Essays on Foreign Exchange and Credit Risk

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ISSN 0906-6934

Print ISBN: 978-87-93579-98-9
Online ISBN: 978-87-93579-99-6

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PhD Series 26-2018

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PhD School in Economics and Management

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The PhD School in Economics and Management is an active national and international research environment at CBS for research degree students who deal with economics and management at business, industry and country level in a theoretical and empirical manner.

Foreword

This thesis is the result of my PhD studies at the Department of Finance at Copenhagen Business School, and it consists of summaries in English, Danish, and three self-contained essays on foreign exchange and credit risk which can be read independently. I gratefully acknowledge the financial support of the Center for Financial Frictions (FRIC), grant no. DNRF102.

I have benefited greatly from advice, discussions, and suggestions from a number of people over the years. In particular, I would like to thank my supervisors David Lando and Christian Wagner. I am indebted to David Lando for being a great supervisor and mentor that helped me grow as an academic. I truly appreciate his encouragement and support over the years. Also, I would like to thank him for a great collaboration on the first essay—I have learned a lot from the process. I would like to give special thanks to Christian Wagner for the support and detailed and honest feedback. His reflections and great comments on my research helped me to sharpen and clarify my ideas. The quality of this thesis has benefited greatly from his help.

A number of people deserve a special acknowledgement. Mike Chernov for sponsoring my visit at UCLA Anderson School of Management and for taking his time to discuss my research; Peter Christoffersen for much appreciated help and feedback; my fellow PhD students and colleagues that made work a pleasant and joyful experience. Finally, but not least, I would like to thank friends and family for their endless support and for bearing with me in difficult and stressful times.

Andreas Bang Nielsen

Copenhagen, April 2018

Summary

Summary in English

Essay 1: Quanto CDS Spreads (co-authored with David Lando)

We investigate how currency denomination affects the price of credit risky securities of the same issuer. We focus on eurozone sovereign quanto spreads, i.e., differences in credit default swap (CDS) premiums denominated in U.S. dollar and Euro of the same reference entity. Quanto spreads of eurozone sovereigns reached unprecedented levels during the European debt crisis and have remained significant ever since. Quanto spreads do not simply reflect differences in contractual terms linked to currency denomination, because CDS contracts trade under the same standardized terms independent of currency denomination, including credit events and recovery rates.

In order to understand which factors drive quanto spreads, we propose a no-arbitrage model that shows in a simple and rigorous manner that quanto spreads arise without any market frictions through two risk channels.

The first channel, currency crash risk, reflects the risk of an adverse jump in domestic versus foreign currency triggered by default of the reference entity. Intuitively, currency crash risk causes the expected recovery payment to be relatively smaller on the domestic CDS compared to the foreign CDS, because the recovery payment on the domestic contract is received in the 'crashed' currency.

The second channel, covariance risk, contributes to quanto spreads through covariance between the exchange rate and default risk of the reference entity. The intuition for how this channel works is as follows. If default risk rises (falls) CDS premiums increase (decrease) in both foreign and domestic currency, i.e., there is a gain (loss) on a long CDS position

in the relevant currency. However, if foreign currency tends to appreciate (depreciate) versus domestic currency when credit risk increases (decreases) then the gain (loss) is largest (smallest) on the foreign CDS. Foreign CDS protection is therefore more valuable than domestic protection since it has larger expected gains and smaller expected losses, implying a positive quanto spread caused by covariance risk.

Guided by the insights of our simple model, we propose an affine term structure model that captures both crash risk and covariance risk. We estimate the model to quanto CDS data for Italy, Spain, Portugal, and Ireland. Our estimations show that the EURUSD is expected to jump more if Spain and Italy were to default compared to if Portugal and Ireland were to default. We document that crash risk accounts for most of the quanto spreads at shorter maturities and that the covariance risk component embedded in quanto spreads increases in maturity. Covariance risk is particularly important in times of distress, when credit risk and exchange rate risk are volatile and co-vary strongly, while crash risk is important throughout the sample period.

Finally, we document that yield spreads between bonds denominated in U.S. dollar and Euro issued by eurozone sovereigns are significantly related to our estimated model-implied quanto yield spreads, especially during the peak of the European debt crisis. Our results indicate that a large portion of the differences in bond yields across currency denominations is caused by crash and covariance risk, and thus not solely by market imperfections, as previous research suggests.

Essay 2: Forward-Looking Currency Betas

This paper proposes a model-free method that uses currency option prices to compute risk exposures (betas) with respect to any currency factor. While traditional currency betas are based on exchange rate covariances estimated from historical data, the option-implied betas that I propose are based on exchange rate covariances derived from the most recent cross-section of currency option prices, without assuming any parametric structure on correlations. Typically, betas are estimated by means of rolling window regressions that are backward-looking, adjust slowly to new information, and the econometrician has to decide on which subset of the data to use for the estimation. In contrast, since the option-implied betas are inferred from the latest cross-section of option prices, they require neither historical data nor choices of estimation window and frequency—they are a market-based measure of betas.

I calculate currency betas by inferring the covariances between exchange rates from options on cross-pair exchange rates. For example, consider three currencies: the Euro, the British pound, and the U.S. dollar. Options exist on each pair-wise combination of these currencies. Specifically, the options on the Euro versus the British pound allow me to pin down the covariance between the Euro versus U.S. dollar and the British pound versus U.S. dollar, without assuming any parametric structure on their covariance. Using the same procedure for any other pair of currencies against the U.S. dollar, I calculate the full exchange rate covariance matrix from which betas with respect to currency portfolios can be derived.

In order to test the empirical properties of the option-implied betas compared to traditional rolling window betas, I use the dollar factor—an equally weighted portfolio of the G10 currencies against the U.S. dollar—as the systematic factor driving currency excess returns. I use the dollar factor because it captures the aggregate level of foreign currencies versus the U.S. dollar, i.e., it is essentially the market portfolio of foreign currencies from the perspective of a U.S. investor and, more importantly, because it has been documented by Lustig, Roussanov, and Verdelhan (2011, 2014) to carry a significant risk premium.

For both types of betas, I separately construct portfolios of currencies sorted by their dollar factor betas. I identify a significant positive relation between option-implied portfolio betas and ex-post portfolio returns, whereas there is an insignificant relation when using rolling window betas. Interestingly, this is because the option-implied betas predict currency spot changes and not because of the interest rate component of the portfolio returns, which is the most typical source of excess returns for currency strategies. Furthermore, I provide evidence that the model prediction errors of portfolio excess returns are significantly smaller when using option-implied betas as inputs in the model compared to using rolling window betas.

Finally, I find that option-implied betas are significantly better predictors of realized betas than rolling window betas at all horizons, both for portfolios and individual currencies. This finding strikes as a likely explanation for why option-implied betas are better in predicting currency excess returns than rolling window betas.

Essay 3: Systematic Currency Volatility Risk Premia

It has been documented in previous research that currency volatility risk premia are significantly negative on individual currencies, indicating that investors are willing to pay high premiums for insuring against currency volatility risk. In this paper, I investigate if currency volatility risk premia are explained by exposure to systematic variance risk. I propose a method for decomposing variances of exchange rates into a systematic component and an idiosyncratic component which I use to investigate the relation between systematic variance risk and returns for providing currency volatility insurance. The main result of the paper is that I uncover a negative relation between volatility excess returns and the proportion of systematic variance, suggesting that investors are more concerned with systematic variance risk vis-à-vis idiosyncratic variance risk.

More specifically, I assume that currency excess returns are driven by exposure to the dollar factor, that is, an equally weighted portfolio of G10 currencies versus the U.S. dollar. This factor structure in currency excess returns implies that currency variances can be decomposed into a dollar factor variance component (systematic variance) and an idiosyncratic variance component. I document that the dollar factor volatility risk premium is negative, on average, with an upward sloping and concave term structure, i.e., systematic volatility risk is particularly expensive to hedge at shorter maturities. Consistent with this pattern, I find that dollar factor variance risk is priced in the cross-section of currency volatility excess returns, but most significantly at shorter horizons.

For each currency, I calculate the systematic variance components and risk exposures using a model-free methodology based on currency options, i.e., the systematic variance risk components are inherently forward-looking. I then build portfolios of volatility swaps and forward volatility agreements (FVAs) constructed based on their share of systematic variances. I find that a systematic volatility factor (SYS factor) that buys (sells) volatility protection on currencies with the smallest (largest) shares of systematic variance delivers significant mean excess returns and high Sharpe ratios, especially at shorter maturities.

For example, the monthly mean excess return of the SYS factor based on 1-month volatility swaps is 4.47% with an annualized Sharpe ratio of 0.71. The SYS factor constructed based on FVAs in which the forward contract and its underlying volatility has a 1-month maturity delivers a monthly mean excess return of 2.73% with an annualized Sharpe ratio of

0.95. At shorter maturities, the excess returns of the SYS factor cannot be attributed to exposure to traditional currency factors, equity factors, or the volatility carry factor proposed by Della Corte, Kozhan, and Neuberger (2017).

Summary in Danish

Essay 1: Quanto CDS-Spænd (med David Lando)

Vi undersøger hvordan valutadenominering påvirker prisen på kreditrisikofyldte aktiver på samme udsteder. Vi fokuserer på quanto-spænd, dvs. forskelle i credit default swap (CDS) præmier denomineret i amerikanske dollar og euro på samme udsteder. Quanto-spændene på europæisk statsgæld nåede hidtil usete niveauer under den Europæiske gældskrise og har været betydelige siden da. Quanto-spænd afspejler ikke blot forskelle i kontraktvilkår knyttet til valutadenominering, fordi CDS-kontrakterne handler under de samme standardiserede vilkår, uafhængig af valutadenominering, herunder kreditbegivenheder og udbetalingsrate per enhed hovedstol i tilfælde af fallit.

For at forstå, hvilke faktorer der driver quanto-spænd, foreslår vi en ingen-arbitrage model, der viser på en simpel og stringent måde, at quanto-spænd opstår uden nogen markedsfriktioner gennem to risikokanaler.

Den første kanal, hopperisiko, afspejler risikoen for et negativt spring i den indenlandske valuta relativt til udenlandsk valuta, der er forårsaget af selve fallithændelsen for udstederen. Intuitivt betyder hopperisikoen, at den forventede udbetaling ved fallit er relativt mindre på de indenlandske CDS i forhold til de udenlandske CDS, fordi udbetalingen ved fallit på den indenlandske kontrakt betales i en devalueret valuta.

Den anden kanal, kovariansrisiko, bidrager til quanto-spændene gennem kovarians mellem valutakurs og udstederens fallitrisiko. Intuitionen for, hvordan denne kanal fungerer, er som følger. Hvis fallitrisikoen stiger (falder), så øges (falder) CDS-præmierne i både udenlandsk og indenlandsk valuta, dvs. der er en gevinst (tab) på en lang CDS-position i den relevante valuta. Men hvis den udenlandske valuta har tendens til at stige (falde) i forhold til indenlandsk valuta, når kreditrisikoen stiger (falder), så er gevinsten (tabet) størst (mindst) på den udenlandske CDS. Udenlandsk CDS-beskyttelse er derfor mere værdifuld end indenlandsk beskyttelse, da den har større forventede gevinster og mindre forventede tab, hvilket

forårsager et positivt quanto-spænd som følge af kovariansrisiko.

Baseret på vores indsigt opnået via den enkle model foreslår vi en affin model, der fanger både hopperisiko og kovariansrisiko. Vi estimerer modellen til quanto CDS data for Italien, Spanien, Portugal og Irland. Vores estimater viser, at EURUSD forventes at springe mere i tilfælde af hvis Spanien og Italien går fallit sammenlignet med tilfældet hvor Portugal og Irland går fallit. Vi dokumenterer, at hopperisikoen tegner sig for det meste af quanto-spændene på kortere løbetider, og at kovariansrisiko-komponenten, der er indlejret i quanto-spændende, stiger i løbetid. Kovariansrisiko er særlig vigtig når der er finansiel uro, dvs. når kreditrisiko og valutakursrisiko er volatile og korrelerer kraftigt, mens hopperisiko er vigtig i hele vores stikprøveperiode.

Endelig dokumenterer vi, at rentespænd mellem obligationer denomineret i amerikanske dollar og euro udstedt af eurozone stater er væsentligt relateret til vores estimerede quanto rentespænd, især på højdepunktet af den Europæiske gældskrise. Vores resultater tyder på, at en væsentlig del af forskellene i obligationsrenter på tværs af valutadenomineringer skyldes hopperisiko og kovariansrisiko, og dermed ikke udelukkende misprisninger i markedet, som tidligere forskning finder.

Essay 2: Fremadskuende Valuta-Betaer

I dette papir foreslås en modelfri metode, der bruger valutaoptionspriser til at beregne risikoeksponeringer (betaer) med hensyn til en hver given valutafaktor. Mens traditionelle valuta-betaer er baseret på valutakovarianser estimeret ud fra historiske data, så er de options-baserede betaer, som jeg foreslår, baseret på valutakovarianser, der stammer fra det seneste tværsnit af valutaoptionspriser uden at antage nogen parametrisk struktur på korrelationer.

Typisk estimeres betaer ved hjælp af rullende vinduesregressioner, der er bagudskuende, justerer langsomt til nye oplysninger, og derudover skal økonometrikeren tage stilling til hvilket data der skal anvendes til estimationen. I modsætning hertil stammer de options-baserede betaer fra det seneste tværsnit af optioner og kræver derfor ikke brug af historisk data eller valg af længden på det vindue og den datafrekvens, der bruges til estimationen.

Jeg beregner valuta-betaer ved at udlede kovarianserne mellem valutakurser fra optioner på krydspar valutakurser. For eksempel betragt tre valutaer: Euroen, det britiske pund og den amerikanske dollar. Der findes optioner på hver parvis kombination af disse valutaer. Specielt giver optionerne på euroen mod det britiske pund mig mulighed for at identificere kovariansen mellem euroen mod amerikansk dollar og det britiske pund mod amerikansk dollar uden at antage nogen form for parametrisk struktur på deres kovarians. Ved at anvende den samme procedure for ethvert andet par af valutaer mod amerikanske dollar beregner jeg hele valutakovariansmatricen, hvorfra betaer med hensyn til enhver valutaportefølje kan udledes.

For at undersøge de empiriske egenskaber ved options-baserede betaer sammenlignet med traditionelle betaer, bruger jeg dollarfaktoren—en ligevægtet portefølje af G10-valutaerne mod amerikanske dollar—som den systematiske faktor der driver valutamerafkast. Jeg bruger dollarfaktoren, fordi den reflekterer det samlede niveau af udenlandsk valuta i forhold til den amerikanske dollar, dvs. vi kan tænke på den som markedsporteføljen for udenlandskevalutaer set udfra en amerikansk investors perspektiv. En endnu vigtigere årsag, der lægger til grund for dette valg er at Lustig et al. (2011, 2014) dokumenterer at dollarfaktoeren bærer en betydelig risikopræmie.

For begge typer af beta konstruerer jeg porteføljer af valutaer sorteret efter deres dollar faktor betaer. Jeg identificerer en signifikant positiv sammenhæng mellem options-baserede betaer på porteføljerne og deres efterfølgende merafkast, mens der er en ubetydelig sammenhæng, når man bruger historiske betaer. Interessant nok skyldes det, at de options-baserede betaer forudsiger valutakursændringer for porteføljerne og ikke på grund af rentekomponenten i porteføljens afkast, hvilket er den mest typiske kilde til merafkast for valutastrategier. Desuden viser jeg, at modelforudsigelsesfejlene for porteføljeafkast er signifikant mindre, når der anvendes options-baserede betaer som input i modellen sammenlignet med hvis historiske betaer er anvendt som input i modellen.

Endelig finder jeg, at options-baserede betaer er betydeligt bedre forudsigere af realiserede betaer end historiske betaer på alle horisonter, både for porteføljer og individuelle valutaer. Dette fund forekommer som en sandsynlig forklaring på, hvorfor options-baserede betaer er bedre til at forudsige valutaafkast end historiske betaer.

Essay 3: Systematiske Valuta Volatilitetsrisikopræmier

Det er blevet dokumenteret i tidligere forskning, at valuta volatilitetsrisikopræmier er signifikante og negative for enkelte valutaer, hvilket indikerer, at investorer er villige til at betale høje præmier for at forsikre mod valutavolatilitet. I dette papir undersøger jeg, om valuta volatilitetsrisikopræmier kan forklares ved eksponering overfor systematisk variansrisiko. Jeg foreslår en metode til dekomponering af valutavarianser i en systematisk komponent og en idiosynkratisk komponent, som jeg bruger til at undersøge forholdet mellem systematisk variansrisiko og merafkast for at sælge forsikring på valutavolatilitet. Hovedresultatet i dette papir er, at jeg finder en faldende sammenhæng mellem volatilitetsmerafkast og andelen af systematisk variansrisiko, hvilket indikerer, at investorer er mere bekymret for systematisk variansrisiko end de er for idiosynkratisk variansrisiko.

Konkret antager jeg, at valutamerafkast er drevet af eksponering overfor dollarfaktoren, som er en ligevægtet portefølje af G10-valutaer i forhold til amerikanske dollar. Denne faktorstruktur i valutamerafkast afkast indebærer, at valutavarianser kan dekomponeres i en dollarfaktor variansekomponent (systematisk varians) og en idiosynkratisk variansekomponent. Jeg dokumenterer, at dollarfaktorens volatilitetsrisikopræmie er negativ i gennemsnit med en stigende og konkav løbetidsstruktur, dvs. systematisk volatilitetsrisiko er særlig dyr at afdække ved kortere løbetider. I overensstemmelse med dette mønster finder jeg, at dollarfaktorvariansrisiko er prissat i tværsnittet af volatilitetsmerafkast, i særlig grad på kortere horisonter.

For hver valuta beregner jeg de systematiske variansekomponenter og risikoeksponeringer ved hjælp af en modelfri metode baseret på valutaoptioner, dvs. de systematiske variansrisikokomponenter er fremadskuende. Jeg bygger derefter porteføljer af volatilitets swaps og forward volatility agreements (FVA'er) bygget ud fra deres andel af systematiske varianser. Jeg dokumenterer, at en systematisk volatilitetsfaktor (SYS-faktor), der køber (sælger) volatilitetsbeskyttelse på valutaer med de mindste (største) andele af systematisk varians, giver betydelige gennemsnitlige merafkast og høje Sharpe-ratios, især på kortere løbetider.

For eksempel er det månedlige gennemsnitlige afkast på SYS-faktoren baseret på 1-måneders volatilitets swaps 4,47% med en årlig Sharpe-ratio på 0,71. SYS-faktoren, bygget udfra FVA'er, hvor forward-kontrakten og den underliggende volatilitet har en løbetid på en måned leverer et månedligt gennemsnitligt afkast på 2,73% med en årlig Sharpe-ratio på 0,95. Ved kortere løbetider kan merafkastet på SYS-faktoren ikke tilskrives eksponering overfor traditionelle valutafaktorer, aktie-faktorer eller volatilitetsfaktoren foreslået af Della Corte, Kozhan, and Neuberger (2017).

Introduction

In the first essay, we investigate how currency denomination affects the pricing of credit risky securities by studying the case of eurozone sovereign quanto CDS spreads, that is, differences in credit default swap (CDS) premiums denominated in USD and EUR of the same issuer. Since the EUR and USD-denominated CDS contracts are issued under the same standardized terms—including identical recovery rates and trigger events—the quanto CDS spread is not due to contractual differences. Quanto CDS spreads therefore represent a clean way to study how currency denomination affects the pricing of credit risky securities and the interaction between foreign exchange rate risk and credit risk.

We develop a no-arbitrage discrete-time model that rationalizes quanto CDS spreads as compensation for risk through two channels. The first channel is currency crash risk, which reflects the risk of a jump in foreign currency (e.g., USD) versus domestic currency (e.g., EUR) in the event of a default. Intuitively, currency crash risk is priced in the quanto CDS spread because the expected recovery payment is larger on the foreign CDS compared to the domestic CDS since the domestic currency is expected to drop at default.

The second channel, covariance risk, reflects compensation for taking exposure to negative correlation between credit risk and foreign exchange rate risk. If credit risk rises (falls), it causes both domestic and foreign CDS premiums to go up (down), that is, a gain (loss) for the protection buyer of CDS in either currency. However, since domestic currency simultaneously tends to decrease (increase) relative to foreign currency when credit risk rises (falls), the gain (loss) is larger (smaller) on the foreign CDS. Therefore, the expected gains are smaller, and the expected losses are greater on the domestic CDS for a protection buyer, implying a positive quanto CDS spread. Moreover, we show that this channel has a larger effect on quanto CDS spreads the larger the expected volatility of currency risk and credit risk are. Our model shows that quanto CDS spreads at shorter maturities are primarily driven by crash risk, while the impact of covariance risk increases in maturity. We can therefore disentangle crash risk from covariance risk using the term structure of quanto CDS spreads.

Guided by the insights of the discrete-time model, we propose an affine term structure model that encompasses crash risk and covariance risk. We estimate the model to sovereign quanto CDS for Spain, Italy, Portugal, and Ireland, at maturities of 1-10 years. Furthermore, to get accurate assessments of the covariance risk components embedded in quanto CDS spreads, we use currency options to estimate forward-looking currency volatility risk.

We find that both covariance and currency crash risk are important contributors to quanto CDS spreads. We estimate the (risk-neutral) expected percent-wise jump in the EURUSD at sovereign default for Spain and Italy to 15.6% and 9.6%, significantly larger than the currency jump size of about 5% in the event of a Portuguese or Irish default. Our estimations show that covariance risk is most pronounced in times of financial distress, i.e., when the exchange rate and credit spreads are volatile and highly correlated. During the most severe period of the European debt crisis, we estimate the covariance components at the 5-year maturity to range from 18.4 bps to 35.6 bps, corresponding to 25%-58% of the average quanto CDS spreads. Without accounting for covariance risk, we would erroneously overestimate the implied jump size in the EURUSD upon sovereign default. Furthermore, consistent with our intuition from the discrete-time model, we find that crash risk accounts for a larger part of quanto CDS spreads at shorter maturities and that the contribution from covariance risk increases in maturity.

