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Pricing anomaly at the first sight: same borrower in different currencies faces different credit spreads—an explanation by means of a quanto option

Andreas W. Rathgeber · David Rudolph · Stefan Stöckl

Abstract Can the credit spreads of one and the same issuer differ in two different currencies? If so, can an investor exploit this situation? To answer these questions and to contribute to the existing literature, we extend the Jarrow/Turnbull-model with a second currency, price a quanto option, and test the theoretical results with an extensive empirical study. A major result of the study was the key insight that the credit spreads, and therefore the cumulated implied default probabilities of nearly all bonds denominated in USD in comparison to EUR denominated bonds, are significantly higher for all terms, and are mostly driven by the correlation between default risk and exchange rate.

Keywords Credit spreads · Efficient market hypothesis · Foreign currency government bonds · Implied default probabilities · Term structure of interest rate

JEL Classification G12 · G13 · G15

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1 Introduction

Credit spreads represent both credit risk for a debtor and opportunity for an investor. What if an investor could choose between two credit spreads in two different currencies from the same debtor?

It is a widely held practice to use credit spreads of one issuer to valuate bonds, for example, without taking into account the existence of bonds from the same issuer denominated in different currencies. However, this approach is only correct if we assume the credit spreads of one issuer do not correlate with the corresponding currency, i.e. continuous credit spreads are equal across all currencies and therefore risk-neutral implied default probabilities are also equal.

However, this is only true in a normally distributed framework and an arbitrage-free market, specifically in the case of non-correlation between credit spread and risk-free interest rate on the one hand and between credit spread and exchange rate on the other hand, as shown by Jankowitsch and Pichler (2005). In case these conditions are not fulfilled an option theoretic approach is required.

Since Garman and Kohlhagen (1983) many publications are pricing currency options in the presence of stochastic interest rates (e.g., Amin and Jarrow 1991) as well as in the presence of stochastic volatility (e.g., Melino and Turnbull 1990). Among these approaches Heston's (1993) model is one special case, in which stochastic interest rates are perfectly correlated with stochastic exchange rates. Bakshi and Chen (1997) relax these assumptions and present a model where stochastic interest rates as well as exchange rates are endogenously determined by modelling the monetary supply and output of real production as driving factors. Hence, it allows any correlation between interest rates and exchange rates as well as the existence of volatility risk. However, these models include no credit risk and cannot be applied directly to the problem inspected here.

Furthermore, there exists another body of literature which analyses the performance of carry trades, a currency speculation strategy in which an investor borrows low-interest-rate currencies and lends high-interest-rate currencies. By the end of the last century, the abnormal performance had been confirmed by many authors, including Fama (1984) and Lothian and Wu (2011) with long time series amongst many others. To explain this performance there are different studies to predict the returns of carry trades in cross section (Lustig and Verdelhan 2007 and Menkhoff et al. 2012) as well in time series (Bakshi and Panayotov 2013). While in the latter case several factors like commodity returns or a global liquidity factor seem to predict returns of carry trades the picture in the first case is not that clear. Yet, another body of literature explains these returns with the help of the so-called Peso problem. Following Burnside et al. (2011) and Jurek (2014) the Peso state causing the Peso problem is a rare event of a sudden huge change in exchange rates coupled with a high stochastic discount factor (SDF) for that state. Adding an out of the money option to the carry trade reduces the loss in the Peso state resulting in a reduced skewness of the return distribution. Insofar the Peso problem inspected by Brunnermeier et al. (2008) and Burnside et al. (2011) is comparable with the situation of the rare event of the default of a borrower inspected in this work. Otherwise, carry trades classically ignore counterparty risk aside from liquidity risk which may coincide with the latter.

Furthermore, there is a large amount of empirical literature assessing the impact of various factors on credit spreads and therefore on the implied default probabilities of sovereign bonds (e.g., Edwards 1986; Chen et al. 2013; Cantor and Packer 1996; Eichengreen and Mody 1998; Min 1998; Peter 2002; Grandes 2003; Jostova 2006; Hilscher and Nosbusch 2007; Pan and Singleton 2006; Remolona et al. 2007; Longstaff et al. 2007, amongst many others).

However, the impact of currency denomination on credit spreads has so far received little attention. Amongst the few to look at this issue, Kamin and von Kleist (1999) find that credit spreads of emerging market sovereign debt denominated in US Dollar (USD) were systematically higher in the 1990s, which is attributed to comparably higher U.S. treasury yields. However, this is not a satisfactory explanation, as will be shown later.

Kercheval et al. (2003) find that corporate credit spread returns of corresponding issuer clusters, i.e., of clusters containing bonds with identical rating, sector, and currency denomination, are somewhat uncorrelated across different currencies, whereas credit spread returns of similar issuer clusters with identical currency denomination are highly correlated. This effect is also found at a single-issuer level, but only based on a very limited sample. The authors attribute their empirical results to market segmentation. However, the study has, according to Jankowitsch and Pichler (2005), a few shortcomings. For example, the data doesn't include foreign currency bonds, so a potential home bias could arise within the different clusters.

McBrady (2003) investigates the empirical determinants (in this particular case: default risk, liquidity premiums and segmented markets) of industrial country sovereign spreads. Among other things, he shows that market segmentation is present even in the largest and most liquid bond markets. He argues that investors prefer local currency bonds due to regulatory constraints. Such restrictions for institutional investors, such as pension funds, local benchmark orientation of funds and limited access of private investors to swaps, prevent them from effectively hedging exchange rate risk.

Ehlers and Schönbucher (2004) investigate whether the same credit risk occurs in different currencies. They show theoretically that their empirically observed differences in these credit spreads are mostly driven by the dependency between the default risk of the debtor and the exchange rate. In their empirical study, they find that a mere diffusion-based correlation between the exchange rate and the default intensity cannot explain the observed differences between JPY (Yen) and USD Credit Default Swaps (CDS) rates for a set of large Japanese debtors.

Jankowitsch and Pichler (2005) also analyze corporate credit spreads across different currencies. Contrary to Kercheval et al. (2003), they do not compare credit spread returns of issuer clusters, but rather credit spread curves at a single-issuer level. They find that credit spread curves at the single-issuer level differ significantly across currencies. Jankowitsch and Pichler (2005) interpret their findings as a rejection of the hypothesis of independence between default risk and the risk-free interest rate or default risk and exchange rate. Again, it is not clear whether the corporate bonds in the sample are comparable in terms of seniority and collateralization.

Van Landschoot (2008) systematically compares the determinants of Euro (EUR) and USD yield spread dynamics in the corporate bond markets. Among other results, she finds the liquidity risk to be higher for USD corporate bonds than for EUR corporate bonds. Her results also indicate that the credit cycle, as measured by region-specific default probabilities, significantly increases USD yield spreads, whereas these findings are not valid for EUR yield spreads.

Sener and Kenc (2008) empirically examine the determinants of the same risk across different currencies and their effects on the valuation of risky debt. Their central result is that only a small fraction of the spread can be explained by default risk, as credit spreads of the same issuer are not only currency-dependent. Other influencing factors, such as a variation in bankruptcy laws, tax regimes and liquidity conditions in domestic markets, also play a significant role in the calculation of credit spreads across different credit markets.

In their empirical study Buraschi et al. (2013) investigate the dynamic properties of arbitrage limits in the sovereign bond market (particularly in Brazil, Mexico and Turkey) around a period of market plight. They link the credit spreads of pairs of sovereign bonds issued in two foreign currencies (in this particular case: EUR and USD), of the above mentioned emitters with the help of a theoretical arbitrage relationship. Ultimately, they find the limits to arbitrage are time-varying and state-dependent. Moreover, the countries Brazil and Mexico pay higher and more volatile risk premiums in EUR. This conclusion partly contradicts the findings of the current empirical study, which may be due to the time periods inspected. In particular, Buraschi et al.'s (2013) time frame comprises the years of 2008–2010, when the financial crises occurred. Furthermore, the authors used a different methodology to measure differences in credit spreads, i.e.: by detecting arbitrage opportunities in a set of US government bonds. Hence, the influence of the correlation plays only a minor role in their empirical study.

To sum up, there exists no study explicitly stating the correlation between default risk and the exchange rate in both a theoretical and empirical analysis.

In a nutshell, the aim of this project is twofold: First, the existing theoretical models were unable to draw a conclusion as to whether the variables default, exchange, and interest rate are interdependent. Therefore, we show by means of the value of a quanto option, derived in a Jarrow/Turnbull-model extended with a second currency, that given a positive correlation between exchange rate (defined as EUR/USD) and the event of default, the credit spreads in USD should theoretically be higher than in EUR, and vice versa. Secondly, the empirical studies published so far often only partially compare credit spreads of different issuers, or the same issuer with different contractual specifications. They also do not take into account a potential home bias. Furthermore, these surveys do not consider the fact that term structure estimation errors may influence the difference in credit spreads. Above all, due to the drawbacks of the applied theoretical models, the influence of the correlation structure cannot be measured. We overcome these shortcomings by choosing a well conditioned data set and by considering appropriate independent variables within the regression analysis in order to examine potential driving factors for different credit spreads.

More condensed this paper contributes to the question: Is this price difference between USD and EUR denominated bonds of the same borrower rational or not.

- Therefore we derive firstly with the help of a quanto option an explicit value for the price difference which implies also a difference for credit spread and implied default probabilities. Thereby, the dependence between exchange rate and default risk plays the major role and determines the sign of the price difference.

- Secondly, we test this result empirically by extending the framework in use by variables measuring the dependence structure as well as controlling for measurement errors. We find that the quanto option can explain the price difference in many cases and we show by means of a random effects model that the correlation as a measure for dependence between exchange rate and default risk is highly significant in explaining those price differences.

The paper is organized as follows: the extension of the theoretical model, including the hypotheses in Sect. 2, is followed by an extensive empirical study in Sect. 3, in which the shortcomings outlined above are avoided. Section 4 presents the findings of the univariate, the multivariate analysis, and the robustness analysis, and includes a discussion of the findings. Finally, Sect. 5 concludes the paper.

