Sovereign credit risk and exchange rates: Evidence from CDS quanto spreads[*]

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Abstract

Sovereign credit default swap quanto spreads tell us how financial markets view the interaction between a country's likelihood of default and associated currency devaluations (the Twin Ds). A no-arbitrage model applied to the term structure of eurozone quanto spreads can isolate the Twin Ds and can gauge the associated risk premiums. Conditional on the occurrence of default, the true and risk-adjusted one-week probabilities of devaluation are 42% (2%) and 90% (55%) for the core (periphery) countries. The weekly risk premium for euro devaluation in case of default for the core (periphery) exceeds the regular currency premium by up to 18 (13) basis points.

JEL classification codes: C1, E43, E44, G12, G15

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1. Introduction

The risk of sovereign default and exchange rate fluctuations are inextricably linked. The depreciation of a country's currency is often a reflection of poor economic conditions. Default events tend to be associated with currency devaluations. Such devaluations may either strategically support the competitiveness of the domestic economy or penalize a country's growth due to increased borrowing costs or reduced access to international capital markets.

Despite the importance of studying the Twin Ds (default and devaluation), the subject has received relatively limited attention in the literature. In our view, one reason for this is that it is exceedingly difficult to measure the interaction between the two. Indeed, both types of events are rare, so few data are available for researchers to use. Disagreement over measurements are a clear manifestation of the problem. For example, Reinhart (2002) estimates the probability of devaluation conditional on default at 84%, while Na et al. (2018) find it to be 48%. Another hard-to-measure dimension of the Twin Ds is whether default has an immediate or long-term impact on the exchange rate. Krugman (1979) argues that default leads to a change in the expected depreciation rate (change in exchange rate), while Na et al. (2018) suggest a one-off drop in the exchange rate at default.

In this study, we take advantage of a recent development in financial markets to offer an asset-pricing perspective on the measurement of the risk premiums associated with the Twin Ds. Specifically, sovereign credit default swap (CDS) contracts are available in different currency denominations starting from August 2010. For example, contracts that protect against a credit event in Germany could be denominated in euros (EUR) or US dollars (USD) (we use the terms default and credit event interchangeably). The difference between the two respective CDS premiums of the same maturity, also known as quanto CDS spread, reflects the market's view about the interaction between the Twin Ds. Quanto spreads of different maturities are informative about the interaction over different horizons.

We focus on quanto spreads in the 17 European countries that share the euro as the common exchange rate. We find the associated CDS markets particularly interesting for two reasons. First, despite the common currency and monetary policy, the term structures of quanto spreads are different, suggesting intriguing implications regarding the different fiscal policies. Second, the contractual arrangements of the European contracts make observed quanto spreads transparent in terms of the implications for the Twin Ds. Specifically, in contrast to contracts for emerging markets (EM), payouts on Western European sovereign CDS of all denominations are triggered irrespective of whether a default is associated with domestic or foreign debt.

We start by explaining the institutional arrangements behind sovereign CDS in different currencies. We then explain how quanto spreads are related to the interaction between the Twin Ds. Next, we develop a model that allows us to characterize the relation between default and devaluation probabilities.

Describing the joint behavior of 17 different term structures and an exchange rate is a daunting task. Model-free analysis is impossible because the nonparametric description of conditional joint distributions is famously unreliable even in bivariate cases. Therefore, we develop a no-arbitrage affine term structure model that allows us to evaluate whether it is possible to reconcile all of this evidence within a tractable specification. Along the way, we encounter a typical problem in the analysis of credit-sensitive financial instruments: as the realization of credit events is rare, we can identify only the risk-adjusted distribution of these events based on asset prices. We exploit (i) the interaction between credit risk and currency risk and (ii) the currency commonality to identify the true distribution of credit events. As a bonus, our approach allows us to identify the loss given default (LGD), which we assume to be constant and the same for all countries. While this assumption is an oversimplification, it nevertheless offers progress on a thorny empirical problem.

The model we propose features the following critical components: a model of the US reference interest rate curve, a model of credit risk, and a model of the spot/forward foreign exchange (FX) curve. We use overnight indexed swap (OIS) rates as a reference curve and construct a two-factor model to capture its dynamics.

The starting point for our credit risk model is a credit event whose arrival is controlled by a doubly stochastic Cox process, a popular modeling device in the literature. The default intensity in each country is controlled by two factors — global and regional — which are weighted differently for each country. Given that our data are from countries in the eurozone during the sovereign debt crisis, we derive an extension of our model to allow for the possibility of credit contagion: the occurrence of a credit event in one country affecting the probability of a credit event in another country.

Last but not least, we model the behavior of the spot USD/EUR FX rate and the associated forward rates. We follow the literature on realistic modeling of the time series on FX rates. We allow for time-varying expected changes in the depreciation rate, heteroskedastic regular shocks to the rates, and jumps.

We connect jumps in the FX rate to sovereign credit risk by requiring them to take place simultaneously with credit events. Presumably, a credit event in Germany would be more impactful on the FX rate than a credit event in Spain. We accommodate this intuition by distinguishing between jumps in response to core or periphery credit events.

That FX jumps coincide with credit events is our identifying assumption that establishes the true distribution of credit events. To enhance the statistical reliability of our estimates, we follow Bai et al. (2015) and complement our identification strategy by associating realized credit events with extreme movements in quanto spreads.

We use joint data on the term structure of the quanto spreads of six countries, which represent the most liquid CDS contracts across the European core and periphery, some data on Greece prior to its credit event, the spot and forward FX rates, and a cross-section of credit events for the six countries. We estimate the model via the Bayesian Markov Chain

Monte Carlo (MCMC) method. The model offers an accurate fit to the data. It also fits the quanto spreads of the remaining 10 eurozone countries that were not used in the estimation.

We find that a more parsimonious model without contagion fits the data just as well and does not differ significantly from the larger model in terms of its likelihood. Therefore, we perform the rest of the analysis using the simpler model, which features only four factors: three credit factors (one global and two regional) and one FX variance factor.

We find a substantial cross-sectional variation in how the credit risks of countries load on the credit factors. For instance, exposure to the global factor varies from a high of 2 times the German level (for Spain) to a low of 1/16 times (for Finland). Finland is unambiguously the least risky country. There are multiple candidates, primarily from Southern Europe, for the most risky. Some countries, both core and peripheral, have significant exposure to the global factor only.

As we conjectured earlier, a large part of this cross-sectional variation is driven by differences in the fiscal policies as measured by a country's debt. A fixed effect regression of hazard rates on the debt-to-GDP ratio implies that a one percentage point increase in the ratio of a given country is, on average, associated with an increase in its CDS premium of approximately 14 basis points (bps).

Our setting allows for the estimation of the true distribution of credit events. As a result, we can characterize the time-varying credit risk premium, which is typically measured by the ratio of the risk-adjusted to the true default intensity. We find that the credit risk premium is about two, on average. This estimate is consistent with earlier studies of corporate credit risk that were estimating a constant risk premium. We also find evidence for a significant amount of variation over time, as the credit risk premium ranges between one and eight during our sample period.

Our estimated model is also informative about the measurement issues highlighted earlier. For instance, we find the true one-week probability of devaluation conditional on default to be 42% and 2% for core and periphery countries, respectively. That is consistent with the view by Na et al. (2018) of a large drop in the exchange rate upon default. In addition, we find that the expected depreciation rate is unrelated to credit factors, contradicting the view of Krugman (1979).

The risk-adjusted probability of devaluation conditional on default is 90% and 55% for core and periphery countries, respectively, suggesting a hefty risk premium for this event. Indeed, the model-implied risk premiums for exposure to the Twin Ds exceed the regular currency risk premiums by as much as 18 bps per week. Thus, the default-contingent currency risk premium is significantly larger than either the default or the currency risk premiums (when considered separately), which is suggestive of the large economic importance of that risk. All of this evidence suggests that default-contingent devaluation, although not highly probable, occurs during the worst states of the economy when the marginal utility of investors is particularly high.

This study is most closely related to the literature on the relation between sovereign credit and currency risks. Della-Corte et al. (2016) empirically show that the common component in sovereign credit risk correlates with currency depreciations and predicts currency risk premiums. Carr and Wu (2007) propose a joint valuation framework for sovereign CDS and currency options with an empirical application to Mexico and Brazil. Du and Schreger (2016) study the determinants of local currency risk as a distinct component of foreign default risk in EM. Buraschi, Sener, and Menguetuerk (2014) suggest that geographical funding frictions may be responsible for persistent mispricing of EM bonds denominated in EUR and USD.

The most recent development in the joint FX-sovereign risk literature is research on CDS quanto spreads. Mano (2013) proposes a descriptive segmented market model that is consistent with nominal and real exchange rate depreciation upon an exogenous default trigger. DeSantis (2015) uses quanto spreads to construct measures of redenomination risk, while Della-Corte et al. (2019) use quanto spreads to forecast FX excess returns. No-arbitrage term structure models for quanto spreads are proposed by Ehlers and Schönbucher (2004) for Japanese corporate CDS and by Brigo, Pede, and Petrelli (2016) for Italian CDS.

In contemporaneous and independent work, Monfort et al. (2017a) develop a model of sovereign quanto CDS spreads in the eurozone. That study focuses on the modeling implications of credit events for quanto spreads using the gamma-zero distribution (often denoted as ARG_0) but does not focus on the Twin Ds. Lando and Nielsen (2017) focus on the contribution of each of the normal and jump risks to the shape of the quanto term structure in the eurozone and estimate a model of quanto spreads on a country-by-country basis.

In contrast, we exploit the entire term structure of CDS quanto spreads across countries, which ought to be consistent with common exchange rate movements, to pin down the time variation in the risk premia associated with expectations of exchange rate depreciation conditional on default. In addition, using an identifying assumption that exploits the common currency, we estimate both the true and risk-adjusted default intensities. Joint estimation is also necessary for identifying the propagation of shocks across countries (that is, contagion), a subject of public and academic debate.

Methodologically, our paper is related to the vast literature on no-arbitrage affine intensity-based modeling of credit-sensitive instruments, prominently summarized in Duffie and Singleton (2003). Table 1 summarizes the specific modeling elements across the key studies of sovereign credit spreads. The table visually highlights the primary differences between the current study and others. The first 11 papers demonstrate the objects that can be estimated without quanto spreads: the risk-free term structure, exchange rate dynamics, various types of factors driving the default probability, and LGD.

Methodologically, our work encompasses most of the existing approaches and contributes to the literature by analyzing the Twin Ds. That effort requires using data on quanto spreads. Moreover, since we study the exchange rate that is connected to many countries,

we have to model the joint conditional distribution of the EUR exchange rate and default probabilities of several countries. Conceptually, our work departs significantly from other studies due to our analysis of the Twin Ds. Using market prices, we shed new light on a number of debates surrounding the impact of sovereign default on exchange rates (Krugman, 1979; Reinhart, 2002; Na et al., 2018). Importantly, since we study both the true and risk-adjusted default intensities, we can identify the risk premiums associated with CDS quantos and the Twin Ds.

2. Sovereign CDS contracts and currencies

First, we discuss sovereign CDS quanto contracts and then explain the critical difference between their cash flows and those of sovereign bonds. Next, we provide a theoretical interpretation of CDS quantos. We conclude with a discussion of the data.

2.1. Cash flows and settlement

Sovereign CDS are contracts that pay off in case of a sovereign credit event. This section reviews what such an event represents. Given the focus on USD/EUR quantos of eurozone countries, we limit the discussion to the legal details associated with European contracts (see Online Appendix A).

We use S_t to denote the nominal USD/EUR FX rate (amount of USD per EUR) at time t. The first row of Fig. 1 depicts the cash flows associated with a EUR-denominated CDS contract (long protection) with a premium of $€C_0^€$ established at time 0 (the time of the contract's maturity T is omitted for brevity). In this example, the notional is $\$1 = €(1/S_0)$, implying an annualized payment of $€(C_0^€/S_0)$. For the purpose of this discussion, we ignore the quarterly or semi-annual frequency of the premium payment settlement and focus on the annual amounts. The second row shows a USD-denominated contract (short protection) with a premium of $\$C_0^\$$. Given the same notional, the annual payment is $\$C_0^\$$, which is equivalent to $€(C_0^\$/S_t)$ at the spot exchange rates on the settlement dates.

We highlight three implications of the rules that are particularly relevant for this paper. First, a credit event that affects all CDS contracts regardless of the currency of denomination could be triggered by a default pertaining to a subset of bonds, such as a sovereign defaulting on domestic debt but not on bonds issued in other jurisdictions. Therefore, a CDS quanto spread would not reflect the risk of selective default. This is in contrast to EM bonds, studied by Du and Schreger (2016), where differences between the credit spreads denominated in USD and local currency could reflect such a risk, and to EM CDS contracts, studied by Mano (2013), for which a credit event is not triggered by default on domestic currency or domestic law bonds.

Second, recovery is free of any exchange rate consideration, a point also made by Ehlers and Schönbucher (2004). That is because of two important features of the settlement

procedure, which are hardwired in all CDS contracts since 2009. First, an obligation is deemed deliverable into the contract settlement regardless of its currency of denomination or that of the CDS contract. This means that one and the same bond could be delivered into the settlements of CDS contracts of different denominations. Second, the bond auction associated with the CDS settlement determines the value of all deliverable bonds in terms of percentage of the notional value. Incidentally, the auction-established bond value is used for cash settlement of CDS contracts. Therefore, physical and cash settlements are economically equivalent. See Chernov, Gorbenko, and Makarov (2013).

As an extreme example, imagine a European sovereign that has 1% of all its debt issued in EUR and the rest issued in USD. Both EUR- and USD-denominated contracts would be triggered in the case of selective default on the small amount of outstanding EUR-denominated debt. Because one can deliver a bond of any denomination into a contract of any denomination, the cash value of payments in case of a credit event is an identical fraction, denoted by L (LGD) in Fig. 1, of a contract's notional amount. We fix L to be a constant, in line with the literature on CDS pricing (Pan and Singleton, 2008). This removes any uncertainty about nominal payments in the respective currencies.

The only uncertainty that arises with a credit event is due to expressing these payments in the same currency. For example, the EUR-denominated contract pays a certain amount of L/S_0, while the EUR value of the payment of the USD-denominated contract is L/S_τ, with τ denoting the time of a credit event. This comparison illustrates the impact of devaluation. If the EUR devalues during the credit event, S_τ is lower than S_0, and as a result, the value of the payment on the EUR-denominated contract is lower than that on the USD-denominated one.