Finally, we use our estimated model to explain quanto bond yield spreads for Italy, Spain, and Portugal, which are differences in yields on USD and EUR-denominated bonds. From 2010-2013, i.e., at the peak of the European debt crisis, we provide evidence that our model-implied quanto bond yield spreads co-vary significantly with the observed quanto bond yield spreads, while in the post-crisis period they seem unrelated. Our results suggest that in times of market turmoil, crash risk and covariance risk are important determinants of yield spreads between EUR and USD-denominated eurozone sovereign bonds, implying that quanto bond yield spreads, at least partly, are attributable to risk and that they do not necessarily reflect market mispricings.

In the second essay, I propose a method for calculating forward-looking betas (risk exposures) with respect to factors constructed from currencies. I make use of a unique feature of currency option markets that allows me compute forward-looking covariances/variances for currencies. In particular, I exploit that there are options traded on each pair-wise combination of the G10 currencies, which I use to infer currency variances and correlations from which I derive currency betas.

The option-implied betas that I propose are inherently forward-looking and measured in real time. Whenever option prices change, the option-implied betas adjust immediately, and since the option prices are forward-looking, the option-implied betas are forward-looking as well. In contrast, betas calculated based on rolling window regressions (which is the most commonly used approach to calculate betas) are slow-moving and may not reflect current expectations about future betas over, say, the next month.

Purely forward-looking betas cannot be obtained in other major asset classes, for example for stocks, since there is no (liquid) market for options that depend on the price of two stocks. My contribution is important because asset prices reflect compensation based on expected future risk exposures, and not historical realizations of risk exposures that traditional methods offer.

In order to test the empirical properties of the option-implied betas compared to traditional rolling window betas, I use the dollar factor—which is an equally weighted portfolio of G10 currencies versus the U.S. dollar—as the systematic factor in currency excess returns. I use the dollar factor because it is well-documented that it carries a significant risk premium and because it reflects the aggregate level of foreign currencies from the perspective of a U.S. investor (Lustig, Roussanov, and Verdelhan (2011, 2014)). However, my methodology can be applied to any currency factor model.

I provide evidence that the option-implied dollar factor betas are significantly better predictors of realized dollar factor betas than rolling window dollar factor betas, both for betas of portfolios and for betas of individual currencies. Having established this fact, we would expect that option-implied betas are better in predicting currency returns, which is indeed what I find support for in the data. In order to compare the cross-sectional properties of the two types of betas, I construct monthly rebalanced portfolios of currencies sorted on betas, for each type of beta separately.

Lustig, Roussanov, and Verdelhan (2014) show that the dollar factor tends to appreciate (depreciate) whenever the average of short-term foreign interest rates is above (below) the short-term U.S. interest rate. Therefore, I construct the portfolios such that the investor goes long (short) in each portfolio whenever the average foreign interest rate is above (below) the U.S. interest rate. When sorting on the basis of option-implied betas, I find a significantly positive relation between ex-ante betas and ex-post portfolio returns, whereas there is an insignificant relation when the rolling window betas are used. Using the option-implied betas, a long-short portfolio that buys the upper tertile beta currencies and shorts the lower tertile beta currencies gives a significant annualized mean excess return of 3.35% (Sharpe ratio of 0.41), whereas it has an insignificant annualized mean excess return of 0.95% (Sharpe ratio 0.11) when sorting on rolling window betas.

Interestingly, the difference in mean excess returns on the long-short portfolio for the two types of beta stems from the spot component and not from the carry component (interest rate differential) of the portfolio excess returns, which is in contrast to the currency carry trade, where the excess returns primarily come from the interest rate component. This implies that option-implied betas outperform the rolling window betas for portfolio construction because they are better predictors of currency spot changes. Furthermore, I show that the model time-series prediction errors are smallest, on average, when using option-implied betas and that rolling window betas tend to underestimate low-beta portfolio returns and overestimate high-beta portfolio returns, while option-implied betas deliver unbiased predictions.

I provide evidence suggesting that a reasonable explanation for why the option-implied betas are better predictors of currency excess returns is because they are better in predicting realized betas, both for portfolios and individual currencies. Moreover, rolling window betas deliver biased forecasts; they underestimate (overestimate) betas for low-beta (high-beta) portfolios, while the option-implied betas deliver virtually unbiased predictions.

In the third essay, I study if risk premia associated with currency volatility risk are attributable to exposure to systematic variance risk. The main objective of the study is to investigate if the large volatility excess returns for individual currencies that have been documented in previous research are driven primarily by systematic variance risk. To this end, I propose a simple method for decomposing variances of exchange rates into systematic and idiosyncratic variance risk, which I use to empirically investigate the relation between systematic variance risk and volatility excess returns. Specifically, I assume that currency excess returns are driven by exposure to the dollar factor which implies that currency variances consist of a variance component stemming from exposure to dollar factor variance risk (systematic variance risk) and an idiosyncratic variance risk component.

Because exposure to dollar factor variance risk is the source of volatility excess returns under my hypothesis, I begin the empirical analysis by establishing a number of stylized facts about the volatility risk premium on the dollar factor. The dollar factor volatility risk premium is, on average, negative and tends to have an upward sloping and concave term structure, i.e., it is steep at the short end and virtually flat at longer maturities. This pattern indicates that investors are willing to pay for hedging systematic volatility risk but that they are more concerned with short-term systematic volatility risk relative to long-term systematic volatility risk.

The factor structure in currency excess returns allows me calculate forward-looking measures of the systematic variance components by using the option-implied dollar factor betas and variances that I proposed in the second essay. Using this methodology for calculating systematic variance risk, I find a negative relation between the (expected) share of systematic variance and realized volatility excess returns, i.e., excess returns on volatility swaps and forward volatility agreements (FVAs). As a consequence, it has been profitable for investors to sell volatility protection on currencies with a high share of systematic variance and buy volatility protection on currencies with a low share of systematic variance.

For example, the monthly mean excess return of a long-short portfolio of 1-month volatility swaps based on the share of systematic variance is 4.47% with an annualized Sharpe ratio of 0.71, and for FVAs, in which the forward contract and volatility have a 1-month maturity, the monthly mean excess return is 2.73% with an annualized Sharpe ratio of 0.95. At shorter maturities, the excess returns of the long-short systematic variance risk portfolios cannot be explained by exposure to traditional currency factors, equity factors, or the volatility carry factor proposed by Della Corte, Kozhan, and Neuberger (2017).

Contents

Essay 1

Quanto CDS Spreads

Quanto CDS Spreads

David Lando and Andreas Bang Nielsen*

Abstract

Quanto CDS spreads are differences in CDS premiums of the same reference entity but in different currency denominations. Such spreads can arise in arbitrage-free models and depend on the risk of a jump in the exchange rate upon default of the underlying and the covariance between the exchange rate and default risk. We develop a model that separates the contribution of these two effects to quanto spreads and apply it to four eurozone sovereigns. Furthermore, using our model estimates, we provide evidence that quanto effects can explain a significant part of the yield spread between eurozone sovereign bonds issued in Euro and U.S. dollar. Our findings suggest that comparing bond yields across currency denominations using standard FX forward hedges misses an important quanto effect component.

Keywords: Sovereign credit risk, CDS premiums, currency risk, systemic risk
JEL Codes: H63, G13, F31, G01

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1.1 Introduction

During the European debt crisis, the European sovereign credit market experienced tremendous distress with sovereign credit spreads widening to unprecedented levels. But not only did the levels of CDS premiums for sovereigns spike; the difference between CDS premiums on European sovereigns denominated in EUR and USD, the so called quanto spread, also increased significantly. The 5-year quanto spread reached 95 bps for Italy, 105 bps for Spain and 145 bps for Portugal and it has continued to be substantial after the crisis. Since the EUR and USD-denominated CDS contracts are issued under the same standardized ISDA terms—including same recovery rate and trigger events—the quanto spread is not due to contractual differences.

It is well known that quanto spreads can arise without any frictions. If there is a risk of a crash in the exchange rate coinciding with default of the reference name of the CDS, then this leads to a quanto spread. It is less obvious, and seemingly less recognized, that correlation between FX-rate fluctuations and the default intensity of the reference name also leads to a quanto spread, and that this contribution to the spread can arise even if there is no depreciation of one currency in the event of default. An accurate assessment of currency crash risk in the event of default from quanto spread requires a correction for this correlation effect.

We propose here a simple two-factor discrete-time model in which the effects can be understood simply and rigorously. The first factor, the FX crash risk factor, captures the market's (risk-neutral) anticipation of a jump in foreign currency (EUR) against domestic currency (USD) in the event of a sovereign default. If crash risk is present, it implies a smaller expected recovery on a EUR contract relative to a similar USD contract and thus causes protection in USD to be more expensive. The second factor, the currency/default risk covariance factor, captures the propensity for the EUR to depreciate (appreciate) against the U.S. dollar when eurozone sovereign credit risk rises (falls). If there is a positive (negative) shock to credit risk, CDS premiums in both EUR and USD increase (decrease). However, if the EUR simultaneously decreases (increases) relative to the USD, the gain (loss) is larger (smaller) on the USD CDS compared to the similar EUR CDS. Therefore, the expected gains are smaller, and the expected losses are greater on the EUR CDS compared to the USD CDS, implying a positive quanto CDS spread.

The model offers a number of important insights on how these two channels affect quanto spreads and how we can distinguish between them. Importantly, we show that short-term quanto spreads are primarily driven by crash risk, as the maturity goes to zero, this is the only factor that drives quanto spreads. Quanto spreads at longer maturities, on the other hand, are impacted by both crash risk and covariance factor—with the latter gaining more significance as time to maturity increases. A key implication of the model is therefore that the term structure of quanto spreads can help to differentiate between crash and covariance risk.

Based on the insights of the discrete-time model, we propose an affine term structure model that captures both time-varying default risk, covariance between the FX-rate and the default intensity and currency jump risk associated with sovereign default. We estimate the model using USD-denominated CDS, quanto CDS spreads, and EURUSD currency options. Currency options are included in the estimation to identify the dynamics of exchange rate risk which is an important contributor to quanto spreads through the covariance risk channel.

We find that the covariance component is highly time-varying and tends to spike in times of crisis, while the crash risk component is persistent over the sample period, and, on average, accounts for the largest fraction of quanto CDS spreads. In essence, the covariance component reflects the distress-related part of quanto spreads; it shoots up in times when volatilities of credit risk and exchange rates are high and when they covary strongly. On the other hand, the crash risk component is of more static nature, because it captures the expected depreciation conditional on default. For example, in a model with no uncertainty surrounding credit risk (e.g., constant default risk) the covariance component is clearly zero, while crash risk causes a quanto spread if the market anticipates a jump in the exchange rate in reaction to a default.

Furthermore, we document that the relative contribution of covariance risk and crash risk to quanto spreads depends on the maturity. The short end of the quanto CDS term structure is almost exclusively driven by crash risk, while the covariance component increases in time to maturity. Intuitively, this is because the crash risk component causes a parallel shift in the term structure of quanto CDS spreads, while the covariance component affects the slope of the quanto CDS term structure. As a consequence, we find that covariance risk is particularly important for the relative pricing across currency denominations for longer-dated credit risky securities.

More specifically, we use our model to decompose the quanto CDS spreads, at maturities from 1-10 years, into a crash risk and a covariance risk component for Italy, Spain, Portugal, and Ireland over the period from August 2010 to April 2016. For Spain and Italy, we estimate the impact of a sudden sovereign default on the EURUSD to 15.6% and 9.6%, respectively. While for Portugal and Ireland, we estimate the currency crash to be significantly smaller at 5.3% and 5.0%, respectively.

Based on our model, we find that for Portugal and Ireland the average covariance components are 15.2 bps and 23.5 bps for the 5-year quanto spreads, corresponding to shares of 35% and 75% of their average quanto spreads. Consistent with our intuition that the covariance component is particularly important in times of distress, we indeed find that covariance risk is largest at the peak of the European debt crisis. For Ireland and Portugal, the covariance components during this period reach up to 60-70 bps which, in fact, exceed the contribution of crash risk to their quanto spreads. Without taking into account covariance risk, we would erroneously interpret the large quanto spreads for Portugal and Ireland as a sign of risk of a large downward jump in the Euro upon the default of these sovereigns.

The covariance components are not only substantial for the peripheral sovereigns, they also account for a large proportion of the quanto spreads for Spain and Italy. We find that the average of the covariance components at the 5-year maturity are 9.42 bps and 16.35 bps, which corresponds to 20% and 35% of their total quanto spreads. However, as is the case for the peripheral sovereigns, their covariance components exhibit strong time-variation and reach 38.51 bps and 55.25 bps at the peak of the European debt crisis, corresponding to 40% and 65% of their total spreads.

Quanto effects also apply to yield spreads of bonds issued by the same entity in different currencies. The advantage of studying quanto spreads from the perspective of CDS contracts is that recovery rates are the same for CDS contracts denominated in different currencies. This eliminates uncertainty related to differences in recovery rates, for example due to legal risk, between local currency and foreign currency denominated bonds, as addressed for example in Du and Schreger (2016).

On this basis, we use the model estimated from CDS data to construct model-implied quanto bond yield spreads, and we investigate if they can explain the observed yield spreads on bonds denominated in EUR and USD issued by Italy, Spain, and Portugal. We find that a significant part of the contemporaneous variation in quanto yield spreads can be explained by our model-implied quanto yield spreads, especially during the peak of the European debt crisis. An implication of our findings is thus that the previous literature that compares bonds across currency denominations using FX forward hedges, without accounting for quanto effects, may potentially miss an important component of yield spreads caused by quanto effects.

1.2 Literature

The unpublished work of Ehlers and Schönbucher (2006) is, to our knowledge, the first to recognize the joint effects of crash risk and covariance risk on CDS premiums in different currencies. While they focus on developing a theoretical framework that can be used to construct models for credit risky securities in different currencies, we focus on understanding and quantifying, both theoretically and empirically, the driving factors of quanto CDS spreads.

There are two closely related papers that study quanto CDS spreads in the eurozone which both focus on using quanto CDS spreads to imply out expected depreciations in the Euro versus the U.S. dollar at different horizons. Mano (2013) uses quanto CDS spreads for eurozone sovereigns to imply out risk-neutral expected depreciations upon default, without distinguishing between crash risk and covariance risk. In more recent and contemporaneous research, Augustin, Chernov, and Song (2018) propose an affine term structure model for eurozone quanto CDS spreads, which they use to estimate objective expected depreciations in the EURUSD conditional on sovereign defaults at different horizons. Our work differs from these papers in its main objective, we focus on what causes quanto CDS spreads and differences in bond yields across currency denominations. We identify two risk factors, covariance risk and currency crash risk, and we estimate their contribution to quanto CDS spreads and their time-series variation. Furthermore, we also use our model to explain what causes yield spread differences for eurozone sovereign bonds issued in Euro and U.S. dollar. Besides this, there are two other relevant papers that study eurozone quanto CDS spreads, De Santis (2015) and Brigo et al. (2016). The former uses quanto CDS spreads for eurozone sovereigns to estimate redenomination risk, that is, compensation for risk that EUR-denominated securities are redenominated into a new devalued currency. The latter focuses on developing a pricing model for quanto CDS spreads and calibrate it to Italian quanto CDS spreads.

Carr and Wu (2007b) provide evidence that sovereign credit risk is priced in the currency option markets for Brazil and Mexico. They obtain inference on the (risk-neutral) jump size in local currency upon sovereign default by estimating a joint model for options and sovereign CDS. Since option prices are driven by numerous factors apart from sovereign credit risk, e.g., macroeconomic news (Chernov et al., 2016), this approach makes it difficult to quantify the effect of sovereign default on local currency. Since the payoff on a quanto CDS is directly linked to currency jump risk at default, we contribute by providing a clean method for estimating the crash risk upon default.

Our paper is related to the vast literature that studies sovereign credit risk through the lens of CDS premiums, e.g., Longstaff, Pan, Pedersen, and Singleton (2011), Aït-Sahalia, Laeven, and Pelizzon (2014), Pan and Singleton (2008), Benzoni, Collin-Dufresne, Goldstein, and Helwege (2015), and Della Corte, Sarno, Schmeling, and Wagner (2016). The latter is, perhaps, the closest related to this paper. They document empirically a significant relationship between sovereign credit risk and returns on currencies and currency option strategies. While their paper is purely empirical, our objective is to develop models that allow us to quantify and understand the interconnection between credit and currency risk.

We contribute to the literature that studies pricing of similar credit risky securities across currency denominations, in particular bonds. There is a growing literature that analyzes deviations in yields for sovereign bonds across currency denominations (Buraschi et al., 2014; Corradin and Rodriguez-Moreno, 2016; Du and Schreger, 2016).

In these papers, the objective is to use the so-called "yield basis", defined as the difference between yields on a domestic and a synthetic domestic bond (which is constructed from foreign currency denominated bonds using FX forwards), to measure violations of the law of one price. Corradin and Rodriguez-Moreno (2016) show that the yield basis for eurozone sovereigns is large and volatile, and they attribute it to differences in collateral value and ECB purchases of EUR-denominated bonds. Buraschi, Menguturk, and Sener (2014) find a substantial yield basis for emerging market bonds during the 2007-2008 crisis and explain it by frictions in banking capital structure and non-conventional policy interventions. However, our theory shows that a yield basis may arise because of crash risk and covariance risk. Our empirical results suggest that this not only a theoretical concern. We provide evidence that indicates that the yield spread between EUR and USD-denominated bonds for eurozone sovereigns reflects compensation for risk related to covariance and crash risk.

1.3 Default and Recovery in Different Currencies

CDS contracts on the same reference entity but denominated in different currencies share a number of characteristics that are important to understand before setting up a model.

A Credit Default Swap (CDS) is an insurance against default on debt of an underlying reference entity. The contract involves two parties: a protection buyer and a protection seller. Every period, if no credit event has occurred of the reference entity, the buyer pays a percent-wise premium (often quarterly) of an agreed notional amount to the seller. If a credit event occurs, the buyer receives a recovery of the notional protected. Credit events are defined by the International Swaps and Derivatives Association (ISDA) and involves different scenarios, including outright bankruptcy, restructuring of debt, or deferred interest payments.

If a credit event occurs, an auction is held to determine the recovery rate based on a pool of bonds delivered into the auction. Importantly, the recovery rate is the same for all CDS contracts, independently of the currency denomination (see below for more details).

The auction is typically conducted between 30-35 days following the event determination date. Once an event has occurred, protection buyers are entitled to settle by physically delivering any of the specified deliverable obligations to settle the contract.

According to the standardized ISDA terms, the deliverable bonds are subject to a number of requirements. The payments of the obligation must be made in one of the *specified currencies* which for reference entities of Western Sovereigns are CAD, CHF, EUR, GBP, JPY, or USD. This means, for example, that a holder of a CDS contract denominated in EUR on Germany can choose to deliver German sovereign bonds denominated in USD. The relevant exchange rates for delivering obligations in a different currency to the CDS contract are fixed the day before the auction at 4pm at the WM/Reuters 4pm London mid-point rate.

1.4 The Quanto Spread in a Discrete Model

The option to choose in which currency to deliver bonds of the defaulted issuer means that the currency denomination becomes important. This can be seen through a very simple example: Consider two CDS contracts on Germany: One EUR-denominated with a notional amount of 1 EUR and one USD-denominated with a notional amount of 1 USD. Imagine for simplicity that the exchange rate is 1 at the initiation of the contract. If a default occurs before maturity, and at the same time the EUR drops to, say, a value of 0.5 USD, then the scale of protection offered by the two contracts differs. The holder of the EUR-denominated CDS can deliver 1 EUR notional and receive 1 EUR, whereas the holder of the USD protection can deliver a notional amount of 2 EUR, since the USD equivalent notional of 2 EUR is now only 1 USD because of the 'crash' of the EUR. Hence the amount of notional protected becomes effectively larger for the USD contract.

A similar mechanism is at play when currency depreciation has a positive correlation with a decrease in credit quality. Again, a simple example can provide the intuition. Imagine, as above, that the time 0 exchange rate is 1, and that the value of 1 USD can become 1.2 Euro or 0.8 Euro with equal probabilities 0.5 (under the USD risk-neutral measure) in the period 1, and that the exchange rate stays put in the second period until the CDS matures at time 2. Assume also for simplicity that the default probability of the reference entity is perfectly correlated with the exchange rate and becomes 3 percent in the state where the exchange rate is 1.2 and 1 percent in the other state. Assume zero interest rate in both currencies, and zero recovery in default. In this case, the USD value of protection of the CDS contract in two states is summarized in the following table:

State/denomination	USD	EUR
1.2/3%	0.03	$\frac{0.03}{1.2} = 0.025$
0.8/1%	0.01	$\frac{0.01}{0.8} = 0.0125$

Since $0.5 \cdot 0.025 + 0.5 \cdot 0.0125 = 0.0187 < 0.5 \cdot 0.01 + 0.5 \cdot 0.03 = 0.02$, we see that the value of the protection leg at time 1 is smaller for the EUR-denominated contract. If we assume (again for simplicity) that default risk is 0 between time 0 and time 1, then we have shown that the effect also applies for correlated default probability and FX-rate.