2 Theoretical analysis in perfect markets

2.1 Completion with the help of a quanto option

In the Jarrow/Turnbull-framework applied here, we assume two complete and perfect markets in two currencies ("Domestic Currency Unit" shortened as DCU and "Foreign Currency Unit" shortened as FCU), with fixed recovery rates δ and payment of the recovery rates upon bond maturity. The risky and riskless discount factors can be expressed as a function of the maturity T and spot rate s:

$$\begin{aligned} &\mathrm{ZB_{DCU}(0,T)} = \exp\left(-s_{\mathrm{DCU}}(\mathrm{T}) \cdot \mathrm{T}\right) \ \ \mathrm{ZB_{FCU}(0,T)} = \exp\left(-\, s_{\mathrm{FCU}}(\mathrm{T}) \cdot \mathrm{T}\right) \\ &\mathrm{RZB_{DCU}(0,T)} = \exp\left(\left(-\, s_{\mathrm{DCU}}(\mathrm{T}) - cs_{\mathrm{DCU}}(\mathrm{T})\right) \cdot \mathrm{T}\right) \ \ \mathrm{RZB_{FCU}(0,T)} = \exp\left(\left(-s_{\mathrm{FCU}}(\mathrm{T}) \right.\right. \\ &\qquad\qquad \left.\left. - \, cs_{\mathrm{FCU}}(\mathrm{T})\right) \cdot \mathrm{T}\right) \end{aligned} \tag{1}$$

Whereas the credit spread cs is defined as the difference between risky and riskless continuously compounded interest rates:

$$\begin{aligned} \mathrm{cs(T)_{DCU}} &= \frac{1}{\mathrm{T}} \ln\left(\mathrm{ZB_{DCU}(0,T)/RZB_{DCU}(0,T)}\right) \\ \mathrm{cs(T)_{FCU}} &= \frac{1}{\mathrm{T}} \ln\left(\mathrm{ZB_{FCU}(0,T)/RZB_{FCU}(0,T)}\right) \end{aligned} \tag{2}$$

According to Jarrow and Turnbull (1995), the so defined credit spreads can be expressed in terms of the recovery rate δ and the implied default probability λ:

$$\mathrm{cs(T)}_{\bullet} = \frac{1}{\mathrm{T}} \ln\left(\frac{1}{\lambda_{\bullet}(0,\mathrm{T})\,(\delta_{\bullet} - 1) + 1}\right) \tag{3}$$

Because we assume that the recovery rate is constant (and the same in different currencies), the credit spread is a bijective function of the implied default probability.

Furthermore, two riskless short term interest rate processes can be defined as

$$d_{DCU}(0,T) = \int_0^T \exp\left(-s_{DCU}(t)\right) dt \quad d_{FCU}(0,T) = \int_0^T \exp\left(-s_{FCU}(t)\right) dt \tag{4}$$

Additionally, the process for the discount factor, based on the interest rate difference in two money market accounts, is defined by a short term interest rate difference process

$$d(0,T) = \int_0^T \exp\left(-s_{DCU}(t) + s_{FCU}(t)\right) dt \tag{5}$$

Because the framework is constructed for a single currency area, we extend the framework to a second currency area, connected via a stochastic exchange rate driven by a single factor. Both credit risk-free interest rates are stochastic, and in contrast to Jarrow and Turnbull (1995), may depend on the stochastic interest rate or the default event. Secondly, there is an exchange rate for transferring one unit of the foreign currency (FCU) in e(0) units of the domestic currency (DCU) at time 0. Because we can trade zero bonds maturing at T in both interest rates, the corresponding forward exchange rate is e(0,T).

To achieve a complete market in the correlated setting, we implement a quanto option, hereafter called an FX conditional forward contract (CFC). With the help of the CFC, differences in implied default probabilities can be traced back to the value of this contract. The CFC binds the parties to exchange of currencies under a certain condition. The condition here is the survival of the issuer.

As one can see in Table 1, buying one unit of a risky domestic currency bond maturing in T $RZB_{DCU}(0,T)$ leads to the same payoffs as the following portfolio: one unit of forward value at T e(0,T) of a risky foreign currency bond at $RZB_{FCU}(0,T)$, which is the amount of the recovery rate δ of an FX forward contract and $1-\delta$ of the conditional forward contract. In the case of survival, the domestic currency bond pays 1 DCU just as the foreign currency bond pays 1 FCU. The latter can be exchanged at the same rate partly by using the FX forward and partly by using the CFC. In the default case, both bonds pay only the recovery rate in DCU and FCU. Now only the recovery rate is exchanged using the FX forward. Because the FX forward is by nature a transaction with no ex ante capital exchange, we can use the value of the CFC to evaluate eventual pricing differences and driving forces.

Insofar the problem set-up here is similar to the problem described in Burnside et al. (2011). Both approaches construct portfolios of long and short currency forward and bond position as well as an option on a rare event. However, in our case the rare event is the credit event, where an explicit measure (here: the implied default probability) of that date can be given.

2.2 Value of the conditional forward contract

If the value of the CFC is zero, the arbitrage-free criterion results in

$$RZB_{DCU}(0,T) = \frac{e(0)}{e(0,T)} RZB_{FCU}(0,T).$$

Table 1 Duplication portfolio for a risky domestic currency government bond

Transaction	Time 0	Survival at T	Default at T
Portfolio FCU			
$\frac{1}{e(0,T)}$ of $RZB_{FCU}(0, T)$	$\frac{-e(0)RZB_{FCU}(0,T)}{e(0,T)}$	$\frac{1e(T)}{e(0,T)}$	$\frac{\delta e(T)}{e(0,T)}$
$\frac{\delta}{e(0,T)}$ of Forward	0	$\frac{\delta(e(0,T)-e(T))}{e(0,T)}$	$\frac{\delta(e(0,T)-e(T))}{e(0,T)}$
$\frac{(1-\delta)}{e(0,T)}$ of Conditional Forward	0	$\frac{(1-\delta)(e(0,T)-e(T))}{e(0,T)} + CFC$	0 + CFC
Altogether	$\frac{-e(0)RZB_{FCU}(0,T)}{e(0,T)}$	1 + CFC	δ + CFC
DCU			
1 of $RZB_{DCU}(0, T)$	$-RZB_{DCU}(0, T)$	1	δ

The table presents the duplication of the payoffs of a domestic currency bond with the help of several transactions in foreign currency instruments. The former can be found in line 6. The latter is depicted in lines 2–5. Hereby, column 2 shows the quantity of the financial instruments traded. Columns 3–5 are dedicated to the cash flows resulting from the transaction at time 0 and time T. In the latter case, the cash flows depend on the state of the bond's issuer. As one can see the cash flows of the duplication portfolio in line 5 differ from the payoffs of a domestic currency bond only by the amount of the fixed value of CFC, which can be discounted to time 0

Keeping in mind that

$$e(0,T) = \frac{ZB_{FCU}(0,T)}{ZB_{DCU}(0,T)} e(0)$$

this yields

$$RZB_{DCU}(0,T) = \frac{ZB_{DCU}(0,T)}{ZB_{FCU}(0,T)} RZB_{FCU}(0,T)$$

and

$$\frac{RZB_{DCU}(0,T)}{ZB_{DCU}(0,T))} = \frac{RZB_{FCU}(0,T)}{ZB_{FCU}(0,T)} \Leftrightarrow \lambda_{DCU} = \lambda_{FCU}. \tag{6}$$

Consequently, if the value of CFC = 0, the default probabilities are the same in both currencies. Therefore a positive value for CFC coincides with a lower implied default and higher survival probability in the domestic currency and vice versa. Because the sign of CFC does not change, if we alter the payment style from cash to future-style, we assume for simplification purposes that the payment is in the future-style for the following formulas.

2.3 Factors affecting the implied default probability

In the next step we analyze the value of CFC if the variables: survival of the issuer and/or exchange rate and/or discount factor, based on the interest rate difference, are stochastically dependent. As shown in Table 2, the value of CFC is always zero if the exchange rate and the survival are independent and conditionally independent, given

Table 2 Value of the contingent forward contract and correlations in case of a conditional independence of exchange rate and money markets discount factor

Dependence/ correlation of Survival and	Survival and	Discount factor (conditional independent of Exchange Rate)		
		Negative	Independent/zero	Positive
Exchange rate DCU/FCU	Negative	CFC(0) $>$ CFC(inspected) $>$ 0 $=>$ $\lambda_{\mathrm{FCU}} - \lambda_{\mathrm{DCU}} > 0$	CFC(+) $>$ CFC(inspected) $>$ 0 $=>$ $\lambda_{\mathrm{FCU}} - \lambda_{\mathrm{DCU}} >$ diff($-$) $>$ 0	CFC(inspected) $>$ 0 $=>$ $\lambda_{\mathrm{FCU}} - \lambda_{\mathrm{DCU}} >$ diff(0) $>$ 0
	Independent/zero	CFC(inspected) $=$ 0 $=>$ $\lambda_{\mathrm{FCU}} = \lambda_{\mathrm{DCU}}$	CFC(inspected) $=$ 0 $=>$ $\lambda_{\mathrm{FCU}} = \lambda_{\mathrm{DCU}}$	CFC(inspected) $=$ 0 $=>$ $\lambda_{\mathrm{FCU}} = \lambda_{\mathrm{DCU}}$
	Positive	CFC(0) $<$ CFC(inspected) $<$ 0 $=>$ $\lambda_{\mathrm{FCU}} - \lambda_{\mathrm{DCU}} < 0$	CFC(+) $<$ CFC(inspected) $<$ 0 $=>$ $\lambda_{FCU} - \lambda_{\mathrm{DCU}} <$ diff($-$) $<$ 0	CFC(inspected) $<$ 0 $=>$ $\lambda_{FCU} - \lambda_{DCU} <$ diff(0) $<$ 0

The table presents the value of the CFC dependent on the correlation between survival and exchange rate as well as on the conditional correlation between survival and discount factor. E.g., in the case of independence between survival and exchange rate the CFC is zero (line 4). Otherwise the CFC(inspected) is either positive (line 3) or negative (line 5). Furthermore the height of the CFC(inspected) is compared to the height of the CFC(0) in case of independence. Additionally the implied default probabilities are given resulting from the value of the CFC. The differences in default probabilities differ according to the table (see e. g., line 3 column 4: diff($-$) means that the difference in default probabilities in case of a negative correlation between survival and discount Factor, diff(0) in case of no correlation)

the discount factor in the martingale measure world. At first glance, this result contradicts the findings of Jankowitsch and Pichler (2005), but is explainable because they use a Duffie/Singleton-framework, in which implied probabilities of credit spreads and/or recovery rates are not temporally deterministic. Furthermore, in contrast to Jankowitsch and Pichler (2005), supporting evidence can be shown if dependence between exchange rate and survival is presumed. Due to negative correlation the exchange rate is low at survival and therefore the value of the CFC is positive. Finally, the implied default probability λ of $\mathrm{RZB}_{\mathrm{FCU}}$ exceeds the implied default probability of $\mathrm{RZB}_{\mathrm{DCU}}$ and vice versa. This is due to the fact that, in the case of survival, a lower exchange rate (DCU/FCU) is more likely in the martingale measure. Consequently, investment in the foreign currency bond cuts the possible cash flows of the contingent position and therefore lowers the value of this position.

These theoretical considerations lead to the first hypothesis (H1), which explains the difference between two credit spreads in two different currencies of the same issuer. Given a positive correlation between exchange rate (defined as EUR/USD) and the event of default, H1 is formulated in the following way:

H1 The higher the correlation between exchange rate and implied survival probability, the higher the theoretical difference between implied default probabilities of USD and EUR denominated foreign currency government bonds.

If the CFC value is non-zero, a correlated discount factor of the interest rate difference influences the value of the CFC, even though discount factors for the interest rate difference and the exchange rate are conditionally independent, given the survival. As shown in Table 2, a positive correlation of the discount factor for the interest rate difference and the survival leads to a higher absolute value of the CFC, leading to a higher absolute difference between the implied probabilities. Due to the fact that the discount factor for interest rate differences is strictly positive, the sign of the CFC cannot change and therefore dumping of the CFC's absolute value through the correlation of discount factors for the interest rate differences and survival cannot completely equalise the effect resulting from the correlation of exchange rate and survival (see Table 2). Therefore, the second hypothesis (H2) can be formulated as follows:

H2 The higher the correlation between the discount factor for the interest rate difference and the implied survival probability, the higher the absolute theoretical difference between implied default probabilities of USD and EUR denominated foreign currency government bonds.

