a better performing allocation strategy.

At each period $t$, our investor will solve the following problem:

$$\max_{w_t} \quad \mathbb{E}_t[r_{p,t+1}] = w_t \mathbb{E}_t[r_{t+1}] + (1 - w_t)\, r_{f,t}^{\$}$$

$$\text{s.t. } \sigma_p^\star = w_t \mathbb{V}_t[r_{t+1}],$$

where $\sigma_p^\star$ is the target volatility of the portfolio returns, $w_t$ is the weight on the risky asset, and $\mathbb{E}_t[r_{p,t+1}]$ is the expected return of the portfolio strategy. This strategy delivers an investor's realized portfolio return given by $r_{p,t+1} = w_t r_{t+1} + (1 - w_t)\, r_{f,t}^{\$}$.

## 6.2 Performance measures

We assess the economic value of exchange rate predictability with a set of standard mean-variance performance measures. We begin our discussion with the Fleming et al. (2001) performance fee, which is based on the principle that at any point in time, one set of forecasts is better than another if investment decisions based on the first set lead to higher average realized utility (West et al., 1993). The performance fee is computed by equating the average utility of the random walk (RW) optimal portfolio with the average utility of the alternative (i.e., CRP-based) optimal portfolio, where the latter is subject to expenses $\mathcal{F}$. Since the investor is indifferent between these two strategies, we interpret $\mathcal{F}$ as the maximum performance fee she will pay to switch from the RW to the CRP-based strategy. In other words, this utility-based criterion measures how much a mean-variance investor is willing to pay for conditioning on better exchange rate forecasts. The performance fee will depend on $\delta$, which is the investor's degree of relative risk aversion (RRA). To estimate the fee, we find the value of $\mathcal{F}$ that satisfies:

$$\sum_{t=0}^{T-1} \left\{ \left(R_{p,t+1}^\star - \mathcal{F}\right) - \frac{\delta}{2\,(1+\delta)} \left(R_{p,t+1}^\star - \mathcal{F}\right)^2 \right\} = \sum_{t=0}^{T-1} \left\{ R_{p,t+1} - \frac{\delta}{2\,(1+\delta)} R_{p,t+1}^2 \right\}, \tag{16}$$

where $R_{p,t+1}^\star$ is the gross portfolio return constructed using the forecasts from CRP, and $R_{p,t+1}$ is the gross portfolio return implied by the benchmark RW model. This quadratic utility allows us to consider nonnormal distributions of returns, while providing a high degree of analytical tractability. Additionally, quadratic utility may be viewed as a second-order Taylor series approximation to expected utility.[21]

We also evaluate performance using the premium return, which builds on the Goetzmann et al. (2007) manipulation-proof performance measure and is defined as:

$$\mathcal{P} = \frac{1}{(1-\delta)} \ln \left[ \frac{1}{T} \sum_{t=0}^{T-1} \left( \frac{R^{\star}_{p,t+1}}{R^{\$}_{f,t}} \right)^{1-\delta} \right] - \frac{1}{(1-\delta)} \ln \left[ \frac{1}{T} \sum_{t=0}^{T-1} \left( \frac{R_{p,t+1}}{R^{\$}_{f,t}} \right)^{1-\delta} \right], \tag{17}$$

which is robust to the distribution of portfolio returns and does not require the assumption of a particular utility function to rank portfolios, in contrast to the Fleming et al. (2001) performance fee that assumes a quadratic utility function. We can interpret $\mathcal{P}$ as the certainty equivalent of the excess portfolio returns and hence can also be viewed as the maximum performance fee an investor will pay to switch from the benchmark to another strategy. In other words, this criterion measures the risk-adjusted excess return an investor enjoys for using the information content of CRP rather than assuming a random walk. We report both $\mathcal{F}$ and $\mathcal{P}$ in basis points (*bps*) per annum.

In the context of the mean-variance analysis, a commonly used measure of economic value is the Sharpe ratio ($\mathcal{SR}$). The realized $\mathcal{SR}$ is equal to the average excess return of a portfolio divided by the standard deviation of the portfolio returns.[22] We also compute the Sortino ratio ($\mathcal{SO}$), which measures the excess return to bad volatility. Unlike the $\mathcal{SR}$, the $\mathcal{SO}$ differentiates between volatility due to up and down movements in portfolio returns. It is equal to the average excess return divided by the standard deviation of negative returns only. In other words, the $\mathcal{SO}$ does not take into account positive returns in computing volatility because these are desirable. A large $\mathcal{SO}$ indicates a low risk of large losses.

### 6.3 Impact of transaction costs

The effect of transaction costs is also an essential consideration in assessing the profitability of dynamic trading strategies. We account for this effect by calculating the break-even proportional transaction cost, $\tau^{be}$, that renders investors indifferent between two strategies (e.g., Han, 2006; Della Corte et al., 2009). We assume that $\tau^{be}$ is a fixed fraction of the value traded in all assets in the portfolio. Then, the cost of the dynamic strategy is $\tau^{be}|w_t - w_{t-1}\frac{1+r_t}{1+r_{p,t}}|$. In comparing a dynamic CRP-based strategy

[21] In an investigation of the empirical robustness of the quadratic approximation, Hlawitschka (1994) finds that a two-moment Taylor series expansion "may provide an excellent approximation" to expected utility and concludes that the ranking of common stock portfolios based on two-moment Taylor series is "almost exactly the same" as the ranking based on a wide range of utility functions.

[22] It is well known that because the $\mathcal{SR}$ uses the sample standard deviation of the realized portfolio returns, it overestimates the conditional risk an investor faces at each point in time and hence underestimates the performance of dynamic strategies (e.g., Marquering and Verbeek, 2004; Han, 2006).

with the benchmark RW strategy, an investor who pays transaction costs lower than $\tau^{be}$ will prefer the dynamic strategy. Since $\tau^{be}$ is a proportional cost paid every time the portfolio is rebalanced, we report $\tau^{be}$ in *bps* per week.

## 6.4 Empirical evidence

We now assess the economic value of exchange rate predictability by analyzing the out-of-sample performance of dynamically rebalanced portfolios. We first convert daily data into weekly observations by sampling on every Wednesday as in Burnside et al. (2007) and then generate one-week ahead forecasts using the following predictive regression

$$\Delta s_t = \alpha + \beta\, CRP_{t-1} + \varepsilon_t. \tag{18}$$

We estimate this regression every week using a 1-year rolling procedure that only conditions upon information available at time $t$. The rolling least-squares estimates $\widehat{\alpha}_t$, $\widehat{\beta}_t$, and $\widehat{\sigma}_{\varepsilon,t}$ of this regression are then used to generate 1-week ahead forecasts for our asset allocation exercise, i.e., $\mathbb{E}_t[r_{t+1}] = \widehat{\alpha}_t + \widehat{\beta}_t\, CRP_t$ and $\mathbb{V}_t[r_{t+1}] = \widehat{\sigma}_{\varepsilon,t}$. As a robustness exercise, we also sample data on every Tuesday and Thursday.[23]

We focus on the Sharpe ratio ($\mathcal{SR}$), the Sortino ratio ($\mathcal{SO}$), the Fleming et al. (2001) performance fee ($\mathcal{F}$), the Goetzmann et al. (2007) premium return measure ($\mathcal{P}$), and the break even transaction cost $\tau^{be}$. Following Della Corte et al. (2009), we focus on the maximum expected return strategy, as this is the strategy most often used in active currency management. We consider a target volatility equal to $\sigma_p^\star = 10\%$ per annum and a degree of RRA set to $\delta = 6$. Different values of $\sigma_p$ and $\delta$ have qualitatively little impact on the asset allocation results.

TABLE 9 ABOUT HERE

Table 9 reports the out-of-sample portfolio performance and shows that there is high economic value associated with using CRP to forecast currency returns. Panel B reports the economic performance of our selected models. The left side panel employs no restrictions on the slope coefficient $\beta$ whereas the right side panel impose a positive sign restriction on the slope coefficient $\beta$ in the spirit of Campbell

[23] We proxy the yield on the short-term bond in USD with the 1-week eurodeposit rate, which we average on a 5-day rolling window to mitigate day-of-the-week effects on interest rates.

and Thompson (2008). The random walk model has a nearly-zero out-of-sample $\mathcal{SR}$ and is clearly outperformed by our theoretically-motivated predictor. In particular, the unrestricted debt-weighted CRP specification displays an annualized $\mathcal{SR} = 0.31$ and a US investor would be willing to pay a performance fee ($\mathcal{F}$) higher than $300$ *bps* per annum for switching from the benchmark RW model to this alternative specification. The performance improves when imposing economic constraints akin to Campbell and Thompson (2008) since we uncover an annualized $\mathcal{SR} = 0.42$ and a performance fee ($\mathcal{F}$) of about $430$ *bps* per annum. The premium return $\mathcal{P}$ also leads to the same conclusion suggesting that quadratic utility characterizing the Fleming et al. (2001) performance fee is not affecting our results. The empirical evidence, moreover, remain qualitatively similar when CRP uses an alternative weighting scheme based on each country's GDP.

If transaction costs are sufficiently high, the period-by-period fluctuations in the dynamic weights of an optimal strategy will render the strategy too costly to implement relative to the static random walk model. We address this concern by computing the break-even transaction cost $\tau^{be}$, expressed in weekly basis points. For the unconstrained CRP, we find a positive and high $\tau^{be}$ meaning that a US investor will switch back to the RW model if she is subject to a proportional transaction cost of about $28$ *bps* per week. The $\tau^{be}$ increases its value when using economic restrictions on the slope coefficient as $\tau^{be} = 49.3$ *bps* per week. By and large, these value remain reasonably high and unlikely to be hit by a professional FX trader. In recent years, the typical transaction cost a large investor pays to trade the USD/EUR exchange rate FX market is few pips and this would roughly corresponds to very small proportional transaction cost.

In Panel A and Panel C, we repeat our asset allocation exercise by sampling data on every Tuesday and Thursday, respectively, thus accounting for intra-week seasonal patterns (e.g. McFarland et al., 1982; Bessembinder, 1994). We find slightly weaker results in Panel A but much stronger results in Panel C. For example, the restricted debt-weighted CRP strategy rebalanced on every Tuesday (Thursday) produces an annualized $\mathcal{SR} = 0.40$ ($\mathcal{SR} = 0.72$), which is lower (higher) than the corresponding value recorded in Panel B. Overall, we find that CRP generates tangible out-of-sample economic gains to an investor that uses exchange rate forecasts within active portfolio strategy.

## 7 Conclusion

This paper uncovers, both theoretically and empirically, a novel source of currency risk premium, which we label the *credit-implied risk premium* (CRP). CRP reflects investors' risk-neutral expectations about currency movements conditional on a severe but rare event, such as a sovereign default. We exploit dual-currency sovereign CDS to derive a market-based measure of the expected currency depreciation conditional on sovereign default, using daily data over the period 2010–2019.

We find that an aggregate measure of CRP for the Eurozone positively predicts future USD/EUR returns at various horizons, even after controlling for the interest rate differential, the quanto-implied risk premium of Kremens and Martin (2019), FX liquidity and volatility, option-based risk measures, and traditional currency factors. The predictability holds in- and out-of-sample. Investors are thus compensated for bearing the risk of a currency depreciation in the case of a sovereign default. Furthermore, CRP generates tangible economic gains to an investor using dynamic forecasts in active portfolio management. We obtain strong economic evidence against the random walk benchmark using a sample of weekly non-overlapping observations, and our results are robust to reasonably high transaction costs. Overall, CRP appears to be a critical driver of exchange rate returns, and we provide evidence that investors can benefit from this new source of information.

## References

Asness, C. S., Moskowitz, T. J., and Pedersen, L. H. (2013). Value and momentum everywhere. *Journal of Finance*, 68:929–985.

Augustin, P. (2018). The term structure of cds spreads and sovereign credit risk. *Journal of Monetary Economics*, 96:53–76.

Augustin, P., Chernov, M., and Song, D. (2019). Sovereign credit risk and exchange rates: Evidence from CDS quanto spreads. *Journal of Financial Economics, forthcoming*.

Bank for International Settlements (2016). *Triennial Central Bank Survey of Foreign Exchange and Derivatives Market in 2016*. Bank for International Settlements Press, Basel.

Barclays (2011). European credit alpha. *Credit Research, Barclays Capital*.

Bénétrix, A. S., Lane, P. R., and Shambaugh, J. C. (2015). International currency exposures, valuation effects and the global financial crisis. *Journal of International Economics*, 96:98–109.

Bessembinder, H. (1994). Bid-ask spreads in the interbank foreign exchange markets. *Journal of Financial Economics*, 35:317–348.

Britten-Jones, M. and Neuberger, A. (2000). Option prices, implied price processes, and stochastic volatility. *Journal of Finance*, 55:839–866.

Buraschi, A., Menguturk, M., and Sener, E. (2015). The geography of funding markets and limits to arbitrage. *Review of Financial Studies*, 28:1103–1152.

Burnside, C., Eichenbaum, M., and Rebelo, S. (2007). The returns to currency speculation in emerging markets. *American Economic Review*, 97:333–338.

Campbell, J. Y. and Thompson, S. B. (2008). Predicting excess stock returns out of sample: Can anything beat the historical average? *Review of Financial Studies*, 21:1509–1531.

Chernov, M., Graveline, J., and Zviadadze, I. (2018). Crash risk in currency returns. *Journal of Financial and Quantitative Analysis*, 53(1):137–170.

Choudhry, M. and Ali, A. (2010). *Structured Credit Products: Credit Derivatives and Synthetic Securitisation, 2nd Edition*. Wiley.

Cieslak, A. and Schrimpf, A. (2019). Non-monetary news in central bank communication. *Journal of International Economics*, 118:293–315.

Clark, T. E. and West, K. D. (2007). Approximately normal tests for equal predictive accuracy in nested models. *Journal of Econometrics*, 138:291–311.

Colacito, R., Croce, M. M., Gavazzoni, F., and Ready, R. (2018). Currency risk factors in a recursive multi-country economy. *Journal of Finance*, forthcoming.

Corradin, S. and Rodriguez-Moreno, M. (2014). Limits to arbitrage: Empirical evidence from Euro area sovereign bond markets. *Working Paper, European Central Bank*.

De Santis, R. A. (2015). A measure of redenomination risk. *ECB Working Paper Series*, No 1785.

Della Corte, P., Ramadorai, T., and Sarno, L. (2016a). Volatility risk premia and exchange rate predictability. *Journal of Financial Economics*, 120:21–40.

Della Corte, P., Riddiough, S. J., and Sarno, L. (2016b). Currency risk premia and global imbalances. *Review of Financial Studies*, 29:2161–2193.

Della Corte, P., Sarno, L., Schmeling, and Wagner, C. (2018). Exchange rates and sovereign risk. *Working Papers, Imperial College London*.

Della Corte, P., Sarno, L., and Tsiakas, I. (2009). An economic evaluation of empirical exchange rate models. *Review of Financial Studies*, 22:3491–3530.

Diebold, F. X. and Mariano, R. S. (1995). Comparing predictive accuracy. *Journal of Business & Economic Statistics*, 13:253–263.

Du, W. and Schreger, J. (2016). Local currency sovereign risk. *Journal of Finance*, 71:1027–1070.

Ehlers, P. and Schoenbucher, P. (2004). The influence of FX risk on credit spreads. *Working Paper, ETH Zürich*.

Engel, C. and West, K. (2005). Exchange rates and fundamentals. *Journal of Political Economy*, 113:485–517.

Fama, E. F. (1984). Forward and spot exchange rates. *Journal of Monetary Economics*, 14:319–338.

Farhi, E., Fraiberger, S. P., Gabaix, X., Ranciere, R., and Verdelhan, A. (2009). Crash risk in currency markets. Technical report, National Bureau of Economic Research.

Fleming, J., Kirby, C., and Ostdiek, B. (2001). The economic value of volatility timing. *Journal of Finance*, 56:329–352.

Gabaix, X. and Maggiori, M. (2015). International liquidity and exchange rate dynamics. *Quarterly Journal of Economics*, 130:1369–1420.

Garman, M. B. and Kohlhagen, S. W. (1983a). Foreign currency option values. *Journal of International Money and Finance*, 2:231–237.

Garman, M. B. and Kohlhagen, S. W. (1983b). Foreign currency option values. *Journal of International Money and Finance*, 2:231–237.

Goetzmann, W., Ingersoll, J., Spiegel, M., and Welch, I. (2007). Portfolio performance manipulation and manipulation-proof performance measures. *Review of Financial Studies*, 20:1503–1546.

Han, Y. (2006). Asset allocation with a high dimensional latent factor stochastic volatility model. *Review of Financial Studies*, 19:237–271.

Herz, B. and Tong, H. (2008). Debt and currency crises: Complements or substitutes? *Review of International Economics*, 16:955–970.

Hlawitschka, W. (1994). The empirical nature of taylor-series approximations to expected utility. *American Economic Review*, 84:713–719.

Hull, J. C. and White, A. (2000). Valuing credit default swaps: No counterparty default risk. *Journal of Derivatives*, 8(1):29–40.

ISDA (2003). 2003 ISDA credit derivative definition. *International Swaps and Derivative Association*.

Jarrow, R. A. and Turnbull, S. M. (1995). Pricing derivatives on financial securities subject to credit risk. *Journal of Finance*, 50:53–85.

Jiang, G. J. and Tian, Y. S. (2005). The model-free implied volatility and its information content. *Review of Financial Studies*, 18:1305–1342.