1.4.1 Model Assumptions and Definitions

We now build a simple discrete-time model that makes these observations rigorous. The model allows us to derive comparative statics and to analyze term structure effects. For the remainder of the paper, we define the exchange rate at time t, X_t, as units of domestic currency per unit of foreign currency, i.e., an increase in X_t implies that the foreign currency has appreciated against the domestic currency. Furthermore, we assume the existence of fixed riskless interest rates in both foreign and domestic currency, which we denote r_d and r_f, and we let $P_i(t,T) = e^{-r_i(T-t)}$ denote the price at time t of a zero-coupon bond paying one unit of currency $i = d, f$ at time T. In a no-arbitrage setting, we can then express the time t forward exchange rate with maturity T, $F(t,T)$, in terms of the foreign and domestic bond prices and the spot exchange rate as

$$F(t,T) = X_t \frac{P_f(t,T)}{P_d(t,T)}$$

Our model has a time horizon of $\bar{t}$ and we subdivide the time horizon into N equidistant time points which we label $t_0 = 0, t_1 = 1, \ldots, t_N = \bar{t}$. In each time period t there is a probability λ_t that the reference entity will default between time t and time $t+1$. We model FX crash risk upon default of the reference entity by assuming that the exchange rate drops by a fixed fraction of δ of the (risk-neutral) unconditional expectation of the exchange rate. Specifically, conditional on default between t and $t+1$, the exchange rate takes two possible values at $t+1$: $\delta \cdot uX_t$ and $\delta \cdot u^{-1}X_t$ with probabilities q and $1-q$, respectively. Conditional on no default, the exchange rate takes the values $C(\lambda_t) \cdot u$ and $C(\lambda_t) \cdot u^{-1}$ with respective probabilities q and $1-q$, where $C(\lambda_t)$ is a compensating factor $C(\lambda_t)$ defined as

$$C(\lambda_t) = \frac{1 - \delta\lambda_t}{1 - \lambda_t}$$

and it is needed to ensure no-arbitrage by compensating the exchange rate movement for crash risk. Had there been no crash risk, the exchange rate would either move up by a factor of u or down by a factor of u^{-1}. We show formally in Appendix 1.11.1 that this model is consistent with no-arbitrage. For tractability, we choose to do the compensation of crash risk through the jump size rather than through the martingale probabilities, which is an alternative option. We assume that the default probability can assume two values (λ^U, λ^D) in each period, and for simplicity we assume that the respective probabilities q^λ and $(1 - q^\lambda)$ do not depend on the current state. To capture the joint dynamics of default risk and exchange rates, we introduce correlation between the movements in the exchange rate and the default probability. Let Q_{ij} denote the one-step probability of the exchange rate to reach state i and the default probability to reach state j (conditional on survival), where $i = 1/j = 1$ correspond to an up move, and $i = 0/j = 0$ to a down move. At any point in time, we specify the joint distribution of the exchange rate and default probability as

$$Q_{11} = q(q^\lambda + A_1), \;\; Q_{10} = q(1 - q^\lambda - A_1) \tag{1.1}$$

$$Q_{01} = (1 - q)(q^\lambda - A_0), \;\; Q_{00} = (1 - q)(1 - q^\lambda + A_0) \tag{1.2}$$

where, $A_1 = \rho\sqrt{\frac{q^\lambda}{q}(1 - q)(1 - q^\lambda)}$ and $A_0 = \rho\sqrt{\frac{q^\lambda}{1-q}q(1 - q^\lambda)}$. The important parameter here is ρ, which is the correlation between the Bernoulli variables controlling the up and down moves of the exchange rate and default probability. Clearly, if $\rho < 0$, then $A_1 < 0$ and $A_0 < 0$, which implies that the exchange rate and the default probability tend to move in the opposite direction compared to the uncorrelated case ($\rho = 0$). Note that it only takes a specification of the unconditional probabilities q and q^λ and the correlation parameter to specify all the relevant quantities. q^λ and ρ can be chosen freely in $(0, 1)$ and $(-1, 1)$, respectively, but q is endogenously determined through the no-arbitrage condition for the currency movement which can be expressed simply in terms of the one-period forward rate $F = F(t, t + 1)$ as

$$q = \frac{F/X_t - u^{-1}}{u - u^{-1}} \tag{1.3}$$

See Appendix 1.11.1 for the derivation. Figure 1.1 illustrates the joint dynamics of the exchange rate and the default probability over two periods. The multi-period dynamics are obtained by repeating this tree from each individual node. After default of the reference entity, the tree terminates.

1.4.2 Pricing the Domestic and Foreign CDS

We model a Credit Default Swap (CDS) contract focusing on the 'fair running premium' that the buyer of protection should pay to obtain credit protection. For a contract with maturity T, we assume that no payment is exchanged at time 0 and that at every period $t_i \leq t_N \equiv T$, the buyer of the CDS contract pays a premium if the reference issuer has not defaulted at this time. If default occurs in the time interval $(t_{i-1}, t_i]$, the seller of insurance pays $1-R$ per unit face value—which we without loss of generality assume to be 1.

In this setting, the CDS premium in domestic currency with maturity T, $S^d(0,T)$, is given by

$$S^d(0,T) = (1-R)\frac{\sum_{i=1}^{N} P_d(0,t_i)Q\left(\tau = t_i\right)}{\sum_{i=1}^{N} P_d(0,t_i)Q\left(\tau > t_i\right)} \tag{1.4}$$

According to the standardized rules of ISDA, the foreign CDS contract is subject to the exact same contractual terms as the domestic contract, apart from currency denomination (CDS premiums are paid in foreign currency, and in the event of default, the recovery is received in foreign currency). The rules imply that the recovery rate is the same regardless of currency denomination of the contract.

Recall, that Q is the risk-neutral pricing measure when using the domestic bank account as numeraire. Defining Q^f as the risk-neutral measure corresponding to having the foreign account as numeraire, we can now express the premium of the same CDS contract denominated in the foreign currency as

$$S^f(0,T) = (1-R)\frac{\sum_{i=1}^{N} P_f(0,t_i)Q^f\left(\tau = t_i\right)}{\sum_{i=1}^{N} P_f(0,t_i)Q^f\left(\tau > t_i\right)} \tag{1.5}$$

where $P_f(0,t)$ denotes the discount factor corresponding to the foreign interest rate. To compare the two expressions we will need to understand the relationship between Q and Q^f.

Let M_t^i denote the pricing kernel for currency denomination $i = d, f$. Starting with the objective measure, P, we can price any foreign-denominated security with a price, Z_t^f, using the foreign pricing kernel:

$$1 = E_t^P \left(\frac{M_T^f}{M_t^f} \frac{Z_T^f}{Z_t^f} \right) = E_t^{Q^f} \left(P_f(t,T) \frac{Z_T^f}{Z_t^f} \right) \tag{1.6}$$

As in, e.g., Backus, Foresi, and Telmer (2001), we construct a domestic security from the foreign security using the exchange rate: $X_t Z_t^f$. Since this claim is denominated in domestic currency, we can price it using the domestic pricing kernel:

$$1 = E_t^P \left(\frac{M_T^d}{M_t^d} \frac{X_T Z_T^f}{X_t Z_t^f} \right) = E_t^Q \left(P_d(t,T) \frac{X_T Z_T^f}{X_t Z_t^f} \right) \tag{1.7}$$

Equations (1.6) and (1.7) hold for any security which implies that there is the following relationship between the domestic and foreign pricing kernels, the exchange rate, and the foreign and domestic risk-neutral measures:

$$\frac{M_T^f}{M_T^d} \frac{M_t^d}{M_t^f} = \frac{X_T}{X_t}, \quad M_T = \frac{X_T}{X_t} \frac{P_d(t,T)}{P_f(t,T)} \tag{1.8}$$

where M_T changes measure from the foreign to the domestic risk-neutral measure (i.e., $M_T = \frac{dQ^f}{dQ}(T)$). We refer to Appendix 1.11.2 and 1.11.2 for the closed-form model expressions of the domestic and foreign CDS premiums as well as their derivation.

1.4.3 Quanto CDS Spreads Comparative Statics

We now discuss how each parameter of the model impacts the quanto spread. First, we show that the quanto spread widens in the expected severity of the crash in foreign currency upon default.

Proposition 1. *The quanto spread, $QS(0,T)$, is decreasing in δ for all T*

Proof. See Appendix 1.11.2 □

To gain some intuition on Proposition 1, we propose a stylized example with a fixed default probability (implying independence between the default probability and the exchange rate), and a crash risk premium of δ. In Appendix 1.11.2, we show that in this case, the

CDS premiums in domestic and foreign currency, of any maturity, are given by

$$S^d = (1 - R)\frac{\lambda}{(1-\lambda)} \tag{1.9}$$

$$S^f = (1 - R)\frac{\lambda\delta}{(1-\lambda\delta)} \tag{1.10}$$

In the case of a fixed default probability, the riskless interest rates do not affect CDS premiums, i.e., the expressions for the CDS premiums in (1.9) and (1.10) hold for any choice of foreign and domestic interest rates. Assume $\delta < 1$, which implies that foreign currency depreciates upon default. Under this assumption, the recovery payment on the foreign CDS, $(1 - R)\delta$, is strictly smaller compared to the domestic CDS. The net present value of the premium leg payments, on the other hand, is larger than on the domestic CDS, because the foreign currency is expected to appreciate vs. domestic currency conditional on survival. Therefore, when $\delta < 1$, the value of the premium leg is greater and the value of the protection leg is smaller than for the domestic CDS, implying a positive quanto spread.

Figure 1.2 shows the CDS premiums denominated in foreign and domestic currency plotted against the expected depreciation upon default. The foreign CDS premium decreases as the risk-neutral expected crash in the currency increases, while the domestic CDS premium is fixed for a given level of the default probability, implying that the quanto spread increases in the severity of the crash.

Proposition 2. *The quanto CDS spread, $QS(0,T)$, is decreasing in ρ for all $T \geq 2$. Furthermore, if $\rho < 0$ ($\rho > 0$) then $QS(0,T)$ is increasing (decreasing) in u and $\lambda^U - \lambda^D$.*

Proof. See Appendix 1.11.2 □

The intuition behind Proposition 2 is that if there is negative correlation between the exchange rate and default risk, it is more likely that default occurs in states in which foreign currency has depreciated relative to its unconditional expectation. This effectively causes the foreign contract (converted into domestic currency) to deliver a smaller expected recovery payment, in the event of a default, compared to the domestic contract. The value of the premium leg, on the other hand, is largest on the foreign contract. This is because the risk-neutral expectation of the exchange rate conditional on survival must be larger than its unconditional expectation, otherwise, the currency forward is not priced consistently with no-arbitrage. The exchange rate thus tends to move unfavourably in both default and non-default states for the buyer of foreign CDS, implying that the fair foreign CDS premium must be smaller than the domestic CDS premium, i.e., a positive quanto spread.

An increase in the volatility of the exchange rate or the default probability, measured by the spread between up and down states (i.e., u and $\lambda^U - \lambda^D$), causes the quanto CDS spread to widen. An intuitive explanation for this is as follows. When credit risk goes up (down), then there are gains (losses) on both the foreign and the domestic CDS in the respective currencies. However, if the exchange rate tends to simultaneously decrease (increase), then the gain (loss) is smaller (larger) on the foreign CDS compared to the domestic CDS. Thus, the larger the moves in the credit risk and the exchange rate, the smaller (greater) the expected gains (losses) on the foreign CDS versus the domestic CDS, causing the quanto CDS spread to widen.

Finally, an important aspect of Proposition 2 is that the one-period quanto CDS spread is exclusively driven by crash risk, while the quanto CDS spread of two periods or more are impacted by both crash risk and covariance risk. Crash risk and covariance risk thus affect the term structure of quanto CDS spreads differently which allows us to distinguish between them by using data for quanto CDS spreads at different horizons.

1.4.4 Calibrating the Quanto CDS Term Structure

In the following, we use the discrete-time model to get a grasp of the magnitude of the crash and covariance risk embedded in quanto CDS spreads. The purpose is to gain intuition on how crash and covariance risk affect quanto spreads and to get an approximate estimate of their effect on observed quanto CDS spreads. Although the model is static, the central intuition gained from the model carries over to a richer dynamic term structure model, which we will analyze further in section 1.7.

We calibrate the model using CDS premiums for Spain, Italy, Portugal, and Ireland over the period August 2010-August 2012, i.e., at the height of the European debt crisis where CDS and quanto CDS spreads peaked. More specifically, the model parameters are calibrated such that they match the average observed 5-year quanto CDS spread, the 5-year CDS spread volatility, the EURUSD FX volatility, and the realized correlation between FX spot and 5-year USD CDS spread changes. We proxy ρ, the default probability/exchange rate correlation, with the correlation between daily percent-wise changes in the 5-year USD-denominated CDS premium and the EURUSD exchange rate. The parameter u is chosen such that the model's FX volatility matches the average 1-year risk-neutral volatility[1]. We compute the risk-neutral volatility from EURUSD currency options using the "model-free" methodology of Bakshi, Kapadia, and Madan (2003) (see section 1.6 for further details on the data). The empirical moments used for the calibration are reported in Table 1.1.

Fixing ρ and u as described above, we calibrate the default probability parameters, $(\lambda^D, \lambda^U, q^\lambda)$, and the currency crash risk parameter, δ, such that the model exactly matches the average 5-year CDS premiums denominated in USD and EUR. The calibration shows that the risk-neutral expected crash in the EURUSD in the event of a default is substantially larger for Spain and Italy relative to Portugal and Ireland. In particular, in the event of default of Spain and Italy, we estimate the risk-neutral expected depreciation in the EURUSD to 16% and 15%, respectively, while for Portugal and Ireland we estimate it to 5% and 7%, respectively. The results seem reasonable; the Euro is expected to take a much larger hit in the event of a Spanish or Italian default as these countries are more important economies for the eurozone.

If we were to ignore covariance risk ($\rho = 0$), the impact of a sovereign default on the EURUSD exchange rate would have been overestimated. In this case, for Spain and Italy, we estimate the crash risk to 21% and 19%, and 7% and 9% for Portugal and Ireland, underlining the importance of including covariance risk in the model to get an accurate assessment of the implied effect of a sovereign default on the exchange rate.

In Figure 1.3, we show the calibrated term structure of quanto spreads for Portugal, Ireland, Italy, and Spain. We see that the quanto spread increases in time to maturity. In the model—as shown explicitly in equations (1.9) and (1.10)—the term structures of foreign and domestic CDS premiums are flat when there is no covariance risk. Hence, the upward sloping quanto CDS curve is caused by covariance risk. The orange graph shows the quanto CDS spread in the case of no crash risk, i.e., the case where the entire quanto spread stems from covariance risk. We see that the curve is upward sloping in maturity, implying that covariance risk accounts for larger share of the quanto spread at longer maturities. Therefore, consistent with the intuition discussed previously, we can infer the magnitude of

[1] Since the one-year FX volatility, σ_{FX}, and the size of the up step, u, in a Cox, Ross, and Rubinstein (1979) tree are related as $u = e^{\sigma_{FX}}$.

covariance risk from the slope of the quanto CDS term structure.

1.4.5 Bond Pricing in Different Currencies

A growing empirical literature studies the pricing of bonds issued by the same issuer denominated in different currencies, e.g., Buraschi, Menguturk, and Sener (2014); Corradin and Rodriguez-Moreno (2016); and Liao (2016). In these papers, they compare yields of domestic bonds with yields on synthetic domestic bonds that are constructed from foreign-denominated bonds using FX forward hedges. However, as we will show below, the yield of a synthetic bond constructed in this manner only has the same yield as the domestic bond if there is no crash or covariance risk.

Consider two coupon bonds, on the same issuer, in foreign and domestic currency with prices $P_C^f(0,T)$ and $P_C^d(0,T)$, with respective coupons C_t^f and C_t^d. To focus on quanto effects, we assume the same coupons on the domestic and the foreign bond, but in different currencies, the exchange rate is 1 at time 0, no recovery payment at default, and that risk-free rates are 0. The price of the domestic and foreign risky bonds are:

$$P_C^d(0,T) = \underbrace{E_t^Q\left(\sum_{i=1}^{N} C_{t_i}^d 1_{(\tau>t_i)}\right)}_{\text{in domestic currency}} \tag{1.11}$$

$$P_C^f(0,T) = \underbrace{E_t^{Q^f}\left(\sum_{i=1}^{N} C_{t_i}^f 1_{(\tau>t_i)}\right)}_{\text{in foreign currency}} \tag{1.12}$$

We construct a synthetic domestic bond, which consists of the foreign bond and a portfolio of currency forward contracts entered at time 0 which converts each foreign-denominated coupon payment into domestic currency. The time 0 price of this synthetic bond in terms of domestic currency is:

$$
\begin{aligned}
P_C^{d,synth}(0,T) = & \underbrace{\sum_{i=1}^{N} C_{t_i}^f E_0(X_{t_i}|\tau > t_i) Q(\tau > t_i)}_{\text{Value of foreign bond in domestic currency}} \\
& + \underbrace{\sum_{i=1}^{N} C_{t_i}^f \left(F(0,t_i) - E_0(X_{t_i}|\tau > t_i)\right) Q(\tau > t_i)}_{\text{Value of forwards conditional on survival}} \\
& + \underbrace{\sum_{i=1}^{N} C_{t_i}^f \Big(F(0,t_i) - E_0^Q\left(X_{t_i}|\tau = t_i, \tau > t_{i-1}\right)\Big) Q(\tau = t_i|\tau > t_{i-1})}_{\text{Value of forwards conditional on default}}
\end{aligned} \tag{1.13}
$$

It is natural to believe that the price of $P_C^{d,synth}(0,T)$ is the same as $P_C^d(0,T)$, since the forward contracts hedge the exchange rate risk inherent in the foreign coupon payments. However, this is only correct if we assume that the last expression is 0, that is, default risk and exchange rate risk are independent. Under this assumption, we get the expression of the synthetic bond price that Buraschi, Menguturk, and Sener (2014); Corradin and Rodriguez-Moreno (2016); and Liao (2016) use to measure deviations from the law of one price, that is, $\sum_{i=1}^{N} C_{t_i}^f F(0,t_i) Q(\tau > t_i)$. In general, however, this price of synthetic domestic bond does not equal the price of the domestic bond, because the value of the forward contracts conditional on default deviates from the (unconditional) value of the forward contracts. Rather, in order for the synthetic bond to have the same value as the domestic bond, the foreign bond payments must be hedged using forward contracts that cancel at default such that the last expression is 0 by construction of the hedge, and not by assumption.

We illustrate this point in Table 1.2 by comparing the payoffs of two risky zero coupon bonds issued in EUR and USD in a one-period model. We assume that the EUR falls by 50% versus the USD at default, risk-free rates are 0, the forward price is 1, and no recovery on the bonds. We see from the table that a strategy that buys the USD bond and sells the synthetic USD bond has zero payoff in survival states since the forward contract hedges any exchange rate risk. However, it has a negative payoff of 0.5 USD in the default state, because the seller of the synthetic bond is obliged to pay 1 USD per 1 EUR from the forward contract, which is now worth only $\frac{1}{2}$ USD, and neither the EUR bond nor the USD bond pay anything. Important to note is that the EUR is expected to appreciate versus the USD in survival states to compensate for the EUR crash, but this gain has been hedged out by the forward contract. As a consequence, the synthetic USD bond must trade at a premium to the "real" USD bond to compensate for the crash in the EUR in default states. This simple example illustrates that at least a part of the observed yield spreads between synthetic and "real" bonds may be caused by currency crash risk, unrelated to any market frictions or imperfections in the international bond markets.

Likewise, covariance risk affects bond yields across currency denominations. We illustrate this in a multi-period model using the discrete-time model with parameters calibrated to 5-year Spanish CDS data (the parameters are reported in Table 1.1). The coupon bonds are assumed to be 1 and the principal is set to 100 (in respective currencies). For simplicity to convey the main idea, we assume 0 recovery rate and interest rates. Table 1.3 shows the results. The first row is the yield of a synthetic coupon bond, including crash risk. The second row shows yields on a long synthetic bond assuming no crash risk, and the third row is the yield on the domestic bond. The synthetic bond is long a foreign coupon bond, which pays coupons of one unit foreign currency and 100 at maturity, and short a portfolio of FX forward contracts that match the bond's payments (conditional on no default). The yield of the synthetic bond is 127 bps lower than the yield of the domestic bond, where 36 bps stems from covariance risk and 91 bps from crash risk. Raising the volatility of the exchange rate to 20.5% (the maximum EURUSD volatility over 2010-2012), the covariance component increases to 51 bps, while the crash risk component is unaltered. Overall, the results show that the synthetic bond trade at a substantially lower yield using realistic parameters to derive the covariance and crash risk components. Furthermore, the model suggests that the difference between the domestic and the synthetic yield is expected to increase in FX volatility. However, this implication must be interpreted with some caution since the model is static. In what follows, we explore more rigorously the driving factors causing the time-series variation in quanto spreads by using a dynamic term structure model.

1.5 A Term Structure Model of Quanto CDS Spreads

The discrete-time model is useful for obtaining the main intuition on how quanto spreads are driven by crash risk and default/currency covariance risk, but the static nature of the model makes it unable to capture time variation in credit and exchange rate risk. To this end, we propose an affine term structure model that captures the salient features of quanto CDS spreads discussed in the discrete-time model.

1.5.1 The Risk-Neutral Dynamics of the Model

In the model, the default risk of a sovereign i is driven by a compound Poisson process with a stochastic arrival rate, $\lambda_{i,t}$. Sovereign i's default intensity consists of two components: a systematic factor, $l_{i,t}$, which is correlated with the exchange rate, and a country-specific idiosyncratic component, $z_{i,t}$, which is orthogonal to the systematic factor

$$\lambda_{i,t} = l_{i,t} + z_{i,t} \tag{1.14}$$

Under the domestic risk-neutral measure, we let the exchange rate follow a Heston (1993) type dynamics with stochastic volatility, v_t, and a jump component driven by the sovereign default risk intensities:

$$dX_t = X_{t-}\left(r_{d,t} - r_{f,t}\right)dt + \sqrt{v_t}X_{t-}\left(\rho dW_{sys,t} + \sqrt{1-\rho^2}dW_{x,t}\right) + X_{t-}\sum_{i=1}^{K}\left(\zeta_i dN_{i,t} + \zeta_i\lambda_{i,t}dt\right) \tag{1.15}$$

The drift of the exchange rate, that is, the difference between domestic and foreign risk-free interest rates, insures that forward contracts are priced consistently with no-arbitrage. The jump component captures jumps in the exchange rate induced by sovereign default: conditional on country i defaulting at time t, the exchange rate depreciates instantly by a percent-wise fraction: $\frac{X_t - X_{t-}}{X_{t-}} = 1 + \zeta_i$, where ζ_i is a fixed country-specific jump size parameter. We then add up all jump components to get the aggregate crash risk component in the exchange rate, i.e., K represents the number of sovereigns included in the model. We specify the domestic risk-neutral dynamics of the state variables for sovereign i as follows:

$$\begin{bmatrix} dv_t \\ dl_{i,t} \\ dz_{i,t} \\ dm_{i,t} \end{bmatrix} = \left(\begin{bmatrix} \kappa_v \theta_v \\ \kappa_{l,i} \theta_l \\ \kappa_{z,i} m_{i,t} \\ \kappa_{m,i} \theta_{m,i} \end{bmatrix} - \begin{bmatrix} \kappa_v v_t \\ \kappa_{l,i} l_{i,t} \\ \kappa_{z,i} z_{i,t} \\ \kappa_{m,i} m_{i,t} \end{bmatrix} \right) dt + \begin{bmatrix} \sigma_v \sqrt{v_t} & 0 & 0 \\ \sigma_{l,i} \sqrt{l_{i,t}} & 0 & 0 \\ 0 & \sigma_{z,i} \sqrt{z_{i,t}} & 0 \\ 0 & 0 & \sigma_{m,i} \sqrt{m_{i,t}} \end{bmatrix} \begin{bmatrix} dW_{sys,t} \\ dW_{zi,t} \\ dW_{mi,t} \end{bmatrix} \tag{1.16}$$

where $W_{sys,t}, W_{zi,t}$, and $W_{mi,t}$ are independent. The systematic Brownian shock, $W_{sys,t}$, causes correlation between the exchange rate and the instantaneous volatility/systematic default risk component, which is assumed fixed and denoted ρ (as in, e.g., Bates (1996) and Carr and Wu (2007b)). The state variable, $m_{i,t}$, induces a central tendency in the idiosyncratic factor, i.e., our model has two state variables capturing the shape (level and slope) of the term-structure of domestic CDS premiums (Balduzzi, Das, and Foresi, 1998). This allows for the systematic component of the default intensity to freely capture the default/currency correlation risk, which is an important feature of our model in order for it to appropriately fit the term structure of quanto CDS spreads.