The case can also be considered in which the discount factor for the interest rate difference is independent and conditionally independent on the survival. Here, a positive correlation between discount factor and exchange rate decreases the signed value of the CFC, and vice versa. Consequently, the value of positive & negative CFC's is lowered when the correlation coefficient is positive. Furthermore, in the case of a positive CFC, the difference in implied probabilities correspondingly decreases, just as it increases in the case of a negative CFC. According to Table 3, it can be stated that the CFC always reacts in the same way as in the independent case. Even if the correlation of exchange rate and survival and the correlation of discount factor and survival contradict each other, the sign of the value will not change. This leads to our last hypothesis H3:

H3 The higher the correlation between discount factor and exchange rate, the lower the theoretical difference between implied default probabilities of USD and EUR denominated foreign currency government bonds.

In a nutshell, the main driver of difference in credit spreads and implied default probability is the correlation between exchange rate and the event of survival or default. The correlation between discount factor and exchange rate or the correlation between discount factor and default can only alter the size of the difference, but not the sign (see Tables 2, 3).

Additionally, as would be expected in an option pricing setting the volatility plays a role in pricing the CFC yielding to a price difference between both bonds. Before we look at the volatility of exchange and interest rates, we focus on the default probability as a measure for default risk. Here it can be easily derived from the pricing formula, that the covariance between exchange rate and implied survival probability is the driving force. Whereas the correlation coefficient determines the sign of the price difference, the default probability is responsible for the absolute height of the price difference. To put it in other words: Higher default probabilities lead to higher price differences, when the correlation between exchange rate and implied survival probability is positive. In case of a negative correlation higher probabilities lead to lower (negative) differences. The latter will have implications for the choice of our data set.

Contrarily for the volatilities of exchange and interest rates it is not as clear cut as in the former case. Even so a higher volatility not only enlarges the absolute price

Table 3 Value of the contingent forward contract and correlations in case of a conditional independence of survival and money markets discount factor

Dependence/ correlation of Survival and	Exchange rate and	Discount factor (conditional independent of Survival)		
		Negative	Independent/zero	Positive
Exchange rate DCU/FCU	Negative	CFC(inspected) > CFC(0) > 0 => $\lambda_{FCU} - \lambda_{DCU} >$ diff(0) > 0	CFC(inspected) < 0 => $\lambda_{FCU} - \lambda_{DCU} >$ diff(+) > 0	CFC(0) > CFC(inspected) > 0 => $\lambda_{FCU} - \lambda_{DCU} > 0$
	Independent/zero	CFC(inspected) = 0 => $\lambda_{FCU} = \lambda_{DCU}$	CFC(inspected) = 0 => $\lambda_{FCU} = \lambda_{DCU}$	CFC(inspected) = 0 => $\lambda_{FCU} = \lambda_{DCU}$
	Positive	CFC(0) < CFC(inspected) < 0 => $\lambda_{FCU} - \lambda_{DCU} < 0$	CFC(inspected) < 0 => $\lambda_{FCU} - \lambda_{DCU} <$ diff(-) < 0	CFC(inspected) < CFC(0) < 0 => $\lambda_{FCU} - \lambda_{DCU} <$ diff(0) < 0

The table presents the value of the CFC dependent on the correlation between survival and exchange rate as well as on the conditional correlation between exchange rate and discount factor. E.g., in the case of independence between survival and exchange rate, the CFC is zero (line 4). Otherwise the CFC(inspected) is either positive (line 3) or negative (line 5). Furthermore the height of the CFC(inspected) is compared to the height of the CFC(0) in case of independence. Additionally, the differences in default probabilities differ according to the table (see e.g. line 3 column 4: diff(+) means that the difference in default probabilities in case of a positive correlation between exchange rate and discount factor, diff(0) in case of no correlation, diff(−) in case of a negative correlation)

difference by increasing the covariance between exchange rate and implied survival probability, but it also lowers the price difference by increasing the covariance between exchange rate and differences in discount factors in the presence of positive price differences. The same applies to the volatility of interest rates. As a consequence, the influence of the volatilities depends on the correlation between exchange rate and implied survival probability.

2.4 Control variables

In addition to correlation, further potential influencing factors were tested, which relate to an imperfect and/or incomplete market, such as liquidity, coupon rate, collective action clauses and pricing error.

Amongst many others, Easley et al. (2002), argue that liquidity is priced because investors maximize expected returns net of transactions (or liquidity) costs. Several studies, including for example Collin-Dufresne et al. (2001) or Longstaff et al. (2005), conclude that (changes in) yield spreads and therefore (the implied default probabilities are significantly affected by liquidity risk (e.g., Driessen 2005; Van Landschoot 2008; Chen et al. 2007). Consequently, we introduce *Liq* as a control variable for the liquidity risk factor.

Generally, differences in income tax regimes between countries or currency areas are the one main reason why credit spreads, and therefore differences in the implied default probabilities, may be affected by taxes: (e.g., Driessen 2005; Sener and Kenc 2008; Van Landschoot 2008). Such an analysis requires knowing the nationality and legal status of the bondholders. Consequently, our analysis is limited to the perspective of German and U.S. investors. For German private investors, bonds with a low coupon have recently been advantageous from a tax point of view, because capital gains were tax-exempt after a one year holding period—whereas interest income is always taxable. For German companies, there is no difference in taxation. Furthermore, capital gains and interest income are taxed differently in the U.S. for private investors. In the U.S. there is no differentiation between sundry income, resulting in the assessment of interest income using a relevant individual tax rate. In comparison, capital gains are liable to a tax rate of 15 % and, unlike in Germany, there is no tax-exemption after a one year holding period. Prior to the tax reform of 1986, U.S. sovereign bonds with high coupons were clearly not preferred (Jordan 1984; Litzenberger and Rolfo 1984; Ronn 1987), whereas this coupon effect could not be found after the reforms (Green and Odegaard 1997; Elton and Green 1998).

In the case of German investors, higher coupon rates were also a disadvantage in the past from a tax point of view. Therefore, it is possible that a (potential) coupon tax effect could explain differences in the implied default probabilities between government bonds denominated in EUR and USD. Hence, we use the coupon rate as an explanatory variable to control for the tax effect (Lim et al. 2004).[1]

As real world default probabilities are equal for all bonds considered from an issuer, differences in default risk can only be due to different recovery rates. But this cannot be attributed to differences in seniority, as all bonds considered from an issuer rank pari passu. Therefore, only different legal procedures in case of default could lead to a different recovery rate. In fact, there are different legal procedures in place precisely because of collective action clauses. Collective action clauses enable a qualified majority of bondholders to change material characteristics of the bond, such as principal, coupon rate or maturity. In this case, the higher credit spreads/implied default probabilities or yields would compensate bondholders for a lower recovery rate. However, even if most studies cannot find a significant impact of collective action clauses on bond yields (Tsatsaronis 1999; Dixon and Wall 2000; Gugiatti and Richards 2003; Becker et al. 2003, amongst many others), we must control for differences in collective action clauses in our analyses.

Generally, a direct comparison of two bonds denominated in EUR and in USD can only be successfully carried out if the two bonds have the same maturity as well as coupon payments. As this rare event is non-existent in the data source, a comparison is achieved, according to Jankowitsch and Pichler (2005) or Sener and Kenc (2008), by comparing credit spread curves, which are estimated from bond prices. The drawback of this methodology is neglect of error in the credit spread estimation process.

To allow for this effect, we use the credit spread curve or, equivalently, the implied default probabilities from USD to calculate bond prices in EUR and vice versa. We then compare this price with the price estimated using implied default probabilities in the opposite currency. We use this result to control for measurement error in the credit spread curve estimation.

[1] The fact that higher coupon rates coincide with higher implied default probabilities (Elton et al. 2004) should generally not affect the results, as this is the case for both currencies.

Table 4 Tabulation of descriptive data for the sample

	Brazil		Mexico		Poland		Turkey	
	USD	EUR	USD	EUR	USD	EUR	USD	EUR
Number of bonds	24	6	17	7	5	10	16	8
CAC	10	1	6	3	3	0	10	0
Average LOT-measure	2.8 %	0.31 %	0.14 %	0.33 %	0.79 %	0.03 %	0.09 %	0.06 %
Average coupon (per country and currency)	9.65 %	9.81 %	7.81 %	7.13 %	5.93 %	4.76 %	9.16 %	6.48 %

The table contains a brief descriptive statistics of the sample inspected. CAC stands for the number of bonds with collective action clauses. LOT-measure stands for the liquidity measure according to Lesmond et al. (1999)

3 Dataset and methods

3.1 Dataset

The data sample contains prices for 93 bonds and approximately 11,000 trading days for sovereign bonds of Brazil, Mexico, Poland, and Turkey, 31 of which are denominated in EUR and the remainder in USD (see Table 4). The daily prices were downloaded from Thomson Reuters Datastream and cover the period of April 1998 to June 2008. These four countries were chosen for analysis because only these countries fulfill all three necessary criteria during the analysis period. The first criterion is that all countries must have at least five outstanding bonds, both in EUR and USD, as five bonds are necessary to estimate stable spot rate curves using established methods. The second criterion is that all bonds must be foreign currency bonds, due to the potential different default risk of local currency sovereign bonds. EUR and USD are the most common foreign currencies in which sovereign bonds are denominated, and have the additional advantage that the yield curve, i.e., the risk-free spot rate curve, can be estimated very accurately for both currencies. The third criterion is that the foreign currency rating of all issuers considered must be below A on the S&P scale in 2006, leaving enough room for the occurrence of significant differences in the credit spreads or implied default probabilities. All bonds have fixed coupon payments once or twice a year and are redeemed at par. None of the bonds contain embedded options or are redeemable or callable prior to maturity. All bonds constitute unsubordinated and unsecured debt, i.e., they rank pari passu, assured by a so-called negative pledge in the prospectuses. This ensures that all bonds from an issuer should have, apart from differences specified in Sect. 2.4, the same default risk, and that their credit spread can be fully attributed to default risk, because the cash-flows are known ex-ante in the case of no default occurring before maturity. To estimate the implied default probabilities for each issuer and each currency according to the reduced-form model of Jarrow and Turnbull (1995), the risky and the risk-free discount factors must be estimated.

For the correlation structure analysis, a EUR/USD exchange rate was required. The applied rate was one fixed in Frankfurt by the German Central Bank.

3.2 Estimation of risky and risk-free discount factors

Three estimation methods were applied to estimate risky discount factors: the method established by Chambers et al. (1984), the method presented by Nelson and Siegel (1987), and the Restricted Integro Basis Spline-Method, proposed initially by Rathgeber (2008). The latter guarantees always positive forward rates, even if only a few bonds are outstanding. Consequently, the Restricted Integro Basis Spline-Method turned out to be superior, both in terms of estimation quality and consistency of the estimated discount factors. Finally, we used the discount factors estimated by this method for further analysis.

German sovereign bonds ("Bundesanleihen") were used to estimate risk-free EUR spot rates. The German Central Bank uses a method proposed by Svensson (1994), which is an extension of the method of Nelson and Siegel (1987), allowing a more flexible estimation of the EUR yield curve. It is important to consider that the German Central Bank does not use the exact discount function developed by Svensson (1994), rather a slightly modified version proposed by Schich (1997), which inconsistently uses discretely compounded interest rates, rather than continuously compounded interest rates.