J.P. Morgan (2010). Trading credit in different currencies via quanto CDS. *Europe Credit Derivatives Research, J.P. Morgan*.

Karnaukh, N., Ranaldo, A., and Söderlind, P. (2015). Understanding FX liquidity. *Review of Financial Studies*, 28:3073–3108.

Kilian, L. (1999). Exchange rates and monetary fundamentals: What do we learn from long-horizon regressions? *Journal of Applied Econometrics*, 14:491–510.

Kremens, L. and Martin, I. (2019). The quanto theory of exchange rates. *American Economic Review*, 109:810–843.

Lando, D. and Bang Nielsen, A. (2018). Quanto CDS spreads. *Working Paper, Copenhagen Business School*.

Lane, P. R. and Milesi-Ferretti, G. M. (2007). The external wealth of nations mark II: Revised and extended estimates of foreign assets and liabilities, 1970–2004. *Journal of International Economics*, 73:223–250.

Londono, J. M. and Zhou, H. (2017). Variance risk premiums and the forward premium puzzle. *Journal of Financial Economics*, 124:415–440.

Longstaff, F. A., Pan, J., Pedersen, L., and Singleton, K. J. (2011). How sovereign is sovereign credit risk? *American Economic Journal: Macroeconomics*, 3:75–103.

Lustig, H., Roussanov, N., and Verdelhan, A. (2011). Common risk factors in currency markets. *Review of Financial Studies*, 24:3731–3777.

Lustig, H., Roussanov, N., and Verdelhan, A. (2014). Countercyclical currency risk premia. *Journal of Financial Economics*, 111(3):527–553.

Mancini, L., Ranaldo, A., and Wrampelmeyer, J. (2013). Liquidity in the foreign exchange market: Measurement, commonality, and risk premiums. *Journal of Finance*, 68:1805–1841.

Mano, R. (2013). Exchange rates upon sovereign default. *Working Paper, International Monetary Fund*.

Mark, N. C. (1995). Exchange rates and fundamentals: Evidence on long-horizon predictability. *American Economic Review*, 85:201–218.

Marquering, W. and Verbeek, M. (2004). The economic value of predicting stock index returns and volatility. *Journal of Financial and Quantitative Analysis*, 39:407–429.

Martin, I. (2017). What is the expected return on the market? *The Quarterly Journal of Economics*, 132(1):367–433.

McCracken, M. W. (2007). Asymptotics for out of sample tests of granger causality. *Journal of Econometrics*, 140:719–752.

McFarland, J. W., Pettit, R. R., and Sung, S. K. (1982). The distribution of foreign exchange price changes: Trading day effects and risk measurement. *Journal of Finance*, 37:693–715.

Meese, R. A. and Rogoff, K. (1983). Empirical exchange rate models of the seventies: Do they fit out of sample? *Journal of International Economics*, 14:3–24.

Menkhoff, L., Sarno, L., Schmeling, M., and Schrimpf, A. (2012). Carry trades and global foreign exchange volatility. *Journal of Finance*, 67:681–718.

Na, S., Schmitt-Grohé, S., Uribe, M., and Yue, V. (2018). The twin Ds: Optimal default and devaluation. *American Economic Review*, 108:1773–1819.

O'Kane, D. and Turnbull, S. M. (2003). Valuation of credit default swaps. *Fixed Income Quantitative Credit Research, Lehman Brothers*.

Pan, J. and Singleton, K. J. (2008). Default and recovery implicit in the term structure of sovereign CDS spreads. *Journal of Finance*, 63:2345–2384.

Reinhart, C. M. (2002). Default, currency crises and sovereign credit ratings. *World Bank Economic Review*, 16:151–170.

Veronesi, P. (2010). *Fixed Income Securities: Valuation, Risk and Risk Management*. J. Wiley & Sons.

Welch, I. and Goyal, A. (2008). A comprehensive look at the empirical performance of equity premium prediction. *Review of Financial Studies*, 21:1455–1508.

West, K. D. (1996). Asymptotic inference about predictive ability. *Econometrica*, 64:1067–1084.

West, K. D., Edison, H. J., and Cho, D. (1993). A utility-based comparison of some models of exchange rate volatility. *Journal of International Economics*, 35:23–45.

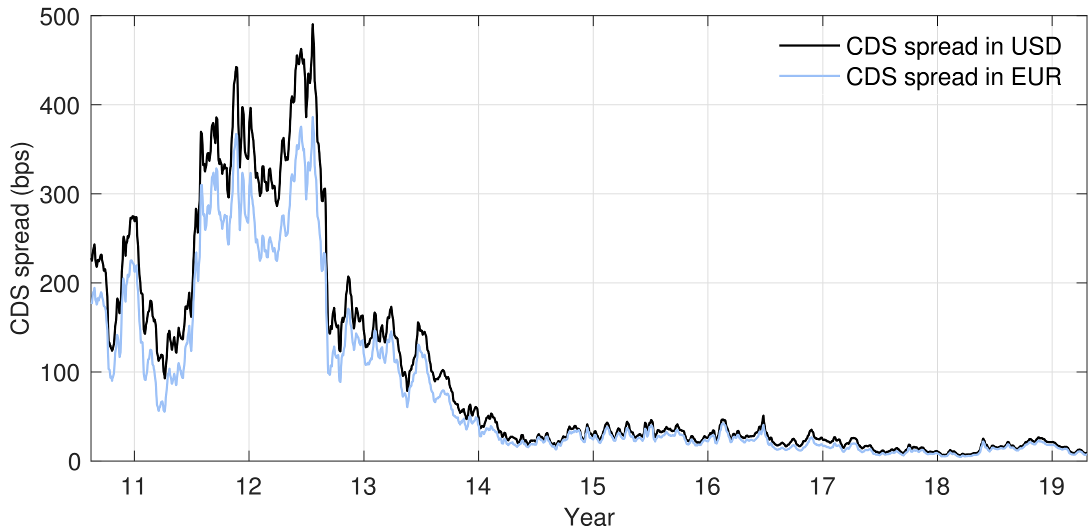


**Figure 1: Spain's CDS spread in EUR versus in USD**
This figure shows Spain's sovereign CDS spreads denominated in EUR and in USD. The sample consists of observations between August 20, 2010, and April 26, 2019.

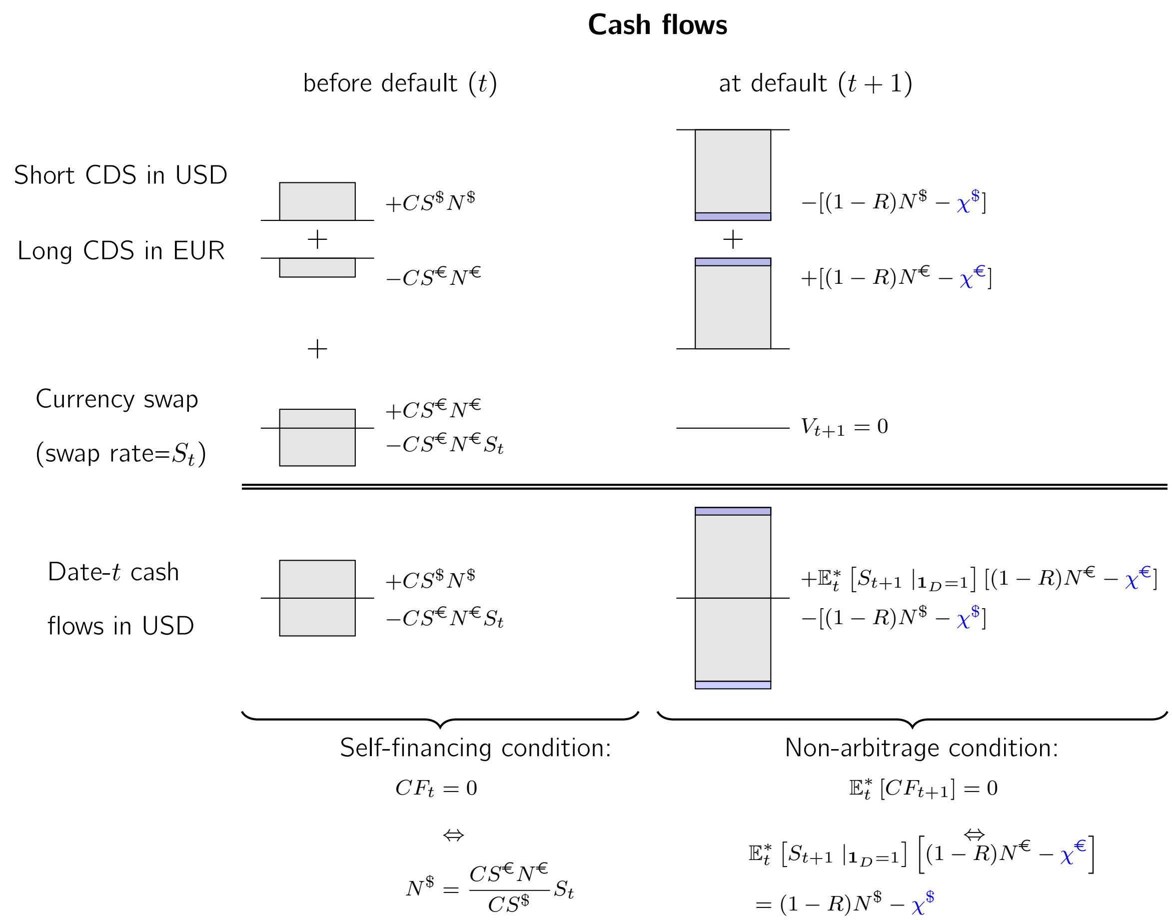


**Figure 2: Cash flows of a long-short CDS strategy**

This figure shows the cash flows of a strategy that involves a long position in a sovereign CDS denominated in EUR and a short position in an identical CDS but denominated in USD. The strategy has two periods: the time of inception $t$ and the potential default time $t+1$. The cash flows in black color are those of the benchmark case. We highlight in blue the additional cash flows related to the accrued premiums. The long-short strategy is discussed in Section 3.2.

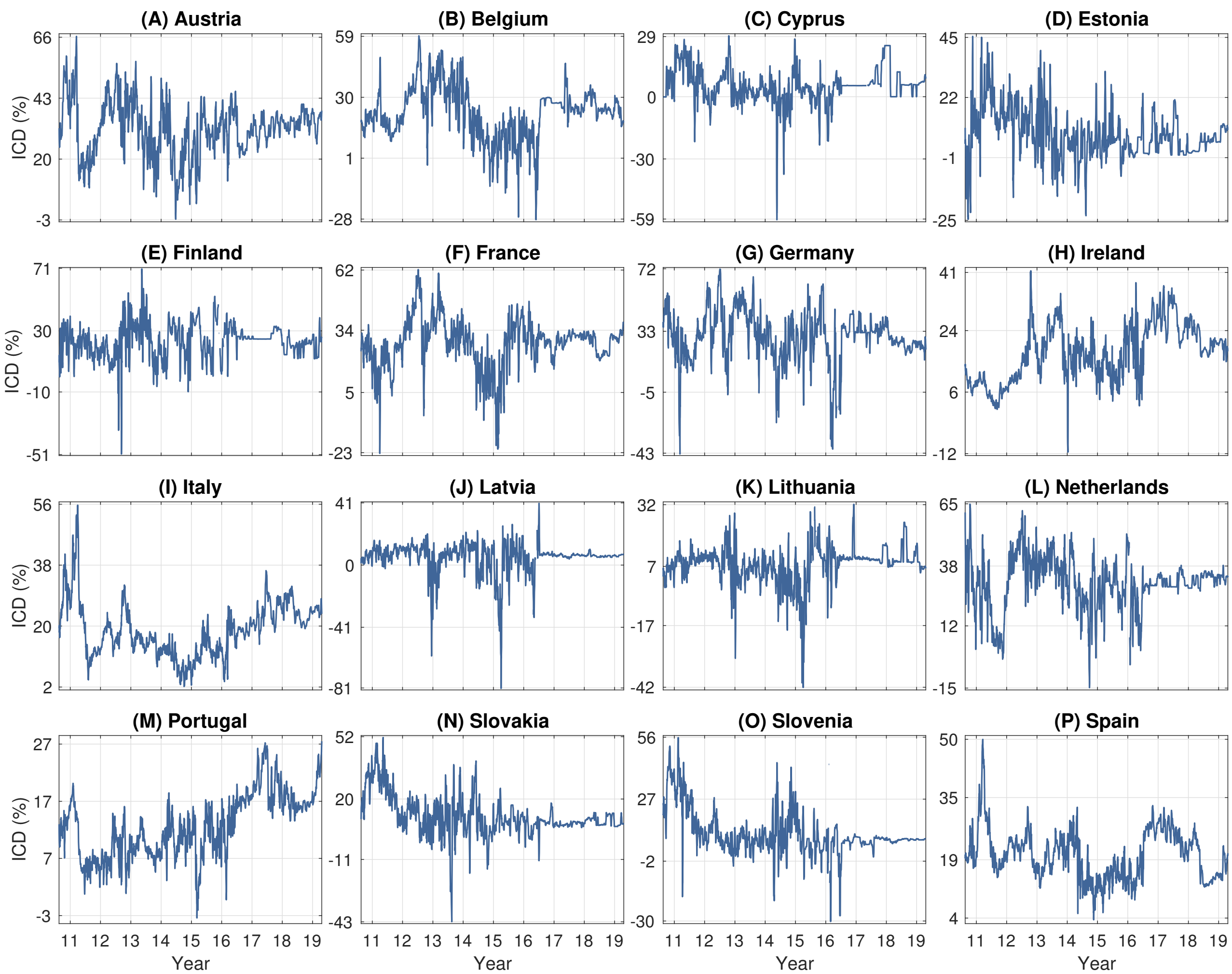


**Figure 3: Implied currency depreciation for individual Eurozone countries**
This figure shows the implied currency depreciation (ICD) of the EUR against the USD for Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Ireland, Italy, Latvia, Lithuania, the Netherlands, Portugal, Slovenia, Slovakia, and Spain. The computation of ICD is detailed in Section 3.2. The sample consists of observations between August 20, 2010, and April 26, 2019.

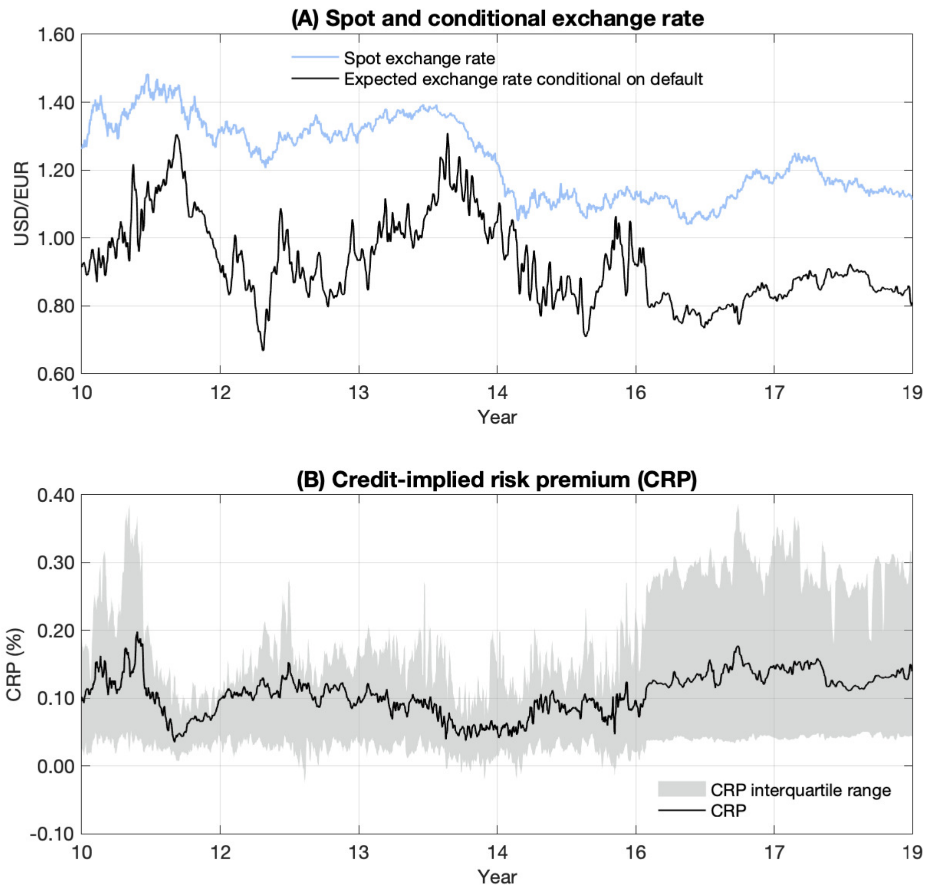


**Figure 4: Time series of the credit-implied risk premium**
This figure illustrates the dynamics of the credit-implied risk premium (CRP) for the Eurozone. Panel A shows the expected USD/EUR exchange rate upon default at a 1-year horizon, which is compared to the spot exchange rate. Panel B shows the time series of CRP. Section 3.3 details the construction of CRP. The countries included in the computation are Belgium, France, Germany, Italy, Ireland, Portugal, Spain, Netherlands, Lithuania, Austria, Slovenia, Slovakia, Cyprus, Estonia, Finland and Latvia. The sample consists of observations between August 20, 2010, and April 26, 2019.