1.5.2 Specification of Pricing Kernels

We use a change of numeraire technique to price the foreign-denominated CDS contract which is no different than the techniques used to price derivatives by changing from the objective measure to the risk-neutral measure. Specifically, $M_t = X_t \frac{P_d(0,t)}{P_f(0,t)}$, is used to change numeraire from the domestic bond to the foreign bond, or put differently, M_t relates the (risk-neutral) parameters that are used to price domestic and foreign CDS contracts. Since the exchange rate in (1.15) jumps in the event of a sovereign default, we must be capable of handling jumps in the process governing the change of measure. Thus, we formulate Lemma 1 in Appendix 1.12 which slightly extends the extended affine risk premium specification of Cheridito, Filipovic, and Kimmel (2007) to jump diffusions. Roughly, Lemma 1 states that diffusions are drift-adjusted under the foreign measure according to their covariance with the exchange rate, i.e., there is no drift-adjustment in the uncorrelated case. Furthermore, the ratio between the default intensity under the foreign and domestic measure equals the jump size in the exchange rate upon sovereign default.

Besides this, we also use Lemma 1 to specify risk premia by relating the objective measure P and the risk-neutral domestic measure Q, which thus completes a triangle that allows us to switch between the domestic, foreign, and objective measure. Equivalent to Cheridito, Filipovic, and Kimmel (2007), Lemma 1 shows that if the square root processes under both P and Q, as characterized by parameters Θ_P and Θ_Q, fulfil the Feller condition [2], then the dynamics governed by Θ_P and Θ_Q are consistent with no-arbitrage. Therefore, to preclude arbitrage opportunities, we assume that the P and Q-dynamics of each state variable follow square root processes that fulfil the Feller condition, but with different parameters.

We do not model a jump to default risk premium between P and Q, as studied extensively in Benzoni, Collin-Dufresne, Goldstein, and Helwege (2015) in the context of eurozone sovereign CDS. They measure the jump to default risk premium as the ratio between the objective and risk-neutral default intensity, which is parallel to our setup where the currency jump size upon default equals the ratio between the foreign and domestic default intensities. An important distinction between the jump to default risk premium and the currency crash risk premium is that CDS premiums in both foreign and domestic currency are observable, which helps us pin down currency crash risk, whereas the jump to default risk premium is not tied to any observable quantity.

1.5.3 CDS Premiums in Domestic Currency

The derivation of the domestic CDS premiums follows the same procedure as in Pan and Singleton (2008) and Longstaff, Pan, Pedersen, and Singleton (2011). Here we briefly go through the main steps that are specific for our case. First, let $S_d(t,T)$ denote the domestic CDS premium at time t at maturity T, $P_d(t,T)$ the domestic discount factor, and R a fixed recovery rate. The state variable vector for country i, $x_{i,t} \equiv \left[l_{i,t}\ z_{i,t}\ m_{i,t}\right]^T$, is affine which

[2] The boundary non-attainment condition is important for square root processes. Let $X_t = (b+\beta X_t)dt + \sigma\sqrt{X_t}dW_t^P$ and consider a risk premium, $\phi(t)$, that preserves the affine structure under Q, i.e., $\phi(t) = \frac{c+dX_t}{\sigma X_t}$. Then it is in general not the case that the Radon-Nikodym, $L_t \equiv \frac{dQ}{dP}$, is a true martingale and the probability measure Q need not exist. However, if we impose the zero boundary non-attainment conditions (the Feller condition) $b^P \geq \frac{\sigma^2}{2}$ and $b^Q \geq \frac{\sigma^2}{2}$ then L_t is indeed a true martingale.

entails that we can compute the following transforms as

$$\psi\left(x_{i,t}, t, T\right) \equiv E_t^Q\left(e^{-\int_t^T \lambda_{i,s} ds}\right) = e^{\alpha_i(t,T)+\beta_i(t,T)\cdot x_{i,t}} \tag{1.17}$$

$$\phi(x_{i,t}, t, T) \equiv E_t^Q\left(\lambda_{i,T} e^{-\int_t^T \lambda_{i,s} ds}\right) = \psi(x_{i,t}, t, T)\big(A_i(t,T) + B_i(t,T)\cdot x_{i,t}\big) \tag{1.18}$$

where $\alpha_i(t,T), \beta_i(t,T), A_i(t,T)$, and $B_i(t,T)$ solve a set of ordinary differential equations (see, e.g., Duffie, Pan, and Singleton (2000)). The exact specification of the ODEs are reported in Appendix 1.12.2. Given a quarterly payment scheme for the premium leg and a fixed recovery rate on the protection leg, we have that their present values are given by

$$\Pi^{prem}(t,T) = S_d(t,T)\frac{1}{4}\sum_{j=1}^{4T} P_d\left(t, t+\frac{j}{4}\right)\psi\left(x_{i,t}, t, t+\frac{j}{4}\right) \tag{1.19}$$

$$\Pi^{prot}(t,T) = (1-R)\int_t^{t+T} P_d\left(t, t+u\right)\phi(x_{i,t}, t, u)du \tag{1.20}$$

The domestic CDS premium, which is consistent with no arbitrage, is then determined such that the present values of the premium leg and the protection leg are equal:

$$S_d(t,T) = \frac{\Pi^{prot}(t,T)}{\Pi^{prem}(t,T)} \tag{1.21}$$

1.5.4 CDS premiums in Foreign Currency

In the discrete-time model, we derive the foreign CDS premium directly by using $M_t = \frac{X_t}{X_0}\frac{P_d(0,t)}{P_f(0,t)}$ to convert each foreign-denominated payment into a domestic payment. In the affine model, this is rather cumbersome. We take a more convenient approach and price the foreign-denominated CDS contract using a change of numeraire technique. Formally, $M_t = \frac{dQ_f}{dQ}$, is the Radon-Nikodym derivative that changes measure from the domestic to the foreign risk-neutral measure. To apply the change of numeraire technique, we need the dynamics of the Radon-Nikodym derivative between Q and Q_f, which is given by:

$$dM_t = M_t\sqrt{v_t}\left(\rho W_{sys,t} + \sqrt{1-\rho^2}dW_{x,t}\right) + M_t\sum_{i=1}^{K}\left(\zeta_i dN_{i,t} + \zeta_i\lambda_{i,t}dt\right) \tag{1.22}$$

By using Lemma 1 with M_t as the pricing kernel, the default intensity under the foreign risk-neutral measure is given by:

$$\lambda_{i,t}^f = \lambda_{i,t}(1 + \zeta_i) \tag{1.23}$$

$$dv_t = \kappa_v^f(\theta_v^f - v_t)dt + \sigma_v\sqrt{v_t}dW_{sys,t}^f \tag{1.24}$$

$$dl_{i,t} = \left(\kappa_{l,i}(\theta_{l,i} - l_{i,t}) + \sigma_{l,i}\rho\sqrt{l_{i,t}v_t}\right)dt + \sigma_{l,i}\sqrt{l_{i,t}}dW_{sys,t}^f \tag{1.25}$$

where $\kappa_v^f = (\kappa_v - \sigma_v\rho)$, $\theta_v^f = \frac{\kappa_v\theta_v}{\kappa_v - \sigma_v\rho}$, and $\lambda_{i,t}$ is the domestic default intensity.

Lemma 1 states that the ratio between the default intensity under the foreign measure and domestic measure equals the jump size conditional on sovereign default: $\lambda_t^f = \lambda_t(1+\zeta)$. For this reason, very short-term quanto CDS spreads are exclusively driven by crash risk because $S^d(t,T) \approx (1-R)\lambda_t$ and $S^f(t,T) \approx (1-R)(1+\zeta)\lambda_t$, when t approaches T. Even in the case of a purely idiosyncratic default intensity (i.e., no covariance risk), a quanto CDS spread emerges solely through the crash risk channel. This is consistent with our intuition from the discrete-time model, where we showed that a quanto CDS spread arises in the case of a constant default probability through crash risk.

Under the foreign measure, each process that is exposed to $W_{sys,t}$ is drift-adjusted via the pricing kernel (1.22). For l_t, the drift adjustment is $\sigma_l\rho\sqrt{l_tv_t}$, i.e., it depends on the instantaneous volatility of the exchange rate, the systematic default component, and their correlation. If there is negative correlation between the exchange rate and the default intensity, then the drift correction is negative which causes the expected default risk to be smaller under the foreign measure than under the domestic measure, implying a positive quanto CDS spread.

The covariance adjustment has less impact at shorter horizons, because the drift adjustment does not affect the instantaneous default risk. An implication of the model is therefore that quanto CDS spreads tend to widen in maturity if there is negative covariance between default and exchange rate risk. This is consistent with the results of our calibration exercise based on the discrete-time model, where we showed that the quanto CDS spread widens in maturity because of covariance risk. To summarize, crash and covariance risk affect the foreign default intensity through different channels; crash risk scales and covariance risk drift-adjusts the default intensity, and this distinction is what allows us to separate the two effects using the term structure of quanto CDS spreads.

In order to fit the model into the affine framework, we approximate the term, $\sqrt{l_t v_t}$, in the systematic default risk's drift with a first-order Taylor expansion around the respective processes' mean reversion levels [3]. The foreign transforms are then computed as in the domestic setting

$$\psi_f\left(x_{i,t}, t, T\right) = e^{\alpha_{f,i}(t,T)+\beta_{f,i}(t,T)\cdot x_{i,t}} \tag{1.26}$$

$$\phi_f(x_{i,t}, t, T) = \psi_f(x_{i,t}, t, T)\big(A_{f,i}(t,T) + B_{f,i}(t,T)\cdot x_{i,t}\big) \tag{1.27}$$

and the foreign premium and protection legs are given by

$$\Pi_f^{prem}(t,T) = S_f(t,T)\frac{1}{4}\sum_{j=1}^{4T} P_f\left(t, t+\frac{j}{4}\right)\psi_f\left(x_{i,t}, t, t+\frac{j}{4}\right) \tag{1.28}$$

$$\Pi_f^{prot}(t,T) = (1-R)\int_t^{t+T} P_f\left(t, t+u\right)\phi_f(x_{i,t}, t, u)du \tag{1.29}$$

From the dynamics of the foreign state variables, i.e., equation (1.25), we see that the currency/default covariance risk introduces v_t as an additional state variable compared to the domestic case, that is, $x_{i,t} \equiv \left[l_{i,t}\ z_{i,t}\ m_{i,t}\ v_t\right]^T$. The exact specification of the ODEs which $\alpha_{f,i}, \beta_{f,i}, A_{f,i}$, and $B_{f,i}$ solve are provided in Appendix 1.12.2.

1.6 Data and Descriptive Analysis

1.6.1 Credit Default Swap Data

We collect CDS premiums from Markit on eurozone sovereign bonds issued by Austria, Belgium, Germany, Finland, Ireland, France, Italy, Netherlands, Portugal, and Spain denominated in EUR and USD. Markit provides us with daily quotes at maturities of 1, 3, 5, 7, and 10 years. We use the complete restructuring clause on the CDS contracts which allows the protection buyer to deliver bonds of any maturity (and currency denomination) into the CDS auction. Markit performs a number of data cleaning procedures on the CDS data that they receive from their contributors, e.g., to avoid stale quotes and outliers, and

[3] The exact form of the Taylor approximation is given by: $\sqrt{l_t v_t} = 1/2\left(v_t\left(\frac{\theta_l}{\theta_v}\right)^{1/2} + l_t\left(\frac{\theta_v}{\theta_l}\right)^{1/2}\right)$.

they only report quotes if there are at least three quotes from different contributors. Before August 2010, Markit aggregated quotes across currency denominations into one quote. As our focus is on the impact of currency denomination on the pricing of CDS contracts, we initiate our analysis in August 2010, and our sample ends in April 2016.

1.6.2 Currency Options Data

One of our main objectives is to estimate the contribution of covariance risk to quanto spreads which essentially depends on three factors: risk-neutral exchange rate volatility, volatility of systematic default risk, and the correlation between credit risk and the exchange rate. The latter two factors can be identified from USD-denominated CDS premiums and quanto CDS spreads, but CDS data are not particularly informative about the first factor. Therefore, in order to pin down the risk-neutral distribution of exchange rate volatility, we include currency options data in our estimation, as in, e.g., Bates (1996); Carr and Wu (2007a,b).

We collect EURUSD currency options data from Bloomberg from August 2010 to April 2016. The data consist of Garman and Kohlhagen (1983) implied volatilities of delta-neutral straddles, 10, 25-delta risk reversals, and 10, 25-delta butterfly spreads which are the common quoting conventions in currency option markets. The maturities are fixed and are 1, 2, 3, 6, 9, and 12 months.

A straddle is a portfolio which is long a call and a put option with the same strike and maturity. The payoff of a straddle is directionless and the buyer of the straddle is long at-the-money volatility.

A risk reversal consists of a long position in an out-of-the money (OTM) put option and a short position in an OTM money call option with symmetric deltas[4]. The long position in the OTM put protects against large depreciations in foreign currency (EUR), and in contrast, the short OTM call loses money when large depreciations in the USD occur. Risk reversals therefore measure the slope of the implied volatility curve against moneyness, also called the skew of the implied volatility curve.

A butterfly spread is the difference between the average IV of and OTM call and an OTM

[4]Sometimes the risk reversal is quoted conversely as a long position in a call option and a short position in a put.

put and the IV of the delta-neutral straddle. If the butterfly spread is positive, it reflects that the market price of hedging large FX movements (in either direction) is more expensive compared to the case in which returns are log-normal, i.e., the risk-neutral distribution of exchange rate changes is fat tailed.

Using the Garman and Kohlhagen (1983) formula for the IVs derived from the straddles, risk reversals, and butterflies, we recover five different strikes, spanning from the strike of a put with a delta of -10 percent to the strike of a call option with a delta of 10 percent. We skip the details on how this procedure works and refer to Della Corte, Sarno, Schmeling, and Wagner (2016) and Jurek (2014) for an elaborate explanation.

1.6.3 Interest Rate Data

For the pricing of CDS denominated in Euro and U.S. dollar, we need to compute discount curves in both currencies. We take the most common approach and build discount curves from overnight index swap rates, OIS for U.S. dollar, and EONIA for Euro. We use overnight index swap rates rather than LIBOR swap rates because it is well-documented that they contain a default risk component. Since 2010, maturities of up to 10 years of overnight index swaps have been traded. We therefore exclusively use overnight index swap rates as proxies for riskless interest rates, since the longest maturity in our CDS data is 10 years. Based on the overnight index swap interest rates, we construct zero-coupon curves in Euro and U.S. dollar using a standard bootstrapping procedure. We collect the data on overnight index swap rates from Bloomberg, and the maturities are 3, 6, 9 months, and 1-10 years, and the data start in August 2010 and end in April 2016.

1.6.4 Descriptive Data Analysis

Table 1.4 reports the averages and standard deviations of eurozone sovereign CDS premiums denominated in EUR and USD, spanning maturities from 1-10 years, over the period August 2010 to April 2016. First, we note that the USD CDS premium is, on average, unambiguously higher than the corresponding EUR CDS premium for all sovereigns. In absolute terms, the average quanto CDS spreads, e.g., at the 5-year maturity, are largest for Ireland, Italy, Portugal, and Spain, ranging from 36-48 bps, while they are the smallest for Finland, Germany, Netherlands, and Austria, ranging from 8-22 bps. In general, the non-GIIPS countries have much smaller average CDS premiums, indicating that the market deemed it unlikely that sovereign defaults would occur for these sovereigns. As an example, the average 5-year USD CDS premium for Portugal is more than ten times larger than for Germany.

In Figures 1.4-1.6, we show the time series of quanto CDS spreads and USD-denominated CDS premiums for all sovereigns at maturities ranging from 1-10 years. The quanto CDS spreads are positive in the entire sample period for all sovereigns. As is the case for the USD CDS premiums, the quanto CDS spreads peak for all sovereigns between the last quarter of 2011 and the Summer of 2012. During this period, the 5-year quanto CDS spreads exceed 100 bps for Spain and Portugal, and almost reach 100 bps for Italy and Ireland as well. From July 2012, in the wake of Mario Draghi's speech in which he insured that the ECB would do whatever it takes to preserve the Euro, the quanto CDS spreads gradually decline, but they stay positive throughout the sample period.

Table 1.5 reports the averages and standard deviations for implied volatilities of straddles, risk reversals, and butterflies for each maturity. The implied volatility for both the 10 and 25-delta risk reversals are, on average, negative, in fact, they are negative throughout our sample period at all maturities. This shows that large downside risk in the Euro has historically been more expensive to insure relative to symmetric downside risk in the U.S. dollar.

The focus of our analysis is the relation between currency risk and credit risk. As a first step in exploring this relation, we proxy aggregate eurozone credit risk by the first principal component of eurozone 5-year USD CDS premiums and investigate its relation to EURUSD implied volatility and spot changes. The principal component analysis shows that there is a strong commonality in CDS premiums for eurozone sovereigns. The first principal component of weekly changes in 5-year USD CDS premiums explains 77% of the common variation of the changes in 5-year USD CDS premiums[5], consistent with Longstaff, Pan, Pedersen, and Singleton (2011), who document strong commonality in global CDS premiums.

Table 1.6 shows results from regressions of weekly innovations in the EURUSD spot exchange rate and the delta-neutral straddle implied volatility on the first principal component

[5]Similar results are obtained when using EUR-denominated CDS.

of the eurozone CDS premiums. Over the entire sample period, there is a significantly negative relation between changes EURUSD spot rate and eurozone credit risk, with a t-statistic of -3.69 and an R^2 of 8.1%. This result suggest that the Euro tends to depreciate when eurozone credit risk rises. Most of the significance, however, stems from the European debt crisis period, i.e., from August 2010 to December 2012. In the post-crisis period (January 2013 to April 2016), there is a negative, but insignificant, relation (t-statistic of -1.34), and a miniscule part of the variation in spot exchange rates is explained by exposure to sovereign credit risk.

The at-the-money implied volatility and eurozone credit risk are significantly positively related over the entire sample period (t-statistic of 3.84), with an R^2 of 12.1%, i.e., increasing forward-looking EURUSD volatility tends to be associated with increasing eurozone credit risk. Our results are consistent with those of Della Corte, Sarno, Schmeling, and Wagner (2016), who document, for a large sample of countries, that exchange rate spot movements and implied volatilities of options are tightly related to sovereign credit risk. The positive relation between EURUSD implied volatility and eurozone credit risk is highly significant in the crisis period, with a t-statistic of 7.70 and an $R^2 = 27.2\%$, but their relation is barely significant in the post-crisis period (t-statistic of 2.17, $R^2 = 3.2\%$). Consequently, the results of our regression analysis indicate that eurozone sovereign credit risk and the currency spot rate and implied volatility primarily co-vary in times of distress.

According to our discrete-time model, the significant covariance between exchange rate risk and sovereign credit risk implies a positive quanto CDS spread for eurozone sovereigns, even without any exchange rate crash risk at default. Moreover, the results of the regressions suggest that the covariance risk components embedded in quanto CDS spreads are most pronounced during the crisis period from 2010-2012. In the next section, we analyze these conjectures using the proposed affine term structure model to decompose quanto CDS spreads into a covariance risk component and a crash risk component.

1.7 Model Results and Estimation

1.7.1 Estimation Approach

We focus on estimating the model for the GIIPS countries: Portugal, Ireland, Italy, and Spain, excluding Greece. We exclude Greece from the analysis because Breuer and Sauter (2012) document that there was virtually no trading activity in the Greek CDS from early 2011, as the market anticipated a Greek default, which, in fact, occurred on March 9, 2012. CDS markets also reflected that a Greek default was anticipated, with elevated CDS premiums on Greek government bonds reaching several thousand bps by the last of quarter of 2011.

We focus on the GIIPS countries (excluding Greece) because they are the least creditworthy in our sample and, arguably, the focal point of the European debt crisis. For example, the 5-year CDS premiums (in USD) for the GIIPS all reached levels exceeding 600 bps, with Portugal and Ireland being the most extreme cases with CDS premiums exceeding 1000 bps. In comparison, the German 5-year CDS barely touched 100 bps, and the French 5-year CDS spiked at about 200 bps.

In the estimation, we use weekly data (each Wednesday) of quanto CDS spreads, USD-denominated CDS premiums, and currency option implied volatilities. Each week, we have 30 option prices (five strikes at six maturities), five CDS premiums denominated in USD, and five quanto CDS spreads at maturities of 1, 3, 5, 7, and 10 years.

If we were to estimate the model in one joint estimation, we would have an unmanageably large set of parameters and a high dimensional state variable vector. For instance, in the case of four sovereigns, the model has 12 state variables and a very large parameter vector containing systematic, country-specific, and measurement error parameters. One approach to reduce the dimension of the state vector is to introduce common factors or to use just one state variable to capture country-specific default risk. However, since we are interested in making accurate assessments of the magnitude of the quanto spreads driven by crash and covariance risk, we need precise estimations. Our estimations suggest that at least two country-specific factors are necessary for the model to accurately fit the cross-section and time-series dynamics of USD CDS and quanto CDS premiums simultaneously.