For the sake of unity, the Svensson method was also used to estimate risk-free USD discount factors. For the estimation, daily prices of 77 treasury bonds and notes were downloaded from Reuters Datascope Select. All bonds have fixed coupon payments twice a year and are redeemed at par. Before we calculate the implied default probabilities, we briefly discuss the results expected from a theoretical point of view.

3.3 Implied default probability

After the estimation of the risky and risk-free discount factors, a recovery rate must be chosen to calculate the implied cumulative default probabilities:

$$\lambda(0, \mathrm{T}) = \frac{\mathrm{RZB(0,T)\ /ZB(0,T)\ \text{-}\ 1}}{\delta - 1} \tag{7}$$

where $\lambda(0, \mathrm{T})$ is the implied cumulative default probability for the period [0,T].

We have chosen a recovery rate of 30 % for all countries, which is equal to the basis recovery rate applied by Fitch Ratings (2006) and is at the lower end of the observed range of historical recovery rates (e.g., Moody's 2008; Merrick Jr. Merrick Jr. 2001). In any case, the results hardly vary with a different recovery rate. Using these probabilities, we were also able to calculate CFC values according to formula (9) in P.3. under independence and conditional independence of default event and discount factors. In the second step, we were able to compare these CFC values to the differences in zero bond prices in the implied default probabilities.

3.4 Dependent and independent variables

To test our hypotheses and other possible influencing factors (see Sect. 2.4), a multiple linear regression analysis was performed which includes the following factors as independent variables (Van Landschoot 2008):

– Correlation $Cor1_{\tau,k}$ at time τ: Pearson's (product-moment) correlation between the exchange rate (EUR/USD) and the implied survival probability for the maturity of bond k used to test hypothesis 1 ($H1$) (2,500 trading days, in case the time series is to short we used the maximum number of trading days minimum 1,000).
– Correlation $Cor2_{\tau,k}$ at time τ: Pearson's (product-moment) correlation between the interest rate difference and the implied survival probability for the maturity of bond k used to test hypothesis 2 ($H2$) (2,500 trading days, in case the time series is to short we used the maximum number of trading days minimum 1,000).
– Correlation $Cor3_{\tau,j}$ at time τ: Pearson's (product-moment) correlation between the interest rate and the exchange rate j used to test hypothesis 3 ($H3$) (2,500 trading days).
– Liquidity risk factor $Liq_{\tau,k}$ at time τ: bid ask spread of a bond k evaluated by LOT-Measure (Lesmond et al. 1999).
– Coupon rate CR_k: coupon rate of a bond k.
– Collective action clauses CAC_k: dummy value, which is 1 in the case of collective action clauses of bond k.
– Measurement error $err_{\tau,k}$ at time τ: defined as relative error between the estimated price $\overline{RB}_{\tau,k}$ and the market price $RB_{\tau,k}$ of a bond k.

$$\begin{aligned}\mathrm{err}_{\tau,k} &= 1 - 1/\mathrm{RB}_{\tau,k}\Bigg(\sum_{t=\tau+1}^{T} \mathrm{CF}_{DCU,k}(t)\cdot \mathrm{ZB}_{DCU}(\tau,t)\cdot(1-\lambda_{DCU}(\tau,t)) \\ &\qquad + \delta\cdot \mathrm{CF}_{DCU,k}(t)\cdot \mathrm{ZB}_{DCU},(\tau,t)\cdot\lambda_{DCU}(\tau,t)\Bigg) \\ &= 1 - 1/\mathrm{RB}_{\tau,k}\overline{\mathrm{RB}}_{\tau,k} \end{aligned} \tag{8}$$

whereby T is the maturity of the risky bond k and $CF_{\bullet}(t)$ are its cash flows.

Skewness: Skewness of the logreturns of exchange rate (EUR/USD) for robustness checks (180 trading days).

Investment currency: Difference of discount factors USD and EUR with maturity 1 year.

Because the liquidity risk factor and the measurement error may be measured with an error and is likely to be correlated with the regression residuals, we performed a Hausmann's specification test to identify potential correlations. In accordance with Griffith, Hill, and Judge (1993, p. 462), we used the lagged variable as instrumental variable. In case of a significant correlation, we used the instrumental variable as an independent variable.

The dependent variable is the relative pricing difference between the two currencies. Assuming the implied default probabilities of the foreign currency are valid in the domestic currency, the relative price difference of a bond k can be expressed as

$$\begin{aligned}\mathrm{diff}_{\tau,\mathrm{k}} = 1 - 1/\overline{RB}_{\tau,\mathrm{k}}\Bigg(\sum_{t=\tau+1}^{\mathrm{T}} \mathrm{CF}_{DCU,\mathrm{k}}(t)\cdot \mathrm{ZB}_{DCU}(\tau,t)\cdot(1-\lambda_{FCU}(\tau,\mathrm{t})) \\ + \delta\cdot \mathrm{CF}_{DCU,\mathrm{k}}(t)\cdot \mathrm{ZB}_{DCU}(\tau,t)\cdot\lambda_{FCU}(\tau,\mathrm{t})\Bigg)\end{aligned} \tag{9}$$

4 Results

4.1 Univariate analysis

Table 5 shows, among other things, the means of the cumulative implied default probabilities per country and currency for year-round maturities between 2 and 10 years at a period of 2 years. The last column shows the foreign currency S&P Rating in 2006.[2] As expected, the means of the implied cumulative default probabilities increase correspondingly as the rating worsens and the maturity lengthens. Much more interestingly, it can be shown that the means of the implied cumulative default probabilities are a long way from being equal for each issuer across both currencies. As Table 5 shows, they are higher for USD denominated bonds across all maturities and issuers. However, Buraschi et al. (2013) ascertain in their empirical study the opposite for Brazil and Mexico: these two countries pay higher and more volatile risk premia in EUR compared to USD. The differences in the implied cumulative default probabilities are economically significant: given a risk-free spot rate of 4 %, today's price difference between two zero-coupon bonds with 10-year maturity is somewhere between 6 and 29 % for the four countries, exceeding usual bid-ask-spreads by a large margin. To establish whether these differences are also statistically significant, the differences were tested with the help of the Wilcoxon signed-rank test. This test was chosen due to its lack of requirements on a specific type of the distribution of the differences (except of symmetry). For each country and trading day, the difference between the implied cumulative default probabilities was calculated for year-round maturities, ranging from 2 to 10 years at a period of 2 years. To test whether USD implied cumulative default probabilities are significantly higher than the results shown, the one-tailed version of the Wilcoxon signed-rank test was applied. Table 5 gives the results in addition to the above mentioned issues (cumulative default probabilities, S&P Ratings). "H_0" means that the null hypothesis of equal expectation values could not be rejected at a 1 % significance level, whereas "H_1" means that the expectation value of the USD implied cumulative default probabilities is higher. It can be seen that this is true for nearly all maturities in Brazil, Mexico, Poland, and Turkey, apart from the two and four year maturities.

As shown in one of the previous sections, the implied default probabilities of one issuer are only equal across different currencies if there is on one hand independence between default risk and the evolution of the risk-free interest rate in the respective currency, and between default risk and the evolution of the exchange rate between the two considered currencies considered on the other hand. Therefore, the finding for different credit spreads or implied default probabilities could be attributed to an incorrect assumption of independence, as Jankowitsch and Pichler (2005) interpreted their results. To analyze a possible dependence between default risk and the evolution of the exchange rate between EUR and USD, the Pearson's (product-moment) correlation coefficient between the exchange rate of the EUR in USD and the USD bonds implied default probability was calculated. For the default risk in EUR, the correlation between the exchange rate of the USD in EUR and the EUR bonds implied survival

[2] The S&P foreign currency rating remained unchanged for all countries through 2006 apart from Brazil where the rating was upgraded from BB- to BB on February 22nd, 2006.

Table 5 Means of the implied cumulative default probabilities per country and currency, the corresponding S&P Ratings per country and the test results of the one-tailed Wilcoxon signed-rank test

	$\lambda(0,2)$		$\lambda(0,4)$		$\lambda(0,6)$		$\lambda(0,8)$		$\lambda(0,10)$		S&P Rating
	EUR (%)	USD (%)	EUR (%)	USD (%)	EUR (%)	USD (%)	EUR (%)	USD (%)	EUR (%)	USD (%)	
Brazil	7.38	8.80***	13.28	15.02***	18.18	19.43***	22.35	22.98***	26.27	26.48***	BB
Mexico	1.05	1.46***	3.10	4.51	5.45	7.95***	8.13	10.76***	11.60	13.35***	BBB
Poland	−0.12	1.36***	1.33	3.37	3.76	5.15***	5.36	6.73***	6.13	8.45***	BBB+
Turkey	2.54	3.69	6.95	9.77	12.62	15.99***	18.18	22.13***	22.08	28.07***	BB-

The table contains the different implied cumulative default probabilities for different maturities in different currencies (Euro and US-Dollar). E.g., $\lambda(0,2)$ represents the probability for a maturity of 2 years. The differences in the probabilities are tested by a one-tailed Wilcoxon signed-rank test. Test results of the Wilcoxon signed-rank test at a 1 % significance level: *** means that the expectation value of the USD implied cumulative default probabilities is higher (sample period 1998–2008) probability as a proxy was also calculated.[3] Therefore, an assumption was introduced that the correlation coefficient is the same in the martingale measure as in the real world measure. The proxy for the default risk serves as the unobservable dichotomy variable.

As can be seen in Table 6 the correlation coefficients are strongly and statistically significantly positive for the USD and negative for the EUR. The effect is persistent among different currencies and maturities. Furthermore, it can be said that the longer the maturity, the higher the correlation coefficients. Due to the fact that the correlation is, with a few exceptions, monotone in maturity, the maturity can be used as an instrumental variable for correlation (see Sect. 3.4).

On the other hand, the data indicates a more or less zero correlation between the risk-free interest rate and the default risk of bonds denominated in the respective currency. As shown in Table 11 (Appendix 1) there is no pattern and the correlation coefficients are frequently insignificant. Therefore an influence is very unlikely. This finding confirms the results found by Eichengreen and Mody (1998), but contradicts the results of Arora and Cerisola (2000).

Furthermore, the extent of the differences in the implied default probabilities (see Table 5), as shown in Table 7 can often be traced to the correlation between exchange rate and implied survival probability, with regard to the different countries and maturities. In addition, a structural difference is apparent between the varying maturities and the respective issuers.

4.2 Multivariate analysis

In our multivariate analysis, we used two methodological approaches to test our hypotheses. Firstly, we followed Buraschi et al. (2013) and Van Landschoot (2008): we used a panel regression for our analysis of all bonds and currencies, as follows

$$\begin{aligned} diff_{\tau,k} = c_0 + c_1 \cdot Liq_{\tau,k} + c_2 \cdot CR_k + c_3 \cdot Cor1_{\tau,k} + c_4 \cdot CAC_k + c_5 \cdot err_{\tau,k} \\ + c_5 \cdot Cor2_{\tau,k} \cdot Cor1_{\tau,k} + c_6 \cdot Cor3_{\tau,j} \cdot Cor1_{\tau,k} \\ + \varepsilon_{\tau,k} \end{aligned} \tag{10}$$

whereby k ∈ {for all j (USD or EUR) denominated bonds}.