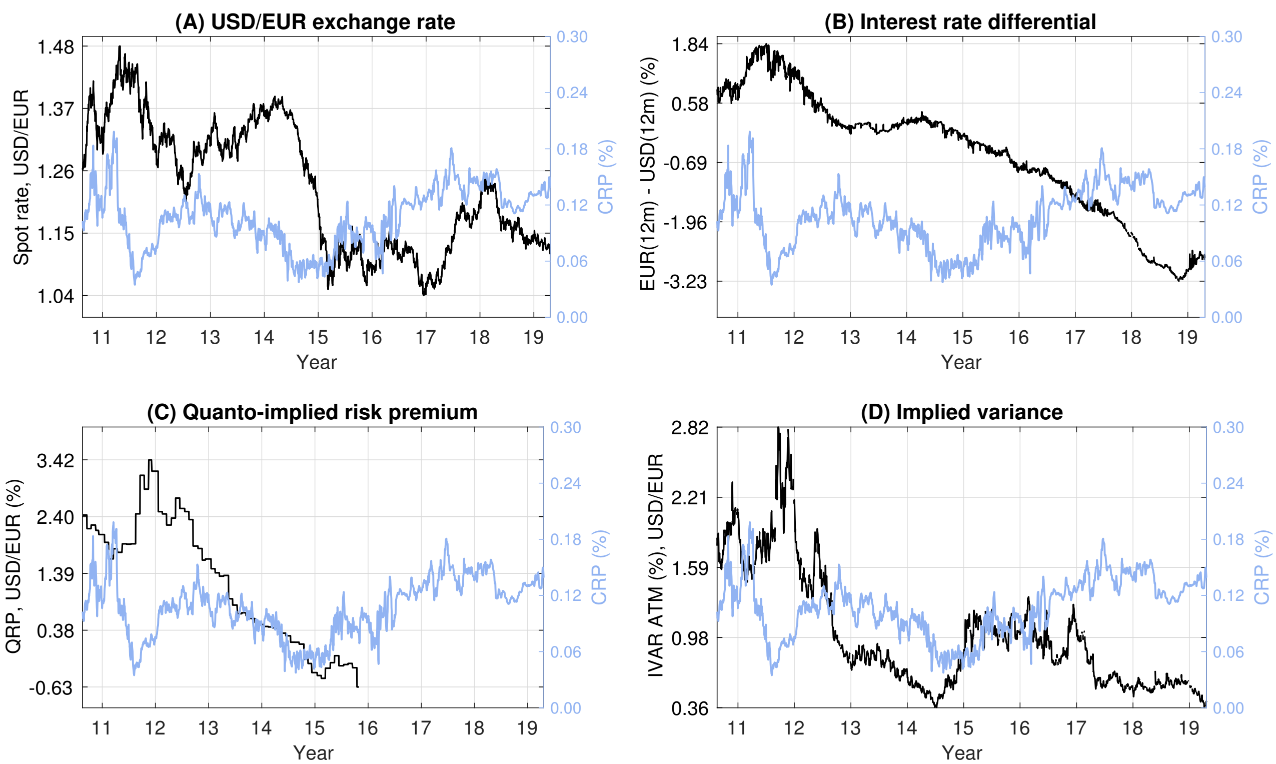


**Figure 5: Time series of the main variables**

This figure illustrates the time series of the main variables used in the benchmark predictability analysis, which we compare with the credit-implied risk premium (CRP) for the Eurozone. We report the spot USD/EUR exchange rate in Panel A, the 1-year interest rate differential in Panel B, the Kremens and Martin (2019)'s measure of quanto-implied risk premium (QRP) in Panel C, and the implied variance on the USD/EUR based on ATM options in Panel D. Section 4.1 describes these variables, while Section 3.3 describes the computation of CRP. The sample consists of observations between August 20, 2010, and April 26, 2019.

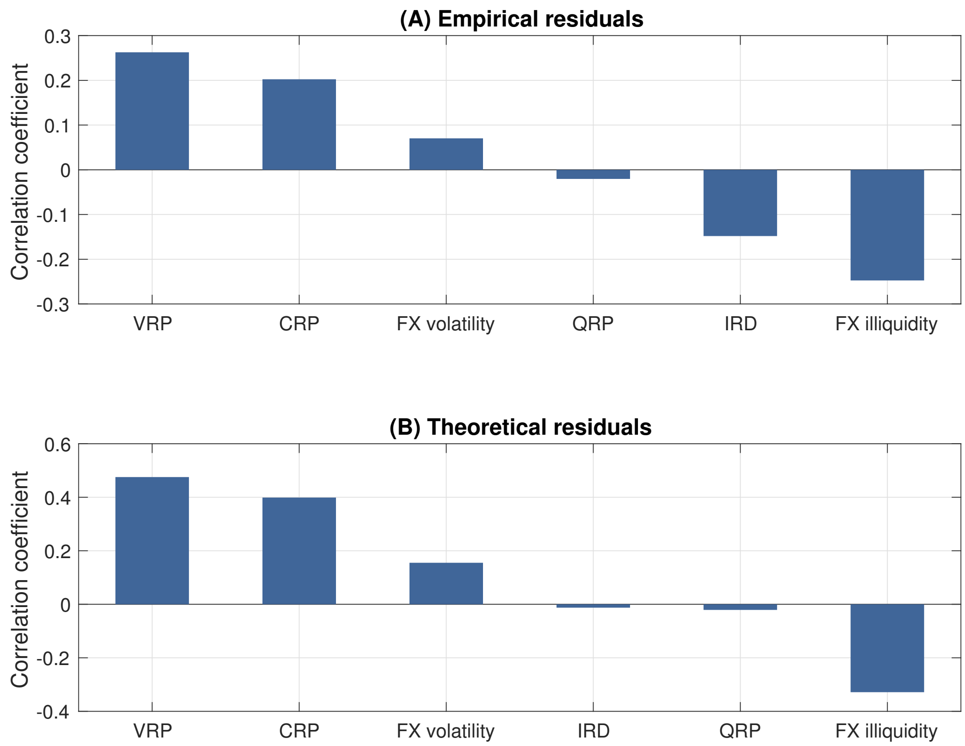


**Figure 6: Analysis of the residual covariance term**
This figure reports the correlations between the residual covariance term and variables of interest. Panel A reports the correlations between each variable and the residuals $\varepsilon_{t+k}$ of the predictive regression (14) at the 3-month horizon, while Panel B uses the theoretical counterparts (i.e., with $a_k = 0$ and $b_k = c_k = d_k = 1$). The figure reports results for the volatility risk premium (VRP) on the USD/EUR, the credit-implied risk premium (CRP), the quanto-implied risk premium (QRP) of Kremens and Martin (2019), the realized volatility on the USD/EUR, the interest rate differential (IRD), and the level of illiquidity in the USD/EUR. Section 4.1 describes these variables, while Section 3.3 describes the computation of CRP. The sample consists of observations between August 20, 2010, and April 26, 2019.

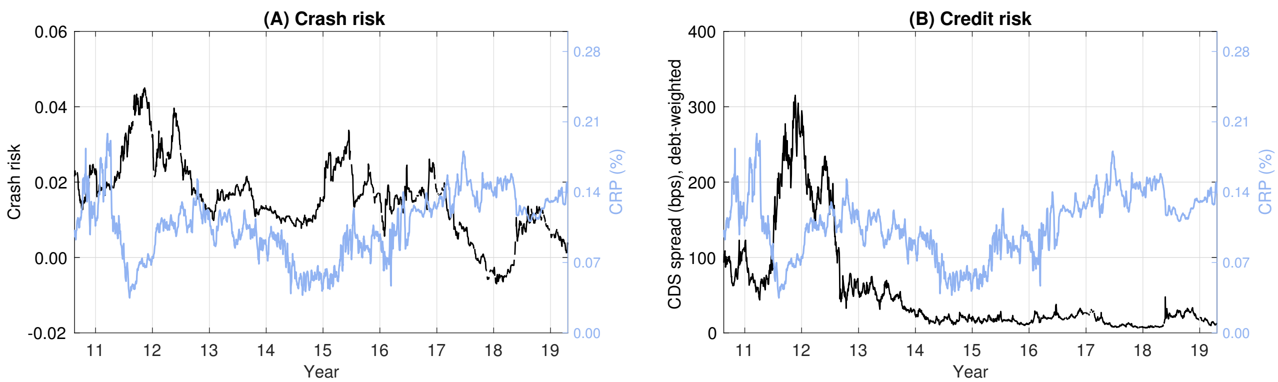


**Figure 7: Crash risk, credit risk, and CRP**
This figure illustrates the time series of crash risk (Panel A) and credit risk (Panel B), which are compared to the credit-implied risk premium (CRP). Crash risk is computed as the implied volatility of 25-delta put options minus the equivalent for call options on the USD/EUR exchange rate. Credit risk is computed as the debt-weighted average of USD-denominated (1-year) CDS spreads of European countries. Section 4.4.5 describes these variables, while Section 3.3 describes the computation of CRP. The sample consists of observations between August 20, 2010, and April 26, 2019.

**Table 1 : Descriptive statistics**

This table reports the descriptive statistics for the credit-implied risk premium (CRP) and its components. Panel A presents country-level statistics of the implied currency depreciation ($ICD_i$), the risk-neutral probability of default ($\mathbb{Q}_i$), the government's debt weight ($\omega_i$), and the credit-implied risk premium ($CRP_i$). Panel B presents the statistics for the unconditional expected currency depreciation (ECD), given by the 1-year interest rate differential in USD and in EUR, as well as for the ICD and CRP at the Eurozone level. The computation of ICD, using 1-year CDS quotes, CRP, and the risk-neutral probability of default are detailed in Sections 3.2, 3.3, and Appendix C respectively. The sample consists of observations between August 20, 2010, and April 26, 2019.

*Panel A: Country measures*

| | $ICD_i$ (%) | | | | $\mathbb{Q}_i$ (%) | $\omega_i$ (%) | $CRP_i$ (%) |
|---|---|---|---|---|---|---|---|
| | Mean | Std | P5 | P95 | Mean | Mean | Mean |
| Austria | 31.30 | 11.45 | 10.44 | 50.47 | 0.316 | 3.132 | 0.060 |
| Belgium | 22.60 | 13.87 | -1.79 | 45.86 | 0.614 | 4.675 | 0.084 |
| Cyprus | 5.25 | 9.60 | -8.30 | 24.33 | 7.783 | 0.139 | 0.264 |
| Estonia | 8.09 | 11.36 | -7.43 | 28.38 | 0.464 | 1.131 | 0.024 |
| Finland | 23.09 | 13.81 | 1.29 | 44.05 | 0.168 | 0.012 | 0.024 |
| France | 27.58 | 13.42 | 3.14 | 47.32 | 0.352 | 20.91 | 0.059 |
| Germany | 29.92 | 17.96 | -2.23 | 57.77 | 0.141 | 26.81 | 0.025 |
| Ireland | 17.54 | 9.10 | 3.87 | 31.83 | 2.855 | 1.855 | 0.308 |
| Italy | 18.80 | 8.41 | 6.66 | 31.85 | 1.691 | 23.78 | 0.194 |
| Latvia | 4.95 | 12.17 | -15.78 | 18.51 | 0.941 | 0.107 | 0.031 |
| Lithuania | 7.68 | 9.30 | -6.82 | 20.98 | 0.916 | 0.130 | 0.045 |
| Netherlands | 30.50 | 13.04 | 4.58 | 50.90 | 0.218 | 4.868 | 0.040 |
| Portugal | 12.59 | 6.10 | 3.35 | 22.73 | 4.703 | 2.223 | 0.367 |
| Slovakia | 11.33 | 11.78 | -3.66 | 34.61 | 0.578 | 0.358 | 0.041 |
| Slovenia | 10.97 | 11.35 | -1.79 | 36.33 | 1.225 | 0.179 | 0.084 |
| Spain | 20.21 | 6.99 | 10.04 | 31.15 | 1.629 | 8.344 | 0.201 |

*Panel B: Eurozone measures*

| | N | Mean | Median | Std | P5 | P95 |
|---|---|---|---|---|---|---|
| $ECD$ (%) | 2266 | -0.505 | -0.081 | 1.325 | -2.933 | 1.471 |
| $ICD$ (%) | 2266 | 24.69 | 25.72 | 7.649 | 9.833 | 36.08 |
| $CRP$ (%) | 2266 | 0.106 | 0.110 | 0.032 | 0.050 | 0.153 |

**Table 2 : Exchange rate return predictability with debt-weighted CRP**
This table presents results on the predictability of the USD/EUR exchange rate return using the credit-implied risk premium (CRP) for the Eurozone. The dependent variable is the annualized average daily log USD/EUR return over a given forecast horizon. CRP is constructed by weighting country-level CRP components by their outstanding debt level. Panel A controls for the 12-month USD/EUR interest rate differential. Panel B controls for the USD/EUR illiquidity, realized volatility, and volatility risk premium. Panel C controls for the global currency factors. Panel D includes all controls. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. The sample consists of daily observations between August 20, 2010, and April 26, 2019.

*Panel A: Benchmark model*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.229*** | 2.407*** | 2.155*** | 1.852*** | 1.009** |
| | (0.009) | (0.004) | (0.001) | (0.006) | (0.038) |
| $R^2$ (%) | 1.08 | 6.24 | 16.26 | 23.35 | 20.26 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel B: Controlling for FX liquidity and volatility*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 1.966** | 2.093** | 1.674*** | 1.355*** | 0.518* |
| | (0.032) | (0.010) | (0.002) | (0.001) | (0.083) |
| $R^2$ (%) | 2.11 | 8.99 | 22.31 | 35.96 | 47.10 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel C: Controlling for global factors*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.267*** | 2.432*** | 2.168*** | 1.858*** | 1.017** |
| | (0.008) | (0.003) | (0.001) | (0.005) | (0.037) |
| $R^2$ (%) | 1.25 | 6.12 | 16.21 | 23.24 | 20.13 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel E: With all controls*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.006** | 2.113*** | 1.685*** | 1.359*** | 0.519* |
| | (0.030) | (0.010) | (0.002) | (0.001) | (0.083) |
| $R^2$ (%) | 2.30 | 8.90 | 22.28 | 35.86 | 47.09 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

**Table 3 : Analysis of the QRP residuals**
This table presents an analysis of the predictability of USD/EUR exchange rate returns with the quanto-implied risk premium (QRP) of Kremens and Martin (2019). Panel A reports the results from the regression $\frac{1}{k}\Delta s_{t+k} = \alpha_k + \beta_k QRP_t + \gamma_k IRD_t + \eta_{t+k}$. Panels B and C report results from the regression $\eta_{t+k} = \phi_k + \omega_k CRP_t + \delta_k X_t + \psi_{t+k}$, without ($\delta_k = 0$) and with ($\delta_t \neq 0$) the full set of controls, respectively. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls. The sample consists of observations between August 20, 2010, and November 9, 2015.

*Panel A: Predictability of USD/EUR returns*

| | 1 month | 6 months | 12 months | 24 months |
|---|---|---|---|---|
| QRP$_t$ | 9.697** | 10.40*** | 11.58*** | 6.645*** |
| | (0.038) | (0.001) | (0.000) | (0.000) |
| R$^2$ (%) | 4.62 | 30.61 | 67.19 | 78.25 |
| N | 1297 | 1192 | 1064 | 814 |

*Panel B: Predictability of residuals $\eta_{t+k}$, without controls*

| | 1 month | 6 months | 12 months | 24 months |
|---|---|---|---|---|
| CRP$_t$ | 2.066* | 1.597*** | 0.827*** | 0.230*** |
| | (0.053) | (0.004) | (0.000) | (0.001) |
| R$^2$ (%) | 4.51 | 21.53 | 20.50 | 10.16 |
| N | 1249 | 1149 | 1028 | 785 |

*Panel C: Predictability of residuals $\eta_{t+k}$, with controls*

| | 1 month | 6 months | 12 months | 24 months |
|---|---|---|---|---|
| CRP$_t$ | 3.016* | 2.221*** | 0.733*** | 0.185*** |
| | (0.051) | (0.000) | (0.000) | (0.000) |
| R$^2$ (%) | 5.83 | 32.29 | 25.72 | 37.68 |
| N | 1249 | 1149 | 1028 | 785 |

**Table 4 : Exchange rate return predictability: Controlling for QRP**

This table presents results on the predictability of the USD/EUR exchange rate return using the credit-implied risk premium (CRP) for the Eurozone and controlling for different measures of the quanto-implied risk premium (QRP). The dependent variable is the annualized average daily log USD/EUR exchange rate return over a given forecast horizon. CRP is constructed by weighting country-level CRP components by their outstanding debt level. Panel A uses the QRP measure of Kremens and Martin (2019) for the USD/EUR whereas Panels B-D employ a synthetic versions of QRP based on the implied variance of the USD/EUR rate return. Panel B employs 1-year over-the-counter currency options and implements the model-free approach of Britten-Jones and Neuberger (2000). Panel C uses the same set of data but employs the simple variance method of Martin (2017). Panel D simply uses the squared implied volatility from at-the-money options. In all panels, we control for the 12-month USD/EUR interest rate differential. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. In Panel A, daily observations for the QRP are obtained by forward filling monthly observations and the sample runs between August 20, 2010, and October 21, 2015. In Panels B-D, the sample consists of daily observations between August 20, 2010, and April 26, 2019.