For this reason, we estimate the model stepwise. In the first step, we estimate a time series of the instantaneous currency volatility, v_t, and its objective and risk-neutral parameters from currency option implied volatilities. We estimate the model using maximum likelihood estimation in conjunction with the unscented Kalman filter. In the next step, now treating v_t as observable and its parameters as fixed, we estimate the parameters for the idiosyncratic and systematic default intensity components, i.e., l_t, z_t, and m_t, using data for USD CDS and quanto CDS spreads for one country at the time. The estimation procedure is described in detail in Appendix 1.13.

1.7.2 Estimation Results

Table 1.7 presents the maximum likelihood estimates of the model, and Figure 1.7 illustrates the estimated state variables l_t, z_t, and m_t for each sovereign. For all sovereigns, the idiosyncratic component of the default intensity, z_t, spikes between the last quarter of 2011 and the Summer of 2012. In the wake of Mario Draghi's (president of the ECB) famous speech in July 2012, in which it was announced that the ECB would do whatever it takes to preserve the Euro within its mandate, the EURUSD exchange rate and the eurozone sovereign credit markets stabilized, which caused both z_t and l_t to decrease rapidly, for all sovereigns.

The systematic component, which captures the part of the default intensity correlated with the foreign exchange rate, l_t, exhibits two peaks (with the exception of Portugal), in early 2011 and by mid-2012. The systematic default component has a more stable path over the sample period compared to the idiosyncratic components that have stronger mean reversion and seem to capture transient credit risk shocks. Clearly, for all the sovereigns, m_t, is highly time-varying, indicating that it is an important feature of our model to allow the mean-reversion level of z_t to be stochastic. Consistent with this, we find considerable improvements in model fits when using a three-factor model instead of a two-factor model. For example, we find that a model in which z_t has a constant mean-reversion level is not sufficiently rich to provide reasonable fits of the USD CDS term structure and the quanto CDS term structure.

Using the estimated parameters and the filtered state variables, we compute model-implied USD CDS premiums and quanto CDS spreads and compare them to their observed counterparts. We show in table 1.8 the summary statistics for the model pricing errors, both in terms of root mean squared errors (RMSEs) and mean absolute pricing errors (APEs) in bps. The time-series fits are illustrated in Figures 1.8-1.9 at maturities of 1, 5, and 10 years.

The average RMSE across the 1-10 years maturities for the USD CDS range from 23.21-26.68 bps for Italy, Spain, and Ireland. The average RMSEs for Portugal, however, are significantly larger at 37.92 bps, especially the 1-year RMSE is comparatively large. Using the APE metric, the Portuguese fit is better, which indicates that large outliers are important contributors to its RMSEs. For all sovereigns, the general pattern is that the pricing errors decline in maturity, i.e., the shorter maturities are the most difficult to capture for the model. A likely explanation for this is that the short end is more volatile/noisy than the long end of the term structure, as shown in Table 1.4.

The model seems to fit the quanto CDS premiums reasonably well, as seen from Figures 1.8-1.9. This is also reflected by relatively small average RMSEs for all sovereigns, with the lowest being 0.98 bps for Ireland and the largest being 4.90 bps for Spain. The RMSEs tend to increase in the maturity of the quanto CDS spread, most notably for Spain. From Figure 1.9, we see that for Spain, the model tends to underestimate the 10-year quanto CDS premium and overestimate the 10-year USD CDS premium. Such a bias, however, is not present for the other sovereigns and does not seem to be a general issue with the model. Overall, considering the large fluctuations in the CDS premiums over a relatively short sample period, we believe that the model performs well in capturing both the USD CDS and the quanto CDS dynamics across all tenors. As an example, to underline the strong time-variation of the CDS premiums over our sample period, the 1-year USD CDS premium for Portugal and Ireland range between 0.23%-23% and 0.07%-14.5%, respectively.

Next, we use the model estimates to decompose quanto CDS spreads for Italy, Spain, Ireland, and Portugal into a currency/default covariance component and a crash risk component. We compute the covariance and crash risk component of the quanto spread as:

$$\text{FX/default covariance risk component} = S^d_{\zeta=1}(t,T) - S^f_{\zeta=1}(t,T) \tag{1.30}$$

$$\text{FX crash risk component} = S^d(t,T) - S^f(t,T) - \left(S^d_{\zeta=1}(t,T) - S^f_{\zeta=1}(t,T)\right) \tag{1.31}$$

where $S^d_{\zeta=1}(t,T) - S^f_{\zeta=1}(t,T)$ denotes the model-implied quanto spread assuming no currency crash at default. Hence, if crash risk accounts for the entire quanto spread, the covariance component is zero. The crash risk component is the residual part of the quanto spread after correcting for covariance risk, i.e., the difference between the total quanto spread and the FX/default covariance component.

Figure 1.10 illustrates the time series of the decompositions at maturities of 1, 5, and 10 years for Spain, Italy, Portugal, and Ireland. Table 1.9 shows descriptive statistics for the decompositions. First, we discuss the estimates of ζ, i.e., the risk-neutral expected percent-wise jump in the EURUSD immediately after sovereign default is announced. For all sovereigns in our estimations, we find that ζ is negative and highly significantly different from zero. This indicates that the Euro is expected to take an immediate hit conditional on the announcement of a sovereign default. The general pattern we find is that the Euro is expected to take a larger downward jump at default of sovereigns that are fundamentally more important for the eurozone economy. Specifically, we estimate ζ for Spain, Italy, Portugal, and Ireland to be -15.6%, -9.6%, -5.3%, and -5.0%, respectively.

Turning to the decompositions of the quanto spreads, we find that the covariance component is economically large and accounts for a large proportion of the quanto spreads for all sovereigns in our estimations. Over the entire sample period, the covariance component of the 5-year quanto spread ranges, on average, from 9.2 – 16.4 bps (20-38% of total spread) for Spain, Italy, and Portugal, and it is 23.5 bps (75% of total spread) for Ireland. Importantly, covariance risk is strongly time-varying and is especially pronounced during the European debt crisis, where credit and exchange rate risk are strongly co-varying and volatile. From August 2010 to December 2012, the average covariance component at the 5-year maturity is 18.4 bps (25% of total spread) for Spain and ranges between 27-36 bps for Portugal, Italy, and Ireland (35%-58% of total spread). During this period, covariance risk reaches as much as 38.5-65.8 bps and accounts for 40-76% of the total 5-year quanto spreads for Spain, Portugal, and Italy and virtually for the entire 5-year quanto spread for Ireland.

We expect that a larger part of quanto spreads at shorter maturities is due to crash risk and that the contribution of covariance risk increases in maturity. The intuition for this is that when the maturity approaches zero, the domestic (USD) CDS premium is well-approximated by $S^d(t,T) \approx (1-R)\lambda_t$, and according to Lemma 1, the foreign (EUR) CDS premium is well-approximated by $S^f(t,T) \approx (1-R)(1+\zeta)\lambda_t$. The longer-term quanto spreads are more exposed to covariance risk, because the covariance between credit risk and exchange rate risk reduces the drift of the Euro default intensity and hence has a larger impact over longer horizons (see section 1.5.4 for an elaborate discussion). Consistent with this reasoning, we indeed find that crash risk accounts, on average, for the largest part of quanto spreads at the 1-year maturity and gradually decreases in maturity. The average term structure of crash risk is particularly steep between the 1-year and 5-year maturity, but almost flat from the 5-year maturity and beyond. Specifically, the crash risk component accounts, on average, for 46% (25%) for Ireland, 80% (65%) for Portugal, 81% (62%) for Italy, and 87% (80%) for Spain of the 1-year (5-year) quanto spreads.

The average quanto spread is steeply upward sloping up to the 5-year maturity and virtually flat at maturities beyond that (see Table 1.4), our estimations suggest that this shape of the quanto spread is because of covariance risk. If only crash risk were present, we would expect a flat quanto spread term structure because crash risk scales the default intensity, i.e., causes parallel-shifts of the quanto spread term structure.

Overall, our findings indicate that covariance between sovereign credit risk and currency risk accounts for a significant share of quanto spreads, especially in times of financial distress. Anecdotal evidence confirms the importance of covariance risk in eurozone credit markets during the European debt crisis. Between 2010-2011, several research notes were released by major investment banks discussing the practicalities of hedging currency/credit risk for eurozone sovereigns and banks (e.g., Barclays Research Note (2011) and J.P. Morgan Research Note (2010)), indicating a large hedging/speculative demand for FX/default covariance risk.

Based on our decompositions, we shed some light on redenomination risk, that is, the risk that a sovereign redenominates its EUR-denominated debt into a new (devalued) domestic currency. According to the standardized ISDA terms, if Spanish (or Portuguese/Irish) sovereign bonds are redenominated into a new currency, i.e., a new "Pesetas", it triggers the Spanish CDS contracts, whereas redenomination is not considered a credit event for Italy. The Euro CDSs for Italy are therefore not protected against a redenomination event, while they are for Spain. Our estimations suggest that redenomination risk is not priced in quanto spreads as a sudden event, because a larger part of the quanto spreads for Spain is caused by crash risk compared to Italy. However, this does not imply that redenomination risk is not a contributing factor to quanto spreads, but rather that it is not priced as a jump event. In support of this finding, articles written by major market participants (e.g., Credit Suisse Research Note (2010)) seemed to share the view that redenomination is legally and practically very difficult to implement "overnight".

Our estimations provide us with the parameters under both the objective and the risk-neutral measure which we can use to calculate the time series of credit risk and quanto credit risk premiums. Longstaff, Pan, Pedersen, and Singleton (2011) argue that a reasonable measure for the credit risk premium—the risk premium associated with holding unpredictable variation in the default arrival rate—is the difference between the CDS premiums based on the risk-neutral parameters (Q-parameters) and the objective parameters (P-parameters). Presumably, since providing credit insurance on eurozone sovereigns is associated with large losses at times of high marginal utility, we expect that credit risk premiums are positive, on average.

In the same spirit, we define a quanto risk premium as the risk premium associated with taking exposure to crash and covariance risk, as defined in equations (1.30)-(1.31). We measure the quanto risk premium as the difference in quanto CDS spreads calculated based on the Q-parameters and the P-parameters. That is, the credit risk premium and the quanto risk premium are defined as:

$$CRP(t,T) = S_d^Q(t,T) - S_d^P(t,T) \tag{1.32}$$

$$QRP(t,T) = S_d^Q(t,T) - S_f^Q(t,T) - (S_d^P(t,T) - S_f^P(t,T)) \tag{1.33}$$

where $S_i^M(t,T)$ is the CDS premium based on parameters under measure $M = Q, P$ in currency i at maturity T. Figure 1.11 illustrates the time series of the quanto and credit risk premiums for each sovereign, and Table 1.10 reports the mean risk premiums in basis points, and the fraction of the risk premiums to total spreads. We find substantial positive risk premiums associated with taking exposure to eurozone sovereign credit risk and quanto risk, especially at the peak of the European debt crisis in 2011-2012. For Spain, Italy, and Portugal, the average 5-year credit risk premiums range from $114-211$ bps, which in relative terms correspond to 59-66% of the total average USD CDS premiums. The large credit risk premiums suggest that investors demand high compensation for providing credit insurance compared to premiums based on objective default risk. In general, the credit risk premiums for Ireland are quite small compared to the other countries and account, on average, for less than 6% of the total USD CDS 5-year spread. For Italy and Spain, the credit risk premiums are positive throughout the sample period, with peaks in 2012, while for Ireland and Portugal, the risk premiums are briefly negative for a period in 2011, but positive for the rest of the sample. At shorter maturities, the risk premium accounts for a smaller part of CDS spreads for all sovereigns, since the unpredictable variation in default risk is smaller.

Finally, we document sizeable and highly time-varying quanto risk premiums for the eurozone sovereigns. The quanto risk premiums are positive at all horizons and account for a significant share of the quanto CDS spreads. The quanto risk premiums are of greatest magnitude for Spain and Italy, both in relative and nominal terms, consistent with the notion that investors demand a larger risk premium for holding quanto risk for more systematically important sovereigns. For example, at the 5-year maturity, the quanto risk premium accounts, on average, for 61% and 73% of the total quanto CDS spreads for Spain and Italy, and for 40% and 15% of the quanto CDS spreads for Portugal and Ireland. The 5-year quanto risk premium is largest in the last part of 2012, where it reaches 28 bps for Spain and 35 bps for Italy. Even though the quanto CDS spreads are of similar order of magnitude for Ireland and Portugal, they have much smaller maximum quanto risk premiums of 6 bps and 18 bps, respectively.

1.7.3 Quanto Effects on Bond Yields

Quanto spreads are not only present in eurozone sovereign CDS, it has also been documented in previous research that the difference in yields on a USD-denominated bond and a EUR-denominated bond tends to be positive, that is, a positive quanto yield spread (Corradin and Rodriguez-Moreno, 2016). In this section, we investigate if the quanto yield spread is attributed to compensation for exposure to crash and covariance risk (quanto effects). To this end, we use the estimated parameters and state variables to compute model-implied yields for EUR and USD-denominated bonds, and we then investigate if they explain those observed in data. There are only a few eurozone sovereigns that have bonds issued USD. Our analysis focuses on Italian, Spanish, and Portuguese government bonds issued in EUR and USD (Ireland has no government bonds issued in USD).

In the presence of no frictions, the model-implied quanto yield spreads should explain all the variation in the observed quanto yield spreads. However, there are many factors unrelated to quanto effects that may cause the observed quanto yield spreads to deviate from zero. First, quanto yield spreads may simply be caused by differences in terms of the bonds, because it is typically not possible to pair EUR and a USD-denominated bonds that have the same maturity, coupon payments, recovery rates, etc. Second, there is evidence for a specialness premium attached to holding EUR-denominated bonds in the eurozone due to favourable regulatory treatment of debt issued in local currency over foreign currency debt. For instance, EUR-denominated bonds tend to have relatively smaller haircuts in repo transactions and carry lower capital weights on banks' balance sheets compared to USD-denominated bonds (Corradin and Rodriguez-Moreno, 2016). Since none of the above-mentioned factors impact quanto CDS spreads, we can use our model to derive cross-currency bond yield spreads caused only by quanto effects.

Constructing the Quanto Yield Spread

We now discuss how to construct quanto yield spreads from observed bond yields, and we then examine how they relate to model-implied quanto yield spreads estimated from CDS data. There are very few bonds issued in USD by eurozone sovereigns which makes it difficult to find matching EUR and USD bonds. We circumvent this issue by constructing a synthetic USD bond from EUR-denominated bonds, which matches the maturity, coupon rate, and coupon frequency of the traded USD bond.

To this end, we calculate the full term structure of riskless zero-coupons in EUR and USD, as well as the zero-coupons in EUR of the risky sovereign. We express the price of the risky zero-coupon in EUR, $P^E(t,s)$, as

$$P^E(t,s) = \frac{1}{\left(1 + r^E(t,s) + s^E(t,s)\right)^{s-t}} \tag{1.34}$$

where $r^E(t,s)$ are the riskless EUR interest rates, and $s^E(t,s)$ are the credit spreads for the risky EUR bonds. From the zero-coupon term structure of risky EUR bonds, we use (1.34) to calculate $s^E(t,s)$ at any maturity. Using $s^E(t,s)$ and the riskless USD interest rates, $r^U(t,s)$, we construct a synthetic USD bond, $PC^U_{synth}(t,T)$, with matching coupons, notional, and maturity of the observed USD bond, $PC^U_{obs}(t,T)$. The prices of the synthetic and the traded USD bonds are given by

$$\begin{aligned} PC^U_{synth}(t,T) &= \sum_s C_s \frac{1}{(1+s^E(t,s)+r^U(t,s))^{s-t}} + N\frac{1}{(1+s^E(t,T)+r^U(t,T))^{T-t}} \\ PC^U_{obs}(t,T) &= \sum_s C_s \frac{1}{(1+s^U(t,s)+r^U(t,s))^{s-t}} + N\frac{1}{(1+s^U(t,T)+r^U(t,T))^{T-t}} \end{aligned} \tag{1.35}$$

We then define the observed synthetic quanto yield spread as the yield differential between the USD bond and its synthetic counterpart:

$$QY_{synth}(t,T) \equiv y^U_{obs}(t,T) - y^U_{synth}(t,T) \tag{1.36}$$

If there are no quanto effects, and no other frictions, the observed quanto yield spread should be zero, since the credit spreads in this case are the same. However, if quanto effects are present, it causes a positive quanto yield spread (i.e., $s^U(t,s) > s^E(t,s)$).

Using the estimated model parameters, we compute model-implied quanto yield spreads. We choose the time to maturity, coupons, and notional amount such that they exactly match those of the traded USD bond. We assume fixed recovery of par value at default and calculate the risky zero-coupon price in currency $i = EUR, USD$ as:

$$P^i(t,T) = E^{Q^i}_t\left(e^{-\int_t^T (r_{i,u}+\lambda_{i,u})du}\right) + R\int_t^T E^{Q^i}_t\left(\lambda_{i,s}e^{-\int_t^s (r_{i,v}+\lambda_{i,v})dv}\right) ds \tag{1.37}$$

In order to emulate the methodology used for constructing the observed quanto yields spreads, we derive a EUR credit spread curve from the risky and riskless EUR zero-coupon prices. We then use this EUR credit spread curve in conjunction with the USD riskless zero-coupon prices to construct a synthetic model-implied USD bond price, exactly as in (1.35). We then compute the model-implied quanto yield spread as the difference in yields between the USD bond and the synthetic USD bond.

Empirical Results

For each sovereign, we obtain the full term structure of risky EUR zero-coupon prices using the benchmark government yield curve provided by Reuters at maturities ranging from six months to 10 years. The riskless zero-coupon prices in EUR and USD are bootstrapped from their respective overnight index swap rates.

For Italy, we study a USD-denominated bond that matures in February 2017, and for Spain we study two USD-denominated bonds with maturities in June 2013 and March 2018, respectively, i.e., the entire sample period from 2010-2016 is covered by a USD-denominated bond for both countries. Portugal, however, only has one USD-denominated bond traded in our sample period with maturity in March 2015. We calculate the observed quanto yield spread as in (1.35), which we refer to as the "synthetic" quanto yield spread. Besides this, we compute a bond quanto yield spread, defined as the yield spread between a USD bond and a EUR bond with similar maturities corrected for the riskless interest rate differential:

$$QY_{bond}(t,T) \equiv y^{U}_{obs}(t,T) - y^{E}_{obs}(t,T) - \big(\bar{r}^{U}(t,T) - \bar{r}^{E}(t,T)\big) \tag{1.38}$$

The bonds that we use are specified in the footnote [6]. One advantage with the measure specified in (1.38) is that it does not involve the extraction of a full term structure of zero-coupon prices and credit spreads. This spread, however, is a cruder measure than the synthetic quanto yield spread, since it does not take into account the term structure of the risky zero coupon prices, differences in coupon schemes, or maturity mismatch.

The justification for this measure is that if there were no quanto effects, or other frictions, only the riskless interest rate differential drives the yield spreads across currency denominations. We would thus expect (1.38) to be close to zero if there are no quanto effects. In the presence of no frictions, the bond quanto yield spread is exactly zero for zero-coupon bonds [7], but it is not necessarily zero for coupon bonds.

However, if the duration of the bond is short, the yield spread between a coupon bond and zero-coupon bond is close to zero, which is the case in our sample, where we consider

[6] The Italian EUR-denominated government bond matures on 1st of February 2017, ISIN: IT0004164775. 4% coupon semi-annual. The USD-denominated Italian government bond has maturity on 12th of June 2017. ISIN: US465410BS63, 5.375% coupon semi-annual. First Spanish bond couple: EUR-denominated government bond matures on January 31 th 2014, ISIN: ES00000121H0, 4.25% coupon semi-annual, and the USD-denominated June 17th 2013. ISIN: XS0363874081, 3.625% coupon semi-annual. Spain bond couple for latter period: EUR bond: 30th of July 2018 4.1% semi-annual coupon rate, and USD bond: maturity 6th of March 2018, 4% semi-annual coupon rate. Portugal bond couple: maturity EUR bond 15th oct 2014 PTOTEOOE0017 and 3.6% coupon rate semi-annual, USD bond maturity 25th march 2015 XS0497536598 and 3.5% coupon rate semi-annual.

[7] To see this, consider two risky zero-coupon bonds in EUR and USD: $P^E(t,T)$ and $P^U(t,T)$ and assume

only bond maturities of less than seven years [8].

In Table 1.11, we report summary statistics of the observed quanto yield spreads. We divide the sample into a crisis period, from August 2010 to March 2013, and a post-crisis period March 2013 to April 2016.

For Italy and Spain, the crisis period is characterized by positive and highly significant quanto yield spreads (i.e., t-statistics exceeding > 5.42), with respective averages of 40.8 bps (59.7 bps) and 62.7 bps (99.0 bps) of the synthetic (bond) quanto yield spread. For these countries, the corresponding average model-implied quanto yield spreads are in the same order of magnitude of 61 bps and 59 bps. The Portuguese observed quanto yield spreads based on the synthetic and the bond method have respective means of 4.3 bps and 28.6 bps, which are both insignificantly different from zero. However, if restrict the sample period to August 2010 to July 2012, i.e., we consider the sample period prior to Draghi's speech, then the synthetic quanto yield spread is significant for Portugal as well. In general, the Portuguese quanto yield spread is more noisy than for Spain and Italy and exhibits larger positive and negative swings.

In the post-crisis period, we only study bonds issued by Italy and Spain, since there are no USD-denominated bond data for Portugal. In this period, the quanto yield spreads are much smaller (albeit still positive) and less significant compared to the crisis period. For Italy, the synthetic bond yield spread has en insignificant average of just 14.0 bps, and the spread is contained within a more narrow range compared to the crisis period, with a 95% percentile of 56 bps relative to 123 bps in the crisis period. Likewise is the average Spanish synthetic quanto yield spread smaller (33.3 bps) in the post-crisis period compared to the crisis period. The corresponding means of the model-implied quanto yield spreads are about 25 bps and 41 bps for Italy and Spain, our model thus captures the falling trend

independence between the exchange rate and the default event ($1_{\tau>T}$):

$$\begin{aligned}
P^E(t,T) &= E_t^{Q_E}\left(exp\left(-\int_t^T r^E(s)ds\right)1_{\tau>T}\right) = E_t^{Q^U}\left(\frac{X_T}{X_t}exp\left(-\int_t^T r^U(s)ds\right)1_{\tau>T}\right)\\
&= E_t^{Q^U}\left(\frac{X_T}{X_t}\right)E_t^{Q^U}\left(exp\left(-\int_t^T r^U(s)ds\right)1_{\tau>T}\right) \Leftrightarrow \frac{P^E(t,T)}{P^U(t,T)} = \frac{exp\left(\int_t^T r^U(s)ds\right)}{exp\left(\int_t^T r^E(s)ds\right)}\\
&\Leftrightarrow y^U(t,T) - y^E(t,T) - \bar{r}^U(t,T) - \bar{r}^E(t,T) = 0
\end{aligned}$$

[8]For example, for Italy, the 7-year EUR spread between the coupon bond and zero-coupon is always negative, with a minimum of -15 bps (3% in relative terms), and since the USD-bond is subject to the same bias, we presume that the bias' affect in the quanto yield spread is small.

in the quanto yield spreads.