Third, outright default is only one of the scenarios that may trigger a credit event. A common concern among observers of the eurozone credit market is that a CDS payout would be triggered by a restructuring of a eurozone member country's liabilities through a redenomination of the principal or interest payments into the country's pre-EUR currency. There are two CDS definitions, those of 2003 and 2014, which treat this event differently. For contracts based on the 2003 definitions, redenomination does not trigger a credit event as long as it involves the currencies of the G7 countries and AAA-rated Organisation for Economic Co-operation and Development (OECD) economies. The newer definitions limit these currencies to the lawful currencies of Canada, Japan, Switzerland, the United Kingdom, the United States, and the eurozone.

After the 2014 definitions were introduced, CDS could be concurrently traded satisfying either definition. We use data on the contracts guided by the 2003 definitions throughout our full sample to ensure internal consistency. Kremens (2018) explores the differences in premiums for the two types of contracts. Regardless of the definition used, some countries might trigger a credit event by redenomination. Our framework automatically accounts for that because our credit factors are designed to capture the probability of a credit event without discerning its type.

2.2. Relation to sovereign bonds

Historically, eurozone sovereign CDS were denominated in USD only. One point worth emphasizing is that EUR-denominated CDS contracts that were introduced in 2010 are not redundant securities. One might think of cash flows as similar to those of a portfolio that consists of a USD-denominated bond and a currency swap (e.g., Du and Schreger, 2016). However, such a strategy does not hedge the behavior of the FX rate conditional on default — the risk that is the focus of this paper.

To see this, consider the third and fourth rows of Fig. 1, which show cash flows to a sovereign par bond and a currency swap of matching maturity. The bond is USD denominated, but the issuer prefers a EUR exposure, hence the additional swap position. The cash flows are presented per \$1 of face/notional values.

At time 0, the issuer sells the bond for \$1 and swaps this amount for the equivalent value in EUR, €$(1/S_0)$. Prior to maturity, the issuer has to pay interest of \$$C_0$ on a bond, which it receives from the swap as interest on the USD value of the notional. In exchange, the issuer has to pay interest of €(F_0/S_0) on the EUR value of the swap, with F_0 denoting the currency swap rate that is determined at time 0.

If there is no credit event, the last transaction takes place at maturity, where the issuer has to repay \$$(1 + C_0)$ on the bond and the EUR/USD notionals combined with the last interest payments that are exchanged in the swap transaction. As a result, the combined position has a pure EUR exposure with predetermined cash flows: €$(1/S_0)$ at inception, $-$€(F_0/S_0) thereafter (including the day of maturity), and $-$€$(1/S_0)$ at maturity. This is conceptually similar to cash flows on the EUR-denominated sovereign CDS if there is no credit event (the difference is in the latter being an unfunded instrument).

If there is a credit event, then the bond-swap combination faces uncertain cash flows between the time of the event τ and maturity T. The bond pays \$$(1 - L)$ at τ and ceases to exist. In the meantime, the swap does not terminate (assuming no counterparty risk) and continues the exchange of cash flows. As a result, the values of the combined cash flows in EUR are $(L + C_0 - 1)/S_\tau - F_0/S_0$ at τ and $C_0/S_t - F_0/S_0$ for $\tau < t \leq T$. In contrast, the EUR-denominated sovereign CDS has a single cash flow of €(L/S_0).

One important implication of this discussion is that quanto CDS spreads contain unique information about the joint distribution of default and devaluation, the Twin Ds. That information can be obtained neither from portfolios of bonds denominated in different currencies (because of selective default), nor from portfolios of bonds denominated in one currency combined with FX swaps, nor single-currency CDS contracts. That is why we use quanto CDS spreads to characterize the properties of the Twin Ds.

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2.3. Interpretation of the quanto CDS spread

Fig. 1 allows us to take a first step toward thinking about quanto CDS spreads. Adding up the two positions described in the first two rows (long EUR-denominated and short USD-denominated protection) gives an exposure to devaluation conditional on default because the only uncertain cash flow is the EUR value of L in the USD contract. Specifically, by buying the EUR-denominated protection and selling the USD-denominated one, we construct a position that benefits if there is no EUR devaluation conditional on default. The payout €$L(S_0^{-1} - S_\tau^{-1})$ is negative if $S_\tau < S_0$.

CDS contracts have a fixed time to maturity. Thus, we observe premiums for new contracts in every period. In other words, we only get to see the difference between the premiums €$(C_0^\$ - C_0^€)/S_0$. Because the choice of notional is arbitrary, the quanto spread $C_0^\$ - C_0^€$ becomes the relevant premium for exposure to the Twin Ds.

To streamline the analytical interpretation of the quanto CDS spread, consider a hypothetical contract that trades all points upfront, meaning a protection buyer pays the entire premium at time t. Further, assume that the risk-free rate is constant. Then, the EUR-denominated CDS premium simplifies to

$$C_0^€ = L \cdot E_0[M_{0,\tau} I\left(\tau \leq T\right) S_{\tau \wedge T}/S_0] \equiv L \cdot E_0^*[e^{-r(\tau \wedge T)} I\left(\tau \leq T\right) S_{\tau \wedge T}/S_0],$$

where M denotes the USD-denominated pricing kernel, $*$ refers to the risk-adjusted probability, $I(\cdot)$ is an indicator variable, and $\wedge$ denotes the smallest of the two variables. The USD-denominated premium is similar but without the terms with S. As a result, the relative quanto spread is

$$
\begin{aligned}
\frac{C_0^\$ - C_0^€}{C_0^\$} &= \frac{E_0^*[e^{-r(\tau \wedge T)} I\left(\tau \leq T\right)\left(1 - S_{\tau \wedge T}/S_0\right)]}{E_0^*[e^{-r(\tau \wedge T)} I\left(\tau \leq T\right)]} \\
&= E_0^*\left[1 - \frac{S_{\tau \wedge T}}{S_0}\right] - cov_0^*\left[\frac{e^{-r(\tau \wedge T)} I\left(\tau \leq T\right)}{E_0^* e^{-r(\tau \wedge T)} I\left(\tau \leq T\right)}, \frac{S_{\tau \wedge T}}{S_0}\right].
\end{aligned}
\tag{1}
$$

The first term reflects the risk-adjusted expected currency depreciation, conditional on a credit event (a positive number corresponds to EUR devaluation). The second term is the risk-adjusted covariance between the depreciation rate up to the contract's maturity or default time and the return on a hypothetical asset that pays \$1 in case of default prior to time T. Intuitively, the covariance should be negative, as currency devalues precisely in the states where the hypothetical asset pays off.

Kremens and Martin (2019) refer to the covariance term as the quanto-implied FX risk premium in the context of predicting FX excess returns. They show that this term represents a fraction of the FX risk premium irrespective of the asset whose returns covary with the depreciation rate. A judicious choice of an asset makes the quanto-implied risk

premium the main fraction. Kremens and Martin (2019) argue for using quanto forward contracts on the S&P 500 as such an asset. Della-Corte et al. (2019) extend their framework and use quanto CDS to forecast FX excess returns. Our focus is not on predicting currency excess returns but on the interaction between FX and default risks.

2.4. Data

As we emphasized, quanto spreads are required for the analysis of the Twin Ds. However, we cannot value quanto spreads without a model that correctly values the term structure of risk-free interest rates, exchange rates, forward currency rates, and assets sensitive to credit risk. That is why we have to introduce a large panel of different types of financial assets into our analysis. In this section we describe our data sources and basic properties of these assets.

2.4.1. CDS

Sovereign CDS contracts became widely available in multiple currencies in 2010. This determines the beginning of our sample, which runs from August 20, 2010 to September 26, 2018. We source daily CDS premiums denominated in USD and EUR from Markit for all 19 eurozone countries. We require a minimum of 365 days of nonmissing information on USD/EUR quanto spreads. This requirement excludes Malta and Luxembourg. Thus, our sample features 17 countries: Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Latvia, Lithuania, Netherlands, Portugal, Slovakia, Slovenia, and Spain. We use weekly data on contracts with maturities of 1, 3, 5, 7, 10, and 15 years.

The five-year contract is the most liquid, but liquidity variation across the term structure is less of a concern for sovereign CDS spreads than for corporate CDS spreads, as trading is more evenly spread across the maturity spectrum (Pan and Singleton, 2008). There might also be differences in liquidity across the currency of denomination because an insurance payment in EUR would probably be less valuable if Germany defaulted. Consistent with this view, USD-denominated contracts tend to be more liquid, as shown in Table 2, which reports the average number of dealers quoting such contracts in either EUR or USD over time; that is, CDS depth (Qiu and Yu, 2012). EUR CDS contracts are quoted by 2.98 to 6.53 dealers, on average, which is economically meaningful given that the CDS market is largely concentrated among a handful of dealers (Giglio, 2014; Siriwardane, 2019). The average difference between the number of USD and EUR dealers ranges between 0.18 and 2.08. That suggests that the difference in liquidity of the two types of contracts is not large. Finally, as the contracts guided by the 2014 definitions are being phased in, decreasing liquidity in contracts guided by the 2003 definitions could be a concern. However, we observe no decrease in the number of dealers quoting the 2003 contracts after 2014.

Notional amounts outstanding, also reported in Table 2, offer a sense of the cross-sectional variation in the size of the market. Regardless of the currency of denomination,

the notionals are converted into USD and are reported on the gross and net basis. To facilitate comparison, we express these numbers as a percentage of the respective quantities for Italy, which has the largest gross and net notionals (Augustin et al., 2018). The amounts for France, Germany, and Spain stand out as a fraction of Italy's amounts, followed by Austria, Belgium, and Portugal.

As a further gauge of the size of the market for sovereign CDS, we compare the gross notional amounts outstanding to the aggregate market size. Statistics from the Bank for International Settlements suggest that sovereign CDS represented $1.458 trillion, about 17% of the entire market, in 2018. The corporate CDS market, with single- and multi-name contracts, accounted for the rest of the market.

To provide a first feel for the data, Fig. 2A displays the time series of the five-year USD-denominated CDS contract for Germany, arguably one of the least likely countries to default. We highlight important events to help frame the magnitudes of CDS premiums. The premium for such a safe country is about 20 bps during the calmest periods, varies over time, and exceeds 100 bps during the sovereign crisis.

2.4.2. Quanto spreads

Table 3 provides basic summary statistics for the quanto spreads. There is a significant amount of both cross-sectional and time-series variation in the spreads. In the cross-section, the average quanto spread ranges from 6 bps for Estonia to 77 bps for Greece, at the five-year maturity. The average quanto slope, defined as the difference between the ten-year and one-year quanto spreads, ranges from -23 bps for Greece to 25 bps for France. Overall, both the level and slope of CDS quanto spreads vary significantly over time in each country.

We limit the estimation of our model to data on six sovereigns because of parameter proliferation. We choose the countries that exhibit the greatest market liquidity and the fewest missing observations. In addition, we incorporate both peripheral and core countries that feature the greatest variation in the average term structure of CDS quanto spreads. This leads us to focus on Germany, Belgium, France, Ireland, Italy, and Spain. Fig. 3A plots the average quanto term spreads of different maturities for these countries. We also use a limited amount of data on Greece, as described later in this section. We use data on the remaining countries to conduct an out-of-sample evaluation of our model.

2.4.3. Exchange rate

We collect the time series of the USD/EUR FX rate from the Federal Reserve Bank of St. Louis Economic Database (FRED) and match it with the quanto data, using weekly exchange rates, sampled every Wednesday. Fig. 2B and 3B display the exchange rate and (log) depreciation rate, respectively. A broad devaluation of the EUR is evident in Fig. 2B.

Fig. 3B suggests that the exchange rate movements were close to being independent and identically distributed (iid) during our sample period. Motivated by that, Online Appendix B shows that the term structure of credit premiums is flat if both the default intensity and depreciation rates are iid but correlated with each other. This result establishes a useful benchmark for interpreting the evidence summarized in Fig. 3A.

CDS premiums and exchange rates are interrelated and move together over time. This is summarized by the cross-correlogram between the first principal component (PC) of changes in five-year CDS premiums and depreciation rates in Fig. 2C. The figure highlights that this interaction is primarily contemporaneous and not readily visible in leads and lags.

Fig. 2D stays with Germany to show an example of time-series variation in (five-year) quanto CDS spreads. As explained in Section 2.3, the spreads reflect the interaction between the Twin Ds and are informative about the covariance risk and the currency jump at default risk. A model is needed to understand the interactions over different horizons and to extract the quanto risk premium.

2.4.4. Interest rates

To develop the model, we also need information on the term structure of US interest rates. Prior to the global financial crisis (GFC) of 2008/2009, it was common practice to use Libor and swap rates as the closest approximation to risk-free lending rates in the interdealer market (Feldhütter and Lando, 2008). Since the GFC, practitioners have shifted toward full collateralization and started using OIS rates as a reference interest rate curve (Hull and White, 2013). We source data on OIS rates for maturities of 0.25, 0.5, 0.75, 1, 3, 5, 7, 10, and 15 years.[1]

Once we have a model of the joint behavior of interest rates and exchange rates, it has implications for forward exchange rates. To discipline our model, we use weekly Thomson Reuters data on forward exchange rates obtained from Datastream. We use the Wednesday quotes for our analysis to match the Wednesday OIS rates and quanto spreads focusing on the maturities of one week and one month. Our analysis does not require the European interest rate values; hence we can avoid addressing the important analysis of covered interest rate parity (CIP) violations in Du, Tepper, and Verdelhan (2018). CIP holds in our model by no-arbitrage. Therefore, the implied foreign interest rate could be viewed as an implicit foreign bank funding rate. Such an interpretation is valid in the light of research focusing on various market frictions leading to violations of CIP in terms of true Libor rates (e.g., Borio et al., 2016).

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[1]The 7- and 15-year OIS rates data start only on May 16, 2012 and September 28, 2011, respectively, so we estimate the term structure of interest rates with partially missing information for these dates.

2.4.5. Credit events

Our last piece of evidence pertains to the true occurrence of credit events. True default information is insufficient for estimating conditional credit event probabilities because realized credit events are rare. This issue is common to the literature on credit-sensitive instruments. When modeling corporate defaults, it is possible to infer information about true conditional default probabilities by grouping companies by their credit rating, as done by Driessen (2005), for instance.

We are considering high quality sovereign names, so we have only the credit event in Greece in our sample. Formal defaults are often avoided because of bailouts, as was witnessed multiple times during the sovereign debt crisis (Greece, Ireland, Portugal, Spain, Cyprus). These bailouts result in large movements in credit spreads, although no formal credit event occurred. At the same time, events that are clearly credit quality related, such as the credit event in Greece or credit ratings changes, are anticipated by movements of CDS premiums. That credit spreads reflect information about default probabilities in a more accurate and timely fashion is formally demonstrated for sovereigns by Ismailescu and Kazemi (2010) and Michaelides et al. (2015). Indeed, Lando (2004), in section 5.13, emphasizes "the event which makes it clear to investors that there is no hope may well come as a surprise and it is this date that is reflected in market prices, not the 'legal' default date."