*Panel A: Controlling for QRP*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.065* | 2.569** | 2.402*** | 1.978*** | 0.917*** |
| | (0.061) | (0.014) | (0.000) | (0.000) | (0.000) |
| $R^2$ (%) | 1.82 | 10.04 | 30.59 | 49.43 | 74.66 |
| N | 1313 | 1297 | 1255 | 1192 | 1066 |

*Panel B: Controlling for $QRP^S_{BJN}$*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.173** | 2.351*** | 2.046*** | 1.689*** | 0.772** |
| | (0.013) | (0.004) | (0.001) | (0.002) | (0.027) |
| $R^2$ (%) | 1.06 | 6.33 | 17.73 | 28.84 | 36.23 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel C: Controlling for $QRP^S_M$*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.148** | 2.332*** | 2.024*** | 1.660*** | 0.742** |
| | (0.014) | (0.005) | (0.001) | (0.002) | (0.035) |
| $R^2$ (%) | 1.08 | 6.38 | 17.85 | 29.18 | 36.26 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel D: Controlling for $QRP^S_{ATM}$*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.199** | 2.382*** | 2.063*** | 1.672*** | 0.751** |
| | (0.012) | (0.003) | (0.001) | (0.002) | (0.031) |
| $R^2$ (%) | 1.17 | 6.77 | 18.91 | 29.68 | 37.28 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

**Table 5 : Exchange rate return predictability with GDP-weighted CRP**

This table presents results on the predictability of the USD/EUR exchange rate return using the credit-implied risk premium (CRP) for the Eurozone. The dependent variable is the annualized average daily log USD/EUR return over a given forecast horizon. CRP is constructed by weighting country-level CRP components by their gross domestic product (GDP). Panel A controls for the 12-month USD/EUR interest rate differential. Panel B controls for the USD/EUR illiquidity, realized volatility, and volatility risk premium. Panel C controls for the global currency factors. Panel D includes all controls. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. The sample consists of daily observations between August 20, 2010, and April 26, 2019.

*Panel A: Benchmark model*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.562*** | 2.717*** | 2.484*** | 2.158*** | 1.176** |
| | (0.008) | (0.003) | (0.001) | (0.004) | (0.025) |
| $R^2$ (%) | 1.12 | 6.21 | 16.82 | 24.56 | 20.86 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel B: Controlling for FX liquidity and volatility*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.278** | 2.364*** | 1.954*** | 1.623*** | 0.641* |
| | (0.027) | (0.009) | (0.001) | (0.001) | (0.060) |
| $R^2$ (%) | 2.14 | 8.95 | 22.78 | 37.06 | 47.60 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel C: Controlling for global factors*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.601*** | 2.744*** | 2.498*** | 2.166*** | 1.185** |
| | (0.007) | (0.003) | (0.001) | (0.003) | (0.024) |
| $R^2$ (%) | 1.29 | 6.09 | 16.78 | 24.46 | 20.73 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel E: With all controls*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 2.319** | 2.385*** | 1.966*** | 1.628*** | 0.643* |
| | (0.026) | (0.008) | (0.001) | (0.001) | (0.061) |
| $R^2$ (%) | 2.33 | 8.86 | 22.74 | 36.96 | 47.59 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

**Table 6 : Exchange rate return predictability: Counterfactual analysis**

This table presents results on the predictability of the USD/JPY exchange rate return using the credit-implied risk premium (CRP) for the Eurozone. The dependent variable is the annualized average daily log USD/JPY return over a given forecast horizon. CRP is constructed by weighting country-level CRP components by their outstanding debt level. Panel A controls for the 12-month USD/JPY interest rate differential. Panel B further controls for the USD/JPY illiquidity, realized volatility, volatility risk premium, and global currency factors. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. The sample consists of daily observations between August 20, 2010, and April 26, 2019.

*Panel A: Benchmark model*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | 0.137 | 0.530 | 0.605 | 0.544 | 0.121 |
| | (0.889) | (0.505) | (0.466) | (0.559) | (0.869) |
| $R^2$ (%) | 0.00 | 0.31 | 1.13 | 1.98 | 4.88 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel B: With all controls*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| $CRP_t$ | -0.826 | -0.344 | -0.384 | -0.342 | -0.338 |
| | (0.418) | (0.638) | (0.583) | (0.692) | (0.616) |
| $R^2$ (%) | 1.76 | 8.73 | 17.13 | 22.30 | 27.06 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

**Table 7 : Exchange rate return predictability: Non-overlapping observations**
This table presents results on the predictability of the USD/EUR exchange rate return using the credit-implied risk premium (CRP) for the Eurozone. The dependent variable is the annualized average daily log USD/EUR return over a given non-overlapping forecast horizon. CRP is constructed by weighting country-level CRP components by their outstanding debt level. Panel A samples the data every Tuesday, Panel B every Wednesday, and Panel C every Thursday. All specifications control for the USD/EUR interest rate differential. The regressions control for the 12-month EUR-USD interest rate differential. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 4.1 describes the computation of CRP, while Section 4.2 discusses the econometric specification. The sample consists of weekly observations between August 20, 2010, and April 26, 2019.

Panel A: Tuesdays

| | 1 week | 2 weeks | 3 weeks | 4 weeks |
|---|---|---|---|---|
| $CRP_t$ | 2.076** | 2.251** | 2.591*** | 2.759*** |
| | (0.036) | (0.029) | (0.001) | (0.000) |
| $R^2$ (%) | 0.62 | 1.57 | 3.82 | 6.79 |
| N | 452 | 226 | 150 | 113 |

Panel B: Wednesdays

| | 1 week | 2 weeks | 3 weeks | 4 weeks |
|---|---|---|---|---|
| $CRP_t$ | 2.231** | 1.925 | 2.669*** | 1.839** |
| | (0.063) | (0.114) | (0.004) | (0.017) |
| $R^2$ (%) | 0.72 | 0.80 | 3.31 | 1.30 |
| N | 451 | 225 | 150 | 112 |

Panel C: Thursdays

| | 1 week | 2 weeks | 3 weeks | 4 weeks |
|---|---|---|---|---|
| $CRP_t$ | 2.375** | 2.494** | 1.974** | 2.120 *** |
| | (0.016) | (0.011) | (0.044) | (0.004) |
| $R^2$ (%) | 0.91 | 1.84 | 1.91 | 2.96 |
| N | 451 | 225 | 150 | 112 |

**Table 8 : Exchange rate return predictability: Statistical analysis**

This table reports the out-of-sample predictive ability of the credit-implied risk premium (CRP) against the random walk for the USD/EUR exchange rate return. The analysis is based on a one-year rolling window. Column I (II) reports results without (with) a sign restriction on the slope coefficient. The sign restriction sets the slope coefficient equal to zero when its estimate is negative. In Column I, the out-of-sample forecasts are generated without any economic constraints. In Column II, the slope coefficient of the predictive regression is subject to an economic sign restriction as in Campbell and Thompson (2008), i.e., we set the slope coefficient equal to zero when its estimate is negative. Panel A (Panel B) presents results based on the debt-weighted (GDP-weighted) CRP, which is described in Section 3.3. The $p$-values, computed based on 1,000 bootstrap replications, are reported in parentheses. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. The sample consists of observations between August 20, 2010, and April 26, 2019.

| | **I. Unconstrained predictability** | | | | | **II. Constrained predictability** | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | 1 week | 1 month | 3 months | 6 months | 12 months | 1 week | 1 month | 3 months | 6 months | 12 months |
| *Panel A: Debt-weighted CRP* | | | | | | | | | | |
| $R^2_{OOS}(\%)$ | -0.538** | 3.975*** | 4.136** | 6.545* | -1.483 | 0.414** | 5.535*** | 9.227*** | 11.972** | 2.900 |
| | (0.035) | (0.007) | (0.036) | (0.075) | (0.312) | (0.014) | (0.002) | (0.007) | (0.022) | (0.187) |
| $\Delta RMSE$ | -0.1865** | 0.589*** | 0.411** | 0.554* | -0.099 | 0.127** | 0.823*** | 0.930*** | 1.028** | 0.196 |
| | (0.035) | (0.007) | (0.036) | (0.070) | (0.323) | (0.014) | (0.004) | (0.009) | (0.016) | (0.169) |
| $MSE\text{-}F$ | -0.005** | 0.041*** | 0.043** | 0.070* | -0.015 | 0.004** | 0.059*** | 0.102** | 0.136** | 0.030 |
| | (0.035) | (0.007) | (0.036) | (0.075) | (0.312) | (0.034) | (0.005) | (0.027) | (0.013) | (0.128) |
| $CW$ | 1.755 | 6.422** | 6.200 | 7.573 | 5.634 | 2.582** | 7.115*** | 8.442** | 10.788** | 7.724 |
| | (0.125) | (0.018) | (0.109) | (0.115) | (0.244) | (0.034) | (0.005) | (0.027) | (0.013) | (0.128) |
| *Panel B: GDP-weighted CRP* | | | | | | | | | | |
| $R^2_{OOS}(\%)$ | -0.579** | 3.526*** | 2.540* | 5.892* | 1.290 | 0.424** | 5.086*** | 7.693** | 10.944** | 5.078 |
| | (0.044) | (0.008) | (0.055) | (0.081) | (0.241) | (0.014) | (0.004) | (0.015) | (0.028) | (0.127) |
| $\Delta RMSE$ | -0.178** | 0.522*** | 0.252* | 0.498* | 0.087 | 0.130** | 0.756*** | 0.772** | 0.937** | 0.345 |
| | (0.042) | (0.009) | (0.055) | (0.075) | (0.236) | (0.014) | (0.006) | (0.016) | (0.024) | (0.111) |
| $MSE\text{-}F$ | -0.006** | 0.037*** | 0.026* | 0.063* | 0.013 | 0.004** | 0.054*** | 0.083** | 0.123** | 0.053* |
| | (0.044) | (0.008) | (0.055) | (0.081) | (0.241) | (0.030) | (0.006) | (0.042) | (0.022) | (0.094) |
| $CW$ | 1.825 | 6.172** | 5.426 | 7.217 | 6.624 | 2.699** | 6.887*** | 7.706** | 10.433** | 8.676* |
| | (0.111) | (0.030) | (0.140) | (0.121) | (0.200) | (0.030) | (0.006) | (0.042) | (0.022) | (0.094) |

**Table 9 : Exchange rate return predictability: Economic evaluation**

This table presents the out-of-sample performance of the credit-implied risk premium (CRP) for the Eurozone against random walk (RW) for the USD/EUR exchange rate return. Each model employs non-overlapping weekly forecasts and builds a maximum expected return strategy subject to a target volatility of 10% per annum for a US investor who holds cash account in euros. For each strategy, we report the percentage mean, percentage volatility, Sharpe ratio ($\mathcal{SR}$), and Sortino ratio ($\mathcal{SO}$) in annual terms. $\mathcal{F}$ denotes the performance fee a risk-averse investor is willing to pay for switching from the benchmark RW strategy to the competing CRP strategy. $\mathcal{P}$ is the premium return generated by the CRP strategy relative to the RW one. $\mathcal{F}$ and $\mathcal{P}$ are computed for a degree of relative risk aversion equal to 6 and are expressed in annual basis points. $\tau^{be}$ is the break-even proportional transaction cost that cancels out the utility advantage of the CRP strategy relative to the RW, and is expressed in weekly basis points. The predictive regressions are re-estimated on every Tuesdays (Panel A), Wednesdays (Panel B), and Thursdays (Panel C) using a one-year rolling window. In column I, out-of-sample forecasts are generated without any economic constraints. In column II, the slope coefficient of the predictive regression is subject to an economic sign restriction as in Campbell and Thompson (2008), i.e., we set the slope coefficient equal to zero when its estimate is negative. Section 3.3 describes the computation of CRP. The sample consists of weekly observations between August 20, 2010, and April 26, 2019.

| | **I. Unconstrained predictability** | | | | | | | **II. Constrained predictability** | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| *Panel A: Tuesdays* | | | | | | | | | | | | | | |
| | $mean$ | $vol$ | $\mathcal{SR}$ | $\mathcal{SO}$ | $\mathcal{F}$ | $\mathcal{P}$ | $\tau^{be}$ | $mean$ | $vol$ | $\mathcal{SR}$ | $\mathcal{SO}$ | $\mathcal{F}$ | $\mathcal{P}$ | $\tau^{be}$ |
| RW | 1.30 | 10.05 | 0.13 | 0.23 | | | | 1.30 | 10.05 | 0.13 | 0.23 | | | |
| Debt-weighted CRP | 3.37 | 10.14 | 0.33 | 0.58 | 201.6 | 201.8 | 23.0 | 3.84 | 9.54 | 0.40 | 0.69 | 284.0 | 284.2 | 32.3 |
| GDP-weighted CRP | 4.28 | 10.14 | 0.42 | 0.74 | 292.4 | 292.5 | 26.6 | 3.37 | 9.55 | 0.35 | 0.60 | 236.6 | 236.6 | 19.6 |
| *Panel B: Wednesdays* | | | | | | | | | | | | | | |
| RW | 0.11 | 10.24 | 0.01 | 0.02 | | | | 0.11 | 10.24 | 0.01 | 0.02 | | | |
| Debt-weighted CRP | 3.20 | 10.25 | 0.31 | 0.52 | 308.5 | 309.3 | 28.1 | 4.18 | 9.88 | 0.42 | 0.71 | 429.3 | 430.4 | 49.6 |
| GDP-weighted CRP | 2.87 | 10.25 | 0.28 | 0.47 | 275.4 | 276.1 | 25.3 | 4.46 | 9.86 | 0.45 | 0.75 | 457.4 | 458.3 | 42.0 |
| *Panel C: Thursdays* | | | | | | | | | | | | | | |
| RW | 1.24 | 10.28 | 0.12 | 0.21 | | | | 1.24 | 10.28 | 0.12 | 0.21 | | | |
| Debt-weighted CRP | 6.33 | 10.35 | 0.61 | 1.11 | 505.1 | 506.5 | 59.3 | 7.01 | 9.80 | 0.72 | 1.22 | 605.7 | 606.5 | 134.2 |
| GDP-weighted CRP | 6.90 | 10.34 | 0.67 | 1.20 | 562.7 | 563.7 | 76.7 | 6.92 | 9.79 | 0.71 | 1.20 | 597.2 | 598.0 | 138.2 |

# APPENDIX

## A Theoretical derivations

This Appendix provides the main derivations of the theory.

### A.1 Proof of the identity

The identity presented in Equation (5) can be derived by expanding the following risk-neutral expectation

$$\mathbb{E}_t^*\left[\frac{S_{t+1}}{S_t}X_{t+1}\right] = cov_t^*\left(\frac{S_{t+1}}{S_t}, X_{t+1}\right) + \underbrace{\mathbb{E}_t^*\left[\frac{S_{t+1}}{S_t}\right]}_{\frac{R_{f,t}^{\$}}{R_{f,t}^{i}}}\underbrace{\mathbb{E}_t^*\left[X_{t+1}\right]}_{R_{f,t}^{\$}}, \tag{A.1}$$

where $\mathbb{E}_t^*\left[X_{t+1}\right] = R_{f,t}^{\$}$ follows from the relation between risk-neutral probability and SDF valuation

$$\frac{1}{R_{f,t}^{\$}}\mathbb{E}_t^*\left[X_{t+1}\right] = \mathbb{E}_t\left[M_{t+1}X_{t+1}\right] = 1. \tag{A.2}$$

The risk-neutral expectation in Equation (A.1) must also be equal to

$$\mathbb{E}_t^*\left[\frac{S_{t+1}}{S_t}X_{t+1}\right] = R_{f,t}^{\$}\mathbb{E}_t\left[M_{t+1}\frac{S_{t+1}}{S_t}X_{t+1}\right]. \tag{A.3}$$

By combining Equations (A.1) and (A.3) and rearranging the terms, we have

$$\mathbb{E}_t\left[M_{t+1}\frac{S_{t+1}}{S_t}X_{t+1}\right] = \frac{1}{R_{f,t}^{\$}}cov_t^*\left(\frac{S_{t+1}}{S_t}, X_{t+1}\right) + \frac{R_{f,t}^{\$}}{R_{f,t}^{i}}, \tag{A.4}$$

which must also be equal to the following expectation expansion:

$$\mathbb{E}_t\left[M_{t+1}\frac{S_{t+1}}{S_t}X_{t+1}\right] = cov_t\left(M_{t+1}X_{t+1}, \frac{S_{t+1}}{S_t}\right) + \underbrace{\mathbb{E}_t\left[M_{t+1}X_{t+1}\right]}_{1}\mathbb{E}_t\left[\frac{S_{t+1}}{S_t}\right]. \tag{A.5}$$

The combination of Equations (A.4) and (A.5) yields

$$\mathbb{E}_t\left[\frac{S_{t+1}}{S_t}\right] = \frac{R_{f,t}^{\$}}{R_{f,t}^{i}} + \frac{1}{R_{f,t}^{\$}}cov_t^*\left(\frac{S_{t+1}}{S_t}, X_{t+1}\right) - cov_t\left(M_{t+1}X_{t+1}, \frac{S_{t+1}}{S_t}\right). \tag{A.6}$$