Next, we test if the observed quanto yield spreads are explained by their model-implied counterparts. We find a significant and positive relation between observed quanto yield spreads and their model-implied counterparts during the peak of the European debt crisis, while they are insignificantly related in the post-crisis period. Our results indicate that a significant portion of the observed yield deviations between EUR and USD-denominated eurozone sovereign bonds is attributable to quanto risk and that quanto yield spreads do not necessarily reflect mispricings. Positive quanto yield spreads persist post-crisis, although much smaller compared to the crisis period, but they are seemingly caused by other factors, such as differences in liquidity and specialness associated with currency denomination (Corradin and Rodriguez-Moreno, 2016).

Table 1.12 shows results from regressions of the observed quanto yield spreads, using both the synthetic (1.36) and the bond method (1.38), on the model-implied counterparts. We also include the 5-year quanto CDS spread in the regressions as an alternative measure for quanto effects. In the crisis period, for Spain and Italy, there is in general a significant positive relationship between the observed quanto yield spreads and the model-implied quanto yield spreads and the quanto CDS spreads.

In particular for Spain, the slope coefficients of the model-implied quanto yield spread and the 5-year quanto CDS spread range between 1.24-1.93, with t-statistics between 2.99 and 6.60, and R^2s ranging from 20.75%-30.54%. Likewise for Italy, there is a positive relation between the synthetic quanto yield spread and its model-implied counterpart and the 5-year quanto CDS spread both with slope coefficients close to unity, with respective t-statistics and R^2s of: 1.43, $R^2 = 8.66\%$ and 2.07, $R^2 = 17.06\%$. Using the bond method to derive the observed quanto yield spreads, we also find significant positive slope coefficients near unity. In this case, a substantial part of the variation in the observed quanto yield spreads are explained by quanto effects, with R^2s between 15.90% and 29.19%. In the post-crisis period, we find an insignificant relation between observed and model-implied quanto yield spreads and quanto CDS spreads for both Spain and Italy.

To conclude, our findings suggest that joint modeling of credit risk and currency risk is a key ingredient in understanding bond yields across currency denominations and that it becomes increasingly important when sovereign bond markets are under distress. In accordance with our above findings for the eurozone, Du and Schreger (2016) construct a quanto yield spread for emerging market sovereign bonds and find that the covariance between currency and credit risk explains a significant part of the quanto yield spread. These findings suggest that our model could be useful for understanding the variation in yield spreads across currency denominations in emerging bond markets, we leave this topic for future research.

1.8 Conclusion

In this paper we analyze quanto spreads in the context of eurozone sovereign CDS contracts. We develop a discrete-time no-arbitrage model, which illustrates how, even in a frictionless setting, quanto CDS spreads arise as a compensation for exposure to two risk factors. The first risk factor is an FX crash risk factor, which captures the market's (risk-neutral) anticipation of a large adverse jump in foreign currency (EUR) against domestic currency (USD) in the event of a sovereign default. The second factor, the currency/default risk covariance factor, captures the propensity for the EUR to depreciate (appreciate) against the U.S. dollar when eurozone sovereign credit risk rises (declines). Our simple model allows for simple comparative statics.

To estimate the relative importance of these factors, we propose an affine term structure model that allows us to distinguish between the two effects and capture their time-variation. We use our model to decompose the quanto spreads for Spain, Italy, Portugal, and Ireland, and find that both covariance and currency crash risk contribute substantially to quanto CDS spreads. The covariance risk factor is highly time-varying and increases in times of distress, when the currency and credit markets are volatile and co-move. However, the implied currency crash risk from sovereign defaults differ greatly in the four cases.

We estimate the (risk-neutral) expected jump in the EURUSD conditional on sovereign default for Spain and Italy to 15.6% and 9.6% which, consistent with our intuition, is significantly larger than the estimated currency jump size of about 5% in the event of a Portuguese or Irish default. We document a significant risk premium associated with currency/default covariance risk and currency crash risk, i.e., a risk premium associated with selling protection in the 'expensive' currency (USD) and buying protection in the

’cheap’ currency (EUR). This risk premium is especially large for Spain and Italy, where it accounts for most of the quanto CDS spread.

Finally, we provide evidence that quanto yield spreads, which are differences in yields on USD and EUR-denominated bonds, are significantly related to quanto effects estimated based on our model. This highlights the importance of taking into account currency crash risk and covariance risk when assessing the relative pricing of bonds across currency denominations.

1.9 Figures

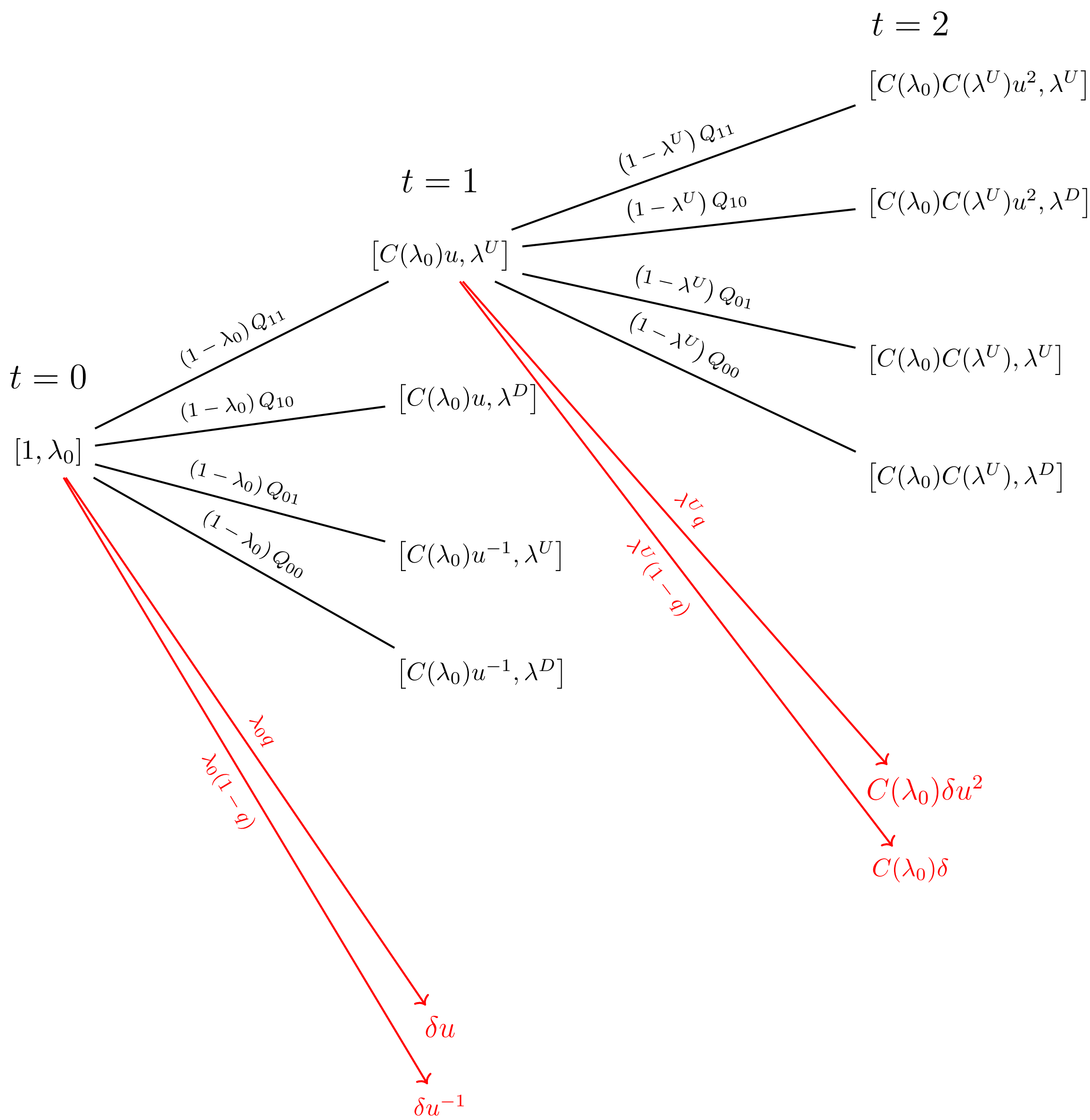

Figure 1.1: **Two-period model of the default probability and the exchange rate.** This figure illustrates the joint dynamics of the default probability and the exchange rate over two periods. At time 0, the exchange rate is 1 and default occurs with a probability of λ_0. If default occurs, the exchange rate is adjusted by δ relative to the state of the exchange rate if there were no crash risk. Conditional on survival, which occurs with probability $1 - \tilde{\lambda}$, the exchange rate is adjusted by the compensating factor $C(\tilde{\lambda})$, where $\tilde{\lambda} = \lambda^U$ or $\tilde{\lambda} = \lambda^D$. Simultaneously, if survival occurs, a new one-period default probability is drawn which takes either a high value λ^U or a low value λ^D, and a relative one-period change of the exchange rate is realized taking two possible values (u, u^{-1}). That is, in total there are four possible outcomes for the default probability and the exchange rate change at each node. The joint probability distribution for reaching each of those four possible states are specified in equations (1.1)-(1.2). There are the same possible states in each survival node. Due to space constraints, we only show the possible states at time 2 starting from the survival node in which the default probability and the exchange rate went up $((\lambda^U, u))$.

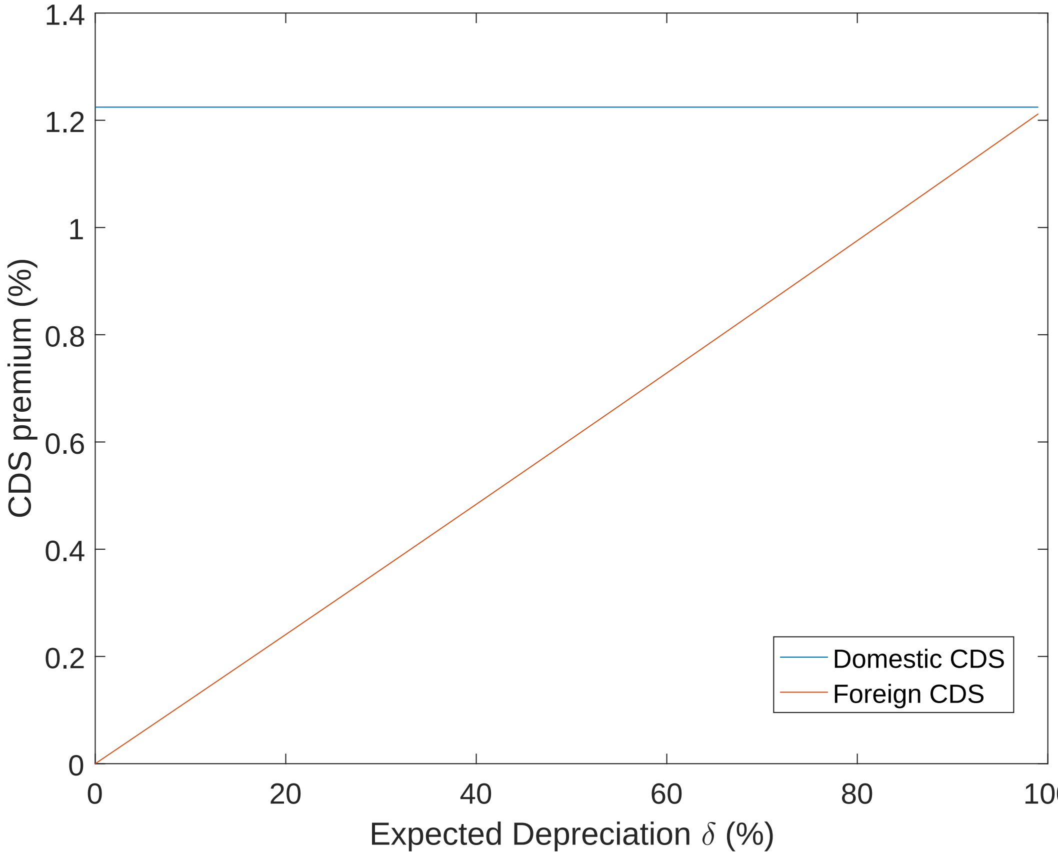

Figure 1.2: **Currency crash risk induced quanto CDS spreads.** This figure illustrates the impact of an expected depreciation upon default, δ, on the premiums of CDS contracts denominated in foreign and domestic currency. The blue graph is the CDS premium in domestic currency, and the red graph is the CDS premium in foreign currency on the same underlying reference entity. The CDS premiums are computed based on a model with fixed default probability and a fixed risk-neutral expected depreciation upon default. Interest rates do not affect CDS premiums in the model when the default probability is constant.

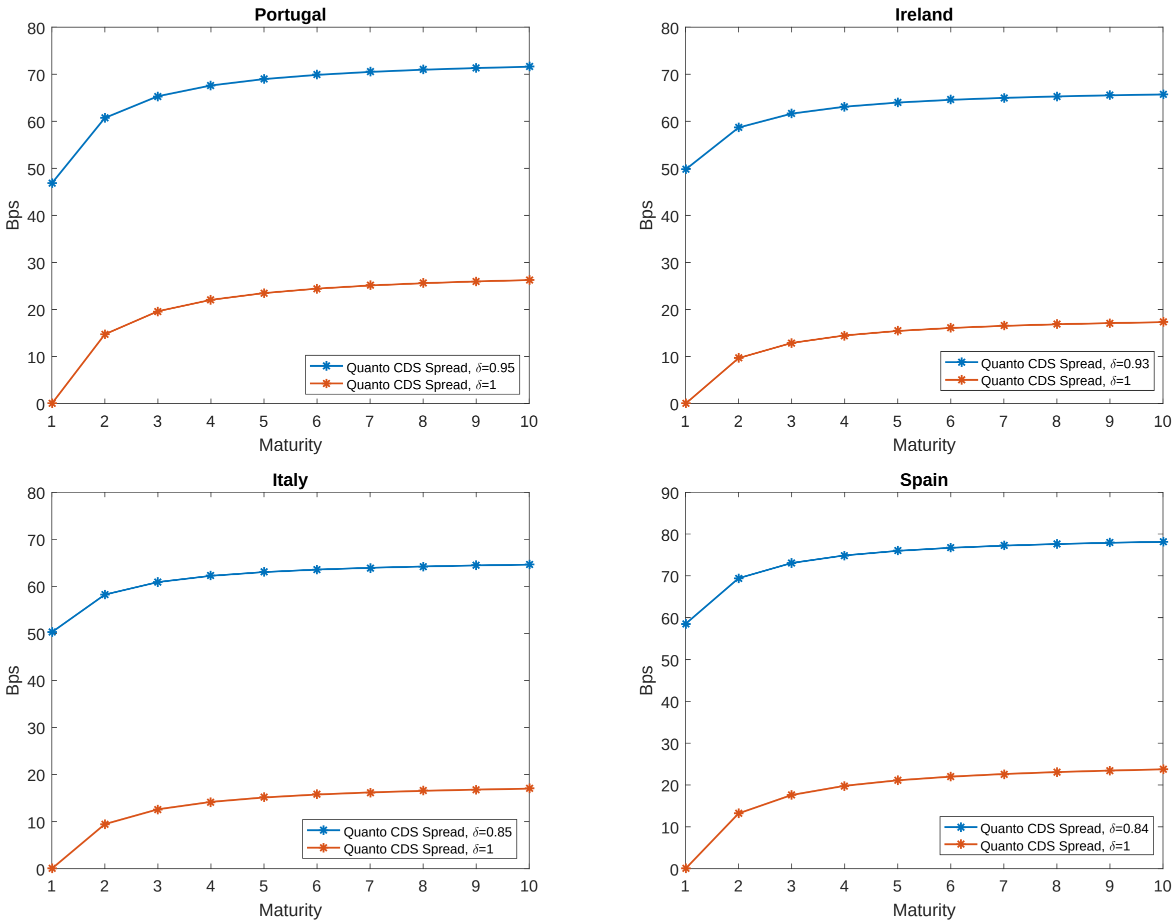

Figure 1.3: **Term structures of calibrated quanto CDS spreads.** This figure illustrates the term structure of model-generated quanto CDS spreads at maturities of one to ten years. The quanto spread is the difference between the CDS premiums on the same reference entity denominated in USD and EUR. The parameters are calibrated to match the empirical average 5-year EUR and USD CDS premiums, the 1-year EURUSD risk-neutral volatility, and the correlation between the 5-year USD CDS premium and the EURUSD spot exchange rate. All model parameters are assumed fixed, and the calibration period is August 2010 to August 2012. The blue graph illustrates the quanto spread at different maturities. The orange graph is the share of the quanto spread stemming from default/currency covariance risk, i.e., the case of $\delta = 1$. The recovery rate is assumed to be 40%, and the choice of foreign and domestic interest rates has no impact on the quanto spread in the model.

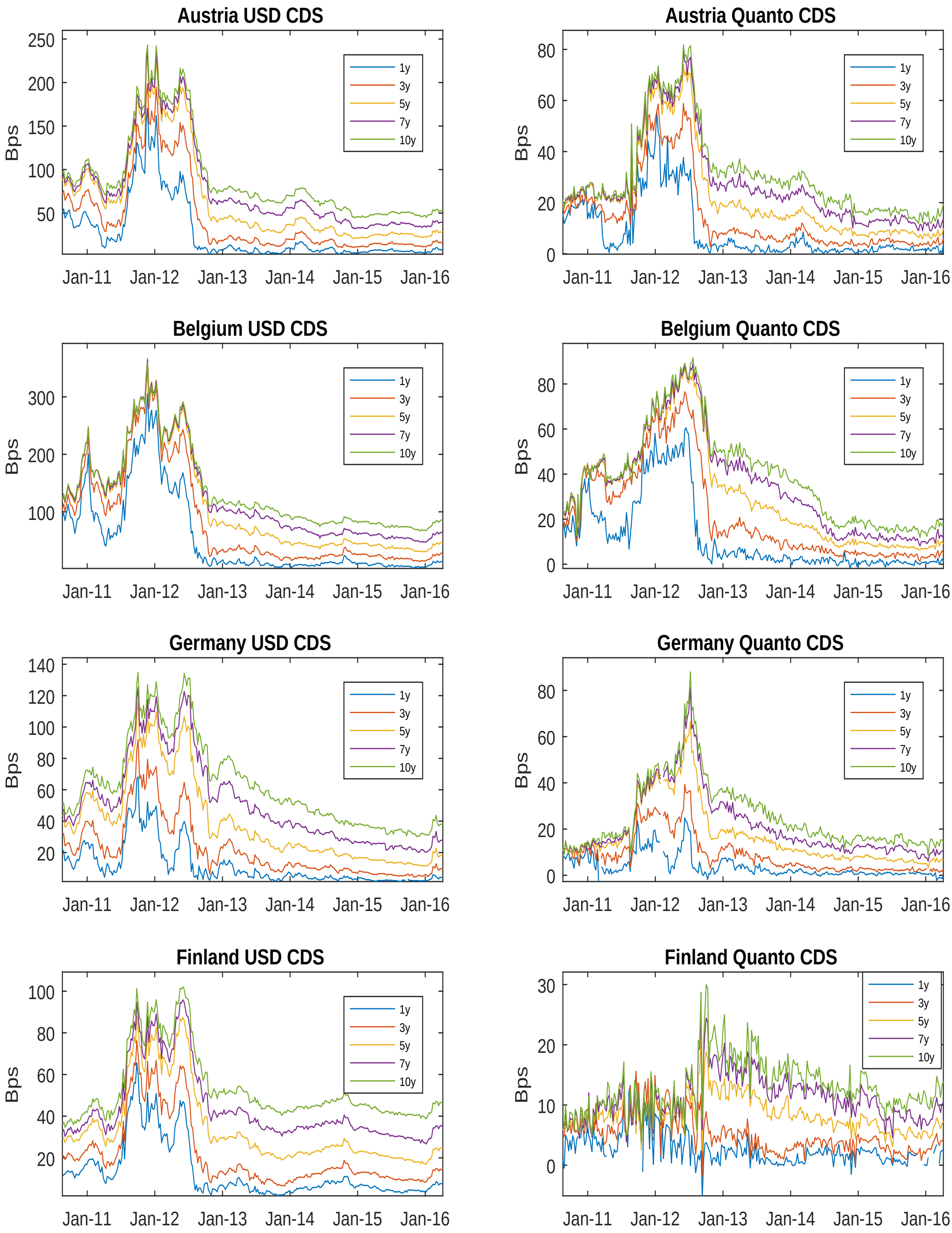

Figure 1.4: **USD CDS and quanto CDS spreads for Austria, Belgium, Germany, and Finland.** This figure shows USD CDS premiums and quanto CDS spreads–defined as the difference between USD and EUR-denominated CDS premiums of the same underlying reference entity–for Austria, Belgium, Germany, and Finland. The sample period is August 2010 to April 2016 and comprises 1402 daily observations obtained from Markit.