Therefore, we associate credit events with extreme movements in quanto spreads (see also Bai et al. (2015) for high quality corporate bond spreads). Specifically, we deem a credit event to have occurred if a weekly (Wednesday to Wednesday) change in the five-year quanto spread is above the 99th percentile of the country-specific distribution of quanto spread changes.[2] A figure in Online Appendix H displays the observed credit events identified in this way for the 16 countries in our sample other than Greece.

Although Greece experienced a formal credit event, it is difficult to use the full available series for pragmatic reasons (see also Ait-Sahalia et al., 2014). As a figure in Online Appendix H shows, the Greek CDS premium jumped to 5,062 bps on September 13, 2011. It exceeded the 10,000 bps threshold, which is equivalent to 100% of the insured face value, on February 15, 2012. All of that happened before the official declaration of the credit event on March 9, 2012.

In our view, this is a manifestation of legal risk associated with the credit event determination for the purposes of CDS settlement. At present, there is poor understanding of the incentives of the determination committee participants, ten from the sell side and five from the buy side, and how that may affect the decision of whether a credit event took place or not.

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[2]We discuss robustness of our conclusions with respect to our definition of credit events in Section 4.3.

Further, trading of Greece CDS spreads was halted between March 8, 2012 and June 10, 2013, so the time series exhibits a long gap in quoted premiums. Assuming the Markit-reported aggregation of quoted spreads was tradable, Greek CDS trading restarted on June 10, 2013 at a level of 978 bps. The corresponding quanto spreads displayed in the same figure exhibit similar swings in magnitudes and gaps in the data. These data problems create severe credit risk identification issues and make it difficult to study the joint behavior of credit factors across countries.

As a result, we designate September 7, 2011 as a credit event instead of the official one. After that day, the available data are unusable. This period is too short to identify the Greek default hazard rate throughout the full sample. Therefore, we also use some data from the period when the trading resumed in June 2013. Out of concern that very large premiums may not be reflective of true traded prices, we use only those premiums that are within 150% of the maximum of premiums with a corresponding maturity among the remaining GIIPS (Greece, Italy, Ireland, Portugal, Spain) countries. The omitted premiums are so large that the estimation results are not sensitive to the choice of the maximum cut-off point.

3. The model

In this section, we present a no-arbitrage model of the joint dynamics of US interest rates, USD/EUR FX rate, forward FX rates, and CDS quanto spreads for eurozone countries. In broad strokes, the key part of the model is the connection between a devaluation of the FX rate and a sovereign credit event. As we highlighted in Section 2.4, modeling of that connection cannot be achieved without capturing the properties of related assets. That is why our model has to accommodate all these other assets. If our purpose was to model risk-free rates, currency rates, and credit risk without regard for Twin Ds, the quanto data would not have been required.

Mathematically, we model the arrival of a credit event via a Poisson process. Credit hazard rates feature one common component that is linked to Germany and regional components. Furthermore, we allow for default contagion effects. To connect devaluation to credit risk, we make the identifying assumption that jumps in the FX rate can take place only if one of the eurozone sovereigns experiences a credit event. This assumption is motivated by Fig. 2C and links sovereign default hazard rates to the FX Poisson arrival rate.

3.1. Pricing kernel

Suppose, $M_{t,t+1}$ is the USD-denominated nominal pricing kernel. We can value a cash flow, $\mathcal{X}_{t+1}$, using the pricing kernel via $E_t(M_{t,t+1}\mathcal{X}_{t+1})$, where the expectation is computed

under the true conditional probability $p_{t,t+1}$. Alternatively, we can value the same cash flow using the risk-adjusted approach

$$E_t(M_{t,t+1}\mathcal{X}_{t+1}) = E_t(M_{t,t+1})E_t\left(\frac{M_{t,t+1}}{E_t(M_{t,t+1})}\mathcal{X}_{t+1}\right) = e^{-r_t}E_t^*(\mathcal{X}_{t+1}),$$

where the expectation is computed under the risk-adjusted conditional probability $p_{t,t+1}^*$ and r_t is the risk-free rate at time t. Thus, the pricing kernel connects the two probabilities via $M_{t,t+1}/E_tM_{t,t+1} = e^{-r_t}p_{t,t+1}^*/p_{t,t+1}$. In this paper, we use both valuation approaches interchangeably.

Implicit in this notation is the dependence of all of the objects on the state of the economy x_t. We can generically write

$$x_{t+1} = \mu_{x,t}^* + \Sigma_{x,t}^*\varepsilon_{x,t+1}.$$

In the sequel we describe various elements of x_t and their dynamics.

3.2. CDS valuation

We start with valuation, as it allows us to introduce the key objects that we model in subsequent sections. A CDS contract with time to maturity T has two legs. The premium leg pays the annual CDS premium $C_{t,T}$ until a default takes place at a random time τ. It pays nothing after the default. The protection leg pays a fraction of the face value of debt that is lost in the default and nothing if there is no default before maturity. Accordingly, the present value of fixed payments of the USD-denominated contract that a protection buyer pays is

$$\pi_t^{pb} = C_{t,T}^{\$}\sum_{j=1}^{(T-t)/\Delta}E_t[M_{t,t+j\Delta}I\left(\tau > t + j\Delta\right)], \tag{2}$$

where Δ is the time interval between two successive coupon periods and $I(\cdot)$ is an indicator function that takes on the value one if the condition inside the brackets is met, and zero otherwise. We have omitted accrual payments for notational simplicity but take them into account in the actual estimation of the model.

A protection seller is responsible for any loss L upon default, and thus the net present value of future payments is given by

$$\pi_t^{ps} = L \cdot E_t[M_{t,\tau}I\left(\tau \leq T\right)].$$

The CDS premium $C_{t,T}^{\$}$ is determined by equalizing the values of the two legs. The premium of a EUR-denominated contract is, similarly,

$$C_{t,T}^{\text{€}} = L \cdot \frac{E_t[M_{t,\tau}I\left(\tau \leq T\right)S_{\tau \wedge T}]}{\sum_{j=1}^{(T-t)/\Delta}E_t[M_{t,t+j\Delta}I\left(\tau > t + j\Delta\right)S_{t+j\Delta}]}. \tag{3}$$

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It is helpful to introduce the concepts of survival probabilities and hazard rates to handle the computation of expectations involving indicator functions. The information set $\mathcal{F}_t$ includes all of the available information up to time t, excluding credit events. Let

$$H_t \equiv Prob\left(\tau = t \mid \tau \geq t; \mathcal{F}_t\right)$$

be the conditional instantaneous default probability of a given reference entity at day t, also known as the hazard rate. Furthermore, let

$$P_t \equiv Prob\left(\tau > t \mid \mathcal{F}_t\right)$$

be the time-t survival probability, conditional on no earlier default up to and including time t. P_t is related to the hazard rate H_t via

$$P_t = P_0 \prod_{j=1}^{t} \left(1 - H_j\right), \quad t \geq 1. \tag{4}$$

Applying the law of iterated expectations to both the numerator and denominator, we can rewrite the CDS premium as

$$C_{t,T}^{\text{\euro}} = L \cdot \frac{\sum_{j=1}^{T-t} E_t[M_{t,t+j}(P_{t+j-1} - P_{t+j})S_{t+j}]}{\sum_{j=1}^{(T-t)/\Delta} E_t[M_{t,t+j\Delta}P_{t+j\Delta}S_{t+j\Delta}]}. \tag{5}$$

A similar expression can be obtained for the USD-denominated contract by setting $S_t = 1$.

3.3. Credit risk

The risk-adjusted default hazard rate of each country $k = 1, \cdots, M_c$ is

$$H_t^{*k} = Prob^*(\tau^k = t|\tau^k \geq t; \mathcal{F}_t),$$

where τ^k is the time of the credit event in country k and M_c is the number of countries. We posit that the hazard rate is determined by the default intensity h_t^{*k} as follows:

$$H_t^{*k} = 1 - e^{-h_t^{*k}}, \quad h_t^{*k} = \bar{h}^{*k} + \delta_w^{*k\top} w_t, \tag{6}$$

such that the default intensity is affine in the credit variables w_t, which are elements of the state vector x_t.

We assume that w_t consists of one global and two regional factors, so each intensity h_t^{*k} is a function of all global factors and one of the regional factors. That implies two credit factors per country (one global, w_{0t}, and one regional, core, $w_{1t,}$ or periphery, w_{2t}). This assumption is motivated by a PC analysis that extracts country-specific components from the quanto spreads. The procedure implies that two factors explain around 99% of the

variation in quanto spreads. Furthermore, the PC analysis of the combination of the first two components across all countries implies that the first PC explains 58% of the variation. By assumption, Germany is exposed to the global factor only, as in Ang and Longstaff (2013).

We also consider the possibility of contagion, i.e., the default of one country affecting the default probability of other countries. Anticipating the estimation results, which do not favor such a model, we relegate its description to Online Appendix C to conserve space.

3.4. FX rate

We model the FX rate S_t as the amount of USD per one EUR. The idea of our model is that the (log) depreciation rate should be a linear function of the state x_t and be exposed to two additional shocks. One is a currency-specific normal shock with varying variance v_t, and the other one is an extreme move associated with devaluation. We posit

$$\Delta s_{t+1} = \bar{s}^* + \delta_s^{*\top} x_{t+1} + (\bar{v} + \delta_v v_t)^{1/2} \varepsilon_{s,t+1} - z_{1t+1} - z_{2t+1}. \tag{7}$$

We assume that $\varepsilon_{s,t+1} \sim Normal(0,1)$ is independent of $\varepsilon_{x,t+1}$, and the variance factor v_t is one-dimensional.

A jump z_{it+1} reflects a large negative shock to the exchange rate in the case of a credit event. The index i lines these jumps up with the notation for the core and periphery credit factors w_{1t} and w_{2t}. The distinction between the two regions captures the idea that the impact of a credit event on the currency could be different depending on the region where the event is emanating from.

Specifically, the jump arrival rate j_{it+1} follows a Poisson distribution with an intensity of λ_{it+1}^*, $j_{it+1} \sim \mathcal{P}(\lambda_{it+1}^*)$ and jump size $z_{it+1}|j_{it+1} \sim Gamma(j_{it+1}, \theta_i^*)$. The minus sign in front of z_{it+1} emphasizes that the EUR is devalued in the case of a eurozone sovereign credit event. Because the jump size has a distribution, some of the realized jumps could be severe, reflecting an outright devaluation, and some could be quite modest and similar to the regular innovations in FX.

We link jumps in the FX to the sovereign default risk by assuming that the FX rate jumps only in case of a credit event. Therefore, the FX jump intensity corresponding to a region is equal to the sum of all country-specific default intensities from that region:

$$\lambda_{it}^* = \sum_{k \in \text{region } i} h_t^{*k} = \sum_{k \in \text{region } i} \bar{h}^{*k} + \sum_{k \in \text{region } i} \delta_w^{*k} \cdot w_t.$$

Here $\delta_{-iw}^{*k} = 0$, where "$-i$" denotes the region other than i.

This connection between jump intensity and country-specific default intensities allows us to identify the LGD, L. The USD CDS contracts are informative only about the products

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Lh_t^{*k} and, therefore, their sum over k (e.g., Duffie and Singleton, 1999). The quanto feature and FX forward rates bring in information about the risk-neutral distribution of exchange rates, which does not depend on the LGD. This allows us to identify $\lambda_t^* = \lambda_{1t}^* + \lambda_{2t}^*$. Given our identification assumption that links currency jumps to credit events, we can recover L by dividing $\sum Lh_t^{*k}$ by λ_t^*. This description is offered for the development of intuition. In practice, L is estimated jointly with the other parameters using our likelihood-based procedure.

Most studies on no-arbitrage modeling of credit-sensitive assets do not estimate the LGD separately from the default intensity because of a joint identification problem. Some papers are able to identify the LGD using CDS data because of the recovery of face value assumption (see Pan and Singleton, 2008; Elkamhi, Jacobs, and Pan, 2014). Doshi, Elkamhi, and Ornthanalai (2018) exploit senior and subordinate CDS to identify the LGD. Last, the most recent studies exploit the insensitivity of equity valuation to the LGD (Kuehn, Schreindorfer, and Schulz, 2017; Li, 2017). Our approach is more closely related to the first and the last studies because we assume recovery of face value and that a jump in exchange rate is unrelated to the LGD.

Our assumptions about the constant mean of the FX jump and the link of its intensity to default intensities are likely to be oversimplifications of reality. Ideally, one would want to separate jumps into default-related and default-unrelated components. While it is certainly possible to specify a model like that, it would be impossible to estimate such a model with any reasonable degree of precision. The evidence in Chernov, Graveline, and Zviadadze (2018) suggests that omitting default-unrelated jumps would not lead to large biases. That is because, according to their model, there were few currency jumps in the 21st century.

3.5. States

We assume that if investors were risk neutral, then an $N \times 1$-dimensional multivariate state vector x_{t+1} would evolve according to

$$x_{t+1} = \mu_x^* + \Phi_x^* x_t + \Sigma_{x,t}^* \cdot \varepsilon_{x,t+1},$$

where Φ_x^* is an $N \times N$ matrix with positive diagonal elements and $\Sigma_{x,t}^*$ is an $N \times N$ matrix that is implied by the specification described below. This implies that $\mu_{x,t}^* = \mu_x^* + \Phi_x^* x_t$. The state x_t consists of two subvectors

$$x_t = (u_t^\top, g_t^\top)^\top,$$

where $g_t = (w_t^\top, v_t)^\top$. We describe the role of u_t and offer details about the properties of elements of x_t in what follows.

3.5.1. US interest rate curve

The factor u_t is an 3×1 vector that follows a Gaussian process:

$$u_{t+1} = \mu_u^* + \Phi_u^* u_t + \Sigma_u \cdot \varepsilon_{u,t+1},$$

and $\varepsilon_{u,t+1} \sim \mathcal{N}(0, I)$. The default-free US dollar interest rate (OIS swap rate) is

$$r_t = \bar{r} + \delta_{u1} u_{1t} + \delta_{u2} u_{2t}. \tag{8}$$

We use $u_{3,t}$ for modeling the expected depreciation rate with implications for forward FX rates (described below).

The price of a zero-coupon bond paying one unit of the numeraire n-periods ahead from now satisfies

$$Q_{t,T} = E_t^* B_{t,T-1}, \tag{9}$$

where $B_{t,t+j} = \exp(-\sum_{u=0}^{j} r_{t+u})$. Given the dynamics of the interest rate defined in Eq. (8), bond prices can be solved using standard techniques. See Online Appendix D.1.