To estimate Eq. (10), we performed a random effects model as well as a fixed effects model. Due to the fact that we had to leave alone the coupon rate CR_k as well as the collective action Clauses CAC_k due to strict multicollinearity in the fixed effects case, we chose to use the random effects model. In order to compare the random effects model with standard OLS regression, we conducted a Lagrange Multiplier test according to Green (2012). Furthermore, we tested the residuals for autocorrelation within the random effects model. To this end, we used a modified Durbin–Watson test according to Bhargava et al. (1982) in association with Baltagi et al. (2003). To

[3] Additionally, the Spearman's rank correlation coefficient, which is proposed in the existing literature, was calculated. Due to the fact that the use of this correlation coefficient revealed approximately to the same results compared to the use of the Pearson's (product-moment) correlation coefficient, the latter was used for the calculations.

Table 6 Correlation coefficients between implied survival probability and exchange rate

	Correlation between	Maturity				
		2 years	4 years	6 years	8 years	10 years
Turkey	USD denominated bonds Implied Survival probability and EUR in USD	0.2858	0.5004	0.6946	0.7699	0.7674
Turkey	EUR denominated bonds Implied Survival probability and USD in EUR	−0.7235	−0.7236	−0.7126	−0.7294	−0.7800
Poland	USD denominated bonds Implied Survival probability and EUR in USD	0.4319	0.4908	0.5550	0.6174	0.6616
Poland	EUR denominated bonds Implied Survival probability and USD in EUR	−0.6241	−0.7363	−0.7498	−0.6506	−0.5634
Mexico	USD denominated bonds Implied Survival probability and EUR in USD	0.2591	0.3167	0.3715	0.4468	0.5256
Mexico	EUR denominated bonds Implied Survival probability and USD in EUR	−0.3688	−0.4813	−0.5727	−0.5676	−0.5579
Brazil	USD denominated bonds Implied Survival probability and EUR in USD	0.6761	0.6929	0.7055	0.7136	0.6911
Brazil	EUR denominated bonds Implied Survival probability and USD in EUR	−0.4629	−0.5043	−0.5407	−0.5692	−0.5856

The table contains the average correlation coefficient between implied survival probability and exchange rate, according to formula (1) in Appendix P.1. Therefore the coefficient differs using direct or indirect quotation. Furthermore, we used a window of 2,500 trading days to estimate the correlation coefficients for the different survival probabilities from 2 to 10 years

Table 7 Differences in implied default probabilities and value of CFP attributed to correlation between exchange rate and implied survival probability

	Maturity				
	2 years (%)	4 years (%)	6 years (%)	8 years (%)	10 years (%)
Brazil					
Difference in implied probability	−1.42	−1.74	−1.25	−0.63	−0.21
Explainable by indirect estimation of covariance	69	97	>100	>100	>100
Mexico					
Difference in implied probability	−0.41	−1.41	−2.50	−2.63	−1.75
Explainable by direct estimation of covariance	16	19	22	35	80
Poland					
Difference in implied probability	−1.48	−2.04	−1.39	−1.37	−2.32
Explainable by direct estimation of covariance	7	15	40	61	51
Turkey					
Difference in implied probability	−1.15	−2.82	−3.37	−3.95	−5.99
Explainable by direct estimation of covariance	16	30	56	73	59

The table contains average differences in implied probabilities between EUR and USD Bonds, and the percentage which can be attributed to the correlation. Therefore, we calculated the value of the CFP using Eqs. (3) resp. (4) in Appendix P.1. and the correlation coefficient from Table 6. In case the theoretical value is in average higher than the inspected difference the table contains >100 %, in case the theoretical value is lower than the inspected difference the table contains the average relation of the the theoretical value and the inspected difference

test for heteroscedasticity, we conducted a robust Langrange Multiplier test according to Montes-Rojas and Sosa-Escudero (2011). This was necessary within our random effects model, as heteroscedasticity existed due to different error terms for different bonds within the covariance matrix.

In order to address the autocorrelation, we used a Prais–Winsten estimation for panel data, as this approach does not modify the length of the time series. After just one iteration, the residual's autocorrelation was reduced to a sufficient extent. The heteroscedasticity of the residuals was addressed by means of robust standard errors according to Rogers (Stock and Watson 2008).

First of all, there is an indirect positive confirmation of $H1$, in that the signed values of the implied default probabilities in USD are significantly higher than the implied default probabilities in EUR. Secondly, the coefficient of the variable correlation in all regressions deviates significantly from zero. Additionally, the sign of the coefficient coincides with the direction proposed by $H1$ according to Table 8.

Contrarily, the coefficients for the interacting variables are only different from zero in half of the cases. When the variable is significant, its regression coefficient often features the right sign. This is the case for the interacting variable (correlation between FX and Survival times correlation between Discount Factor and Survival). Here a positive correlation coefficient (Discount Factor and Survival) should increase

Table 8 Key findings panel regression

	Model 1	Model 2	Model 3	Model 4	Model 5
Constant	−0.01 (0.025; 0.007)	−0.13 (<0.001; 0.009)	−0.47 (<0.001; 0.028)	−0.54 (<0.001; 0.029)	−0.52 (<0.001; 0.028)
Corr. Surv. Prob. FX	−2.80 (<0.001; 0.021)	−3.45 (<0.001; 0.033)	−3.05 (<0.001; 0.028)	−3.58 (<0.001; 0.039)	−3.64(<0.001; 0.038)
Corr. Surv. Prob. FX × Corr. Disc Fac. FX		0.45 (0.007; 0.142)		0.19 (0.110; 0.155)	0.11 (0.221; 0.149)
Corr. Surv. Prob. FX × Corr. Disc. Fac. Surv.		−1.47 (<0.001; 0.059)		−1.39 (<0.001; 0.064)	−1.41 (<0.001; 0.061)
Coupon			0.10 (<0.001; 0.006)	0.10 (<0.001; 0.006)	0.09(<0.001; 0.006)
Coll. action clauses			0.25 (<0.001; 0.033)	0.28 (<0.001; 0.033)	0.19 (<0.001; 0.032)
Liquidity	2.29 (<0.001; 0.120)	1.88 (<0.001; 0.126)	2.89 (<0.001; 0.130)	2.46 (<0.001; 0.140)	2.32 (<0.001; 0.135)
Measurement error	0.27 (<0.001; 0.012)	0.29 (<0.001; 0.012)	0.27 (<0.001; 0.013)		0.29 (<0.001; 0.013)
Number of observations	24,462	24,462	24,462	24,462	24,462
R^2	29.42 %	30.90 %	30.94 %	30.00 %	32.31 %
Durbin–Watson	2.08	2.10	1.98	1.95	2.00
LM-test (OLS)	3.98**	3.98**	3.98**	3.98**	3.98**
LM heterosc.	0.58	0.74	0.97	0.90	0.97

The dependent variable in all columns is the price difference according to Eq. (9). The dependent variables are defined in Sect. 3.4, esp. Eq. (8) (*p* value; standard errors in brackets). The regression equation is given in formula (10). The sample (over the period 1998–2008) is drawn from databases described in Sect. 3. In the second to last line the modified Durbin–Watson statistic after correcting for autocorrelation using Prais–Winsten statistic is depicted. In the last line, the value of the LM-test is presented to compare the random effects model to an OLS regression. The second LM test statistic is the test of Montes-Rojas and Sosa-Escudero (2011) for heteroscedasticity. Standard errors clustered at the firm level are applied to determine significance

** 5% significance level

the absolute difference in default probabilities according to hypothesis $H2$. A high correlation coefficient (Discount Factor and Survival) therefore coincides in the case of a low correlation (FX and Survival) with a low interacting variable. The same appears in case of a low correlation coefficient (Discount Factor and Survival) in combination with a high correlation (FX and Survival). The differences should be high in both cases, according to Table 8. The opposite should appear in the other two cases. Consequently, the regression coefficient is predicted to be negative as observed.

In case of the variable (correlation between FX and Survival times correlation between Discount Factor and FX) the result is not as clear. But one would also expect a positive regression coefficient in the case of a negative correlation (FX and Survival), whereby the coefficient most of the times is positive but insignificant. The reason for this may lie in those rare cases where the correlation (FX and Survival) is positive, which would invert the theoretical prediction.

Regarding the control variables we see an ambivalent picture. The coefficient for liquidity, as an additional possible factor of influence, is always significantly different from zero, whereby the regression coefficient often features the wrong sign. One possible explanation may be that the LOT measure is very low and does not capture fully the liquidity risk of the different bonds. The results, in terms of the effect of taxes, are comparable to liquidity as factor. Again, the coefficient is significant and shows the wrong sign. This may be due to the coupon effect still being in existence in the eurozone for a long time. Furthermore, the purchase of collective action clauses plays a significant role. The regression coefficient for the dummy variable deviates positively and significantly from zero. Hence, it contradicts the theoretical predictions.

Moreover, the effect of the term structure estimation error can only be proved to be influential in every case. Additionally, the regression coefficient is positive as expected and shows a huge influence in an univariate analysis (not reported). In fact, three of four control variables have the wrong sign. The results with wrong signs persist even when the error variable is excluded, and seems to be the result of the dominant influence of correlation between FX and survival.

To sum up, the most important factor of influence on the different credit spreads and therefore the implied default probabilities seems to be the correlation between the exchange rate and the event of default. All other possible factors of influence, which were tested in the context of this empirical study, are more or less able to explain the empirically established differences between the implied default probabilities.

4.3 Robustness analysis

First of all, we used a Hausman specification test to analyze the influence of potential errors in the independent variables. Such an effect may be true for the correlation coefficients as well for the measurement error. We used the lagged independent variables as instrumental variables. Here we discovered that measurement errors in these variables seem to not be a severe problem (not reported). Furthermore, we used a shorter evaluation period (2007–2008) and again received the same results (not reported).

As a robustness check, we followed the methodology of Ehrhardt et al. (1995) or Eom et al. (1998) and performed a regression separately for each point in time, and consequently avoided the observed autocorrelation. In doing so, we avoided a correlation within the time series, as a difference in the implied default probabilities on one day is normally followed by a similar difference on the next trading day. Altogether, this methodology leads to 2,500 regressions per category in the base case.

In addition, we had to focus our attention on a regression model which could be consistent with the sampling process underlying the two samples per day, defined by two currencies. According to Griffith et al. (1993, p. 546), three possibilities can be applied. Hence, we performed three different regression models, where c_0 to c_5 represents the coefficients to be estimated and ε is the residual:

1. Each currency and each issuer were separately incorporated in the respective regression equation (Single Model).

$$\begin{aligned} diff_{\tau,k} &= c_0 + c_1 \cdot Liq_{\tau,k} + c_2 \cdot CR_k + c_3 \cdot Cor1_{\tau,k} + c_4 \cdot CAC_k + c_5 \cdot err_{\tau,k} + \varepsilon_{\tau,k} \\ diff_{\tau,h} &= c_0 + c_1 \cdot Liq_{\tau,h} + c_2 \cdot CR_h + c_3 \cdot Cor1_{\tau,k} + c_4 \cdot CAC_h + c_5 \cdot err_{\tau,h} + \varepsilon_{\tau,h} \end{aligned} \tag{11}$$

   where k, h ∈ { USD denominated bonds}.