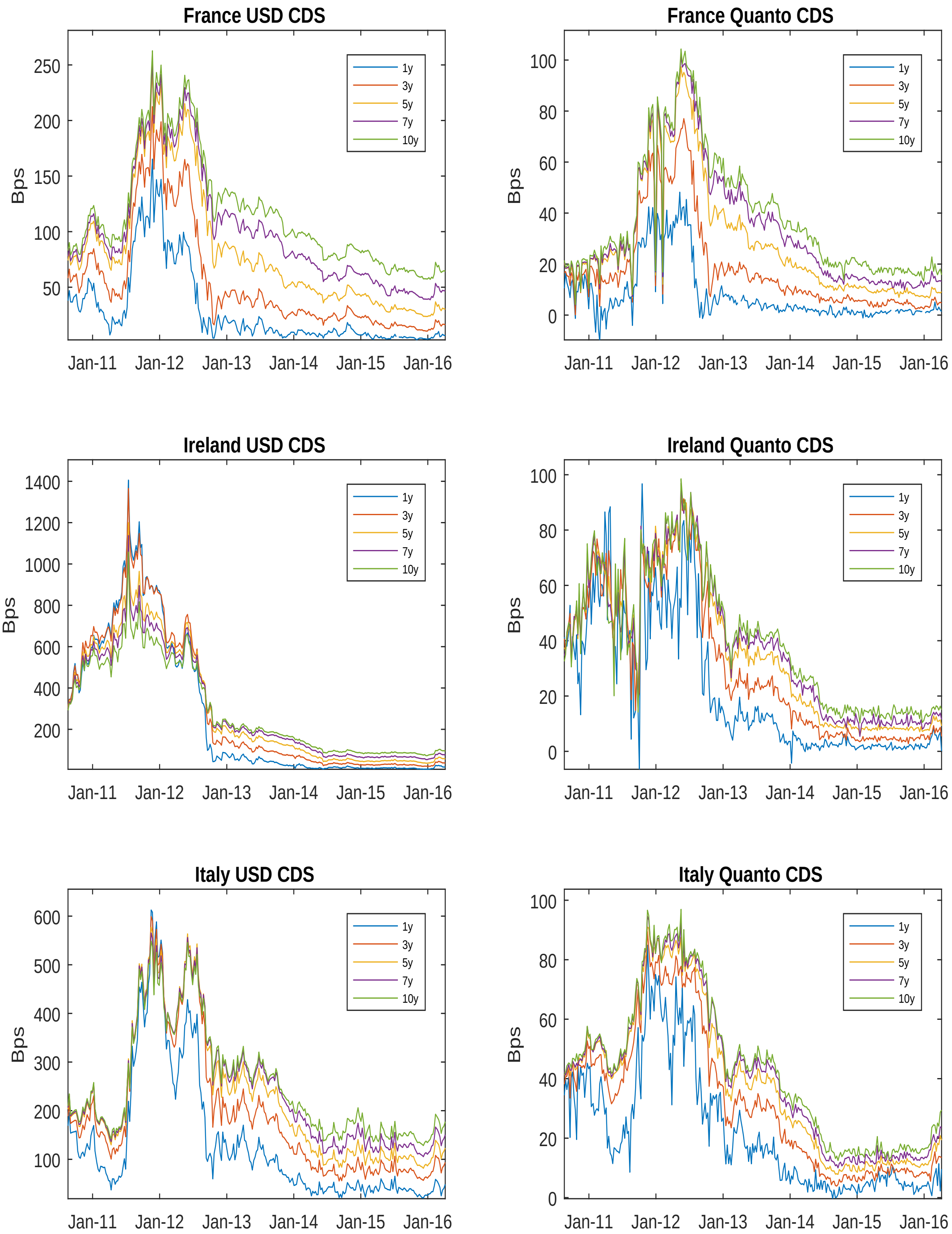

Figure 1.5: **USD CDS and quanto CDS spreads for France, Ireland, and Italy.** This figure shows USD CDS premiums and quanto CDS spreads–defined as the difference between USD and EUR-denominated CDS premiums of the same underlying reference entity–for France, Ireland, and Italy. The sample period is August 2010 to April 2016 and comprises 1402 daily observations obtained from Markit.

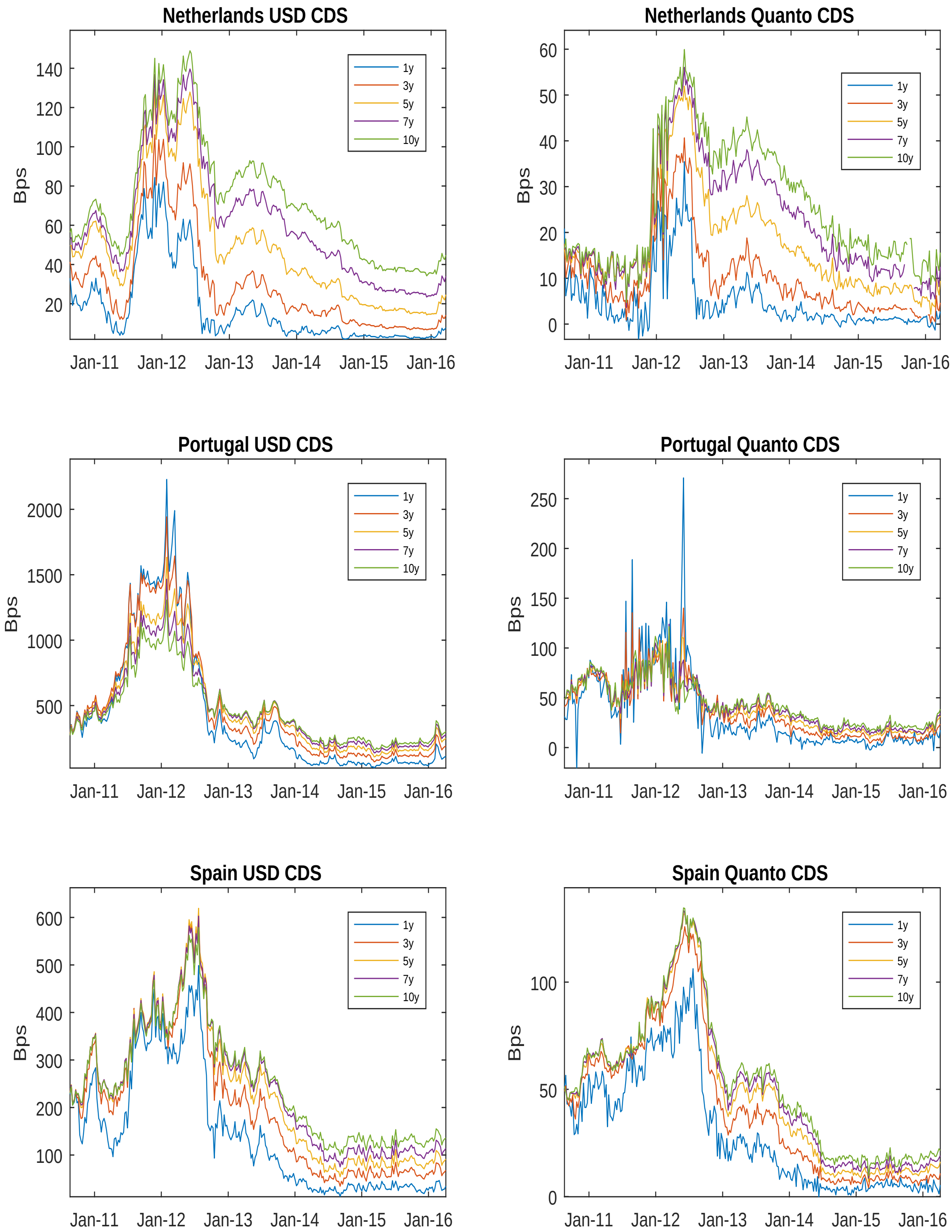

Figure 1.6: **USD CDS and quanto CDS spreads for Netherlands, Portugal, and Spain.** This figure shows USD CDS premiums and quanto CDS spreads–defined as the difference between USD and EUR-denominated CDS premiums of the same underlying reference entity–for Netherlands, Portugal, and Spain. The sample period is August 2010 to April 2016 and comprises 1402 daily observations obtained from Markit.

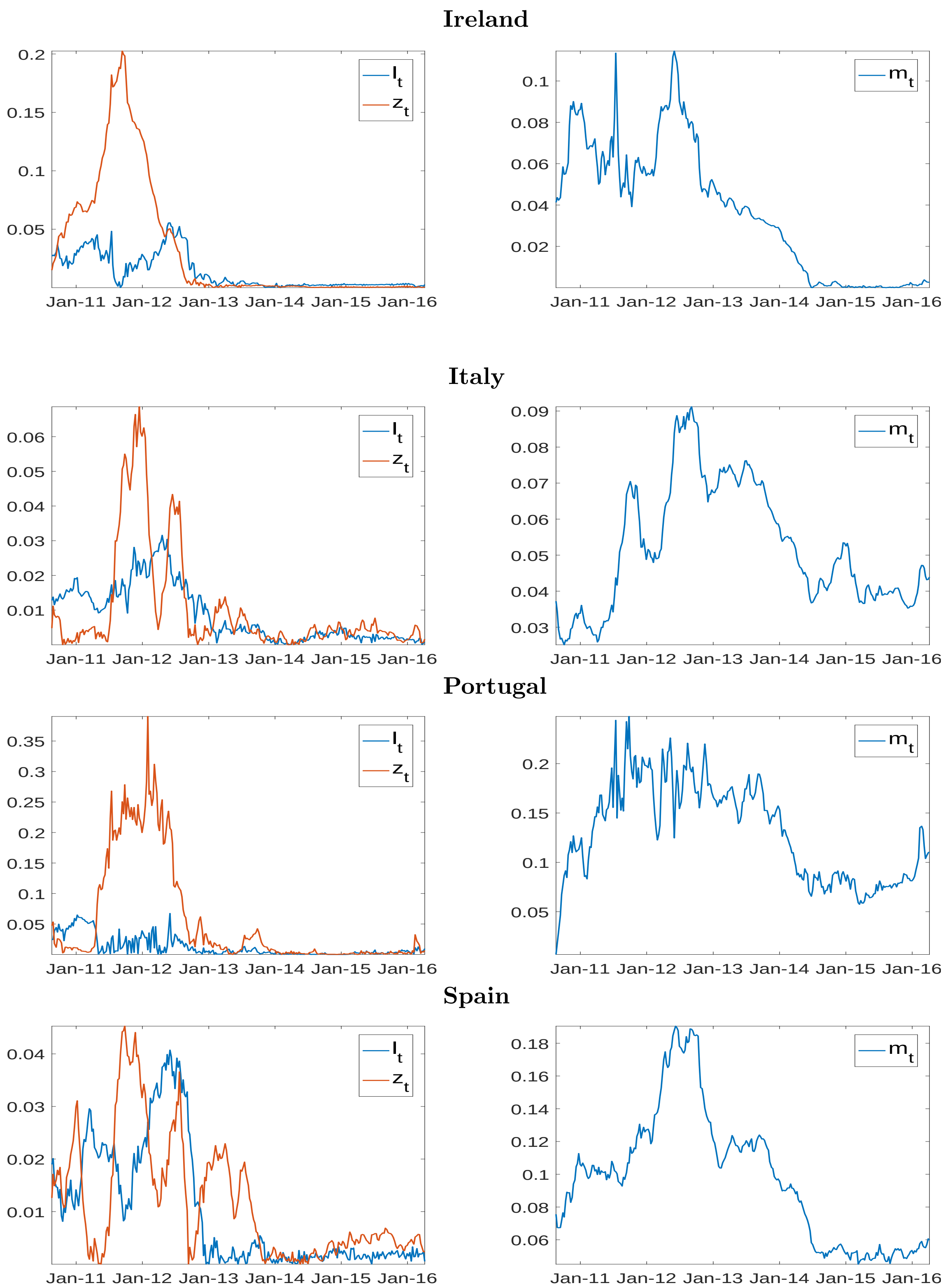

Figure 1.7: **Estimated time series of state variables.** This figure shows the time series of the estimated state variables. The left panel shows the state variables l_t and z_t and the right panel shows m_t. The model is estimated via maximum likelihood estimation in conjunction with the unscented Kalman filter. The sample period is August 2010 to April 2016 and each time series consists of 281 weekly observations.

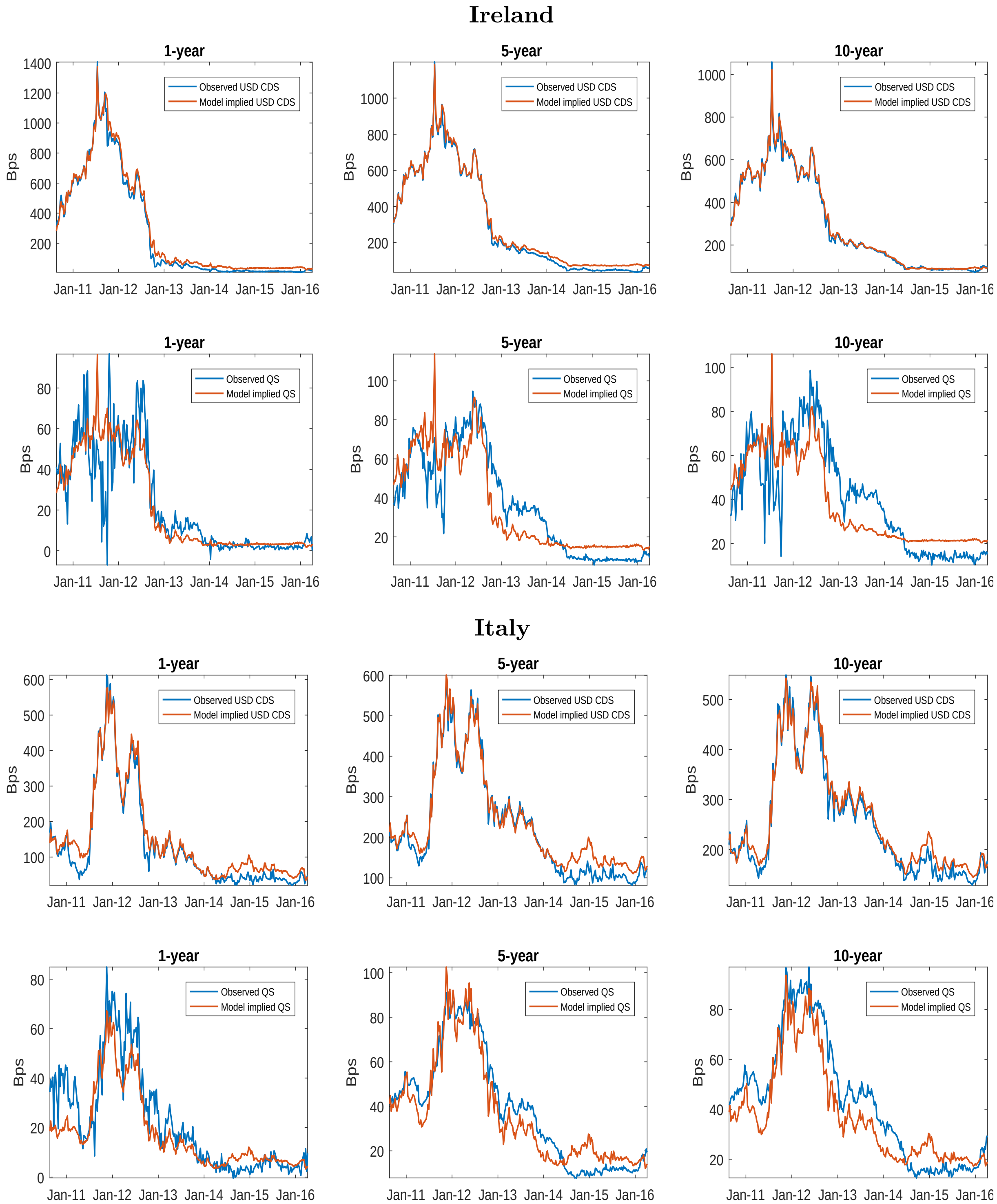

Figure 1.8: **Model fit for Ireland and Italy.** This figure shows the time series of the model-fitted versus the observed USD CDS premiums and quanto CDS spreads for Ireland and Italy. The illustrated maturities are 1, 5, and 10 years. The model is estimated via maximum likelihood estimation in conjunction with the unscented Kalman filter. The sample period is August 2010 to April 2016 and each time series consists of 281 weekly observations.

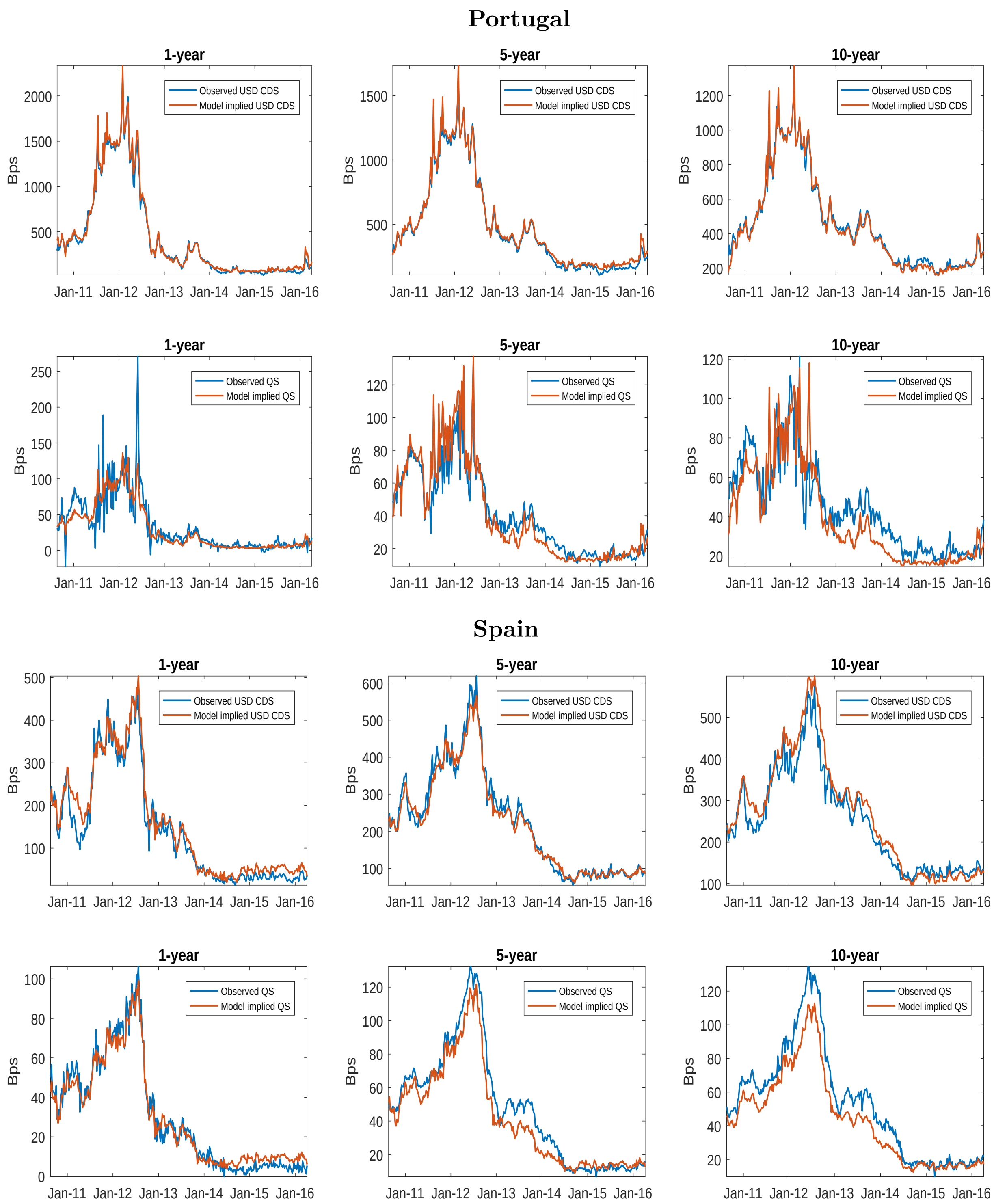

Figure 1.9: **Model fit for Portugal and Spain.** This figure shows the time series of the model-fitted versus the observed USD CDS premiums and quanto CDS spreads for Portugal and Spain. The illustrated maturities are 1, 5, and 10 years. The model is estimated via maximum likelihood estimation in conjunction with the unscented Kalman filter. The sample period is August 2010 to April 2016 and each time series consists of 281 weekly observations.

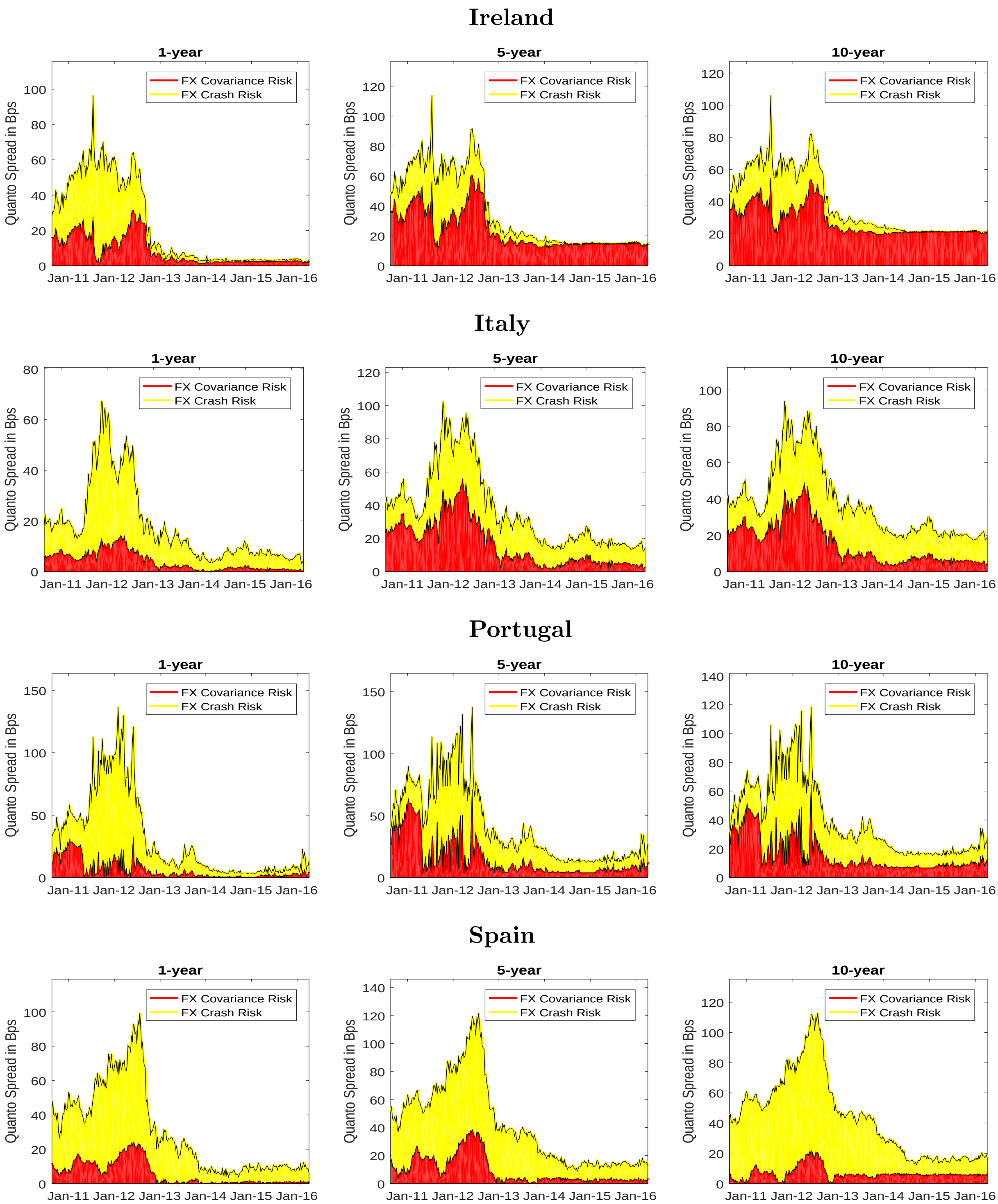

Figure 1.10: **Quanto spreads decomposed into covariance risk and currency crash risk.** This figure illustrates model decompositions of quanto CDS spreads—defined as the difference between USD and EUR-denominated CDS premiums—into a component driven by covariance between the exchange rate and default risk (orange) and a EURUSD jump risk component triggered by sovereign default (yellow). The illustrated maturities are 1, 5, and 10 years. The model is estimated via maximum likelihood estimation in conjunction with the unscented Kalman filter. The sample period is August 2010 to April 2016 and each time series consists of 281 weekly observations.