3.5.2. Currency forward curve

The factor $u_{3,t}$ allows flexibility in the model to match the forward exchange rates, $F_{t,T} = E_t^* S_T$. Given the dynamics of the exchange rate defined in Eq. (7), forward exchange rates can be solved using standard techniques. See Online Appendix D.2.

3.5.3. Quanto curve

The factor g_t is an autonomous multivariate autoregressive gamma process of size M_g. Each component $i = 1, \cdots, M_g$ follows an autoregressive gamma process, $g_{i,t+1} \sim ARG(\nu_i, \phi_i^{*\top} g_t, c_i^*)$, which can be described as

$$g_{i,t+1} = \nu_i c_i^* + \phi_i^{*\top} g_t + \eta_{i,t+1},$$

where ϕ_i^* is a $M_g \times 1$ vector and $\eta_{i,t+1}$ represents a martingale difference sequence (mean zero shock), with conditional variance given by

$$\mathrm{var}_t \eta_{i,t+1} = \nu_i c_i^{*2} + 2 c_i^* \phi_i^{*\top} g_t,$$

where $c_i^* > 0$ and $\nu_i > 0$ define the scale parameter and the degrees of freedom, respectively. To ensure that the multivariate autoregressive gamma process is stationary, we impose parameter restrictions on the persistence matrix ϕ^* such that the eigenvalues have a modulus smaller than one. See Gourieroux and Jasiak (2006), Le, Singleton, and Dai (2010), and Monfort et al. (2017b). See Online Appendix D.3. for the derivation of the CDS premium from Eq. (5).

3.6. Prices of risk

We have articulated all of the modeling components that are needed for security valuation. To estimate the model, we need the behavior of state variables under the true probability of outcomes. Online Appendix E demonstrates that there exists a pricing kernel that supports a flexible change in the distribution of variables involved in the valuation of securities. Most parameters could be different under the two probabilities. One may recover the evolution of state variables under the objective probability by dropping the asterisks $*$ in the expressions of Section 3.5.

Given the focus on credit events, we highlight how the prices of default risk work in our model. All of the variables that are related to credit events have true (h_t^k, λ_t) and risk-adjusted (h_t^{*k}, λ_t^*) versions because the event risk premium could be time varying. In particular, the true and risk-adjusted counterparts may have a different functional form and a different factor structure. In addition, each of these variables may have different true and risk-adjusted distributions that are related to the respective distributions of the factors that drive them.

The risk-adjusted and true distributions of h_t^{*k} and λ_t^* can be identified from the cross-section and time series of quanto spreads, respectively. The true event frequencies h_t^k and λ_t can be identified only from the realized credit events themselves — a challenge for financial assets of high credit quality. As mentioned earlier, we circumvent this difficulty by associating credit events with extreme movements in quanto spreads.

3.7. Implementation

We jointly use data on the USD/EUR FX rate, forward FX rates, the term structure of OIS interest rates, the term structures of CDS quanto spreads for a cross-section of six countries, a cross-section of credit events for six countries, and some data on Greece to estimate the model. The model is estimated using Bayesian MCMC. See Online Appendix F. The outputs of the procedure are the state variables and parameter estimates.

We make the following identifying restrictions. For the model of the OIS term structure, we follow Dai and Singleton (2000) and Hamilton and Wu (2012) and restrict $\mu_u = 0$, $\Sigma_u = I$, $\delta_{u1} \geq 0$, Φ_u^* lower triangular with real eigenvalues, and $\phi_{u11}^* \geq \phi_{u22}^*$. For the credit and variance models, we impose the mean of the state variables g_t to be equal to one under the true probability to avoid scaling indeterminacy. This implies restrictions on parameters $c_i : c_i \nu_i = 1 - \phi_i^\top \iota$, where ι is a vector of ones.

Because of the large number of parameters, we also impose overidentification restrictions. We allow the global credit factor w_{0t} to affect regional factors but not vice versa. This restriction affects elements ϕ_i. We assume the volatility factor v_t to be autonomous under both probabilities. Both sets of restrictions translate into identical restrictions under the risk-adjusted probability because of the functional form of the risk prices for these factors.

Further, we infer the true jump intensities and means from their risk-adjusted counterparts by assuming that their relative magnitudes do not change with probability: $\lambda_{1t}/\lambda_{2t} = \lambda_{1t}^*/\lambda_{2t}^*$ and $\theta_1/\theta_2 = \theta_1^*/\theta_2^*$. We simply do not have enough observations to reliably identify the two types of rare events separately under the true probability.

4. Results

The results in Sections 4.1 and 4.2 are not specific to our use of quanto spreads. We describe them to show that they are consistent with the extant literature, thereby validating our model. Section 4.3 describes implications of our model for the Twin Ds, something that is critically dependent on our use of quantos in the estimation, as discussed in Section 2.2.

4.1. Model selection and fit

A table in Online Appendix G displays the estimated parameters of the OIS term structure. The model is standard and the parameter values are in line with the literature. The key for this paper, as indicated in Table 4, is that the pricing errors are reasonably small. Therefore, we can use the model to discount cash flows using the risk-adjusted valuation method.

We estimate two versions of the credit model: with and without contagion. The results for the former are discussed in Online Appendix C, as we find only weak statistical support for the presence of contagion. We present the estimated parameters in Online Appendix G.

The peripheral countries in our sample have a much larger weight on the global credit factor than the core countries. France has a smaller loading on the core credit factor than Belgium (Germany's weight is zero by assumption). Spain's weight on the peripheral factor is larger than that of Italy and Greece and similar to that of Ireland.

The loss given default parameter, L, is estimated at 59% with a 90% confidence interval of 48% to 73%. As a reference point, in CDS valuation, it is commonly assumed that this value is 60% for corporate and 75% for EM sovereign bonds. The most comparable estimated numbers are provided in Pan and Singleton (2008) at 23% for Mexico and Turkey and 83% for South Korea. We guess that our estimate would have been lower for core countries, as L is affected by events in the periphery (according to the Greek CDS settlement auction, the true loss was 78.5%).

We find that the expected depreciation rate does not load on credit factors under either probability and the same goes for shocks. The loading of the expected depreciation rate on the variance factor is positive under the true probability and negative under the risk-adjusted one. This implies that, ignoring jumps, the EUR is expected to appreciate/depreciate under the corresponding probabilities.

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The FX jump magnitude for core (peripheral) countries is approximately 3% (1%) under the true probability. Thus, the impact from a default of core countries on the exchange rate is more important than that from a default of peripheral countries. The overall low estimates under the true probability are consistent with the visually mild observed movements in Fig. 3B. Under the risk-adjusted probability, it is 26% and 7% for core and periphery, respectively. That implies a huge risk premium associated with currency devaluation upon a sovereign credit event.

Continuing with Table 4 we see that the overall fit is good, with the root mean square error (RMSE) ranging from two to eight bps. To visualize the fit, we plot the time series of the observed and fitted spreads in Fig. 4. The overall high quality of the fit is evident.

As a final and out-of-sample test of the model's quality, we take the estimated factors and reestimate only the country-specific parameters of the default intensities using data for the remaining ten countries. The resulting fit is displayed in Fig. 5. The overall quality of the fit remains high, suggesting that our model is quite reasonable. The selection procedure of the main sample implies that these countries represent less liquid markets. Therefore, fitting these quanto spreads perfectly is not required. The observed spreads for Cyprus are quite erratic, manifesting the illiquidity and riskiness of this sovereign, so the fit here is the worst.

4.2. Default probabilities

In this Section, we discuss the filtered credit factors, the dynamics and cross-section of risk-adjusted default intensities, and their associated default risk premiums.

4.2.1. Credit factors

Fig. 6 displays the filtered credit factors w_t. Recall that one of the identifying restrictions is that the means of these factors are set to one. This explains the similarity in scale. The time-series patterns are quite different. All factors exhibit large movements in 2012, around the height of the sovereign debt crisis. This period was marked by the downgrades of individual sovereigns, the creation of the European Financial Stability Facility, and the political instability in Greece. The crisis reached its peak when Greece officially defaulted in March 2012. The sharp drop in the factors shortly thereafter is associated with the famous speech by Mario Draghi in July 2012, vowing to do "whatever it takes" to save the EUR. All of the factors persistently decrease thereafter.

The peripheral factor w_{2t}, associated with Italy, Ireland, and Spain, also exhibits substantial variation prior to 2012 and, to a lesser degree, for the rest of the sample period. The core factor w_{1t}, which corresponds to Belgium and France, starts to pick up shortly after the marked increase in the periphery factor w_{2t} but much earlier than the pronounced increase in the common factor w_{0t}. Interestingly, w_{0t} and w_{2t} pick up again at the very end of the sample period, a likely association with Italy's fiscal concerns in 2017-2018.

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4.2.2. Time series of risk-adjusted default intensities

Various combinations of these factors deliver risk-adjusted default intensities for the respective countries via Eq. (6). Fig. 7A displays these intensities. Given the common variation in credit factors during 2012, it is not surprising to see elevated default probabilities during that period, irrespective of the country. The largest one-period default probability is almost 2.5% (default probabilities are approximately equal to intensities for such small numbers). Ireland, Italy, and Spain have clearly distinct patterns of intensities in the post-2012 period, a manifestation of their exposure to the factor w_{3t}. The same three countries exhibit some elevation at the end of 2016 — a result of their extensive exposure to the global factor, which also picks up. Overall, Spain is the riskiest country. Germany does not load on the regional components at all. Hence, Germany has the lowest default risk, and its dynamic properties are distinct from those of other countries. These differences in dynamics manifest themselves in the differences between the average quanto curves that we noted in Fig. 3.

4.2.3. Cross-section of risk-adjusted default intensities

Table 5 summarizes how the default intensities of all countries, including the ten out-of-samples ones, load on the factors. The countries are sorted by their loadings on the global credit factor, δ_{w1}^{*k}. Germany is roughly in the middle. Spain has the largest exposure at slightly more than two times that of Germany, and Finland the smallest at 16^{-1} times.

Overall, the countries exhibit large cross-sectional variation in their exposure to credit factors. Among the core countries, Belgium and France have a sizable exposure to the core factor, while others have nearly zero exposure. Finland appears to be the safest country with tiny exposures to both global and core factors.

On the peripheral side, the relatively recent additions to the eurozone, such as Estonia and Slovakia, have a tiny exposure to region-specific risk, so they are largely indistinguishable from some core countries, including Germany itself. The Southern European countries have a large regional component. Spain and Greece stand out in this respect.

Eurozone quanto spreads are driven by common variation in the exchange rate and monetary policy. Thus, cross-country differences in model-implied default intensities ought to be related to cross-country differences in fiscal health. To test that conjecture, we collect information on quarterly debt-to-GDP and FX reserves-to-debt ratios and relate them to the model-implied hazard rates. We source these data from the Bank for International Settlements and the International Monetary Fund International Financial Statistics database. The data span is from the start of our sample, that is, the third quarter of 2010 up to the second quarter of 2015. All data are measured in USD for comparability.

We report the relation between these fiscal variables and model-implied hazard rates in Table 6. The relation between debt-to-GDP and default intensities is positive and statistically significant at the 1% level, with an R^2 of 17% for this univariate regression. The

coefficient of 0.23 implies that a one percentage point increase in a country's debt-to-GDP ratio (+0.01) is, on average, associated with an 0.0023 increase in intensity. This is economically meaningful. If hazard rates are constant, CDS premiums may be approximated by $L \cdot h_t^*$. Given the estimated L of 59%, this implies an increase in CDS premiums of approximately 14 bps.

For the specification in column (2), we add time fixed effects to further absorb any remaining common variation in hazard rates. Given the strong factor structure in CDS premiums in our model, this adjustment should absorb a significant amount of cross-sectional variation. The R^2 increases to 56%, and the regression coefficient is stable, suggesting a meaningful cross-sectional relation between the model-implied hazard rates and each country's fiscal health.

As an alternative measure of fiscal health, we use a country's FX rate reserves to debt outstanding in columns (3) and (4). The negative regression coefficient implies countries that have debt that is better collateralized by FX rate reserves have lower hazard rates. The economic effect is smaller. With $L = 0.59$, the coefficient of -0.04 implies a two bps lower CDS premium for each additional percentage point in FX reserves relative to debt outstanding. The coefficient is unaltered if we absorb additional common variation through time fixed effects in column (4).

We examine the joint effect of both fiscal variables in columns (5) and (6). Debt-to-GDP appears to be the driving variable. Its coefficient remains stable and significant at the 1% level.

4.2.4. Credit risk premiums

While we cannot characterize the default risk premiums for individual countries, we can do so for the overall credit risk. Fig. 7B and Fig. 7C display the aggregate true and risk-adjusted default intensities $\lambda_{i,t}$ and $\lambda_{i,t}^*$, respectively, for $i \in \{1, 2\}$. We show them on different scales because the true default intensity is much smaller. Visually, they track each other quite closely, suggesting a constant risk premium. However, this is not the case. To illustrate this, we characterize the corresponding risk premium, in line with the literature, via the ratio $\lambda_{i,t}^*/\lambda_{i,t}$, displayed in Fig. 7D. Recall that, by assumption, this ratio is the same for both values of i.

On average, this number is 2.3 and is drifting downward toward the end of the sample. Starting from 2014, the premium is insignificantly different from one. As a benchmark, Driessen (2005) assumes a constant default premium in the context of corporate debt and estimates it to be 2.3. Combining CDS-implied default intensities with Moody's KMV expected default frequencies, Berndt et al. (2018) also find an average default premium of around two for a sample of 93 firms in three industries. The ratios of risk-adjusted to true default intensities, however, exhibit substantial time variation, between one and eight.

Bai et al. (2015) and Gouriéroux, Monfort, and Renne (2014) emphasize that if a model of default intensity is missing the contagion effect, then the ratio of intensities might be overstating the true premium for credit risk. In our case, a version of the model with contagion generates the same average premium of 2.3 and by and large similar dynamics.

4.3. Devaluation contingent on default — the Twin Ds

This section focusses on the magnitude of the probability of the Twin Ds, the duration of their impact, and the corresponding risk premiums.

4.3.1. Probability

The estimation results allow us to weigh in on the debate about whether default has an immediate or long-term impact on the exchange rate. Krugman (1979) suggests that default leads to a change in the drift of the depreciation rate, with a persistent and continuous depreciation of the local currency. Na et al. (2018), in contrast, argue that the FX rate suffers a significant one-time drop upon default.