2. Each issuer, without a distinction relating to currency, was separately entered into the respective regression equation (Total Model).

$$diff_{\tau,k} = c_0 + c_1 \cdot Liq_{\tau,k} + c_2 \cdot CR_k + c_3 \cdot Cor_{\tau,k} + c_4 \cdot CAC_k + c_5 \cdot err_{\tau,k} + \varepsilon_{\tau,k} \tag{12}$$

   where k {USD or EUR denominated bonds}.

3. Seemingly unrelated regressions model (SUR).

$$diff_{\tau,k} = c_0 + c_1 \cdot Liq_{\tau,k} + c_2 \cdot CR_k + c_3 \cdot Cor_{\tau,k} + c_4 \cdot CAC_k + c_5 \cdot err_{\tau,k} + \varepsilon_{\tau,k} \tag{13}$$

   where k {USD denominated bonds}.

$$diff_{\tau,h} = c_0 + c_1 \cdot Liq_{\tau,h} + c_2 \cdot CR_h + c_3 \cdot Cor_{\tau,k} + c_4 \cdot CAC_h + c_5 \cdot err_{\tau,h} + \varepsilon_{\tau,h} \tag{14}$$

   where k ∈ {EUR denominated bonds} and $\varepsilon_{\tau,k}$ and $\varepsilon_{\tau,h}$ are correlated

In terms of the findings of the three regression models, the proportion of highly significant regression coefficients (5 % level) as well as the average coefficient over all 2,500 regressions in the base case are listed in tabular form in Table 12 (Appendix 1) (Ehrhardt et al. 1995). The results are corrected for an eventual estimation error in the independent variables correlation liquidity and error. In addition, the standard errors are corrected for eventual heteroscedasticity by using the weighted least square estimation according to Griffith et al. (1993, p. 502). Lastly, we had to exclude several independent variables in the Polish regression due to multicollinearity.

The coefficient of the variable correlation in most regressions deviates significantly from zero. Additionally, the sign of the coefficient coincides with the direction proposed by $H1$.

Liquidity, as an additional possible factor of influence, can be proved only very occasionally, especially in the context of the SUR in the case of nearly all considered countries. When the variable is significant, its regression coefficient often features the wrong sign, as in the panel regression, which might again be due to the quality of measurement by meands of LOT.

The results, in terms of the effect of taxes, are ambivalent. It can be confirmed tendentially for Poland and Turkey by using the total model, but it is very likely confirmed by the single regression. By using the SUR it can be confirmed for Mexico and Turkey (see Table 12 in Appendix 1).

Furthermore, the purchase of collective action clauses plays only a minor role. Nevertheless, we find an effect in the case of Poland and Turkey, due to the fact that both Turkish and Polish EUR government bonds contain no collective action clauses. The dummy variable, by which collective action clauses as a factor of influence were tested, therefore functions as an indicator variable relating to the respective currencies.

Moreover, the effect of the term structure estimation error can only be proved in the case of Brazil and marginally inter alia for Turkey in the context of the SUR. The reason for this is that more government bonds are available for Brazil and therefore a higher pricing error results.

To sum up, the most important factor of influence on the different credit spreads and therefore the implied default probabilities seems to be the correlation between the exchange rate and the event of default. All other possible factors of influence, which were tested in the context of this empirical study, are more or less able to explain the empirically established differences between the implied default probabilities.

Furthermore, we took the duration of bonds expressed in years as an instrumental variable for correlation in order to test hypothesis H1. The correlation coefficient is monotone in maturity (see Table 6) and approximately the same results were received (see Table 13 in Appendix 1). Again, the coefficient of the variable duration deviates significantly from zero in most regressions.

The body of literature (see Burnside et al. 2011) related to the so-called Peso problem reports negatively skewed returns for carry trades. Even if the skewness is (in our sample for exchange rate returns (0.21) and changes in interest rates (EUR) -0.0557 and (USD) 0.1328) not as highly negative as reported in the literature, skewness deviating from zero in subsamples might change our results. This is due to the fact, that our setting neglects higher moments of the underlying's distribution. Therefore, we differentiated our time series of exchange rate returns, where the skewness is positive and negative. We analyzed implied default probabilities in these different periods and run the panel regressions separately. Besides, we included the skewness (measurement half a year, rolling window) as an additional explaining variable in the panel regression.

The reason why skewness might be important is as follows: Skewness seems to be important for carry trades but otherwise it detoriates the dependence as well as the risk measurement in a way, which is not represented in our theoretical model, where the risk is measured by volatility. Consequently, periods with high skewness imply a

Table 9 Investment and funding currencies

Issuer	Investment currency	2 years (%)	4 years (%)	6 years (%)	8 years (%)	10 years (%)
Brazil	EUR	0.09	0.12	0.03	−0.23	−0.01
	USD	6.41	9.68	10.79	11.43	12.18
Mexico	EUR	−0.12	−0.16	−0.25	−0.34	−0.45
	USD	4.29	7.59	10.15	11.76	11.97
Poland	EUR	−0.18	−0.32	−0.38	−0.39	−0.26
	USD	2.86	5.23	7.10	8.69	10.30
Turkey	EUR	−0.49	−0.41	−0.15	0.30	0.67
	USD	1.61	3.97	6.68	9.58	12.36

The table contains average differences in implied probabilities between EUR and USD bonds for different maturities (first line) and issuers (first column) (1998–2008). Therefore, we calculated the differences like in Table 4, whereby we differentiated between periods where EUR and USD is the investment currency (second column). We defined those periods, where EUR is the funding currency, as periods where the interest rates in EUR are higher than those in USD for the given maturity (first line). Periods where the USD is the investment currency are defined vice versa

Peso state for a falling Dollar and given a negative correlation the value of the CFC and in the end the differences raise.

When splitting the data set in two subsamples there was only a small difference in implied default probabilities (not reported). The same is true estimating the panel regression separately. Also the skewness shows only low explanatory power, if it is used as a further independent variable (see Table 10). This result is different from the results found in the case of carry trades (see Burnside et al. 2011). This might be due to the fact, that the skewness is only one of many different factors driving the value of the CFC.

Additionally, the role of carry trade investors might be influential to the difference. Carry traders could eventually invest in the bonds investigated here instead of investing in the credit risk free rate. This might cause a change in the price differences. If in the funding currency the implied default probabilities are higher as in the investing currency the price difference should be enlarged. If the opposite is the case, the price difference might decrease.

To test this we split our data sample into periods, when the short term interest rates in USD are higher as in EUR and vice versa (see e.g., Bakshi and Panayotov 2013 for the definition). We found that in periods, where EUR is the investment currency the differences in implied default probabilities (USD–EUR) are high or zero. In periods, when EUR is the funding currency the differences are highly negative (see Table 9).

To further investigate this effect we included the differences in interest rates (difference in discount factors: USD–EUR, maturity 1 year) in our regression analyses (see Table 10). Again, the explanatory power of this additional variable was low and the regression coefficient was only at a low significance level different from zero. To sum up, there might be some pressure from carry traders on the price differences, but the pressure is not that influential.

Furthermore, we checked our dataset for a possible selection bias, regarding the question of why debtors issue bonds in different currencies. However, no indications

Table 10 Robustness test panel regression

	Model 5a	Model 5b
Constant	−0.51 (<0.001; 0.028)	−0.51 (<0.001; 0.028)
Corr. Surv. Prob. FX	−3.64 (<0.001; 0.038)	−3.64 (<0.001; 0.038)
Corr. Surv. Prob. FX × Corr. Disc Fac. FX	0.27 (0.0634; 0.176)	0.21(0.0931; 0.161)
Corr. Surv. Prob. FX × Corr. Disc. Fac. Surv.	1.42 (<0.001; 0.061)	1.42 (<0.001; 0.061)
Coupon	0.09 (<0.001; 0.006)	0.09 (<0.001; 0.006)
Coll. action clauses	0.19 (<0.001; 0.032)	0.19 (<0.001; 0.032)
Liquidity	2.34 (<0.001; 0.136)	2.33 (<0.001; 0.136)
Measurement error	0.29 (<0.001; 0.013)	0.29 (<0.001; 0.013)
Discount factor (1 year) Dollar–EURO	−1.41 (0.051; 0.856)	
Skewness (0.5 years, historically)		−0.07 (0.0542; 0.046)
Number of observations	24,462	24,462
R^2	32.30 %	32.31 %
Durbin–Watson	1.99	2.00
LM-test (OLS)	3.98**	3.98**
LM heterosc.	0.92	0.97

The dependent variable in all columns is the price difference according to Eq. (9). The dependent variables are defined in Sect. 3.4, esp. Eq. (8) (*p* value; standard errors in brackets). The regression equation is given in formula (10) extended by the robustness variables (model 5a and 5b). The sample (over the period 1998–2008) is drawn from databases described in Sect. 3. In the second to last line the modified Durbin–Watson statistic after correcting for autocorrelation using Prais–Winsten statistic is depicted. In the last line, the value of the LM-test is presented to compare the random effects model to an OLS regression. The second LM test statistic is the test of Montes-Rojas and Sosa-Escudero (2011) for heteroscedasticity. Standard errors clustered at the firm level are applied to determine significance

** 5% significance level

for sequential issues of bonds in the respective currencies were found. To sum up, the results seems to be robust.

5 Conclusions

First, our innovation is the derivation of an explicit pricing formula for the price differences of Euro and USD denominated bonds of one issuer. Secondly, we were able to show that the theoretical derived most important determinant the correlation between exchange rate and default risk is highly significant in explaining those price differences.

To be more detailed, we firstly bridged an existing research gap where existing theoretical models were unable to draw a conclusion on the question of whether the variables of default, exchange and interest rate are dependent (Jankowitsch and Pichler 2005). By means of a quanto option priced in a Jarrow/Turnbull-model, we showed that, given a positive correlation between exchange rate and the event of default, the credit spreads in USD are statistically, significantly higher than in EUR and vice versa. Furthermore, we avoided the existing shortcomings related to the consideration of contractual specifications and the influence of the term structure estimation error in the empirical studies published so far (e.g., Kercheval et al. 2003; McBrady 2003), by taking into account different contractual specifications and the influence of the term structure estimation error. Moreover, due to the drawbacks of the applied theoretical models, we also considered the influence of the correlation structure. The influence of the correlation structure turned out to be highly significant. Additionally, we tested the influence of the correlation between default and interest rates and the correlation between interest rates and FX rate. Whereas the latter was most of the time insignificant, the former shows significantly enlarged differences in bond prices as well as in implied default probabilities.

Additionally, following the example of Elton et al. (2001), Driessen (2005), Sener and Kenc (2008) and Van Landschoot (2008), we found out that other factors of influence are (slightly) responsible for the different credit spreads. In this context, we took the influence of taxes (e.g., Elton et al. 2004) and liquidity risk (e.g., Driessen 2005; Van Landschoot 2008; Chen et al. 2007) as control variables and obtained results similar to those in the cited literature. However, it must be added that the effects of liquidity and taxes could only be occasionally proved. Furthermore, we initially considered different contractual specifications and the term structure estimation error, but in fact these two factors only seem to play a minor role.