We now decompose the risk-neutral covariance between the exchange rate return and the total portfolio gross return as

$$cov_t^*\left(\frac{S_{t+1}}{S_t}, X_{t+1}\right) = cov_t^*\left(\frac{S_{t+1}}{S_t}, r_{t+1}\right) + cov_t^*\left(\frac{S_{t+1}}{S_t}, d_{t+1}\right), \tag{A.7}$$

where the second term can be expanded as follows

$$\begin{aligned} cov_t^*\left(\frac{S_{t+1}}{S_t}, d_{t+1}\right) &= cov_t^*\left(\frac{S_{t+1}}{S_t} - 1, a - \mathbf{1}_D b\right) && \text{(A.8)} \\ &= -b\,cov_t^*\left(\frac{S_{t+1}}{S_t} - 1, \mathbf{1}_D\right) && \text{(A.9)} \\ &= -b\mathbb{E}_t^*\left[\left(\frac{S_{t+1}}{S_t} - 1\right)\mathbf{1}_D\right] + b\mathbb{E}_t^*\left[\frac{S_{t+1}}{S_t} - 1\right]\underbrace{\mathbb{E}_t^*\left[\mathbf{1}_D\right]}_{\mathbb{Q}}. && \text{(A.10)} \end{aligned}$$

Using the law of total expectation,[24] we have then

$$\begin{aligned} \mathbb{E}_t^*\left[\left(\frac{S_{t+1}}{S_t} - 1\right)\mathbf{1}_D\right] &= \underbrace{\mathbb{E}_t^*\left[\mathbf{1}_D = 1\right]}_{\mathbb{Q}}\mathbb{E}_t^*\left[\frac{S_{t+1} - S_t}{S_t} \,|_{\mathbf{1}_D=1}\right] && \text{(A.11)} \\ &= -\mathbb{Q}\mathbb{E}_t^*\left[\frac{S_t - S_{t+1}}{S_t} \,|_{\mathbf{1}_D=1}\right]. && \text{(A.12)} \end{aligned}$$

Combining Equations (A.10) and (A.12) yields

$$cov_t^*\left(\frac{S_{t+1}}{S_t}, d_{t+1}\right) = \mathbb{Q}b\mathbb{E}_t^*\left[\frac{S_t - S_{t+1}}{S_t} \,|_{\mathbf{1}_D=1}\right] + \mathbb{Q}b\left[\frac{R_{f,t}^{\$}}{R_{f,t}^{i}} - 1\right]. \tag{A.13}$$

Finally, after combining Equation (A.6) and (A.13) and rearranging the terms, we obtain the fol-

[24] Let $\mathbf{1}_A$ be an indicator function that equals 1 or 0 at time $t+1$, and $x_{t+1}$ be a random variable. We can write the following decomposition:

$$\begin{aligned} \mathbb{E}_t\left[x_{t+1}\mathbf{1}_A\right] &= P(A)\mathbb{E}_t\left[x_{t+1}\mathbf{1}_A \,|_{\mathbf{1}_A=1}\right] + (1 - P(A))\,\mathbb{E}_t\left[x_{t+1}\mathbf{1}_A \,|_{\mathbf{1}_A=0}\right] \\ &= P(A)\mathbb{E}_t\left[x_{t+1} \times 1 \,|_{\mathbf{1}_A=1}\right] + (1 - P(A))\,\mathbb{E}_t\left[\underbrace{x_{t+1} \times 0 \,|_{\mathbf{1}_A=0}}_{0}\right] \\ &= P(A)\mathbb{E}_t\left[x_{t+1} \,|_{\mathbf{1}_A=1}\right], \end{aligned}$$

where $P(A)$ is the probability of that $\mathbf{1}_A = 1$.

lowing identity:

$$\mathbb{E}_t\left[\frac{S_{t+1}}{S_t}\right] = \frac{R^{\$}_{f,t}}{R^{i}_{f,t}} + \frac{1}{R^{\$}_{f,t}} cov^*_t\left(\frac{S_{t+1}}{S_t}, r_{t+1}\right) + \frac{\mathbb{Q}b}{R^{\$}_{f,t}}\left(\mathbb{E}^*_t\left[\frac{S_t - S_{t+1}}{S_t} |_{\mathbf{1}_D=1}\right] - \mathbb{E}^*_t\left[\frac{S_t - S_{t+1}}{S_t}\right]\right) + A_t, \tag{A.14}$$

where

$$A_t = -cov_t\left(M_{t+1}X_{t+1}, \tfrac{S_{t+1}}{S_t}\right). \tag{A.15}$$

## A.2 Alternative case

We now consider a variant of the theory. Consider a gross-dollar return $X_{t+1}$ that is given by

$$X_{t+1} = 1 + r_{t+1} + d_{t+1}, \tag{A.16}$$

where the first component, $1 + r_{t+1} = R^i_{f,t}\frac{S_{t+1}}{S_t}$, captures the dollar-return of a risk-less foreign bond denominated in the foreign currency, which delivers the gross risk-free rate $R^i_{f,t}$. The second component, $d_{t+1} = a - \mathbf{1}_D b$, continues to be the payoff of a dollar-denominated contingent claim.

In this case, the risk-neutral covariance between the exchange rate return and the total portfolio gross return becomes

$$cov^*_t\left(\frac{S_{t+1}}{S_t}, X_{t+1}\right) = cov^*_t\left(\frac{S_{t+1}}{S_t}, R^i_{f,t}\frac{S_{t+1}}{S_t}\right) + cov^*_t\left(\frac{S_{t+1}}{S_t}, d_{t+1}\right) \tag{A.17}$$

$$= R^i_{f,t} var^*_t\left(\frac{S_{t+1}}{S_t}\right) + cov^*_t\left(\frac{S_{t+1}}{S_t}, d_{t+1}\right). \tag{A.18}$$

After rearranging the terms, we obtain the following identity:

$$\mathbb{E}_t\left[\frac{S_{t+1}}{S_t}\right] = \frac{R^{\$}_{f,t}}{R^{i}_{f,t}} + \frac{R^{i}_{f,t}}{R^{\$}_{f,t}} var^*_t\left(\frac{S_{t+1}}{S_t}\right) + \frac{\mathbb{Q}b}{R^{\$}_{f,t}}\left(\mathbb{E}^*_t\left[\frac{S_t - S_{t+1}}{S_t} |_{\mathbf{1}_D=1}\right] - \mathbb{E}^*_t\left[\frac{S_t - S_{t+1}}{S_t}\right]\right) + A_t, \tag{A.19}$$

which indicates that the quanto-risk premium simplifies into the risk-neutral implied variance of the exchange rate, whereas the credit-implied risk premium remains unchanged.

# B Derivation of the credit-implied risk premium

In this section, we derive the expected USD/EUR exchange rate conditional on a default, the implied currency depreciation, and the credit-implied risk premium.

## B.1 Notation

$CS^a_{t,T}$: Date-$t$ sovereign CDS spread of maturity $T$ denominated in currency $a$.

$N^a_t$: Date-$t$ notional of a sovereign CDS contract in currency $a$.

$S_t$: Date-$t$ spot USD/EUR exchange rate.

$\mathbb{E}^*$: Expectation operator under the risk-neutral measure.

$R$: Bond recovery rate at default.

$V_{t,T}$: Date-$t$ value of the USD/EUR currency swap with maturity $T$.

$\chi^a_t$: Date-$t$ accrued premium in currency $a$.

$F^a_{t,t+j,t+k}$: Date-$t$ forward discount factor in currency $a$ for the period between time $t+j$ and $t+k$, with $k > j$.

$Z^a_{t,t+j}$: Date-$t$ price of a zero-coupon bond in currency $a$ with maturity $t+j$.

$\Gamma_{t+j}$: Risk-neutral survival probability between time $t$ and $t+j$.

$t_D$: Random default time between time $t$ and $T$.

## B.2 General case

Consider a US investor who implements a long-short strategy at time $t$ by trading sovereign CDS contracts in EUR and USD with annual spread payments.[25] We denote the random default time by $t_D \in (t, T)$.

At date $t$, the investor enters a long position in a sovereign CDS denominated in EUR with notional $N^€_t$ and a short position in an identical sovereign CDS but denominated in USD with notional $N^\$_t$. Denote by $CS^€_{t,T}$ and $CS^\$_{t,T}$ the date-$t$ CDS spreads denominated in EUR and USD with maturity $T$. The long position in the CDS denominated in EUR costs $CS^€_{t,T}N^€_t$ annually, while the short position in the CDS denominated in USD pays $CS^\$_{t,T}N^\$_t$ annually. To convert all flows in USD and eliminate the currency risk involved in the payment of the CDS premium in EUR, the investor enters into a fixed-for-fixed currency swap in which EUR are received and USD are paid at the constant exchange rate $\bar{S}$. The

[25] The consideration of annual CDS payments offers the advantage of exploiting the term structure of CDS spreads, as CDS quotes in USD and EUR are available for annual maturities only.

strategy is self-financing and satisfies the following condition

$$CS^{€}_{t,T} N^{€}_t \bar{S} \;=\; CS^{\$}_{t,T} N^{\$}_t, \tag{B.1}$$

which implies that the notional of the CDS denominated in USD equals

$$N^{\$}_t \;=\; \frac{CS^{€}_{t,T} N^{€}_t}{CS^{\$}_{t,T}} \bar{S}. \tag{B.2}$$

### B.2.1 Cash flows at default

At default time $t_D$, the long position in the CDS denominated in EUR pays the investor the amount $(1-R)N^{€}_t S_{t_D}$, while the short position in CDS denominated in USD implies a delivery of $(1-R)N^{\$}_t$ to the protection buyer, where $R$ is the bond recovery rate in default. The investor is also left with a currency swap with remaining maturity $T - t_D$ that must be closed. We denote the date-$t_D$ residual value of this currency swap by $V_{t_D,T}$, which we derive in Section B.2.2. In addition, the investor receives and pays accrued premiums, which we denote by $\chi^{€}_t$ and $\chi^{\$}_t$ for the CDS positions in EUR and USD, respectively.[26] Overall, the date-$t_D$ net cash flow in USD that the investor receives/pays at default, denoted by $CF_{t_D}$, for a strategy implemented at time $t$ with notionals $N^{€}_t$ and $N^{\$}_t$, is given by:

$$CF_{t_D} \;=\; \underbrace{S_{t_D}\left[(1-R)N^{€}_t - \chi^{€}_t\right] + V_{t_D,T}}_{\text{Long CDS in EUR}} - \underbrace{\left[(1-R)N^{\$}_t - \chi^{\$}_t\right]}_{\text{Short CDS in USD}}, \tag{B.3}$$

where $\chi^{€}_t = \frac{1}{2} CS^{€}_{t,T} N^{€}_t$ and $\chi^{\$}_t = \frac{1}{2} CS^{\$}_{t,T} N^{\$}_t$ represent the accrued premiums in EUR and USD, respectively. Hence, in the case of a default, half of the annual CDS premium (determined at the inception of the strategy, i.e., time $t$) is paid to the protection seller at the next payment date.

### B.2.2 Currency swap value at default

We here determine the date-$t_D$ residual value of the currency swap, which corresponds to the discounted value of the remaining net payments (between $t_D$ and $T$) of the fixed EUR and USD legs. The value of the swap is expected to decrease as the time of default approaches the maturity of the CDS contracts.

[26] We account for accrued premiums as in Pan and Singleton (2008) and Longstaff et al. (2011) to avoid a "free lunch" for the protection buyer in the default year: the CDS contract ends at the default time (e.g., in the first half of the year), while the annual CDS premium is paid at the end of the protection year. Because the exact time of future defaults is unknown, we assume that, on average, defaults happen in the middle of the year such that half of the CDS premium is accrued, as in Choudhry and Ali (2010).

To see that, an investor who is long EUR with a 5-year maturity USD/EUR currency swap will have 4 remaining annual payments if a default happens in the first year, while she will only have 2 remaining annual payments if it occurs during the third year. The swap has no residual value if default time coincides with swap maturity, i.e., $t_D = T$.

To compute the value of the fixed EUR leg, we first discount all EUR flows at time $t_D$ and convert the obtained date-$t_D$ value at the corresponding USD/EUR exchange rate level $S_{t_D}$. Similarly, we compute the value of the fixed USD leg by discounting all USD flows at time $t_D$. The date-$t_D$ value of the currency swap $V_{t_D,T}$ is given by the difference in the discounted values of the two legs:

$$V_{t_D,T} = \left[ S_{t_D} \sum_{j=1}^{T-t_D} \left( CS^{€}_{t,T} N^{€}_t F^{€}_{t,t_D,t_D+j} \right) - \sum_{j=1}^{T-t_D} \left( CS^{\$}_{t,T} N^{\$}_t F^{\$}_{t,t_D,t_D+j} \right) \right] \mathbf{1}_{T>t_D}, \quad \text{(B.4)}$$

where the indicator function $\mathbf{1}_{T>t_D}$ is equal to 1 if $T > t_D$, i.e., when the currency swap is still open at default, and zero otherwise. We denote by $F^{€}_{t,t_D,t_D+j}$ and $F^{\$}_{t,t_D,t_D+j}$ the forward prices observed at time $t$, respectively in EUR and USD, that discount risk-free cash flows between time $t_D$ and $t_D + j$ (see Section B.4 for the computation of the forward prices).

### B.2.3 Non-arbitrage condition

In the absence of arbitrage, the date-$t$ value of the expected cash flow at default $CF_{t_D}$ must be null. Because the time of default is unknown, we have to consider all potential dates of default between time $t$ and $T$ and use the term structure of default probabilities to weigh each scenario.[27] We define $\Gamma_{t+k}$ as the risk-neutral survival probability (i.e., probability of no default) between time $t$ and time $t + k$ with $k \geq 1$, so that $\Delta\Gamma_{t+k} = \Gamma_{t+k-1} - \Gamma_{t+k}$ captures the risk-neutral probability of a default at time $t + k$ conditional on no default until time $t + k - 1$. The date-$t$ value of the expected cash flow at default, accounting for any default time $t_D = t + k$, satisfies the following non-arbitrage condition:

$$\mathbb{E}^*_t(CF_{t_D}) = \sum_{k=1}^{T} \left[ \Delta\Gamma_{t+k} Z^{\$}_{t,t+k} CF_{t+k} \right] = 0, \quad \text{(B.5)}$$

where $Z^{\$}_{t,t+k}$ is the time-discount factor in USD between $t$ and $t + k$, i.e., the price of a zero-coupon risk-free bond at time $t$ with maturity $t + k$.

[27] Appendix C provides details on the default probabilities, which we derive from the term structure of CDS spreads in USD.

We now combine Equations B.3, B.4, and B.5, replace $N_t^{\$}$ by $\frac{CS_{t,T}^{€}N_t^{€}}{CS_{t,T}^{\$}}\bar{S}$, according to Equation B.2, and divide all terms by $N_t^{€}$, which yields:

$$\sum_{k=1}^{T}\left[\Delta\Gamma_{t+k}Z_{t,t+k}^{\$}\left(\begin{array}{c}\mathbb{E}_t^*(S_{t_D}|_{t_D=t+k})\left(1-R-\frac{1}{2}CS_{t,T}^{€}\right)\\ +\sum_{j=1}^{T-k}\left(\begin{array}{c}CS_{t,T}^{€}\mathbb{E}_t^*(S_{t_D}|_{t_D=t+k})F_{t,t+k,t+k+j}^{€}\\ -CS_{t,T}^{€}\bar{S}F_{t,t+k,t+k+j}^{\$}\end{array}\right)\mathbf{1}_{T>k}\\ -(1-R)\frac{CS_{t,T}^{€}}{CS_{t,T}^{\$}}\bar{S}+\frac{1}{2}CS_{t,T}^{€}\bar{S}\end{array}\right)\right] = 0, \quad \text{(B.6)}$$

where the indicator function $\mathbf{1}_{T>k}$ is equal to 1 if $T > k$, i.e., when there is a non-zero residual swap value in default, and zero otherwise.

In the case of a strategy with 1-year CDS contracts ($T = 1$), a default can only happen at time $t+1$ and the currency swap has no residual value. Equation B.6 thus simplifies to

$$\Delta\Gamma_{t+1}Z_{t,t+1}^{\$}\left[\begin{array}{c}\mathbb{E}_t^*(S_{t_D})\left(1-R-\frac{1}{2}CS_{t,1}^{€}\right)\\ -(1-R)\frac{CS_{t,1}^{€}}{CS_{t,1}^{\$}}\bar{S}+\frac{1}{2}CS_{t,1}^{€}\bar{S}\end{array}\right] = 0. \quad \text{(B.7)}$$

Isolating the expected USD/EUR exchange rate conditional on a default, $\mathbb{E}_t^*(S_{t_D})$, in the above expression yields

$$\mathbb{E}_t^*(S_{t_D}) = S_t\frac{\Delta\Gamma_{t+1}Z_{t,t+1}^{\$}\left((1-R)\frac{CS_{t,1}^{€}}{CS_{t,1}^{\$}}-\frac{1}{2}CS_{t,1}^{€}\right)}{\Delta\Gamma_{t+1}Z_{t,t+1}^{\$}\left((1-R)-\frac{1}{2}CS_{t,1}^{€}\right)}. \quad \text{(B.8)}$$

where with set the currency swap rate to be equal to the exchange rate at issuance, i.e., $\bar{S} = S_t$.