The aforementioned lack of a significant effect of credit factors on the expected deprecia-tion rate in Eq. (7) implies that default is unlikely to have a long-term effect on currencies. Thus, the estimation results imply that changes in the FX rate are linked to credit risk only through jumps. The jump component is significant, implying an acute and short-term interaction between the Twin Ds.

How likely is the EUR to devalue if one of the sovereigns in the eurozone defaults? The answer requires a more formal definition of devaluation. We associate it with an extreme movement in exchange rates and rely on our model to measure the probability of such an event. Specifically, we define the EUR devaluation as a drop in S_t that is larger than three conditional standard deviations.

The model implies that the true one-week probabilities of devaluation conditional on default are 41.8% and 2.3% for the core and periphery, respectively. The risk-adjusted probabilities of devaluation conditional on default are 89.9% and 55.0%, respectively. These numbers suggest a large risk premium for this event. The risk premium would be high if currency devaluation occurs at particularly bad states of the economy that tend to coincide with a high marginal utility of the representative agent. This interpretation is similar to the equilibrium models of Augustin and Tédongap (2016) and Chernov, Schmid, and Schneider (2019), who characterize the large risk premia demanded by risk-sensitive investors for selling CDS protection on sovereigns in developed economies. In the next section, we directly measure the risk premiums associated with default-contingent devaluation.

The true one-year probability is around 0.01% for both core and periphery, which is smaller than the one-week probability. Thus, devaluation comes primarily through a sharp,

short-term decline in the EUR. This decline in probabilities with the horizon is consistent with a one-off impact of default on the FX rate, as discussed above. This number is also much smaller than those reported by Na et al. (2018) and Reinhart (2002). Part of this difference may be explained by the sample (developed versus emerging economies) and by our primary use of market data rather than realized defaults. The corresponding one-year risk-adjusted probabilities of devaluation conditional on default are 1.2% and 1.0% for the core and periphery, respectively.

We also measure the sensitivity of the probability of devaluation to changes in the probability of default, which we represent by the default intensity. We can view the previous exercise as a particular case in which the default intensity is equal to infinity. Over the one-week horizon, the true (risk-adjusted) devaluation probability increases by 0.31% and 0.02% (0.82% and 0.50%) for every 1% increase in the default intensity for the core and periphery. The small sensitivity of the true probability is a manifestation of the small jump size. Thus, a large increase in the intensity is needed for the jump component to affect devaluation.

Fig. 7B and Fig. 7C show that the largest movements in the true (risk-adjusted) default intensity were equal to 0.3% and 2% (1.8% and 6%), which occurred between December 2011 and June 2012 during the eurozone debt crisis. These increases translates into an increase of 0.09% and 0.04% (1.48% and 3%) in the true (risk-adjusted) probability of devaluation for the core and periphery.

4.3.2. Risk premiums

We start with the case that ignores the timing of default and compute both the risk-adjusted expectation of the depreciation rate, $E_t^*[S_T/S_t]$, and its true counterpart, $E_t[S_T/S_t]$. The first row of Fig. 8 shows that the true expectations indicate expected EUR devaluation, and the average term structure of such expectations is downward sloping (EUR is expected to devalue more at longer horizons). The risk-adjusted expectation shows that the USD is expected to devalue, on average, at horizons of up to three years.

To understand this result, consider covered interest parity, which implies $E_t^*[S_T/S_t] = F_{t,T}/S_t = \exp[(T-t)(y_{t,T} - \widehat{y}_{t,T})]$ with $\widehat{y}_{t,T}$ denoting the EUR benchmark yield. Thus, the expectation is positive whenever $y_{t,T} > \widehat{y}_{t,T}$. The displayed relation is thus pinned down by the currency forward rates $F_{t,T}$ that we used in estimation.

One implication is that we observe the standard currency premium relation, $E_t S_T / E_t^* S_T$, often expressed as the carry trade. Borrowing in the high-interest rate USD, converting to EUR, and then lending in this currency, leads, on average, to negative (log) excess returns at short horizons. At long horizons, (log) excess returns are not significantly different from zero, consistent with the evidence in Lustig, Stathopoulos, and Verdelhan (2016).

The second (third) row of Fig. 8 includes a condition on the timing τ of any credit event within core (periphery) countries rather than the one in a specific country. In this

case, the arrival rate is controlled by $\lambda_{1,t}$ ($\lambda_{2,t}$), and we can compute both the risk-adjusted expectation and its true counterpart. Default risk has almost no impact in the short term, so the plot for $T = 1$ year is similar to that for the no-default case (it is literally identical if $T = 1$ week). The level of risk-adjusted expectations moves upwards with longer maturities in this case, indicating the risk-adjusted expected USD depreciation conditional on default. This happens because early termination leads to a loss in the expected USD appreciation.

In risk premium terms, we observe a drastic difference in relation to the no-default case. Carry returns to EUR are still negative in the short run and continue to be so as the horizon expands. That is a manifestation of the interaction with default risk. No amount of interest earned on a long-term debt instrument can compensate for currency losses experienced in case of default.

Indeed, we can quantify how much more risk premium is required for such a scenario. Expressing risk premiums in logs, $(T - t)^{-1} \log[E_t S_T / E_t^* S_T]$ (no default) and $(T - t)^{-1} \log[E_t S_{t \wedge T} / E_t^* S_{t \wedge T}]$ (default), we can compute the difference between the two. It reflects the extra compensation one earns for the currency loss in case of default, compared with the compensation for the regular currency risk. The fourth row of Fig. 8 displays the average term structure for the core and periphery.

At short horizons this "excess risk premium" is not significantly different from zero for both core and periphery. The same average effects persist for maturities of up to four years.[3] After that, the excess premium for EUR depreciation for the core (periphery) increases to 0.18% (0.13%) at the 15-year horizon and levels off at 0.22% (0.15%) per week at the 30-year horizon (not shown).

4.3.3. Magnitude of risk premiums

We have shown a remarkably large risk premium for the exposure to the Twin Ds. It either manifests itself via large differences between the risk-adjusted and true probabilities of devaluation conditional on default or expected excess returns on hypothetical currency forwards that terminate in case of default. These numbers are large, and liquidity effects could be, in part, responsible for such a magnitude. But even one half of these premiums would still be economically significant. Further, our statements about risk premiums depend on (i) the true probability of outcomes, that is, on our assumptions about what constitutes a realized credit event; and (ii) the risk-adjusted probability of outcomes, that is, on our modeling assumptions.

The definition of credit events introduced in Section 2.4 may bias downwards the coincidence of large drops in the EUR value and credit events. That would result in an

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[3] At first blush, this result appears to be at odds with the large differences in true and risk-adjusted one-week probabilities of devaluation conditional on default that we reported in the previous section. The risk premiums reflect compensation for any adverse movement in EUR. In contrast, the probabilities reflect an extreme event of at least a three standard deviation move.

overstatement of the risk premium's magnitude. We argue that our conclusions are robust in this respect.

The main reason is conceptual. First, we have painstakingly shown that other implications of our model are in line with the existing research. That should give overall credibility to the model. Second, the joint incidence of a stress in the credit markets and sharp movements in the exchange rate is quite low. Thus, variation in the definition cannot really make the true probability of devaluation conditional on default large and thereby lower the risk premium.

We also explicitly investigate an alternative definition of realized credit events that was proposed by our readers. Our main definition assigns the same number of credit events to all countries because we use percentiles. In the new definition, we use an absolute cut-off: occurrences when the log change of a five-year USD-denominated CDS premium is above 0.15 constitute a credit event. The cut-off point is defined as two standard deviations based on log changes of five-year USD-denominated CDS premiums for Germany, Belgium, France, Ireland, Italy, and Spain. This alternative definition for credit events does not materially affect the estimated probabilities of the coincidence of these two types of events.

To summarize, variations in assumptions about actual outcomes do not appear to be important for the magnitude of risk premiums. Another possibility is that the risk-adjusted probability of the Twin Ds is biased upwards because of the modeling assumptions. In this respect, the model we present herein represents a robustness check that supplants the original model.

In an earlier version of the model, we did not distinguish between jumps in core and periphery countries. That model had the counterintuitive implication that the EUR devalues regardless of the type of the country experiencing a credit event. Furthermore, that model implied a much larger gap between the probabilities of devaluation contingent on default: 5% and 77% for the true and risk-adjusted probabilities, respectively. Also, the risk premium for the EUR devaluation in case of default was 50% larger.

4.3.4. Expected devaluation and relative quanto spreads

Several authors, such as Du and Schreger (2016) and Mano (2013), use observed relative quanto spreads, in the context of EM, to measure the anticipated currency devaluation in case of a credit event. As we show in Eq. (1), the relative spread consists of two components: the risk-adjusted expected depreciation conditional on default and the quanto-implied FX risk premium.

To facilitate the discussion, we switch to the ratio of the local and the foreign currency CDS spreads, which has a more straightforward connection to the depreciation rate,

$$\frac{C_{t,T}^{\text{€}}}{C_{t,T}^{\$}} = E_t^* \left[\frac{S_{\tau \wedge T}}{S_t} \right] + cov_t^* \left[\frac{e^{-r(\tau \wedge T - t)} I\left(\tau \leq T\right)}{E_t^* e^{-r(\tau \wedge T - t)} I\left(\tau \leq T\right)}, \frac{S_{\tau \wedge T}}{S_t} \right].$$

This expression makes it explicit that a direct reading of the data, as expected depreciation upon default, is prone to be biased, as it may overstate the importance of the impact of sovereign default on the local currency depreciation. Our model allows us to gauge the relative importance of both terms, both across countries and over time. We illustrate these decompositions in Fig. 9.

As a start, focus on the last column that depicts the time averages of relative quanto spreads across the term structure for the six countries that we use in the main estimation. The black-circled lines display the observed relative CDS spreads, $C_{t,T}^{\text{€}}/C_{t,T}^{\$}$. If one were to take the view that the spread reflects the risk-adjusted expectation of the depreciation rate only, then the figure would tell us that, for one-year contracts, the euro is expected to depreciate, on average, by about 30% for Germany, 28% for France, and up to about 15% for Ireland. For longer horizon contracts, the magnitude of the expected impact is increasing. For the five-year maturity, for example, the expected drop in the local currency given default is as large as 44% in the case of Germany.

The gray lines depict the model-implied default-contingent risk-adjusted expected depreciation rate $E_t^*[S_{\tau \wedge T}/S_t]$. The gap between the two lines emphasizes that the view of that expectation as the observed relative quanto spread is biased. Generally, the model-based risk-adjusted expectation is above the relative quanto spread, implying a negative cov_t^* term in the decomposition, on average. This is intuitive, as one would expect the event of default and the value of the EUR to move in the opposite directions. Quantitatively, the difference ranges from a high of 27% for Germany to a low of 10% for Ireland at short horizons and declines at longer horizons. The decline suggests that market participants perceive the sensitivity of the FX rate to default risk to weaken in the long term.

The cross-country variation in the cov_t^* term suggests that the impact of a credit event in one of the core countries is more important than that in the peripheral ones. The time series plots of relative quanto spreads and expected depreciation rates at different horizons, in the first three columns of Fig. 9, convey more detail about this observation. We focus on the five-year horizon in the second column. The gap between the gray and black lines is increasing in the build-up of the European debt crisis in 2011 and 2012. This suggests that the covariance risk became of greater importance during that period. A sharp change of this pattern is visible in July 2012, when the gap between the two plotted lines narrows significantly in particular for the peripheral countries (Ireland, Italy, and Spain) but not for the core countries (Belgium, France, Germany). This episode coincides with Draghi's speech, vowing to do whatever it takes to save the EUR.

One interpretation of the data is that Draghi was, to some extent, successful in changing the market's perception about the interaction of FX and default risks. This perception has nearly collapsed to zero for peripheral countries but not so for core countries. That difference highlights the systemic role of the core countries and the fear of the impact of their defaults on the EUR currency, irrespective of how unlikely that event may be. Consequently, markets may still expect a sell-off of the EUR in the case of default. Cross-country differences over

time consistently suggest that Germany is expected to have the greatest impact on the exchange rate if it were to default.

5. Conclusion

Sovereign CDS quanto spreads offer a market perspective on the Twin Ds. We study their interactions and the corresponding risk premiums from an asset-pricing perspective. To that end, we propose an affine no-arbitrage model of the joint dynamics of the quanto spreads of 17 eurozone countries and the USD/EUR exchange rate. Controlling for exchange rate risk, credit risk has a parsimonious three-factor structure (global, core, and peripheral), indicating a lot of common movement. Cross-country differences in credit risk are related to a measure of debt, indicating the connection between cross-sectional variation in default risk and variation in fiscal policies.

We find that the probability of devaluation conditional on default is low. But the risk premium for this event is large, much larger than either the credit risk premium or the carry trade return individually. These results pinpoint the economic importance of the Twin Ds, as their occurrence coincides with the worst states of the economy when investors' marginal utility is high.

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Table 1. Affine term structure models of sovereign credit spreads

This table summarizes the main affine term structure models proposed for the pricing of sovereign credit spreads using intensity-based frameworks. We describe the focus of the paper, which can encompass the term structure (TS), foreign exchange rates (FX), and CDS quantos (Quanto). We also indicate the main model components of the default intensity and, if applicable, of the depreciation rate dynamics. We refer to the presence of homoskedastic or heteroskedastic shocks, extreme events, and contagion. We further describe the type of risk factors, which can be country specific, regional, or common. We indicate whether the estimation is done jointly for all countries or on a country-by-country basis. + indicates that the estimation is done pairwise for two countries. The $*$ refers to the fact that the estimation is performed on the short end of the term structure, up to the five-year maturity. $**$ denotes that the estimation considers only two maturity segments, five and ten years. LGD indicates whether a paper estimates the loss given default.