As a large part of the results could be explained by the dependence between the event of default and exchange rate, the findings of the empirical study lend support to the validity of the efficient market hypothesis. Therefore, our results partially counter the findings of Buraschi et al. (2013). Maybe these differences can be attributed to a different time span being considered.

For future research, it would be very interesting to analyse more intensively the influence of the volatilities of exchange and interest rates as well as the role of the default risk. For the latter an extended data set would be needed, which contains issuers with low as well as high credit risk.

Furthermore, it would also be interesting to consider corporate bonds denominated in different (foreign) currencies from one issuer and to take more contingent factors of influence into account. In addition, further research should deal with a possible economic explanation for the ascertained correlation between the event of default and the exchange rate. Possible explanations, for example, can be found by considering the balance of trades for the respective countries; however, in our context, these reflections do not provide an appropriate explanation.

Acknowledgments This paper has already been presented—among others—at the following conferences: 18th Annual Conference of the Multinational Finance Society; 24th International Conference of the European Financial Management Association; 28th International Conference of the French Finance Association; 43rd International Conference on Operations Research. Furthermore, the authors would like to thank Günter Bamberg, Gonzalo Cortazar, Silva Florinda, Rainer Jankowitsch, William F. Sharpe, Manfred Steiner, Martin Wallmeier and Marco Wilkens for their helpful comments.

Appendix 1

Table 11 Correlation coefficients between implied survival probability and cross country differences in discount factors for riskless money market accounts as well as exchange rates and differences in discount factors for riskless money market accounts

	Correlation between	Maturity				
		2 years	4 years	6 years	8 years	10 years
Turkey	USD denominated bonds ISP and Cumulative riskless money market sdf	−0.1803	−0.1180	−0.0660	−0.0087	−0.0413
Turkey	EUR denominated bonds ISP and Cumulative riskless money market sdf	−0.0822	−0.0540	−0.0304	−0.0415	−0.2662
Poland	USD denominated bonds ISP and Cumulative riskless money market sdf	0.2583	0.2665	0.2558	0.2232	0.1333
Poland	EUR denominated bonds ISP and Cumulative riskless money market sdf	−0.0585	−0.0615	0.0292	−0.4225	0.3124
Mexico	USD denominated bonds ISP and Cumulative riskless money market sdf	0.1952	0.2345	0.2616	0.3055	0.3391
Mexico	EUR denominated bonds ISP and Cumulative riskless money market sdf	−0.0470	−0.0986	−0.1756	−0.2228	−0.2531
Brazil	USD denominated bonds ISP and Cumulative riskless money market sdf r	0.1855	0.1856	0.1937	0.2185	0.2526
Brazil	EUR denominated bonds ISP and Cumulative riskless money market sdf	−0.0310	−0.0558	−0.0841	−0.1146	−0.1427
	Dollar and Cumulative riskless money market sdf	0.4021	0.3732	0.3301	0.2547	0.1298
	Euro and Cumulative riskless money market sdf	−0.4021	−0.3732	−0.3301	−0.2547	−0.1298

The table contains the average correlation coefficient between the implied survival probability and cross country differences in discount factors for formula (5) in Appendix P.2. Therefore, the formula differs using direct or indirect quotation. Furthermore, we used at most a window of 2,500 trading days to estimate the correlation coefficients for different survival probabilities from 2 to 10 years. Additionally this table contains the average correlation coefficient between the exchange rate (indirect and direct quotation) and cross country differences in discount factors (last two lines)

Table 12 Key findings of single regressions (theoretical direction +/−)—(I)

	Constant			Coupon (−)			Correlation (−)			Liquidity (−)			CAC (−)			Error (+)			R^2
	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.
Single model ("two currencies")																			
Brazil ($)	21.63	46.0	6.7	−0.05	2.4	8.7	−31.91	6.2	50.4	0.68	23.2	13.0	−0.03	1.4	0.2	0.59	66.7	0.2	87.8
Brazil (€)	−21.61	3.4	36.7	0.68	85.9	4.8	−33.69	4.8	39.1	1,407.96	5.8	12.2	Infeasible			0.91	82.2	1.2	97.8
Mexico ($)	−1.24	18.5	22.1	−0.03	5.0	9.9	0.15	17.0	24.9	791.15	17.8	0.7	−0.36	2.1	20.4	0.12	9.3	4.8	70.7
Mexico (€)	−3.92	1.0	31.4	−0.78	2.2	13.0	−16.35	0.2	39.3	−217.86	5.7	10.5	−5.03	0.2	26.4	−1.58	4.8	7.0	94.2
Poland ($)	−7.23	13.4	30.9	Infeasible			−8.63	17.3	25.4	Infeasible						0.33	5.3	0.0	83.1
Poland (€)	−2.14	0.5	20.6	0.15	12.5	0.2	−5.42	0.5	31.2	−2,954.19	4.2	21.1	Infeasible			0.33	1.5	3.1	72.6
Turkey ($)	2.74	31.7	0.7	−0.13	2.9	20.9	−6.16	0.3	93.3	−506.73	2.6	14.1	−0.11	1.5	6.0	0.96	43.2	0.9	96.0
Turkey (€)	3.58	8.9	1.0	−0.14	0.7	7.2	1.91	7.7	2.2	4,078.54	10.8	1.5				0.38	3.3	4.1	66.9
Total model ("one currency")																			
Brazil	−1.82	0.0	1.0	0.25	4.6	0.2	−4.08	0.0	88.0	−3.36	15.4	2.7	−1.91	0.2	14.4	0.46	22.3	0.0	73.2
Mexico	−0.09	0.9	1.2	−0.01	1.0	1.5	−3.36	0.9	84.7	−57.00	0.2	1.7	−0.29	0.3	2.9	0.33	9.9	0.7	73.9
Poland	−0.94	24.7	37.7	0.21	36.0	27.1	−2.44	0.0	93.3	−3.81	54.4	17.5	−0.12	8.9	5.1	−0.06	5.0	4.6	93.5
Turkey	−0.17	5.5	6.2	−0.02	1.9	10.1	−3.61	0.0	98.8	−586.83	8.9	3.1	−0.11	0.3	1.2	0.63	18.2	1.2	94.0
SUR																			
Brazil ($)	16.78	57.0	19.8	−0.03	23.5	38.0	−27.68	17.4	61.5	−1.12	48.5	26.8	Infeasible			0.44	78.2	6.2	95.7
Brazil (€)	−21.60	8.9	81.6	0.60	78.7	8.9	−40.85	16.2	73.2	2,054.13	37.3	40.5	Infeasible			0.79	57.7	18.2	95.7
Mexico ($)	−1.15	28.0	38.3	−0.03	34.0	44.7	−2.00	24.4	28.4	603.68	30.3	45.7	−0.38	64.9	9.1	0.12	29.6	23.9	92.5

Table 12 continued

	Constant			Coupon (−)			Correlation (−)			Liquidity (−)			CAC (−)			Error (+)			R^2
	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.
Mexico (€)	−3.95	18.9	61.9	−0.26	17.4	54.3	−14.34	1.5	76.3	−61.08	33.0	34.7	−3.22	6.9	70.6	−0.39	14.6	48.8	92.5
Poland ($)	−4.85	14.8	44.8	Infeasible			−3.74	17.9	28.4	Infeasible						1.09	21.1	1.7	74.4
Poland (€)	−1.53	0.2	15.8	Infeasible			−4.18	0.3	26.6	Infeasible						0.77	5.3	0.2	74.4
Turkey ($)	1.90	66.2	9.1	−0.08	16.7	55.3	−5.49	1.5	94.2	−382.40	22.0	44.7	Infeasible			0.91	69.1	5.8	98.1
Turkey (€)	−0.47	31.4	28.4	−0.11	10.0	41.9	−4.74	24.4	38.3	1,818.68	44.7	17.9	Infeasible			0.39	33.2	22.9	98.1

The table contains the average height of regression coefficients for all regressions respective to trading days (evaluation period: 2006–2008). Furthermore, the table depicts the proportion of significant (significance level: 5 %) regression coefficients among the number of total regressions respective to trading days, wherein the results are differentiated between positive and negative coefficients. The sign of the coefficients derived by the hypotheses are given in the first line. Lines 4–11 represent the results of the single model according to Eq. (11). Lines 13–16 represent the results of the total model according to Eq. (12). Lines 18–25 represent the results of the SUR model according to Eqs. (13) and (15). The last column depicts the average R^2 over all regressions. Some variables were dropped due to collinearity. At least tests for the estimation error regarding the independent variable error and correlation: Hausmann Specification Test; Heteroscedasticity: Breusch–Pagan test; Multicollinearity: Variance Inflation Factor

Table 13 Key findings of single regressions with duration (theoretical direction +/−)—(II)

	Constant			Coupon (−)			Duration ($ −/€ +)			Liquidity (−)			CAC (−)			Error (+)			R^2
	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver. (%)
Single model (“two currencies”)																			
Brazil ($)	0.52	6.5	8.2	−0.05	2.4	4.3	−0.38	6.9	58.8	0.69	25.0	7.5	−0.03	1.2	0.3	0.65	71.7	0.3	90.9
Brazil (€)	−7.63	11.4	35.6	0.64	85.7	1.2	0.60	83.3	7.7	1,337.47	47.8	11.5	Infeasible			0.92	84.2	1.7	97.8
Mexico ($)	−1.59	1.4	18.9	−0.03	4.5	8.7	−0.01	16.0	25.0	752.04	19.9	0.3	−0.34	1.2	17.0	0.04	8.9	5.0	69.6
Mexico (€)	−4.21	0.9	5.5	0.85	7.6	0.3	−0.45	0.2	8.3	−492.53	1.2	6.0	4.92	7.6	0.3	3.00	7.1	0.5	87.5
Poland ($)	−1.79	36.8	47.0	Infeasible			−0.003	40.1	51.9	Infeasible						1.12	91.8	0.7	99.5
Poland (€)	−0.63	1.2	8.2	0.22	19.4	0.7	0.55	72.2	0.2	−1,339.08	11.1	3.4	Infeasible			0.13	15.3	5.0	98.4
Turkey ($)	2.33	9.1	1.4	−0.20	0.5	11.5	−0.61	0.3	72.7	−401.22	4.1	10.1	−0.30	0.7	5.8	1.25	36.7	1.5	91.2
Turkey (€)	−0.76	2.4	17.5	0.01	11.1	3.1	1.13	72.2	0.9	−42.93	7.5	9.8	Infeasible			0.93	41.5	0.3	96.8
Total model (“one currency”)																			
Brazil	1.24	0.9	0.5	−0.16	1.2	1.2	0.49	31.0	0.5	−7.76	0.7	39.8	−1.82	0.2	16.8	−0.64	2.6	36.5	54.3
Mexico	−0.71	0.0	0.2	−0.001	0.0	1.0	0.03	0.2	0.0	238.05	0.0	0.2	−0.27	0.0	1.4	1.27	1.5	0.2	11.9
Poland	4.36	66.7	0.2	−0.56	1.4	57.3	−0.10	0.7	20.2	−64.70	39.5	4.8	−3.60	0.0	94.2	0.97	3.8	4.8	79.7
Turkey	8.67	92.3	0.0	−0.73	0.0	97.3	−0.35	0.0	25.7	−1,399.82	1.0	7.4	−2.63	0.0	93.1	−0.53	1.0	12.0	73.4
SUR																			
Brazil ($)	0.36	34.7	29.0	−0.04	15.3	37.1	−0.4	14.8	67.4	169.39	48.4	21.6	Infeasible			0.55	86.9	2.9	95.7
Brazil (€)	−6.31	10.3	78.6	1.20	77.2	12.6	0.86	75.2	15.3	−333.26	35.2	41.9	Infeasible			0.73	59.8	16.2	95.7
Mexico ($)	−1.92	6.2	49.7	0.04	0.2	0.3	−0.05	15.0	38.5	646.53	91.9	2.1	−0.09	29.2	40.9	0.1	23.3	13.3	92.7
Mexico (€)	−3.43	7.4	40.9	0.76	56.6	4.5	−0.36	30.2	39.0	−627.16	3.6	56.5	3.55	47.2	8.1	1.8	42.0	14.9	92.7