Eventually, the date-$t$ expected currency depreciation conditional on a default, defined by $ICD_{t,1} = \frac{S_t-\mathbb{E}_t^*(S_{t_D})}{S_t}$, is given by

$$ICD_{t,1} = 1-\frac{\Delta\Gamma_{t+1}Z_{t,t+1}^{\$}\left((1-R)\frac{CS_{t,1}^{€}}{CS_{t,1}^{\$}}-\frac{1}{2}CS_{t,1}^{€}\right)}{\Delta\Gamma_{t+1}Z_{t,t+1}^{\$}\left((1-R)-\frac{1}{2}CS_{t,1}^{€}\right)}. \quad \text{(B.9)}$$

## B.3 Special case

We now derive the implied currency depreciation $ICD_{t,1}$ when there are no accrued premiums (i.e., $\chi_t^{€} = \chi_t^{\$} = 0$). In this situation, Equation B.6 becomes

$$\Delta\Gamma_{t+1} Z_{t,t+1}^{\$}(1-R)\left(\mathbb{E}_t^*(S_{t_D}) - \frac{CS_{t,1}^{€}}{CS_{t,1}^{\$}} S_t\right) = 0, \tag{B.10}$$

which implies that the expected USD/EUR exchange rate conditional on a default within 1 year, $\mathbb{E}_t^*(S_{t_D})$, equals

$$\mathbb{E}_t^*(S_{t_D}) = S_t \frac{\Delta\Gamma_{t+1} Z_{t,t+1}^{\$} \frac{CS_{t,1}^{€}}{CS_{t,1}^{\$}}}{\Delta\Gamma_{t+1} Z_{t,t+1}^{\$}} \tag{B.11}$$

$$= S_t \frac{CS_{t,1}^{€}}{CS_{t,1}^{\$}} \tag{B.12}$$

and the date-$t$ implied currency depreciation over the 1-year horizon is

$$ICD_{t,1} = \frac{CS_{t,1}^{\$} - CS_{t,1}^{€}}{CS_{t,1}^{\$}} \tag{B.13}$$

with $\bar{S} = S_t$. The same formula applies to the case of longer maturities ($T > 1$) when the residual currency swap value at default is assumed to be negligible and if we expect the currency depreciation upon default to be independent on default time, i.e., $\mathbb{E}_t^*(S_{t_D}) = \mathbb{E}_t^*(S_{t_D}|_{t_D = t+k})\ \forall\ k \leq T$. For the more general case, $\mathbb{E}_t^*(S_{t_D})$ and $ICD_{t,T}$ can be obtained from Equation B.6.

## B.4 Zero-coupon and forward prices

In this section, we determine the price of the risk-free zero-coupon bond $Z_{t,t+j}$ and the forward price $F_{t,t+j,t+k}$ for $k > j \geq 1$, following the approach of Veronesi (2010). We assume that the yield curves of interest rate swaps in USD and in EUR are available for all maturities up to $T$. At inception, an interest rate swap value is null and the yield $y_{t+j}$, with $j \in [1, T-t]$, is the constant interest rate of the fixed leg, which is akin a plain vanilla bond. Consider a set of risk-free coupon bonds with maturities ranging from $t+1$ to $T$. We define $Z^a = \begin{pmatrix} Z^a_{t,t+1} \\ Z^a_{t,t+2} \\ \vdots \\ Z^a_{t,T} \end{pmatrix}$ as a $T \times 1$ vector of zero-coupon bond prices, $P^a = \begin{pmatrix} 1 \\ 1 \\ \vdots \\ 1 \end{pmatrix}$ is the $T \times 1$ vector of principal payments set to 1, and

$$C = \begin{pmatrix} Y_{t+1}+1 & 0 & \cdots & 0 \\ Y_{t+2} & Y_{t+2}+1 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ Y_T & Y_T & \cdots & Y_T+1 \end{pmatrix} \tag{B.14}$$

is the $T \times T$ lower triangular matrix of coupons payments (including principal repayments in the diagonal), where $Y_{t+j}$ is the annual coupon of the fixed leg.

At the inception of the interest rate swap, we should have the following equality:

$$\begin{pmatrix} 1 \\ 1 \\ \vdots \\ 1 \end{pmatrix} = \begin{pmatrix} Y_{t+1}+1 & 0 & 0 & 0 \\ Y_{t+2} & Y_{t+2}+1 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ Y_T & Y_T & \cdots & Y_T+1 \end{pmatrix} \begin{pmatrix} Z^a_{t,t+1} \\ Z^a_{t,t+2} \\ \vdots \\ Z^a_{t,T} \end{pmatrix}, \tag{B.15}$$

which implies that the zero-coupon prices are given by

$$\begin{pmatrix} Z^a_{t,t+1} \\ Z^a_{t,t+2} \\ \vdots \\ Z^a_{t,T} \end{pmatrix} = \begin{pmatrix} Y_{t+1}+1 & 0 & 0 & 0 \\ Y_{t+2} & Y_{t+2}+1 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ Y_T & Y_T & \cdots & Y_T+1 \end{pmatrix}^{-1} \begin{pmatrix} 1 \\ 1 \\ \vdots \\ 1 \end{pmatrix}. \tag{B.16}$$

The date-$t$ forward price $F^a_{t,t+j,t+k}$ that discounts risk-free cash flows in currency $a$ between time $t+j$ and $t+k$, with $k > j$, can be expressed in terms of the zero-coupon prices. Under non-arbitrage conditions, we have $Z^a_{t,t+k} = Z^a_{t,t+j}F^a_{t,t+j,t+k}$, which implies that

$$F^a_{t,t+j,t+k} = \frac{Z^a_{t,t+k}}{Z^a_{t,t+j}}. \tag{B.17}$$

# C Risk-neutral default probabilities

We now derive the risk-neutral default probabilities from the term structure of CDS spreads in USD. Our approach builds on Jarrow and Turnbull (1995) and follows the reduced-form method of Hull and White (2000) and O'Kane and Turnbull (2003).

At inception (time $t$), the value of a CDS is null and satisfies

$$\underbrace{CS^{\$}_{t,T}\left[\sum_{k=1}^{T} Z^{\$}_{t,t+k}\left(\mathit{\Gamma}_{t+k} + \frac{\Delta\Gamma_{t+k}}{2}\right)\right]}_{\text{Premium leg: } L_{P,T}} - \underbrace{\sum_{k=1}^{T}\Delta\Gamma_{t+k}Z^{\$}_{t,t+k}(1-R)}_{\text{Recovery leg: } L_{R,T}} = 0, \tag{C.1}$$

where $\Gamma_{t+k}$ is the risk-neutral survival probability between time $t$ and $t+k$, $\Delta\Gamma_{t+k} = \Gamma_{t+k-1} - \Gamma_{t+k}$ is the risk-neutral probability of default at time $t+k$ conditional on no previous default, with $T \geq k \geq 1$. The first term of Equation C.1, denoted by $L_{P,T}$, is the premium leg of the CDS, while the second term of Equation C.1, denoted by $L_{R,T}$, is the recovery leg.

Building on O'Kane and Turnbull (2003), we derive a formula to compute the risk-neutral survival probability of any maturity $T$.[28] We first calculate the risk-neutral probability of a default within 1 year, $\Gamma_{t+1}$, from the 1-year CDS spread and then generalize the computation for longer maturities. In the case of a 1-year CDS contract (i.e., $T = 1$), Equation C.1 simplifies to

$$CS^{\$}_{t,1}Z^{\$}_{t,t+1}\left(\mathit{\Gamma}_{t+1} + \frac{(\Gamma_t - \Gamma_{t+1})}{2}\right) - (\Gamma_t - \Gamma_{t+1})Z^{\$}_{t,t+1}(1-R) = 0, \tag{C.2}$$

which yields

$$\mathit{\Gamma}_{t+1} = \frac{(1-R) - \frac{CS^{\$}_{t,1}}{2}}{(1-R) + \frac{CS^{\$}_{t,1}}{2}} \tag{C.3}$$

using $\Gamma_t = 1$, as a CDS is never triggered at issuance time $t$.

[28] Based on the assumption that CDS contracts pay annual premiums and that CDS spreads quotes and zero coupon prices are available, we obtain the default probabilities analytically. O'Kane and Turnbull (2003) consider multiple payments per year, which implies using a root-searching algorithm to calculate the default probabilities.

We now determine the survival probability between time $t$ and $t+k$, denoted by $\Gamma_{t+k}$, for $k > 1$. We use the CDS spread of maturity $T = k$ and the survival probability estimated between time $t$ and $t+k-1$, which determines the value of $L_{P,t+k-1}$ and $L_{R,t+k-1}$. Again, we exploit the fact that a CDS has no value at inception date, i.e., $L_{P,t+k} - L_{R,t+k} = 0$, which we can decompose as follows

$$L_{P,t+k-1} - L_{R,t+k-1} + Z^{\$}_{t,t+k}\left[CS^{\$}_{t,t+k}\Gamma_{t+k} + \left(\frac{1}{2} - (1-R)\right)(\Gamma_{t+k-1} - \Gamma_{t+k})\right] = 0, \quad \text{(C.4)}$$

which yields the following expression for $\mathit{\Gamma}_{t+k}$

$$\mathit{\Gamma}_{t+k} = \frac{-L_{P,t+k-1} + L_{R,t+k-1} + Z^{\$}_{t,t+k}\mathit{\Gamma}_{t+k-1}\left((1-R) - \frac{1}{2}CS^{\$}_{t,t+k}\right)}{Z^{\$}_{t,t+k}\left(\frac{1}{2}CS^{\$}_{t,t+k} + (1-R)\right)}. \quad \text{(C.5)}$$

To sum up, we first use Equation C.3, the 1-year CDS spread in USD, and the price of a 1-year zero-coupon bond to obtain the risk-neutral survival probability over year 1, $\mathit{\Gamma}_{t+1}$. We then use this probability in Equation C.5 along the 2-year CDS spread and zero-coupon bond to compute the risk-neutral survival probability over the years 1 and 2, $\mathit{\Gamma}_{t+2}$. We resume the procedure until we attain the desired horizon $k = T$.[29]

[29] In additional robustness analyses, we compute CRP using CDS contracts with longer maturities (e.g., 5 years). For these cases, we use the iterative process described above for the computation of the conditional survival probabilities. As CDS contracts with maturities 2 and 4 years are not available, we first compute the synthetic credit spreads for those contracts using a linear interpolation. We then compute the conditional survival probabilities, the unconditional default probabilities, and the implied currency depreciation for each maturity.

## D Benchmark vs. general case

This Appendix examines how the ICD and CRP measures vary whether or not we account for the accrued premiums. Following Appendix B, the date-$t$ level of ICD based at the 1-year horizon is given by

$$ICD_{t,1} \quad = \quad 1 - \frac{\Delta\Gamma_{t+1} Z^{\$}_{t,t+1} \left( (1-R) \frac{CS^{€}_{t,1}}{CS^{\$}_{t,1}} - \frac{1}{2} CS^{€}_{t,1} \right)}{\Delta\Gamma_{t+1} Z^{\$}_{t,t+1} \left( (1-R) - \frac{1}{2} CS^{€}_{t,1} \right)} \tag{D.1}$$

in the general case, which includes the accrued premiums for the EUR and USD CDS positions, while it is given by

$$ICD_{t,1} \quad = \quad \frac{CS^{\$}_{t,1} - CS^{€}_{t,1}}{CS^{\$}_{t,1}} \tag{D.2}$$

in the special case without accrued premiums, which is our benchmark scenario. For each case, we compute CRP using Equation (F.2) and the corresponding ICD measure.

Panel A of Table I reports the descriptive statistics of ICD and CRP in both cases. The mean (median) CRP is 0.106% (0.110%) in the benchmark case and 0.107% (0.111%) in the general case. The standard deviations are respectively 0.032 and 0.033. Overall, the differences between the two cases are economically negligible. We thus focus on the benchmark case throughout the paper, although the empirical results remain very similar when computing CRP under the general case.

**Table I : Descriptive statistics – with and without accrued premiums**

This table reports descriptive statistics of the Eurozone's ICD and CRP. Panel A presents results in benchmark case, which excludes accrued premiums. Panel B reports the results for the general case, which includes the accrued premiums. ICD and CRP are computed with 1-year CDS contracts. The computation of the ICD measure is detailed in Section 3.2 and the computation of CRP in Section 3.3. The countries included in the aggregated measure are Belgium, France, Germany, Italy, Ireland, Portugal, Spain, Netherlands, Lithuania, Austria, Slovenia, Slovakia, Cyprus, Estonia, Finland and Latvia. The reported ICD and CRP measures are the average of these countries' ICD weighted by their respective public debt. The sample consists of observations between August 20, 2010, and April 26, 2019.

| Variables | N | Mean | Median | Std | Min | Max |
|---|---|---|---|---|---|---|
| *Panel A: Benchmark case (without accrued premiums)* | | | | | | |
| $CRP$ (%) | 2266 | 0.106 | 0.110 | 0.032 | -0.042 | 0.225 |
| $ICD$ (%) | 2266 | 24.697 | 25.723 | 7.649 | -7.449 | 50.163 |
| *Panel B: General case (with accrued premiums)* | | | | | | |
| $CRP$ (%) | 2266 | 0.107 | 0.111 | 0.033 | -0.042 | 0.243 |
| $ICD$ (%) | 2266 | 24.761 | 25.799 | 7.667 | -7.463 | 50.262 |

# E Description of the data

This Appendix provides information on the data used in the paper. Table II presents the data sources of the variables that we either use in the computation of CRP or as controls, while Table III reports some descriptive statistics of the main variables.

**Table II : Data sources**
This table provides information on the main variables considered in the paper. We report the frequency and the source of the series, with Bloomberg tickers when appropriate.

| Variables | Frequency | Source | Methodology |
|---|---|---|---|
| *Panel A: Dependent and control variables* | | | |
| USD/EUR exchange rate return | Daily | Computed with Bloomberg (Ticker: EURUSD) | |
| EUR-USD interest rate differential (1y) | Daily | Computed with 12m swaps from Bloomberg | |
| Quanto-implied risk premium on USD/EUR | Monthly | Obtained from Kremens and Martin (2019) | |
| Implied variance on USD/EUR (1y) | Daily | Computed using options' quotes from JP Morgan | |
| Realized volatility on USD/EUR (1y) | Daily | Computed with options' quotes from JP Morgan | Appendix E.1 |
| Volatility risk premium on USD/EUR | Daily | Computed with options' quotes from JP Morgan | Appendix E.1 |
| USD/EUR iliquidity | Monthly | Obtained from Angelo Ranaldo, computed as per Karnaukh et al. (2015) | |
| Markit iTraxx senior financials (5y) | Daily | Obtained from Datastream | |
| *Panel B: Global factors* | | | |
| Carry trade factor | Daily | Computed with spot and forward exchange rates from WM/Reuters | Appendix E.2 |
| Dollar factor | Daily | Computed with spot and forward exchange rates from WM/Reuters | Appendix E.2 |
| Momentum factor | Daily | Computed with spot and forward exchange rates from WM/Reuters | Appendix E.2 |
| Value factor | Daily | Computed with spot and forward exchange rates from WM/Reuters | Appendix E.2 |
| Risk-reversal factor | Daily | Computed with spot and forward exchange rates from WM/Reuters and options' quotes from JP Morgan | Appendix E.2 |
| Global imbalance factor | Daily | Computed with spot and forward exchange rates from WM/Reuters and Foreign asset/liabilities from Philip Lane's website and IMF Intl. financial statistics | Appendix E.2 |
| *Panel C: Components of CRP* | | | |
| Implied currency depreciation (1y) | Daily | Computed with Markit's sovereign CDS | Appendix B.2 and B.3 |
| Risk-neutral probability of default (1y) | Daily | Computed with Markit's sovereign CDS and Bloomberg swaps (Tickers: EUSA1to30, USSA1to30) | Appendix C |
| USD risk-free rate (1y) | Daily | Computed with Bloomberg swaps | |
| EUR risk-free rate (1y) | Daily | Computed with Bloomberg swaps | |
| GDP in Euro by country | Yearly | Retrieved from the ECB's website | |
| Sovereign debt-to-GDP by country | Yearly | Retrieved from Bloomberg (Tickers: GDDB*CTRY*) | |
| *Panel D: Drivers of CRP* | | | |
| VSTOXX | Daily | Retrieved from Bloomberg (Ticker: V2X) | |
| Industrial production (YoY) | Monthly | Retrieved from Bloomberg (Ticker: EUIPEMUY) | |
| Citigroup inflation surprise index | Monthly | Retrieved from Datastream (Ticker: EKCSIIINR) | |
| German bond yield (2y) | Daily | Retrieved from Bloomberg (Ticker: GDBR2) | |
| ECB announcement dates | Daily | Obtained from the ECB website | |

## E.1 Option-based measures

We here discuss the construction of the various option-based measures.