Study	Model focus			Default intensity components			FX components			Factors			Joint est.	LGD
	TS	FX	Quanto	Homoskedastic	Heteroskedastic	Contagion	Homoskedastic	Heteroskedastic	Extr. evts	Country	Common	Region		
Monfort and Renne (2013)				✓						✓				
Duffie, Pedersen, and Singleton (2003)				✓	✓					✓				
Hoerdahl and Tristani (2012)	✓			✓	✓					✓	✓		✓	
Pan and Singleton (2008)	✓			✓						✓				✓
Longstaff et al. (2011)	✓*			✓						✓				
Doshi, Jacobs, and Zurita (2017)	✓			✓						✓	✓		✓	
Zhang (2008)	✓			✓	✓					✓				
Ang and Longstaff (2013)	✓*			✓	✓					✓	✓			
Ait-Sahalia, Laeven, and Pelizzon (2014)	✓**					✓				✓			+	
Benzoni et al. (2015)				✓		✓				✓	✓		✓	
Carr and Wu (2007)	✓*	✓		✓	✓		✓	✓	✓	✓				
Ehlers and Schönbucher (2004)		✓	✓	✓			✓		✓	✓				
Brigo, Pede, and Petrelli (2016)	✓**	✓	✓	✓			✓		✓	✓				
Monfort et al. (2017a)	✓	✓	✓	✓	✓		✓	✓		✓	✓			
Lando and Nielsen (2017)	✓	✓	✓	✓	✓		✓	✓	✓	✓				
The present study	✓	✓	✓	✓	✓	✓	✓	✓	✓		✓	✓	✓	✓

Table 2. Descriptive statistics of sovereign CDS liquidity — weekly

This table reports summary statistics (N, mean, min, max, median) for the depth measure of five-year sovereign CDS spreads denominated in EUR and USD for 17 eurozone countries that have a minimum of 365 days of nonzero information on USD-EUR quanto CDS spreads. A simple t-test for differences in means suggests that differences in means are statistically significant. Depth is defined as the number of dealers used in the computation of the daily mid-market CDS quote. The sample period is August 20, 2010 until September 26, 2018. All statistics are based on weekly (Wednesday) data. We also report the gross and net notional amounts of CDS outstanding in billion USD as well as the ratio of net and gross notional amounts of CDS outstanding to the same quantities of Italy, which represents the most liquid sovereign CDS market. These data are based on weekly information from August 20, 2010 until June 24, 2015. Sources: Markit and Depository Trust and Clearing Corproation (DTCC).

| | N | EUR | | | | USD | | | | USD notionals | | | |
		Mean	Min	p50	Max	Mean	Min	p50	Max	Gross	Gross%	Net	Net%
Austria	418	4.61	2	4	9	5.95	2	6	9	54.37	15.51	5.04	23.35
Belgium	418	4.76	2	5	8	5.41	2	5	10	54.96	15.55	4.50	20.72
Cyprus	418	2.98	2	3	6	3.72	2	3	9	1.88	0.48	0.25	1.24
Estonia	418	3.73	2	3	10	4.18	3	4	10	2.51	0.74	0.27	1.25
Finland	418	3.34	2	3	9	5.43	2	5	9	17.48	4.98	2.17	10.28
France	418	4.90	2	5	9	5.65	2	6	9	136.17	37.96	15.72	73.51
Germany	418	4.06	2	4	8	5.09	2	5	9	121.24	33.97	14.99	70.54
Greece	344	3.35	2	3	7	3.89	2	4	7	48.75	16.97	3.43	13.89
Ireland	418	5.14	2	5	10	6.28	2	6	10	44.36	12.72	2.80	12.82
Italy	418	6.53	2	7	10	6.71	2	7	10	356.35	100.00	21.43	100.00
Latvia	418	4.32	2	4	10	5.34	2	5	10	8.96	2.58	0.50	2.34
Lithuania	418	3.52	2	3	7	4.14	2	4	9	6.53	1.86	0.48	2.24
Netherlands	415	3.63	2	3	8	4.94	2	5	9	28.38	7.94	3.00	14.28
Portugal	418	6.11	3	6	11	7.26	2	7	11	69.48	19.91	4.41	20.14
Slovakia	418	4.38	2	4	11	5.65	3	6	11	10.73	3.06	0.75	3.53
Slovenia	418	3.30	2	3	7	4.13	2	4	9	6.96	1.93	0.68	3.19
Spain	418	5.99	2	6	10	6.33	2	6	10	171.38	48.47	12.52	57.98
All	7,029	4.40	2	4	11	5.32	2	5	11	69.05	19.56	5.64	26.24

Table 3. Descriptive statistics of CDS quanto spreads — weekly

This table reports summary statistics (mean, sd) for the sovereign CDS quanto spreads (difference between the USD- and EUR-denominated CDS spreads) for the 17 eurozone countries that have a minimum of 365 days of nonzero information on USD-EUR CDS quanto spreads. We report values for maturities of 1y, 3y, 5y, 7y, 10y, 15y, and 30y, as well as the slope, defined as the difference between the 10y and 1y quanto spreads. The sample period is August 20, 2010 until September 26, 2018. The data frequency is weekly, based on Wednesday quotes. Source: Markit.

Maturity		1y		3y		5y		7y		10y		15y		30y		10y-1y	
Country	Obs	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
AT	418	7	10	11	14	17	17	22	17	25	17	28	17	30	16	19	11
BE	418	9	15	16	20	23	22	27	22	31	22	32	22	34	22	22	14
CY	418	23	73	24	47	25	37	26	36	28	37	27	36	32	50	5	65
EE	418	3	5	5	5	6	5	6	5	6	5	6	6	7	6	3	4
FI	418	2	2	4	3	7	3	10	4	12	5	13	5	13	6	10	5
FR	418	7	10	14	17	22	22	27	23	31	23	33	23	34	24	25	16
DE	418	3	4	7	8	13	12	17	14	21	14	22	14	23	15	18	12
GR	344	103	393	77	229	77	198	74	194	68	199	60	140	45	104	-23	275
IE	418	17	23	24	25	28	26	30	25	32	24	33	24	34	24	16	12
IT	418	18	18	28	22	33	23	36	22	39	22	41	22	42	22	21	9
LV	418	5	8	8	8	10	8	11	8	13	8	16	9	15	9	9	7
LT	418	5	8	8	8	11	8	14	8	16	8	19	8	20	10	10	5
NL	418	4	6	8	8	14	11	18	12	22	13	23	14	24	15	18	10
PT	418	28	34	32	26	36	23	38	21	40	21	38	20	39	23	12	22
SK	418	5	9	6	7	7	7	8	7	9	8	11	9	12	9	4	8
SI	418	8	12	12	13	14	13	16	13	19	14	20	15	23	17	10	11
ES	418	22	26	32	32	37	33	40	32	43	31	46	30	46	30	21	11
All	7,032	12	64	15	92	18	56	22	50	24	49	26	49	28	36	28	30

Table 4. Model fit

In this table, we report results for the model fit in terms of root mean squared errors (RMSE). For the term structure model, we do not allow for measurement errors for one-year and seven-year maturity bonds in the estimation. We report the RMSE in basis points. The sample period is August 25, 2010 to September 26, 2018. The data frequency is weekly, based on Wednesday rates.

| | Bond yield | Quanto spread | | | | | | |
		Germany	Belgium	France	Ireland	Italy	Spain	Avg across countries
1y	–	1.8	4.6	3.6	9.9	7.8	8.1	6.0
3y	2.3	2.4	6.1	5.1	8.4	6.7	6.3	5.9
5y	1.6	3.3	7.1	5.0	7.6	5.9	6.6	6.0
7y	–	3.4	5.2	4.6	6.9	6.7	6.1	5.5
10y	2.5	4.6	5.2	5.8	6.1	8.1	6.5	6.0
15y	2.6	5.6	5.8	6.9	7.2	9.9	7.6	7.2
Avg across maturity	2.3	3.6	5.7	5.2	7.7	7.6	6.8	–

Table 5. Credit factor loadings

In this table, we report the hazard rate loadings of each country, sorted in the descending order for global loadings. Notation: $**p < 0.05$, $*p < 0.1$. A dash, -, indicates a loading restricted to zero.

Country	Global, δ_{w0}^{*k}	Core, δ_{w1}^{*k}	Periphery, δ_{w2}^{*k}
Spain	0.0036**	–	0.0043**
Italy	0.0022**	–	0.0030**
Cyprus	0.0024*	–	0.0021*
Greece	0.0029*	–	0.0057*
Slovenia	0.0018**	–	0.0010**
Ireland	0.0018**	–	0.0043**
Austria	0.0027**	0.0000	–
Slovakia	0.0004**	–	0.0003
Germany	0.0016**	–	–
Netherlands	0.0002*	0.0004*	–
Lithuania	0.0004*	–	0.0008*
Latvia	0.0003*	–	0.0008*
Portugal	0.0013**	–	0.0046**
Belgium	0.0018**	0.0013**	–
Estonia	0.0002**	–	0.0002**
France	0.0008**	0.0012**	–
Finland	0.0000*	0.0001*	–

Table 6. Default intensities and fiscal health

In this table, we report the regression output from the panel regression of model-implied risk-adjusted default intensities of 17 eurozone countries against their corresponding debt-to-GDP and foreign exchange reserves-to-debt ratios. The data frequency is quarterly. Debt/GDP and FX-Debt ratios are sourced from the Bank for International Settlements and the International Monetary Fund International Financial Statistics database. We have the data from the start of our sample, i.e., third quarter of 2010, up to the second quarter in 2015. All data are measured in USD for comparability. Standard errors are reported in parentheses; ***, **, and * defines statistical significance at the 1%, 5%, or 10% level.

Variable	(1)	(2)	(3)	(4)	(5)	(6)
Debt/GDP	0.23***	0.26***			0.25***	0.28***
	(0.03)	(0.02)			(0.03)	(0.02)
FX res/debt			−0.04***	−0.04***	0.00	0.01
			(0.01)	(0.01)	(0.01)	(0.01)
Constant	0.08***	0.05***	0.28***	0.28***	0.07**	0.05**
	(0.02)	(0.02)	(0.01)	(0.01)	(0.03)	(0.02)
Time FE		✓		✓		✓
N	324	324	307	307	307	307
Observations	324	324	307	307	307	307
R^2	0.17	0.56	0.03	0.37	0.18	0.56

Fig. 1. Cash flows to various sovereign instruments

Dashed boxes emphasize uncertainty of credit events. S_t is \$/€ rate. The notional is $\$1 = €1/S_0$ for all instruments. Loss given default, L, is expressed as a percentage of notional. The first row depicts a €-denominated CDS contract (long protection) with a premium of $€C_0^€$ established at time 0, implying an annual payment of $€C_0^€/S_0$. In case of a credit event, the long position receives $€L/S_0$. The second row shows a \$-denominated CDS contract (short protection) with a premium of $\$C_0^\$$. A quarterly payment is $\$C_0^\$$, which is equivalent to $€C_0^\$/S_t$ at the spot exchange rates. In case of a credit event, the short position pays out $\$L = €L/S_\tau$. The third row shows a short position in the \$-denominated par bond. It pays out coupons $\$C_0$, and in case of a credit event it pays out recovery $1 - L$. The fourth row shows a currency swap with a rate F_0. After exchanging the same notional expressed in different currencies at inception, one party pays interest on the €-denominated notional, $€F_0/S_0$, and the other party pays interest on the \$-denominated notional $\$C_0$. At maturity, the parties swap back the notionals and pay the last interest.
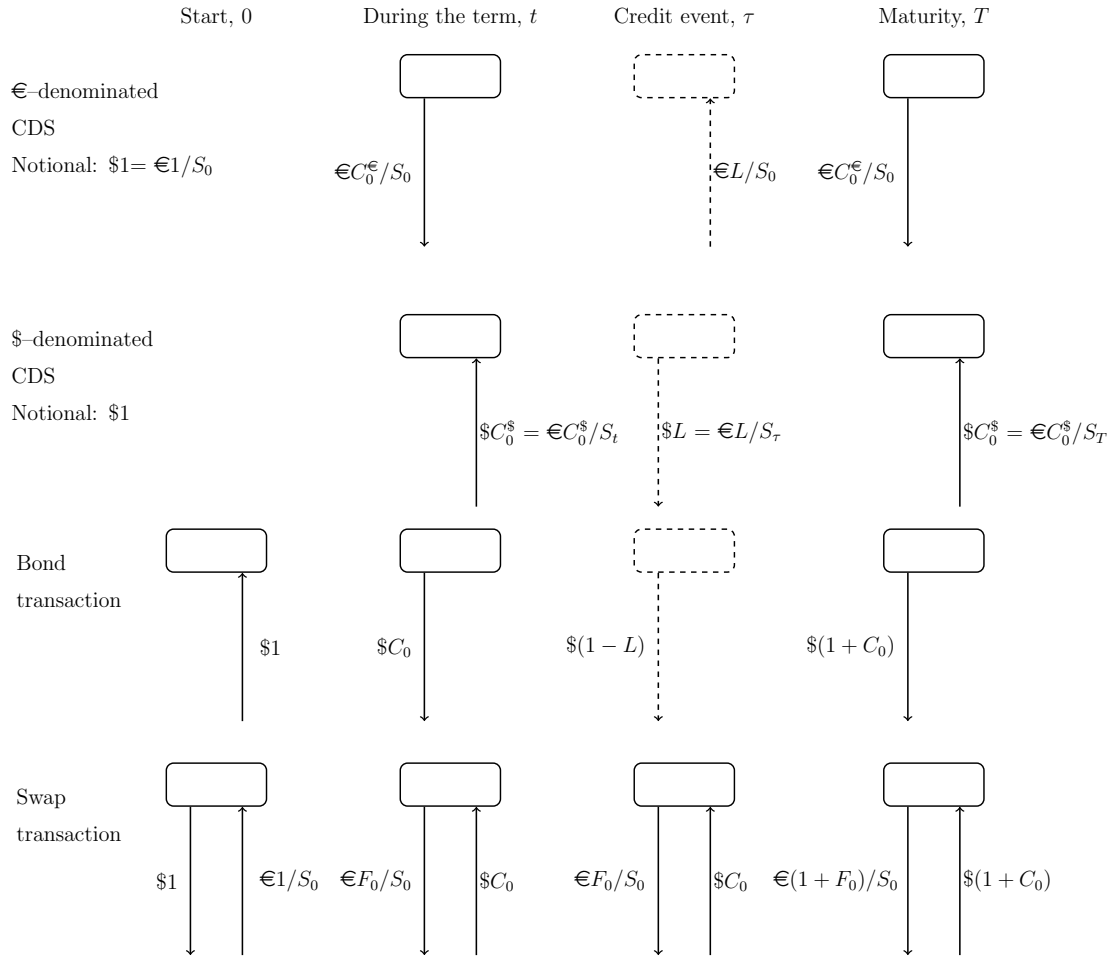
Fig. 2. Data: A first look