Table 13 continued

	Constant			Coupon (−)			Duration ($ −/€ +)			Liquidity (−)			CAC (−)			Error (+)			R^2
	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver.	Pos. (%)	Neg. (%)	Aver. (%)
Poland ($)	−1.99	11.8	38.5	Infeasible			−0.05	31.0	32.1	Infeasible						1.26	82.6	9.6	97.5
Poland (€)	0.33	70.4	22.9	Infeasible			0.55	99.1	0.4	Infeasible						1.07	89.5	4.9	97.5
Turkey ($)	0.52	23.4	11.2	−0.07	8.6	27.7	−0.51	2.6	90.2	−127.47	25.8	29.0	Infeasible			1.48	59.8	7.4	97.5
Turkey (€)	1.47	33.3	5.3	−0.09	13.4	14.4	0.22	60.5	3.6	1,059.71	29.2	18.9	Infeasible			0.83	32.6	11.7	97.5

The table contents the average height of the regression coefficients for all regressions respectively trading days (Evaluation Period: 1998–2008). Furthermore, the table depicts the proportion of significant (significance level: 5 %) regression coefficients among the number of total regressions respectively trading days, whereby the results are differentiated between positive and negative coefficients. The sign of the coefficients derived by the hypotheses are given in the first line. Lines 4–11 represent the results of the single model according to Eq. (11). Lines 13–16 represent the results of the total model according to Eq. (12). Lines 18–25 represent the results of the SUR model according to Eqs. (13) and (15). The last column depicts the average R^2 over all regressions. Some variables have to be dropped due to collinearity. At least tests for the estimation error regarding the independent variable error and correlation: Hausmann Specification Test; Heteroscedasticity: Breusch–Pagan test; Multicollinearity: Variance Inflation Factor

Appendix 2: Proofs

Proof 1 (P.1.) Estimation of the covariance between the variables exchange rate e and survival ā:

$$Cov(e, \bar{a}) = \iint (e - \bar{e})\,(I_{\bar{a}} - E(I_{\bar{a}}))q(e, \bar{a})ded\bar{a}$$

$$= \iint eI_{\bar{a}}q(e, \bar{a})ded\bar{a} - \bar{e}E(I_{\bar{a}}) = \iint e1q(e\,|\bar{a})q(\bar{a})ded\bar{a} + 0 - \bar{e}E(I_{\bar{a}})$$

$$Cov(e, \bar{a}) > 0 \Leftrightarrow \int eq(e\,|\bar{a})deQ(\bar{a}) > \bar{e}E(I_{\bar{a}}) = e_f Q(\bar{a}) \tag{15}$$

with

e = exchange rate at T (DCU/FCU),
ē = the mean of e at T,
I = Indicator function,
ā = no default (survival) and
q = the martingale measure.

Contingent forward contract in case of independence between the exchange rate and survival:

$$e_f Q(\bar{a}) - \int eq(e\,|\bar{a})deQ(\bar{a})$$

$$= e_f Q(\bar{a}) - \int eq(e)deQ(\bar{a}) = e_f Q(\bar{a}) - e_f Q(\bar{a}) = 0 \tag{16}$$

with

e_f = the forward price in case of fixed interest rates
ē = the mean e under the martingale measure and
Q = the martingale probability of survival.

The price at T of the conditional forward contract in case of dependence between the variables exchange rate e and survival ā:

$$e_f Q(\bar{a}) - \int eq(e\,|\bar{a})deQ(\bar{a}) > 0 \Leftrightarrow Cov(e, \bar{a}) < 0 \tag{17}$$

Conditional forward contract in case of dependence between the variables exchange rate e and survival ā (Conditional independent interest rates):

$$e_f Q(\bar{a}) - \iiint eq^T(d)q^T(e\,\big|\bar{a})q^T(\bar{a})dddedā = e_f Q(\bar{a})$$

$$- \iint \frac{d(T)}{E_Q(d(T))}q(d)eq(e\,|\bar{a})dddeQ(\bar{a})$$

$$= e_f Q(\bar{a}) - \int \frac{d(T)}{E_Q(d(T))}q(d)dd \int eq(e\,|\bar{a})deQ(\bar{a}) = e_f Q(\bar{a})$$

$$-\int eq(e\,|\bar{a})deQ(\bar{a})\ >0 \Leftrightarrow Cov<0 \tag{18}$$

with q^T = the forward measure for time T and $\frac{dQ}{dQ^T} = \frac{d(T)}{E_Q(d(T))}$= the Radon-Nikodym-derivative of the martingale measure in relation to the forward measure. □

Proof 2 (P.2.) Estimation of the covariance between the variables discount factor for interest rate differences d and survival a:

$$\begin{aligned}
&Cov(d,\bar{a}) = \iint (d(T) - E_Q(d(T))\,(I_{\bar{a}} - E(I_{\bar{a}}))q(d,\bar{a})ddd\bar{a}\\
&= \iint d(T)I_{\bar{a}}q(d,\bar{a})ddd\bar{a} - E_Q(d(T))E(I_{\bar{a}}) = \iint d(T)1q(d\,|\bar{a}\,)q(\bar{a})ddd\bar{a}\\
&+0 - B(T)E(I_{\bar{a}})\\
&Cov(d,\bar{a}) > 0 \Leftrightarrow \int d(T)q(d\,|\bar{a}\,)ddQ(\bar{a}) > B(T)Q(\bar{a})
\end{aligned} \tag{19}$$

B(T) = price of zerobond (rate = difference of interest rates)

Conditional forward contract in case of dependence between the variables Interest Rate and Survival (independent variable: exchange rates) and conditional independence.

$$\begin{aligned}
&\iint e_f q^T(d\left|\bar{a}\right.)q^T(\bar{a})ddd\bar{a} - \iiint eq^T(e)q^T(d\left|\bar{a}\right.)q^T(\bar{a})deddd\bar{a}\\
&= \int \frac{d(T)}{E_Q(d(T))}e_f q(d\,|\bar{a})dd\,Q(\bar{a}) - \iint \frac{d(T)}{E_Q(d(T))}q(e)eq(d\,|\bar{a})dedd Q(\bar{a})\\
&= e_f \int \frac{d(T)}{E_Q(d(T))}q(d\,|\bar{a})ddQ(\bar{a}) - \int eq(e)de \int \frac{d(T)}{E_Q(d(T))}q(d\,|\bar{a})ddQ(\bar{a})\\
&= e_f \int \frac{d(T)}{E_Q(d(T))}q(d\,|\bar{a})ddQ(\bar{a}) - \frac{\bar{e}}{E_Q(d(T))}\int d(T)q(d\,|\bar{a})ddQ(\bar{a})\\
&= \frac{e_f}{E_Q(d(T))}\int d(T)q(d\,|\bar{a})ddQ(\bar{a}) - \frac{e_f}{E_Q(d(T))}\int d(T)q(d\,|\bar{a})ddQ(\bar{a}) = 0
\end{aligned} \tag{20}$$

The last equation uses the result, that the mean under the martingale measure is the forward price, if the variables fx and interest rates are independent. □

Proof 3 (P.3.) Conditional independence between the variables exchange rate e and discount factor for interest rate differences d:

$$\begin{aligned}
&\iint e_f q^T(d\left|\bar{a}\right.)q^T(\bar{a})ddd\bar{a} - \iiint eq^T\,(e|\,\bar{a})\,q^T\,(d|\,\bar{a})\,q^T(\bar{a})deddd\bar{a} =\\
&\int \frac{d(T)}{E_Q(d(T))}e_f q(d\,|\bar{a})dd\,Q(\bar{a}) - \iint \frac{d(T)}{E_Q(d(T))}eq\,(e|\,\bar{a})\,q\,(d|\,\bar{a})\,deddQ(\bar{a}) =\\
&e_f \int \frac{d(T)}{E_Q(d(T))}q(d\,|\bar{a})ddQ(\bar{a}) - \int eq\,(e|\,\bar{a})\,deQ(\bar{a}) \int \frac{d(T)}{E_Q(d(T))}q(d\,|\bar{a})dd =
\end{aligned}$$

$$Cov(e,\bar{a})\int \frac{d(T)}{E_Q(d(T))} q(d\,|\bar{a})dd \tag{21}$$

With

$$Cov(d,\bar{a}) > 0 \Leftrightarrow \int d(T)q(d\,|\bar{a}\,)ddQ(\bar{a}) > B(T)Q(\bar{a})$$

we get

$$Cov(e,\bar{a})\int \frac{d(T)}{E_Q(d(T))} q(d\,|\bar{a})dd \; > Cov(e,\bar{a})$$

in the independent case.

$$\begin{aligned}
Cov(d,e) &= \iint (d(T) - E_Q(d(T))\,(e-\bar{e})q(d,e)ddde \\
&= \iint d(T)eq(d,e)ddde - E_Q(d(T))\bar{e} = \iint d(T)eq(d\,|e\,)q(e)ddde - B(T)\bar{e} \\
Cov(d,e) &> 0 \Leftrightarrow \iint d(T)eq(d\,|e\,)q(e)ddde > B(T)\bar{e} \Leftrightarrow e_f > \bar{e}
\end{aligned} \tag{22}$$

Conditional independence between the variables survival a and discount factor for interest rate differences d:

$$\begin{aligned}
&\iint e_f q^T(d) q^T(\bar{a})ddd\bar{a} - \iiint eq^T\,(e|\,\bar{a})\,q^T\,(d|\,e)\,q^T(\bar{a})dedd d\bar{a} \\
&= \int \frac{d(T)}{E_Q(d(T))} e_f q(d)ddQ(\bar{a}) - \iint \frac{d(T)}{E_Q(d(T))} eq\,(e|\,\bar{a})\,q\,(d|\,e)\,deddQ(\bar{a}) \\
&= e_f Q(\bar{a}) - \left(\frac{Cov(d,e\,|\bar{a}\,)}{E_Q(d(T))} + E\,(e\,|\bar{a}\,)\right) Q(\bar{a}) \\
&= Q(\bar{a})\left(e_f - E\,(e\,|\bar{a}\,) - \frac{Cov(d,e\,|\bar{a}\,)}{B(T)}\right) = Q(\bar{a})\left(e_f - E\left(e_f\,|\bar{a}\right)\right)
\end{aligned} \tag{23}$$

If $Cov(d,e\,|\bar{a}\,) < 0$, the value is increasing. □

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