**Table III : Descriptive statistics of the main variables**

This table reports the descriptive statistics of the main variables used in the predictability analysis. All variables are annualized. The sample consists of observations between August 20, 2010, and April 26, 2019.

| Variables | N | Mean | Median | Std | Min | Max |
|---|---|---|---|---|---|---|
| USD/EUR exchange rate return (%) | 2154 | -0.0118 | 0.0186 | 1.4065 | -6.0030 | 7.5997 |
| EUR-USD interest rate differential (1y) | 2154 | 1.0051 | 1.0008 | 0.0132 | 0.9820 | 1.0324 |
| Realized volatility on USD/EUR (1y) | 2154 | 0.0901 | 0.0881 | 0.0206 | 0.0530 | 0.1228 |
| Volatility risk premium (1y) | 2154 | -0.0133 | -0.0102 | 0.0158 | -0.0685 | 0.0170 |
| USD/EUR illiquidity | 107 | -0.3417 | -0.4099 | 0.5479 | -1.4974 | 1.7846 |
| USD risk-free rate (1y) | 2154 | 0.0092 | 0.0051 | 0.0080 | 0.0025 | 0.0298 |
| EUR risk-free rate (1y) | 2154 | 0.0042 | 0.0030 | 0.0072 | -0.0038 | 0.0224 |
| | | | | | | |
| CRP (1y), debt-weighted | 2154 | 0.1068 | 0.1104 | 0.0328 | -0.0416 | 0.2410 |
| CRP (1y), GDP-weighted | 2154 | 0.0998 | 0.1030 | 0.0286 | -0.0219 | 0.2112 |
| | | | | | | |
| Quanto-implied risk premium on USD/EUR | 64 | 0.0122 | 0.0138 | 0.0113 | -0.0063 | 0.0342 |
| Implied variance on USD/EUR, $\text{IVAR}_{\text{ATM}}$ | 2154 | 0.0100 | 0.0084 | 0.0053 | 0.0035 | 0.0287 |
| Implied variance on USD/EUR, $\text{IVAR}_{\text{M}}$ | 2154 | 0.0109 | 0.0090 | 0.0057 | 0.0039 | 0.0312 |
| Implied variance on USD/EUR, $\text{IVAR}_{\text{BJN}}$ | 2154 | 0.0114 | 0.0093 | 0.0063 | 0.0039 | 0.0343 |

### E.1.1 Currency option data

We collect over-the-counter (OTC) data on currency options from JP Morgan. OTC options are quoted in terms of Garman and Kohlhagen (1983a) implied volatility on baskets of constant maturity plain vanilla options for fixed deltas ($\delta$). From these data, one can recover the implied volatility smile ranging from a $10\delta$ put to a $10\delta$ call option. To convert deltas into strike prices and implied volatilities into option prices, we employ exchange rates and zero-yield rates obtained by bootstrapping money market rates and interest rate swap data from Bloomberg.

According to market convention, an option written on a major currency pair (and up to a one-year maturity) is quoted using the spot $\delta$. In this case, the strike prices at time $t$ is given by

$$
\begin{aligned}
K_{t,\tau,atm} &= F_{t,\tau} \cdot \exp\left\{\frac{\tau}{2} \cdot \sigma^2_{t,\tau,atm}\right\},\\
K_{t,\tau,25c} &= F_{t,\tau} \cdot \exp\left\{\frac{\tau}{2} \cdot \sigma^2_{t,\tau,25c} - N^{-1}[0.25 \cdot \exp(i^{\star}_{t,\tau} \cdot \tau)] \cdot \sigma^2_{t,\tau,25c} \cdot \sqrt{\tau}\right\},\\
K_{t,\tau,25p} &= F_{t,\tau} \cdot \exp\left\{\frac{\tau}{2} \cdot \sigma^2_{t,\tau,25p} + N^{-1}[0.25 \cdot \exp(i^{\star}_{t,\tau} \cdot \tau)] \cdot \sigma^2_{t,\tau,25p} \cdot \sqrt{\tau}\right\},
\end{aligned}
$$

where $F_{t,\tau}$ is the forward exchange rate at time $t$ with delivery date $t+\tau$, $K_{t,\tau}$ is the strike price at time $t$ of an option expiring at time $t+\tau$, $\sigma_{t,\tau}$ is the implied volatility at time $t$ of an option expiring at time $t+\tau$, $i^{\star}_{t,\tau}$ is interest rate at time $t$ with maturity $\tau$, $atm$ refers to an at-the-money option, $25c$ stands for a $25\delta$ call option, and $25p$ denotes a $25\delta$ put option. Formulae for $10\delta$ call and put options can be then easily derived.

### E.1.2 Model-free implied variance

The risk-neutral expectation of the integrated variance between two dates $t$ and $t+\tau$ can be calculated by integrating over an infinite range of the strike prices from European call and put options expiring on these dates as

$$\mathbb{E}_t^*[RV_{t,\tau}^2] = \frac{2}{B_{t,\tau}}\Big\{\int_0^{F_{t,\tau}} \frac{P_{t,\tau}(K)}{K^2}dK + \int_{F_{t,\tau}}^{\infty} \frac{C_{t,\tau}(K)}{K^2}dK\Big\}, \tag{E.1}$$

where $P_{t,\tau}(K)$ and $C_{t,\tau}(K)$ are the put and call option prices at time $t$ with strike price $K$ and maturity date $t+\tau$, respectively. $B_{t,\tau}$ is the price of a domestic bond at time $t$ with maturity date $t+\tau$, and $\mathbb{E}_t^*[\cdot]$ denotes the expectation under the risk-neutral measure. $\tau$ is set to one month in our analysis.

This is the model-free approach of Britten-Jones and Neuberger (2000) which is based on no-arbitrage conditions and requires no specific option pricing model. In our implementation, we follow Jiang and Tian (2005) and use a cubic spline around the available implied volatility points. This interpolation method is standard in the literature and has the advantage that the implied volatility smile is smooth between the maximum and minimum available strikes. The simple implied variance of Martin (2017) replaces $K^2$ with $S_t^2$ in Equation (E.1). Finally, we compute the option values using the Garman and Kohlhagen (1983b) valuation formula and solve the integral in Equation (E.1) via trapezoidal integration. The implied volatility is simply computed by taking the square root of the implied variance, i.e., $\mathbb{E}_t^*\left[RV_{t,\tau}\right] = \sqrt{\mathbb{E}_t^*\left[RV_{t,\tau}^2\right]}$.

### E.1.3 Realized volatility

The realized volatility of the log returns of the USD/EUR exchange rate is from now on simply referred to as realized volatility. Following Della Corte et al. (2016a), we proxy $\mathbb{E}_t\left[RV_{t,\tau}\right]$ by simply using the lagged realized volatility, i.e.,

$$\mathbb{E}_t\left[RV_{t,\tau}\right] = RV_{t-\tau,\tau} = \sqrt{\frac{365}{\tau}\sum\nolimits_{i=0}^{\tau} r_{t-i}^2},$$

where $r_t$ is the daily log return on the underlying security and $\tau$ is defined as one year.

### E.1.4 Volatility risk premium

The volatility risk premium is the difference between the physical and the risk-neutral expectation of the future realized volatility. Formally, the $\tau$-period volatility risk premium at time $t$ is defined as

$$VRP_{t,\tau} = \mathbb{E}_t \left[ RV_{t,\tau} \right] - \mathbb{E}_t^* \left[ RV_{t,\tau} \right],$$

where $\mathbb{E}_t [\cdot]$ is the conditional expectation operator at time $t$ under the physical measure. We proxy $\mathbb{E}_t [RV_{t,\tau}]$ by simply using the lagged realized volatility based on daily log exchange rate returns, i.e., $\mathbb{E}_t [RV_{t,\tau}] = RV_{t-\tau,\tau}$. This approach is widely used for forecasting exercises. It makes $VRP_{t,\tau}$ directly observable at time $t$, requires no modeling assumptions, and is consistent with the stylized fact that realized volatility is a highly persistent process. Thus, at time $t$, we measure the volatility risk premium over the $[t, t+\tau]$ time interval as the difference between the ex post realized volatility over the $[t-\tau, t]$ interval and the ex ante risk-neutral expectation of the future realized volatility over the $[t, t+\tau]$ interval.

## E.2 Tradable currency factors

This section briefly outlines the construction of the currency factors used in the main analysis. At the end of each month $t$, we first allocate currencies into five baskets and then construct a long-short strategy that is short (long) the first (last) portfolio. Hence, we use the currency composition of this strategy and track its daily exchange rate return. We repeat this exercise for a variety of currency strategies, which we summarize below.

### E.2.1 Data for portfolio construction

We collect daily spot and one-month forward exchange rates relative to the US dollar (USD) from WM/Reuters via Datastream. Exchange rates are defined as units of US dollars per unit of foreign currency so that an increase in the exchange rate indicates an appreciation of the foreign currency. Monthly data are obtained by sampling end-of-month rates. Our sample includes the most liquid developed and emerging market currencies, i.e., the currencies of Australia, Brazil, Canada, Czech Republic, Denmark, Euro Area, Hungary, Japan, Mexico, New Zealand, Norway, Poland, Singapore, South Africa, South Korea, Sweden, Switzerland, Taiwan, Turkey, and United Kingdom. We also consider a subset of developed currencies, i.e., the currencies of Australia, Canada, Denmark, Euro Area, Japan, New Zealand, Norway, Sweden, Switzerland, and United Kingdom.

Turning to macro data, we obtain end-of-year series on foreign assets and liabilities, and gross domestic product (GDP) from Lane and Milesi-Ferretti (2007), available on Philip Lane's website. We extend the dataset until the end of 2016 using the IMF's International Financial Statistics database. We also use end-of-year series on the proportion of external liabilities denominated in domestic currency from Bénétrix et al. (2015). For these data, we retrieve monthly observations by keeping end-of-period data constant until a new observation becomes available. Finally, we collect monthly year-on-year inflation data from Datastream.

**Dollar and carry factors**. At the end of each month $t$, we allocate currencies to five portfolios on the basis of their forward premia: $20\%$ of all currencies with the highest forward premia are assigned to Portfolio 1, whereas $20\%$ of all currencies with the lowest forward premia are assigned to Portfolio 5. We then compute the return for each portfolio as an equally weighted average of individual currency returns within that portfolio. Following Lustig et al. (2011), the $DOL$ factor is computed as an equally weighted average of these portfolios and the $CAR$ factor as a long-short portfolio formed by going long Portfolio 5 (high-yielding currencies) and short Portfolio 1 (low-yielding currencies).

**Global imbalance factor**. At the end of each month $t$, we first group currencies into two baskets using the net foreign asset position relative to GDP and then rank the currencies within each basket using the percentage share of external liabilities denominated in domestic currency. Hence, we allocate them to five portfolios as in Della Corte et al. (2016b). Portfolio 1 corresponds to creditor countries whose external liabilities are primarily denominated in domestic currency (safest currencies), whereas Portfolio 5 comprises debtor countries whose external liabilities are primarily denominated in foreign currency (riskiest currencies). We then compute the return for each portfolio as an equally weighted average of individual currency returns within that portfolio. We construct the global imbalance factor $IMB$ as return difference between Portfolio 5 and Portfolio 1. The construction of these is theoretically motivated by the work of Gabaix and Maggiori (2015) and Colacito et al. (2018).

**Momentum portfolios**. At the end of each month $t$, we form five portfolios based on exchange rate returns over the previous three-months. We assign $20\%$ of all currencies with the highest lagged exchange rate returns to Portfolio 5, and $20\%$ of all currencies with the lowest lagged exchange rate returns to Portfolio 1. We then compute the return for each portfolio as an equally weighted average of the currency returns within that portfolio. A strategy that is long in Portfolio 5 (winner currencies) and short in Portfolio 1 (loser currencies) is then denoted as $MOM$.

**Value portfolios**. At the end of each month $t$, we form five portfolios based on currency value measured as the 5-year change in log real exchange rate (e.g., Asness et al., 2013). We assign $20\%$ of all currencies with the lowest (highest) currency value to Portfolio 5 (Portfolio 1). We then compute the return for each portfolio as an equally weighted average of the currency returns within that portfolio. A strategy that is long in Portfolio 5 (most undervalued currencies) and short in Portfolio 1 (most overvalued currencies) is denoted as $VAL$.

**Risk-reversal portfolios**. At the end of each month $t$, we form five portfolios based on out-of-the-money options. For each currency in each time period, we compute the risk-reversal, which is the implied volatility of the $10$ delta call less the implied volatility of the $10$ delta put. We then assign $20\%$ of all currencies with the lowest (highest) risk reversal to Portfolio 5 (Portfolio 1). We then compute the return for each portfolio as an equally weighted average of the currency returns within that portfolio. A strategy that is long in Portfolio 5 (high-skewness currencies) and short in Portfolio 1 (low-skewness currencies) is then denoted as $RR$.

# F Additional results

This Appendix provides an additional set of empirical results for robustness.

## F.1 Synthetic Replication of QRP

We provide evidence that QRP is tightly related to our synthetic version of QRP, which we denote by $\text{QRP}^{\text{S}}$. We have $\text{QRP}^{\text{S}} = \text{IVAR} \times R^{€}_{f,t}/R^{\$}_{f,t}$, where IVAR is the option-implied variance on the USD/EUR and $R^{€}_{f,t}$ and $R^{\$}_{f,t}$ are the one-year gross interest rate for the EUR and the USD, respectively. As illustrated by Figure I, for example, the model-free implied variance of currency options on the USD/EUR is closely related to QRP. The correlation between the two series is above 80%.

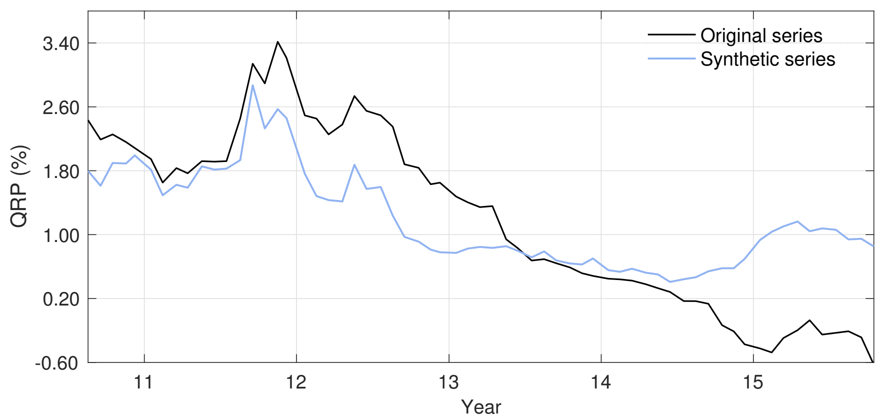


**Figure I: Original and synthetic QRP**
This figure illustrates the time series of the quanto-implied risk premium (QRP) of Kremens and Martin (2019) against its synthetic version. The synthetic version of QRP is computed as $\text{IVAR} \times R^{€}_{f,t}/R^{\$}_{f,t}$, where IVAR is the model-free implied variance of currency options on the USD/EUR. The sample consists of monthly observations between August 2010 and November 2015.

We also test empirically the validity of $\text{QRP}^{\text{S}}$ using a panel of 11 exchange rates akin to Kremens and Martin (2019). Table IV presents panel estimates of QRP regressed on $\text{QRP}^{\text{S}}$ and uncovers a slope coefficient that is statistically indifferent from one. These results are robust to using different versions of the implied variance, i.e., the implied variance from ATM options, model-free implied variance of Britten-Jones and Neuberger (2000), or simple implied variance of Martin (2017), and to considering currency and time fixed effects.

**Table IV : Relation between synthetic and original QRP**
This table presents regression estimates for a panel of exchange rates. The dependent variable is the quanto-implied risk premium (QRP$_i$) of Kremens and Martin (2019) for currency $i$. In Panel A, the independent variable is the implied variance (IVAR$_{i,j}$) from one-year options on currency $i$ with $j = \{BGN, M, ATM\}$, where $BGN$ denotes the model-free implied variance constructed as in Britten-Jones and Neuberger (2000), $M$ refers to the simple implied variance computed as in Martin (2017), and $ATM$ is the at-the-money implied variance. In Panel B, the independent variable is IVAR$_{i,j} \times R^i_{f,t}/R^\$_{f,t}$, where $R^i_{f,t}$ and $R^\$_{f,t}$ are the one-year gross interest rate for currency $i$ and the USD, respectively. We include currency and time (month) fixed effects (*FE*). Standard errors are clustered by currency and time dimensions and are reported in parentheses. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. The sample includes monthly observations from December 2009 to October 2015 for a cross-section of 11 currency pairs relative to the US dollar as in Kremens and Martin (2019). Data are from Bloomberg and Markit.

*Panel A: QRP$_i$ regressed on IVAR$_{i,j}$*

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| IVAR$_{BJN}$ | 1.111*** | 1.013*** | 1.013*** | | | | | | |
| | (0.114) | (0.115) | (0.114) | | | | | | |
| IVAR$_{M}$ | | | | 1.146*** | 1.088*** | 1.087*** | | | |
| | | | | (0.110) | (0.158) | (0.157) | | | |
| IVAR$_{ATM}$ | | | | | | | 1.254*** | 1.190*** | 1.189 *** |
| | | | | | | | (0.120) | (0.165) | (0.163) |
| Constant | -0.001 | 0.161 | 0.161 | 0.039 | 0.131 | 0.132 | 0.044 | 0.136 | 0.137 |
| | (0.469) | (0.187) | (0.189) | (0.409) | (0.244) | (0.246) | (0.452) | (0.232) | (0.233) |
| $R^2$ (%) | 39.2 | 88.0 | 87.9 | 33.0 | 86.6 | 86.6 | 34.2 | 86.8 | 86.8 |
| N | 656 | 656 | 656 | 656 | 656 | 656 | 656 | 656 | 656 |
| Currency FE | | × | × | | × | × | | × | × |
| Time FE | | | × | | | × | | | × |

*Panel B: QRP$_i$ regressed on IVAR$_{i,j} \times R^i_{f,t}/R^\$_{f,t}$*

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| QRP$_{BJN}$ | 1.082*** | 0.989*** | 0.988*** | | | | | | |
| | (0.118) | (0.110) | (0.109) | | | | | | |
| QRP$_{M}$ | | | | 1.126*** | 1.066*** | 1.065*** | | | |
| | | | | (0.107) | (0.150) | (0.148) | | | |
| QRP$_{ATM}$ | | | | | | | 1.229*** | 1.165*** | 1.164 *** |
| | | | | | | | (0.123) | (0.157) | (0.155) |
| Constant | 0.017 | 0.175 | 0.175 | 0.044 | 0.140 | 0.141 | 0.052 | 0.146 | 0.148 |
| | (0.477) | (0.181) | (0.183) | (0.420) | (0.235) | (0.236) | (0.461) | (0.225) | (0.226) |
| $R^2$ (%) | 40.3 | 88.1 | 88.1 | 34.3 | 86.9 | 86.8 | 35.4 | 87.0 | 87.0 |
| N | 656 | 656 | 656 | 656 | 656 | 656 | 656 | 656 | 656 |
| Currency FE | | × | × | | × | × | | × | × |
| Time FE | | | × | | | × | | | × |

## F.2 Predictability with country-level CRP

We now analyze the predictability of CRP for the USD/EUR exchange rate at the country level. The aim is to ensure that our results are driven by the economically impactful countries and not by a few small European countries with negligible credit risk. Figure II reports the results when we estimate the predictive regression (11) using country-level CRP. Panel A reports the regression coefficient $b_k$ for each country, when we consider all controls, and the corresponding 90% confidence intervals. Countries are ranked by the size of their regression coefficient. Panel B reports the predictive power ($R^2$) of each country's CRP obtained from univariate predictive regressions.