We offer a first look at some elements of data: CDS premiums (Panel A) and quanto spreads (Panel D), exchange rate (Panel B), and cross-correlogram between the first principal component of changes in five-year CDS premiums across all countries and the depreciation rate (Panel C).
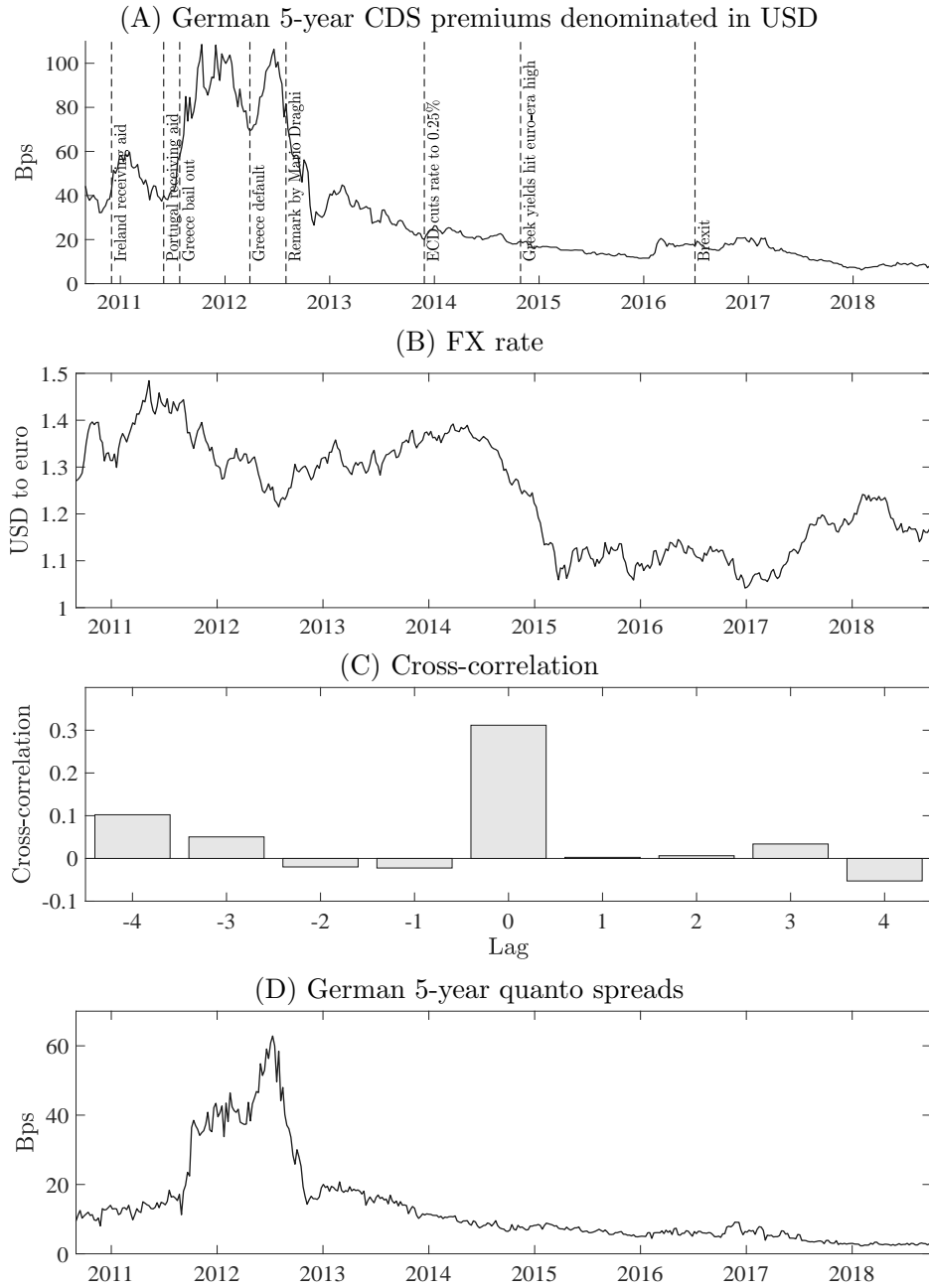
(A) German 5-year CDS premiums denominated in USD

(B) FX rate

(C) Cross-correlation

(D) German 5-year quanto spreads

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Figure 3. Quanto term spreads and exchange rate

Panel A displays the term structure of CDS quanto spreads, defined as the difference between the USD- and EUR- denominated CDS spreads, for 6 eurozone countries: Germany , Belgium, France, Ireland, Italy, and Spain. We compute average term spreads of maturities 3 years, 5 years, 7 years, 10 years, 15 years, and 30 years relative to the one-year quanto spread. All spreads are expressed in basis points. Panel B shows weekly Wednesday-to-Wednesday changes in the USD/EUR exchange rate, expressed in percentage point changes. The sample period is August 25, 2010 to September 26, 2018.

(A) Average term structures of quanto spreads

(B) Depreciation rate (log change in FX rates)
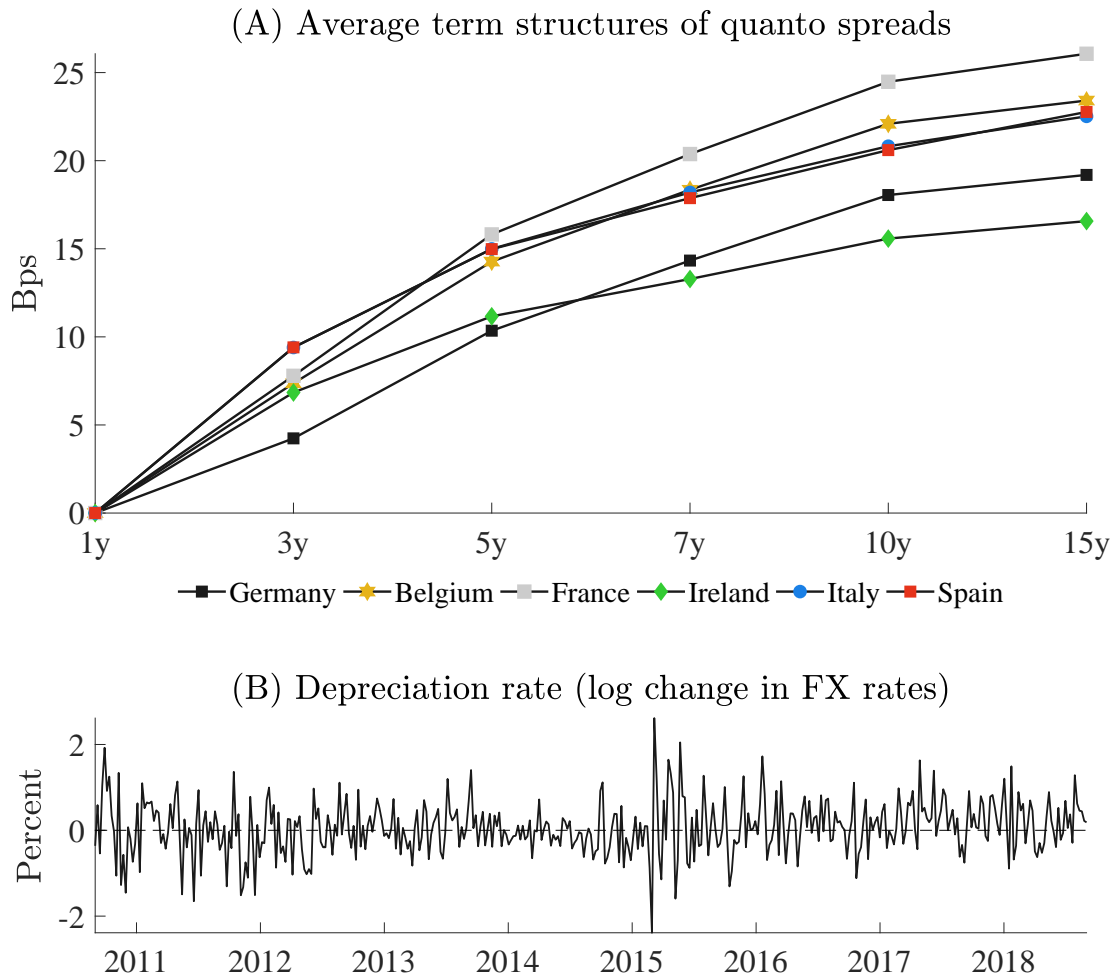
Fig. 4. Quanto spreads: main sample

In these figures, we plot the observed and model-implied USD/EUR quanto spreads for Germany, Belgium, France, Ireland, Italy, and Spain. We report values for maturities of 1y, 5y, 10y, and 15y. Gray lines represent posterior medians of quanto spreads, and gray-shaded areas correspond to 90% credible intervals. The true quanto spreads are plotted with black-circled lines. The sample period is August 25, 2010 to September 26, 2018.
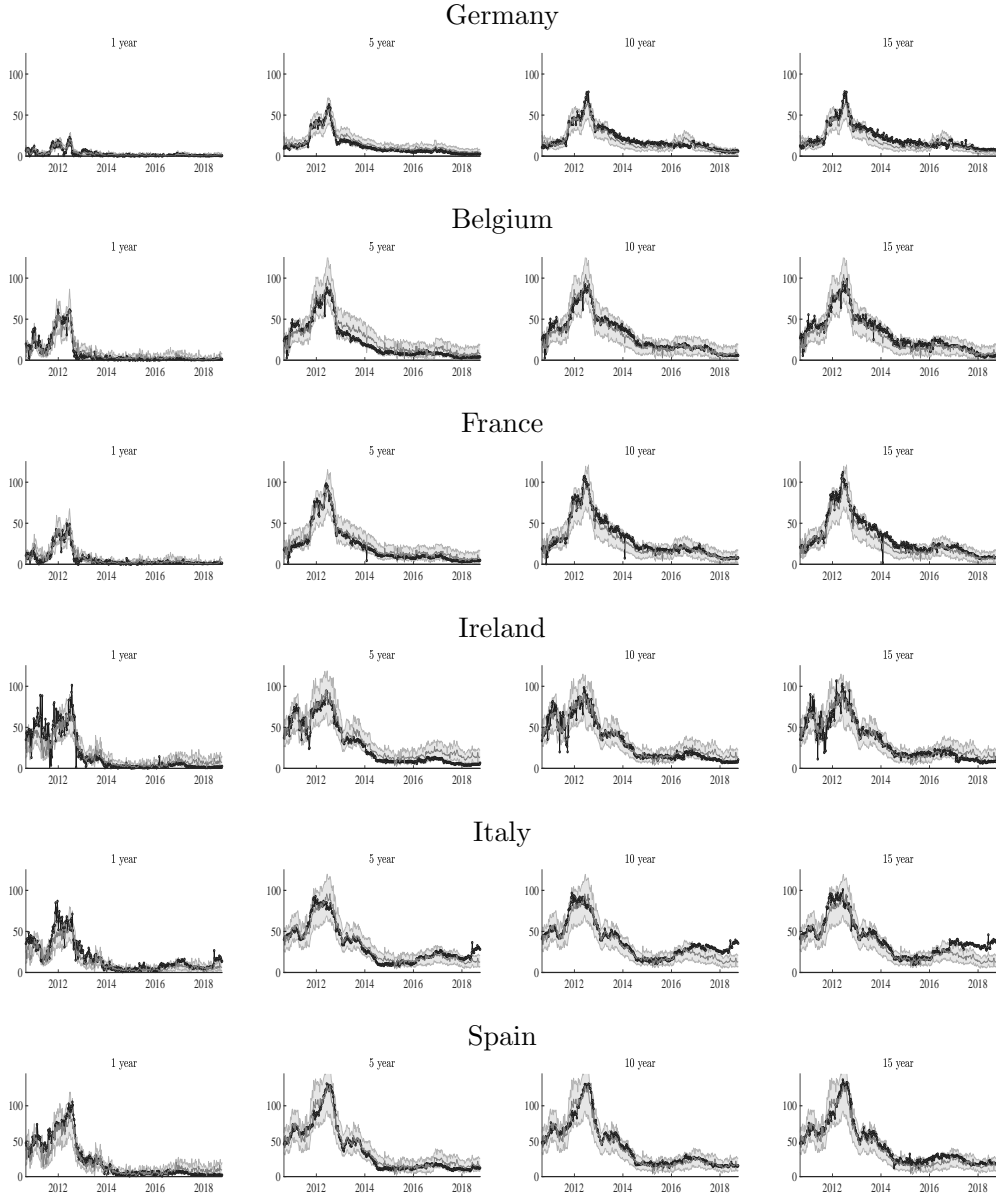

Germany

Belgium

France

Ireland

Italy

Spain

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Fig. 5. Quanto spreads: auxiliary sample

In these figures, we plot the observed and model-implied USD/EUR quanto spreads for Austria, Cyprus, Estonia, Finland, Latvia, Lithuania, Netherlands, Portugal, Slovakia, and Slovenia. We report values for maturities of 1y, 5y, and 10y. Black lines represent posterior medians of quanto spreads and the true quanto spreads are plotted with white-circled lines. The sample period is August 25, 2010 to September 26, 2018.
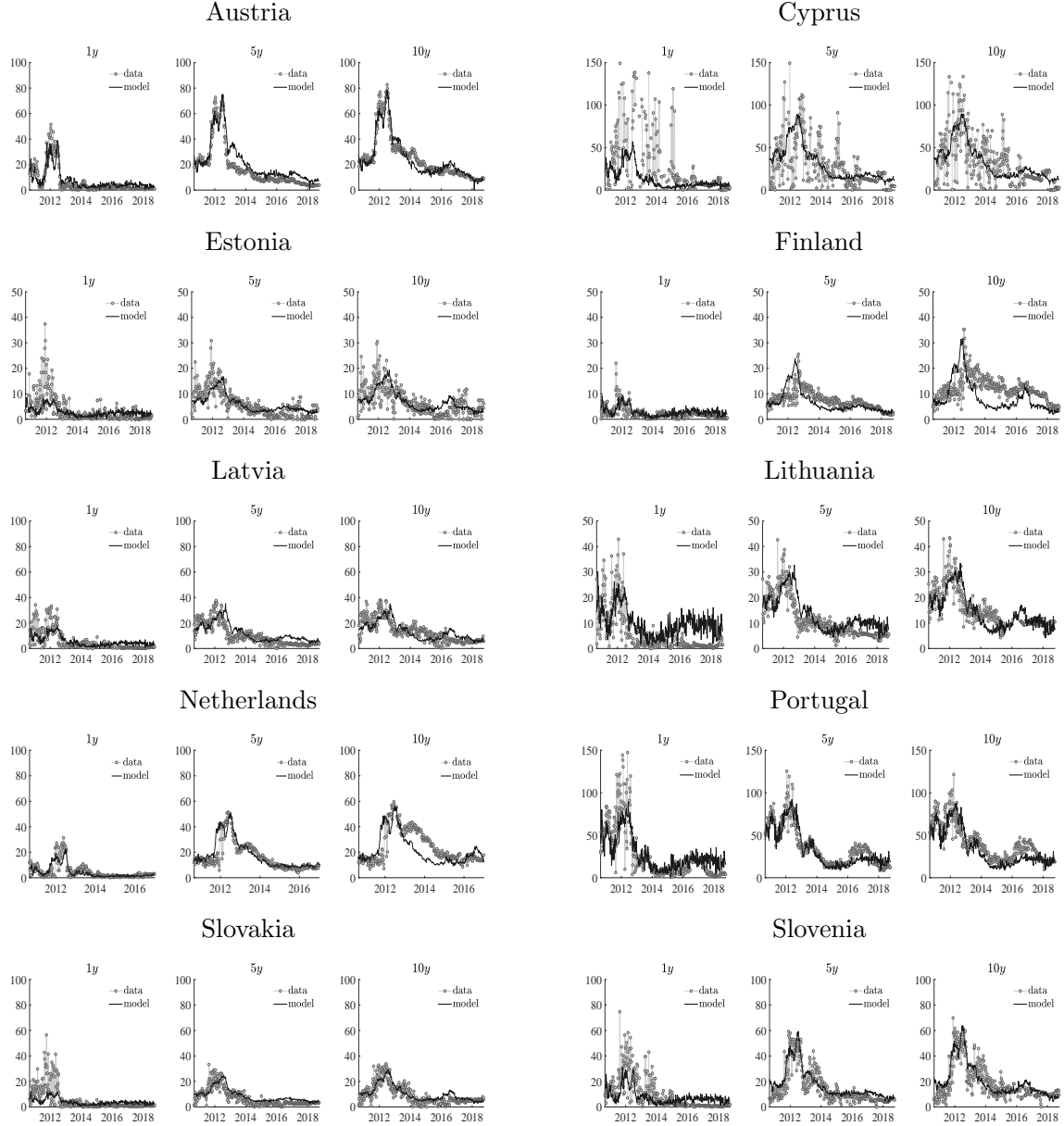

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Fig. 6. States

These figures depict the filtered latent credit and volatility factors implied by the model. w_{1t} is the global credit factor, w_{2t} is a regional credit factor corresponding to the core countries in the eurozone (Belgium and France), and w_{3t} is a regional credit factor corresponding to the peripheral countries (Ireland, Italy, and Spain). The sample period is August 25, 2010 to September 26, 2018. The data frequency is weekly.
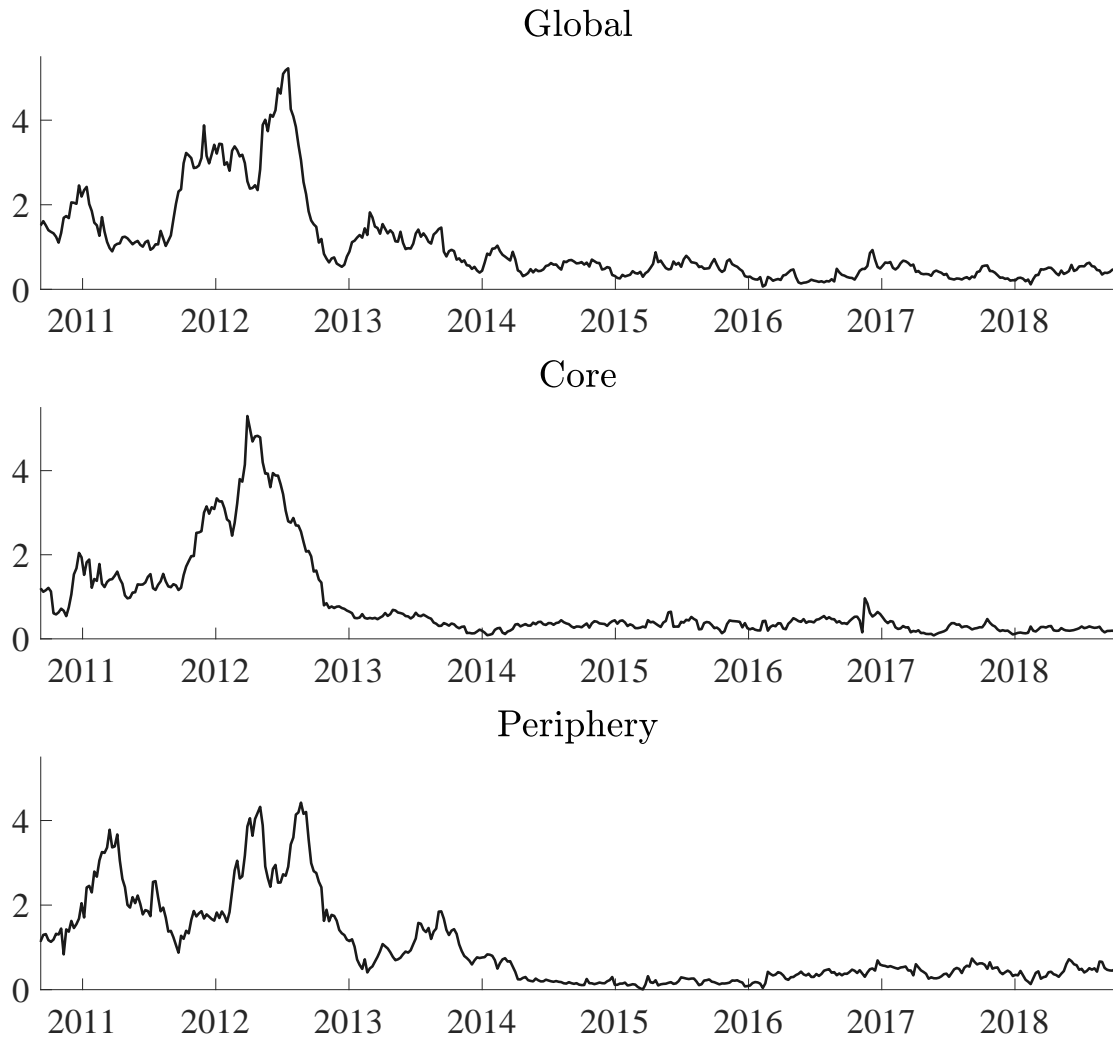

Global

Core

Periphery

Fig. 7. Default intensities

Panel A depicts the estimated risk-adjusted default intensities for Germany, Belgium, France, Ireland, Italy, and Spain. In Panel B (Panel C), we plot the aggregate true and risk-adjusted default intensities for core (periphery) countries. In Panel D, we plot the default risk premium. The dashed horizontal line indicates a zero premium (ratio of intensities is equal to 1). The gray-shaded area represents 90% credible intervals. The average premium is around 2.5. The sample period is August 25, 2010 to September 26, 2018. The data frequency is weekly, based on Wednesday quotes.

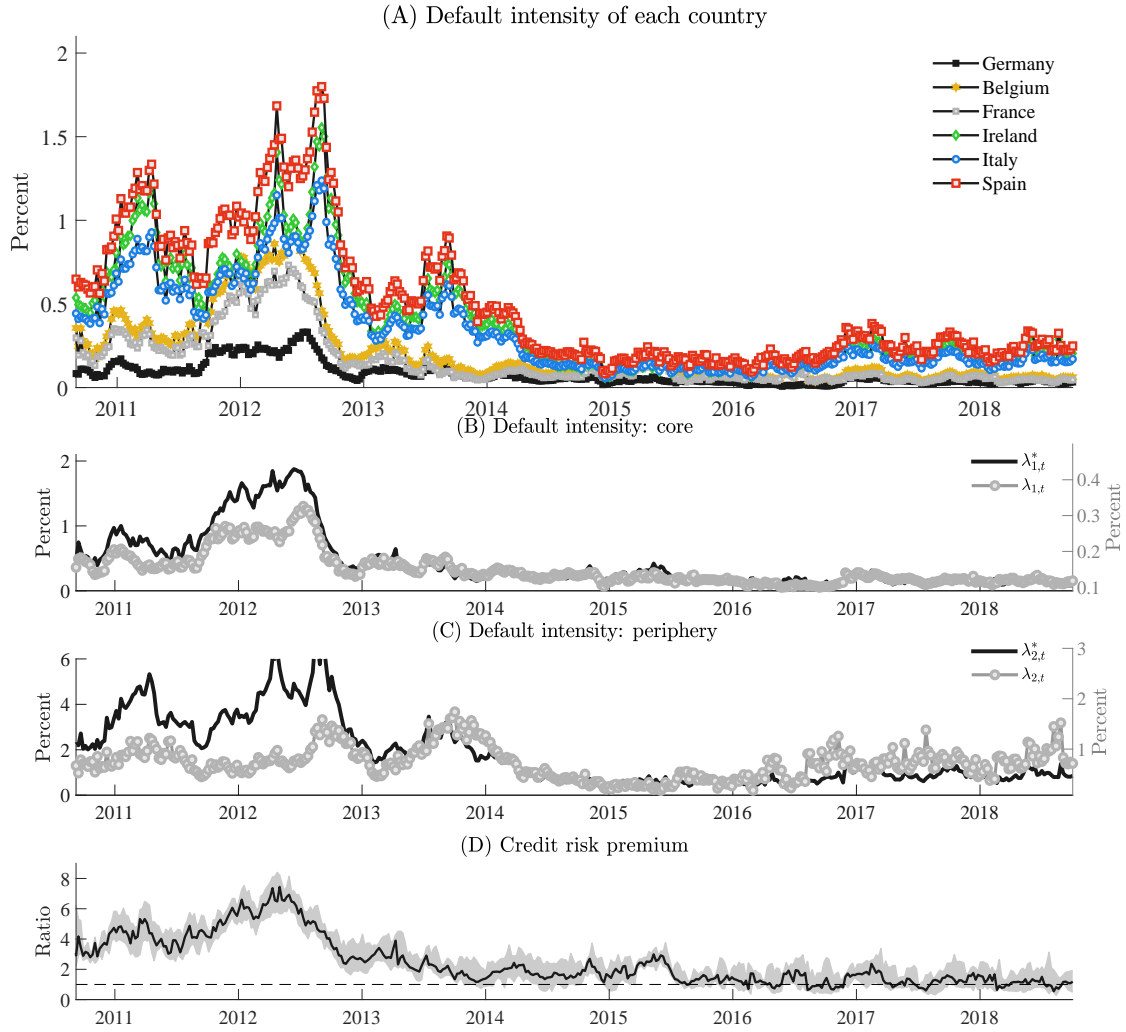(A) Default intensity of each country

Fig. 8. True and risk-adjusted expectations of the depreciation rate

The figures depict the true (light gray) and risk-adjusted (black) expectations of the depreciation rate. The first row shows the expectations ignoring default, $E_t[S_T/S_t]$ and $E_t^*[S_T/S_t]$. The second row accounts for default, $E_t[S_{\tau \wedge T}/S_t]$ and $E_t^*[S_{\tau \wedge T}/S_t]$. The credit event time τ is triggered by a credit event in any of the countries. The left and middle panels show the time series for $T = 1$ year and 5 years, respectively. The right panels display time-series averages of these quantities for a variety of T's. The last row computes the difference between excess log expected returns corresponding to the two scenarios: $(T - t)^{-1} \log[E_t S_{\tau \wedge T}/E_t^* S_{\tau \wedge T} \cdot E_t^* S_T / E_t S_T]$.

Expected depreciation without default

Expected depreciation with default (Core)

Expected depreciation with default (Periphery)

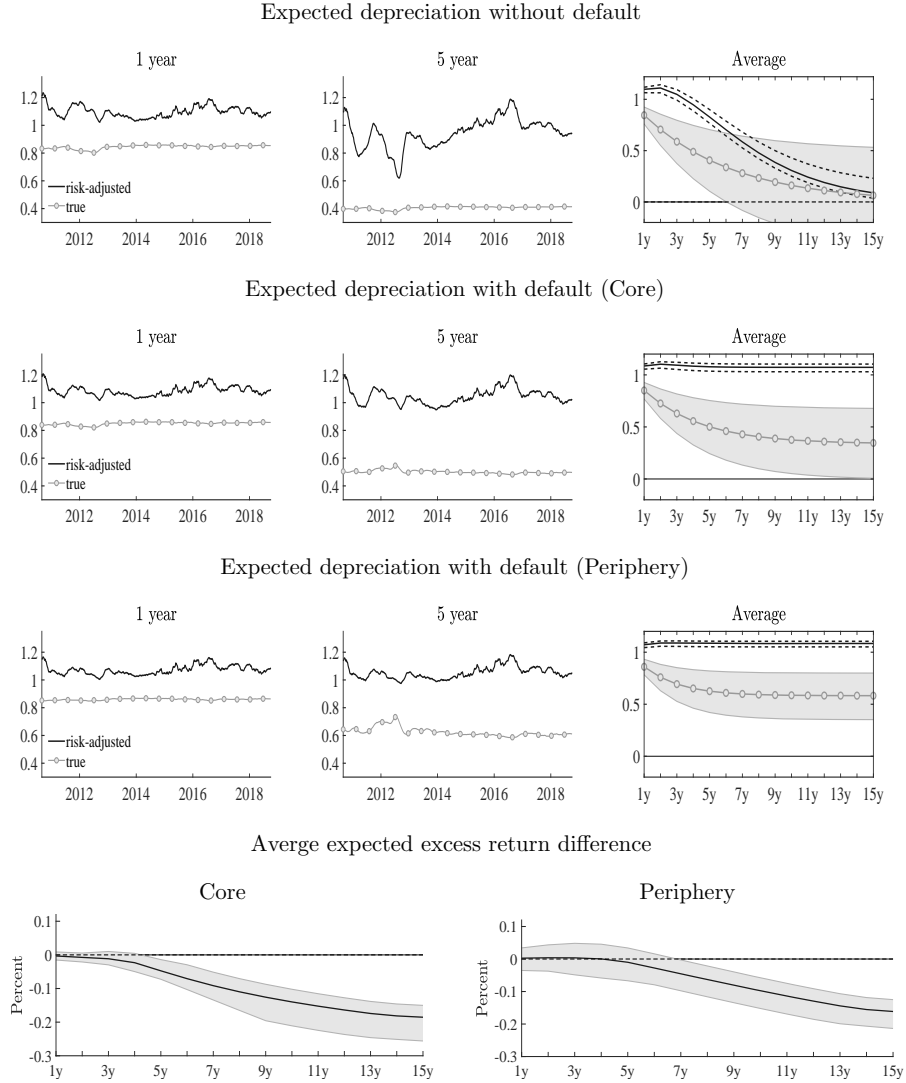Averge expected excess return difference

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Fig. 9. Relative quanto spread and expected depreciation rate

In this figure, we plot the observed relative quanto spreads (black-circled lines) and the model-implied expected depreciation rate $E_t^*[S_{\tau \wedge T}/S_t]$ for $T = 1, 5,$ and 10 years, together with their sample averages in the last column (gray-shaded areas correspond to 90% credible intervals). The results are for Germany, Belgium, France, Ireland, Italy, and Spain.

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