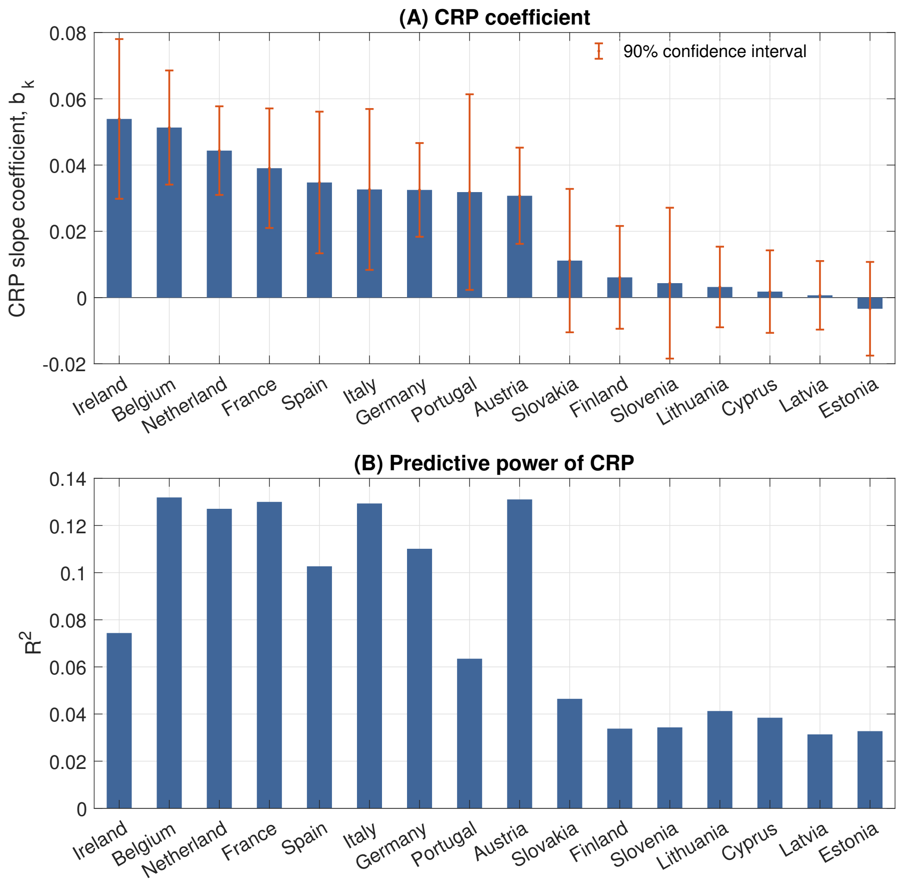


**Figure II: Exchange rate return predictability with country-level CRP**
This figure illustrates the predictability of the USD/EUR exchange rate returns with the credit-implied risk premium (CRP) by country. Panel A shows the regression coefficient of each country's CRP $b_k$ at the 6-month horizon, using the full set of controls. Panel B presents the corresponding $R^2$ of the univariate regression with the country-level CRP measures. The red lines display 90% confidence intervals. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. The sample consists of observations between August 20, 2010, and April 26, 2019.

The countries exhibiting the highest predictability of USD/EUR excess returns are typically the risky countries that contributed largely to the European debt crisis, such as Ireland, Italy, Portugal, and Spain, or the two economically large creditworthy entities, which are France or Germany. Austria, Belgium, and Netherlands also naturally inherit from a high predictive power, given their strong interconnection with the French and German economies. By contrast, the countries that do not seem to contain any predictability are those playing a minor economic role within the Eurozone, such as Cyprus, Estonia, Latvia, or Lithuania. This analysis confirms that our risk premium measure for the Eurozone adequately captures the interaction between currency depreciation and default, as the main contributors are CRP of countries with a high risk of default and/or countries whose economic conditions have a strong impact on the USD/EUR exchange rate.

### F.3 Counterparty risk

An alternative explanation of our findings could be dealers' counterparty risk. Suppose, for example, that CDS in EUR are largely quoted by European banks while CDS in USD are mostly quoted by US banks. It then follows that the percentage spread between EUR- and USD-denominated CDS on the same underlying entity could be partly attributed to dealers' credit risk, as opposed to exchange rate depreciation due to sovereign risk. To shed light on this concern, we use the 5-year Markit iTraxx Europe Senior Financials Index as a proxy of counterparty risk and use it as an additional control variable in our core specification. The Markit iTraxx Europe Senior Financials Index aggregates the 25 most liquid single names CDS spreads of major financial entities in Europe and is generally used to monitor the credit risk exposure of financial institutions in Europe. Table V shows that counterparty risk is not driving the results of this paper, as CRP continues to be a highly significant predictor of USD/EUR excess returns after controlling for credit risk of the major European banks.

### F.4 Counterfactual CRP measure

In this section, we verify that the predictive content of CRP comes from the time variation in implied currency depreciation $ICD_{i,t}$, as predicted by the expression for country $i$:

$$CRP_{i,t} = \frac{b\mathbb{Q}_i}{R^{\$}_{f,t}} \left(ICD_{i,t} - ECD_t\right), \tag{F.1}$$

where the risk-neutral default probability $\mathbb{Q}_i$ is constant.

As a counterfactual, we construct an alternative measure of CRP:

$$CRP^{*}_{i,t} = \frac{b\mathbb{Q}_{i,t}}{R^{\$}_{f,t}} \left(ICD_i - ECD_t\right), \tag{F.2}$$

where $ICD_i$ is now constant but the default probability $\mathbb{Q}_{i,t}$ varies over time. Table VI shows that this

**Table V : Controlling for counterparty risk**
This table presents results on the predictability of the USD/EUR exchange rate return using the credit-implied risk premium (CRP) for the Eurozone and controlling for the counterparty risk of European Banks. The dependent variable is the annualized average daily log USD/EUR return over a given forecast horizon. CRP is constructed by weighting country-level CRP components by their outstanding debt level. Panel A controls for the 12-month USD/EUR interest rate differential. Panel B additionally controls for the USD/EUR illiquidity, realized volatility, volatility risk premium, and global currency factors. Both panels control for counterparty risk with the Markit iTraxx Europe Senior Financials Index, which reflects the average level of credit risk of the major financial entities in Europe. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. The sample consists of daily observations between August 20, 2010, and April 26, 2019.

*Panel A: Benchmark model*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| CRP$_t$ | 2.225** | 2.439*** | 2.066*** | 1.722*** | 0.765*** |
| | (0.010) | (0.004) | (0.001) | (0.003) | (0.010) |
| R$^2$ (%) | 1.03 | 6.24 | 17.15 | 26.42 | 33.95 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel B: With all controls*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| CRP$_t$ | 2.099** | 2.204*** | 1.688*** | 1.349*** | 0.471* |
| | (0.024) | (0.008) | (0.001) | (0.002) | (0.072) |
| R$^2$ (%) | 2.44 | 9.75 | 22.25 | 35.88 | 49.31 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

alternative CRP measure cannot predict USD/EUR excess returns at any horizon, which confirms that $ICD_{i,t}$ is the primary source of predictability.

**Table VI : Predictability with counterfactual CRP**
This table presents results on the predictability of the USD/EUR exchange rate return using a counterfactual credit-implied risk premium (CRP) for the Eurozone. The dependent variable is the annualized average daily log USD/EUR return over a given forecast horizon. CRP is based on a constant implied currency depreciation and a time-varying default probability. Panel A controls for the 12-month USD/EUR interest rate differential. Panel B controls for the USD/EUR illiquidity, realized volatility on the USD/EUR, volatility risk premium, and global currency factors. We report $p$-values in parenthesis, computed using Newey and West (1987) standard errors with the lag equal to the forecasting horizon. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and ***, respectively. Section 3.3 describes the computation of CRP, Section 4.1 presents the controls, while Section 4.2 discusses the econometric specification. The sample consists of daily observations between August 20, 2010, and April 26, 2019.

*Panel A: Benchmark model*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| CRP$_t$ | -0.065 | -0.055 | 0.057 | 0.148 | 0.330 |
| | (0.816) | (0.827) | (0.781) | (0.547) | (0.103) |
| R$^2$ (%) | 0.00 | 0.09 | 1.41 | 4.11 | 19.83 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

*Panel B: With all controls*

| | 1 week | 1 month | 3 months | 6 months | 12 months |
|---|---|---|---|---|---|
| CRP$_t$ | -0.651 | -0.512 | -0.308 | -0.241 | 0.015 |
| | (0.140) | (0.203) | (0.248) | (0.307) | (0.905) |
| R$^2$ (%) | 1.97 | 6.14 | 15.87 | 27.96 | 44.44 |
| N | 2154 | 2138 | 2096 | 2033 | 1907 |

## F.5 Drivers of CRP

We explore how CRP differs from crash risk, credit risk, and QRP with an analysis of their determinants. Table VII reports the results. CRP decreases with the level of financial uncertainty, as measured by the 30-day implied volatility of the EURO STOXX 50 (VSTOXX), and increases with economic activity, as measured with the year-on-year industrial production growth in the Eurozone. Therefore, CRP is pro-cyclical. By contrast, crash risk, credit risk, and QRP (or $\mathrm{QRP}^{\mathrm{S}}$) are all countercyclical, as they are positively linked to the level of uncertainty in financial markets, in line with traditional sources of currency risk premia, such as the countercyclical dollar factor (Lustig et al., 2014).[30]

Nominal conditions, as measured with the 2-year German Bund yield, do not explain CRP but strongly impact the alternative measures of risk. Following a growing literature on the role of central bank communication for asset prices, we also investigate the impact of European Central Bank (ECB) announcements. We find that the type of ECB announcements affects the relationship between the Bund yield and CRP or crash risk. The relation between CRP and the Bund yield becomes negative when ECB announcements essentially reflect monetary news rather than news about economic growth, which we classify based on the direction of the comovement between the stock market and Bund yield on the day of the news (Cieslak and Schrimpf, 2019). In comparison, the relation between crash risk and the Bund yield becomes negative when ECB announcements are viewed as negative news by market participants, as indicated by a fall of the stock market on the day of the news. Hence, CRP is sensitive to the informational content of central bank communications, while crash risk rather depends the direction of the news. By contrast, the type of ECB news does not help explain fluctuations of credit risk or $\mathrm{QRP}^{\mathrm{S}}$.

[30]However, all measures vary similarly with monetary conditions of the Eurozone, as captured by the Citigroup index of inflation surprises for Europe.

**Table VII : Determinants of CRP and alternative risk measures**

This table reports results of the regression of the credit-implied risk premium (CRP), and various other risk measures, on economic, financial, and monetary determinants. Models (5-8) reproduce the baseline model (4) using the daily crash risk for the USD/EUR, the daily (debt-weighted) 1-year credit spread in USD for the Eurozone, the daily synthetic quanto-implied risk premium ($QRP^{S}_{ATM}$) based on at-the-money (ATM) options, and the monthly quanto-implied risk premium (QRP) of Kremens and Martin (2019) as the dependent variable. The computation of CRP and crash risk are detailed in Sections 3.3 and 4.4.5 respectively. Section 4.1 provides detailed on the implied volatility and QRP. The sample consists of observations between August 20, 2010, and April 26, 2019.

| | **Determinants of CRP** | | | | **Crash risk** | **Credit risk** | **$QRP^{S}_{ATM}$** | **QRP** |
|---|---|---|---|---|---|---|---|---|
| | Economic growth | Financial uncertainty | Inflation news | ECB communication | | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Uncertainty (VSTOXX) | -0.208*** | -0.202*** | -0.207*** | -0.210*** | 0.092*** | 0.050*** | 0.045*** | 0.047*** |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.001) |
| Economic growth (IP YoY) | | 0.141*** | 0.193*** | 0.186*** | -0.067*** | -0.066*** | 0.019*** | -0.065 |
| | | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.130) |
| Inflation surprise | | | 0.050*** | 0.051*** | 0.002*** | 0.008*** | 0.005*** | 0.019*** |
| | | | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.001) |
| Bund yield (2y) | | | | 0.120 | 0.520*** | 0.487*** | 0.398*** | 1.094*** |
| | | | | (0.395) | (0.000) | (0.000) | (0.000) | (0.000) |
| 2y*ECB monetary news | | | | -1.268** | 0.177 | -0.005 | -0.040 | |
| | | | | (0.028) | (0.287) | (0.964) | (0.470) | |
| 2y*ECB positive news | | | | 1.076 | -0.576*** | -0.037 | -0.067 | |
| | | | | (0.125) | (0.000) | (0.769) | (0.515) | |
| ECB announcement dummy | | | | -0.287 | 0.066 | 0.029 | 0.021 | |
| | | | | (0.349) | (0.278) | (0.489) | (0.417) | |
| Constant | 15.080*** | 14.744*** | 14.833*** | 14.929*** | -0.230*** | -0.382*** | 0.059*** | 0.112 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.007) | (0.000) | (0.746) |
| $R^2$ (%) | 18.3 | 18.8 | 35.1 | 35.3 | 62.6 | 69.4 | 78.4 | 69.3 |
| N | 2225 | 2225 | 2225 | 2225 | 2225 | 2225 | 2225 | 67 |

## G Bootstrap algorithm

Our bootstrap algorithm follows Mark (1995) and Kilian (1999) and imposes the null of no predictability to generate the critical values for our out-of-sample test statistics. This procedure consists of the following steps:

1. Given the sequence of observations for $\{r_t, x_t\}$, define the out-of-sample window and generate $M$ out-of-sample forecasts by running the predictive regressions

$$r_{t+k} = a_k + b_k x_t + \varepsilon_{t+k}$$

both under the null (i.e., $b_k = 0$) and the alternative. For each horizon $k$, compute the statistic of interest $\widehat{\tau}_k$.

2. The data generating process for $\{r_t, x_t\}$ under the null is assumed to be

$$\begin{aligned} r_t &= a + u_{1,t} \\ x_t &= \phi_0 + \phi_1 x_{t-1} + \ldots + \phi_p x_{t-p} + u_{2,t}, \end{aligned}$$

where the lag order $p$ is determined by a suitable lag order selection criterion such as the Bayesian information criterion (BIC). Estimate this specification using the full sample of observations via least-squares, and store the estimates $\widehat{a}$, $\widehat{\phi}_0, \ldots, \widehat{\phi}_p$, and the residual residuals $\widehat{u}_t = (\widehat{u}_{1,t}, \widehat{u}_{2,t})'$.

3. Generate a sequence of pseudo-observations $\{r_t^\star, x_t^\star\}$ of the same length as the original data series $\{r_t, x_t\}$ as follows:

$$\begin{aligned} r_t^\star &= \widehat{a} + u_{1,t}^\star \\ x_t^\star &= \widehat{\phi}_0 + \widehat{\phi}_1 x_{t-1}^\star + \ldots + \widehat{\phi}_p x_{t-p}^\star + u_{2,t}^\star, \end{aligned}$$

where the pseudo-innovation term $u_t^\star = (u_{1,t}^\star, u_{2,t}^\star)'$ is randomly drawn with replacement from the set of observed residuals $\widehat{u}_t = (\widehat{u}_{1,t}, \widehat{u}_{2,t})'$. The initial observations $\left(x_{t-1}^\star, \ldots, x_{t-p}^\star\right)'$ are randomly drawn from the actual data. Repeat this step $B = 1,000$ times.

4. For each of the $B$ bootstrap replications, generate $M$ out-of-sample forecasts by running the

predictive regressions

$$r^{\star}_{t+k} = a^{\star}_{k} + b^{\star}_{k} x^{\star}_{t} + u^{\star}_{1,t+k}$$

both under the null and the alternative. For each horizon $k$, construct the test statistic of interest $\widehat{\tau}^{\star}_{k}$.

5. Compute the one-sided $p$-value as follows

$$p\text{-value} = \frac{1}{B} \sum_{j=1}^{B} I(\widehat{\tau}^{\star}_{k} > \widehat{\tau}_{k}),$$

where $I(\cdot)$ denotes an indicator function, which is equal to $1$ when its argument is true and $0$ otherwise.