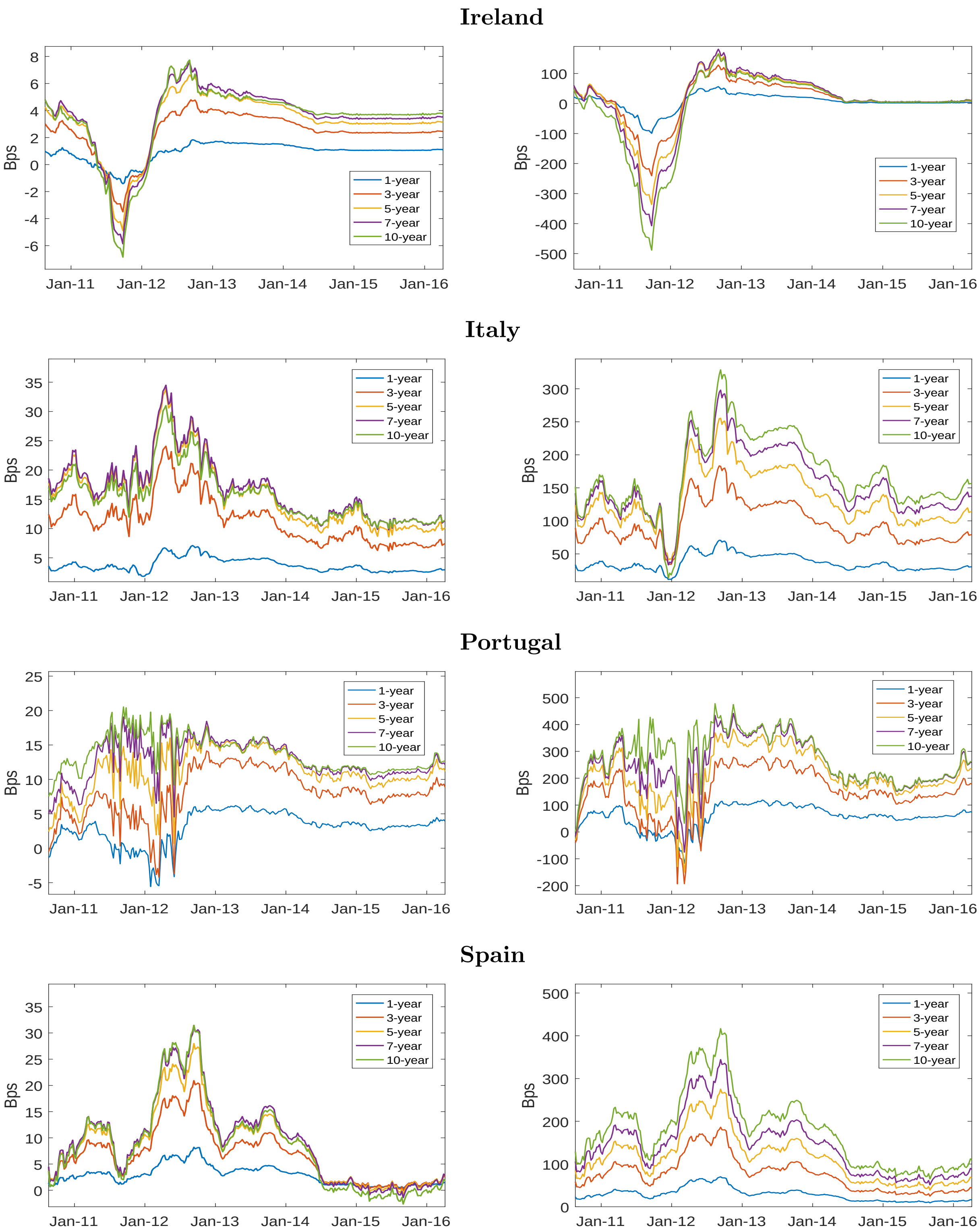

Figure 1.11: **Risk premiums for USD CDS and quanto CDS.** This figure shows the risk premiums associated with selling USD-denominated CDS (right panel) and the risk premiums associated with selling quanto CDS—defined as the difference between USD and EUR-denominated CDS premiums (left panel). The model is estimated via maximum likelihood estimation in conjunction with the unscented Kalman filter. The sample period is August 2010 to April 2016 and each time series consists of 281 weekly observations.

1.10 Tables

Table 1.1: **Model parameters calibrated to moments for the EURUSD and CDS premiums.** This table shows parameter values for the model calibrated to average 5-year quanto CDS spreads for Portugal, Ireland, Italy, and Spain. First column reports the calibrated values for ρ, which is estimated for each sovereign as the correlation between percent-wise changes in the 5-year USD-denominated CDS premium and the EURUSD exchange rate. The second column reports the value for e^u which equals the average risk-neutral volatility derived from EURUSD options maturing in one year. The third column shows annualized standard deviations of daily percent-wise changes in the USD-denominated CDS premiums, and the fourth and fifth columns report the average 5-year CDS premiums of the USD and EUR-denominated contracts, respectively. All moments are estimated over the period August 2010 to August 2012.

	$\rho = \mathrm{Corr}\big(\Delta S^U(t,5y), \Delta X_t\big)$	$e^u = \sigma_{FX}$	$\mathrm{Std}\big(\Delta S^U(t,5y)\big)$	$\mathrm{mean}\big(S^U(t,5y)\big)$	$\mathrm{mean}\big(S^E(t,5y)\big)$
Portugal	−36%	14.6%	57%	8.51%	7.81%
Ireland	−38%	——	51%	6.48%	5.84%
Italy	−56%	——	68%	3.30%	2.67%
Spain	−57%	——	73%	3.44%	2.68%

Table 1.2: **One-period example with crash risk in synthetic bond price.** This table shows the payoffs for a long position in a USD zero-coupon bond and a short position in a synthetic USD zero-coupon bond—which is short a EUR zero-coupon bond and long a forward contract. There are no recovery payments on the bonds. All contracts are initiated at time 0 and expire at time 1. The riskless interest rates are 0, the exchange rate is 1 at time 0, and the forward exchange rate is 1. The default states are assumed to be associated with a 50% depreciation in the EUR against the USD.

	$t = 0$	No default at $t = 1$	Default at $t = 1$
Long USD Bond	$-P^{USD}$	1 USD	0 USD
Short Synthetic USD Bond	$P^{USD,synth}$	-1 EUR + 1 EUR - 1 USD	1 EUR -1 USD
Cash Flow L/S in USD	0	0	-0.5 USD

Table 1.3: **Crash risk and currency/default covariance risk in bond yields.** This table compares yields on domestic and synthetic domestic coupon bonds derived via the discrete-time model. The synthetic domestic bond consists of a long position in a foreign bond that pays 1 at $t = 1, \ldots 5$ and 100 at maturity in foreign currency, and a short position in currency forward contracts that match those payments. The yield of the synthetic bond is reported in the first row with crash risk, and in the second row under the assumption of no crash risk. The third row shows the yield on a domestic coupon bond which pays 1 at $t = 1, \ldots 5$ and 100 at maturity. Rows 4-6 show the corresponding prices of the coupon bonds and the prices for each of the coupon payments. All bond payments are conditional on no default, and there are no recovery payments. The parameters used in the model are calibrated to 5-year EUR and USD CDS data for Spain and EURUSD moments (as reported in Table 1.1).

	$t = 1$	$t = 2$	$t = 3$	$t = 4$	$t = 5$
Yield Synthetic Coupon Bond, ($\delta = 0.84$)					4.46 %
Yield Synthetic Coupon Bond, ($\delta = 1$)					5.37 %
Yield Domestic Coupon Bond					5.73 %
Price Synthetic Domestic Coupon, ($\delta = 0.84$)	0.95	0.91	0.88	0.84	80.41
Price Synthetic Domestic Coupon, ($\delta = 1$)	0.95	0.90	0.85	0.81	76.99
Price Domestic Coupon	0.95	0.89	0.85	0.80	75.67

Table 1.4: **Summary statistics for USD CDS premiums and quanto CDS premiums.** This table reports sample estimates of the means and standard deviations of the USD-denominated CDS premiums and quanto CDS premiums for Austria, Belgium, Germany, Finland, France, Ireland, Italy, Netherlands, Portugal, and Spain. For each sovereign, the quanto CDS premium is defined as the difference in premiums on a USD and a EUR-denominated CDS contract at the same maturity. Panel A reports the time-series means of the premiums of the USD-denominated CDS contracts and the quanto CDS contracts in basis points at maturities of 1-10 years. Panel B reports the standard deviations of the premiums on the USD-denominated CDS contracts and the quanto CDS contracts in percentages at maturities of 1-10 years. The sample consists of daily quotes obtained from Markit from August 2010 to April 2016 (1402 observations for each series).

Panel A: Mean in bps										
	USD CDS					Quanto CDS				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
AUS	26.74	43.96	65.09	78.38	89.88	8.75	14.64	22.20	27.07	30.79
BEL	53.78	82.84	108.98	124.08	136.27	12.12	21.96	30.55	35.24	39.25
GER	11.26	21.72	39.26	51.72	63.08	3.70	9.07	16.90	21.48	25.51
FIN	12.87	21.05	34.27	43.98	52.93	2.54	5.00	8.49	11.10	12.96
FRA	29.01	54.02	82.49	100.45	115.90	8.39	18.60	28.90	34.17	38.42
IRE	271.14	303.88	297.09	298.27	291.82	22.59	31.35	36.46	38.64	40.55
ITA	132.99	193.37	224.06	240.09	250.72	21.84	32.45	38.39	41.01	43.11
NET	17.50	29.83	48.56	61.00	72.15	5.06	10.24	17.51	22.31	25.97
POR	437.00	489.10	478.21	471.72	456.50	34.43	37.71	41.04	42.72	44.56
SPA	142.89	198.90	225.70	239.26	247.96	30.16	41.60	47.68	50.36	52.84
Panel B: Standard Deviation in %										
	USD CDS					Quanto CDS				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
AUS	0.35	0.45	0.53	0.50	0.49	0.12	0.15	0.17	0.17	0.17
BEL	0.71	0.85	0.85	0.78	0.70	0.17	0.21	0.23	0.22	0.21
GER	0.13	0.19	0.27	0.27	0.27	0.05	0.08	0.13	0.14	0.15
FIN	0.13	0.16	0.19	0.18	0.17	0.03	0.03	0.03	0.04	0.05
FRA	0.35	0.47	0.55	0.52	0.50	0.12	0.18	0.23	0.24	0.24
IRE	3.48	3.37	2.82	2.52	2.20	0.25	0.27	0.26	0.25	0.24
ITA	1.34	1.39	1.33	1.22	1.11	0.20	0.24	0.25	0.25	0.24
NET	0.20	0.25	0.31	0.31	0.30	0.07	0.09	0.11	0.13	0.14
POR	5.03	4.55	3.59	3.08	2.61	0.39	0.29	0.25	0.23	0.22
SPA	1.27	1.45	1.41	1.30	1.17	0.28	0.33	0.34	0.33	0.32

Table 1.5: **Summary statistics for currency options data.** This table reports the means and standard deviations of implied volatilities for EURUSD delta-neutral straddles (STR), EURUSD 10 and 25-delta risk reversals (RR10 and RR25, respectively), and EURUSD 10 and 25-delta butterfly spreads (BF10 and BF25, respectively). All quantities are reported in percentages. The data are obtained from Bloomberg and the sample consists of daily quotes from August 2010 to April 2016 (1402 observations for each series).

	Mean (%)						Std (%)					
	1 mo	2 mo	3 mo	6 mo	9 mo	1 yr	1 mo	2 mo	3 mo	6 mo	9 mo	1 yr
STR	9.83	9.93	10.01	10.25	10.42	10.56	2.74	2.68	2.64	2.57	2.52	2.47
RR10	-1.70	-2.26	-2.70	-3.20	-3.47	-3.61	1.41	1.52	1.62	1.60	1.62	1.61
RR25	-1.02	-1.30	-1.51	-1.78	-1.90	-1.97	0.82	0.85	0.87	0.85	0.84	0.83
BF10	11.08	11.54	11.93	12.66	13.10	13.42	3.19	3.26	3.33	3.41	3.42	3.41
BF25	10.27	10.47	10.62	10.97	11.20	11.36	2.85	2.82	2.80	2.77	2.72	2.68

Table 1.6: **Regressions of FX spot and implied volatility changes on eurozone sovereign credit risk.** This table presents estimates from regressions of contemporaneous weekly changes in the EURUSD spot exchange rate and the EURUSD implied volatility on eurozone sovereign credit risk:

$$\Delta X_t = \alpha + \beta \Delta PC1_t^{CDS} + \varepsilon_t, \quad \Delta IV_t = \alpha + \beta \Delta PC1_t^{CDS} + \varepsilon_t$$

EURUSD volatility is proxied by the 1-month implied volatility of a delta-neutral straddle, and eurozone credit risk is measured as the first principal component of weekly 5-year CDS premiums for 10 eurozone sovereigns. Columns 1-2 show the results of the regressions using the full sample, columns 3-4 show the results from the crisis period (August 2010 to December 2012), and columns 5-6 show the results for the post-crisis period (January 2013 to April 2016). Newey and West (1987) t-statistics are reported in brackets, and the superscripts *, **, and *** indicate statistical significance at 10 %, 5 %, and 1 %, respectively. The currency spot and implied volatility data are from Bloomberg, and the CDS data are from Markit. The sample period is from August 2010 to April 2016 (281 weekly observations).

	2010-2016		2010-2012		2013-2016	
	ΔX	ΔIV	ΔX	ΔIV	ΔX	ΔIV
α	-0.0001	-0.0000	0.0005	0.0007	-0.0003	-0.0003
	[-0.27]	[-0.06]	[0.64]	[0.97]	[-1.38]	[-0.80]
β	-0.0988^{***}	0.2330^{***}	-0.1720^{***}	0.5046^{***}	-0.0191	0.0672^{**}
	[-3.69]	[3.84]	[-4.01]	[7.70]	[-1.34]	[2.17]
R^2	0.081	0.121	0.162	0.272	0.002	0.032

Table 1.7: **Parameter estimates for the proposed affine model.** This table reports parameter estimates of the affine model specified in equation (1.16). The numbers in parentheses are standard errors of the estimates. The parameters are estimated using maximum likelihood estimation in conjunction with the unscented Kalman filter, using premiums on USD-denominated CDS and quanto CDS contracts with maturities of 1-10 years, and EURUSD option-implied volatilities at five strikes and six maturities spanning 1-12 months. Each time series consists of 281 weekly observations (each Wednesday) from August 2010 to April 2016.

	Intensity Parameters				**FX Parameters**	
	Ireland	Italy	Portugal	Spain		
κ_l	0.0326	0.2533	0.0308	0.0907	κ_v	1.2129
	(0.0027)	(0.0367)	(0.0133)	(0.0597)		($0.5890 \cdot 10^{-3}$)
θ_l	0.1760	0.0018	0.0608	0.0188	θ_v	0.0183
	(0.0051)	(0.0011)	(0.0019)	(0.0069)		($0.9021 \cdot 10^{-4}$)
σ_l	0.4392	0.2892	0.3484	0.4525	σ_v	0.1452
	(0.0099)	(0.0154)	(0.0081)	($2.56 \cdot 10^{-5}$)		($0.3740 \cdot 10^{-3}$)
κ_l^P	0.0469	0.0005	0.0007	0.0001	κ_v^P	1.5935
	(0.0136)	(0.8737)	(0.0101)	(0.0119)		($0.1305 \cdot 10^{-3}$)
θ_l^P	0.1295	0.0092	(0.0059)	0.0336	θ_v^P	0.0174
	(0.0086)	(0.0068)	(0.0007)	(0.0267)		($0.9532 \cdot 10^{-2}$)
κ_z	0.2620	0.2460	0.2450	0.1283	ρ	-0.6817
	(0.0073)	(0.0298)	(0.0053)	(0.0208)		($0.9032 \cdot 10^{-3}$)
σ_z	0.0000	0.0013	0.3660	0.0445	σ_O	$0.8512 \cdot 10^{-4}$
	(0.0093)	0.0025	0.0031	(0.0617)		($0.5815 \cdot 10^{-3}$)
κ_z^P	0.0037	0.0041	0.0000	0.0010		
	(0.0159)	(0.1480)	(0.0060)	(0.1539)		
κ_m	0.0035	0.0012	0.0241	0.0023		
	(0.0164)	(0.0027)	(0.0057)	(0.0035)		
θ_m	0.0000	0.0000	0.0043	0.0116		
	(0.0071)	(0.0092)	(0.0139)	(0.0052)		
σ_m	0.2200	0.1099	0.2059	0.1201		
	(0.0051)	(0.0100)	(0.0015)	(0.0063)		
κ_m^P	0.0010	0.1641	0.0009	0.0639		
	(0.1690)	(0.0967)	(0.0037)	(0.4256)		
θ_m^P	0.0000	0.0000	0.2615	0.1002		
	(0.0027)	(0.0300)	(0.0123)	(0.6966)		
ζ	-0.0502	-0.0960	-0.0543	-0.1559		
	(0.0041)	(0.0050)	(0.0018)	(0.0012)		
l_0	(0.0098)	(0.0105)	(0.0042)	(0.0033)		
	(0.0003)	(0.0035)	(0.0008)	(0.0032)		
z_0	0.0015	0.0029	0.0272	0.0000		
	(0.002)	(0.0107)	(0.0040)	(0.1965)		
m_0	0.0435	0.0459	(0.0022)	(0.0793)		
	(0.0023)	(0.0093)	(0.0040)	(0.0120)		
σ_U	$3.24 \cdot 10^{-6}$	$1.06 \cdot 10^{-6}$	$1.91 \cdot 10^{-6}$	$5.16 \cdot 10^{-6}$		
	($4.40 \cdot 10^{-6}$)	($3.53 \cdot 10^{-6}$)	($2.91 \cdot 10^{-6}$)	($2.51 \cdot 10^{-6}$)		
σ_{UE}	$4.12 \cdot 10^{-5}$	$2.47 \cdot 10^{-5}$	$0.74 \cdot 10^{-5}$	$0.45 \cdot 10^{-5}$		
	($4.97 \cdot 10^{-6}$)	($1.549 \cdot 10^{-6}$)	($2.22 \cdot 10^{-6}$)	($6.22 \cdot 10^{-6}$)		

Table 1.8: **Summary statistics of model pricing errors for USD CDS and quanto CDS premiums.** This table reports the root mean squared errors and mean absolute pricing errors for model-implied USD-denominated CDS premiums and quanto CDS premiums at maturities from 1-10 years. Both are reported in basis points (bps). The pricing error is defined as the difference between the observed CDS premium/quanto CDS premium and the model-implied CDS premium/quanto CDS premium (using the updated state variable). The model is estimated using maximum likelihood estimation in conjunction with the unscented Kalman filter using USD CDS data, quanto CDS data (both from Markit), and currency options data from Bloomberg. The sample consists of 281 weekly observations from August 2010 to April 2016.

Panel A: Root Mean Squared Errors (bps)												
	USD CDS						Quanto CDS					
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	Mean	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	Mean
IRE	37.61	36.93	24.55	18.93	15.35	26.68	0.65	0.67	0.44	1.04	2.09	0.98
ITA	26.65	25.23	25.18	21.58	17.41	23.21	3.89	2.16	1.48	2.41	5.36	3.06
POR	50.50	45.82	36.81	28.35	28.12	37.92	3.48	1.77	0.43	1.42	4.70	2.36
SPA	26.27	23.61	20.42	18.70	34.04	24.61	0.15	1.92	5.79	7.66	8.99	4.90
Panel B: Mean Absolute Pricing Errors (bps)												
	USD CDS						Quanto CDS					
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	Mean	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	Mean
IRE	28.44	32.48	18.92	13.52	9.07	20.49	7.24	8.06	9.86	11.01	11.27	9.49
ITA	19.44	18.72	18.76	15.58	13.97	17.29	6.82	6.31	6.90	7.31	8.80	7.23
POR	30.93	36.70	25.75	15.71	18.12	25.44	11.02	7.22	6.74	7.61	9.43	8.40
SPA	19.81	17.18	15.11	13.58	26.70	18.48	3.81	5.13	7.20	8.29	9.13	6.71

Table 1.9: **Summary statistics for decompositions of quanto CDS spreads.** This table reports summary statistics for model decompositions of quanto CDS spreads into a covariance risk component and a crash risk component. Panel A reports the mean and the maximum of the covariance component in basis points (bps) over the full sample period. Panel B reports the mean and maximum share for the covariance component of the total quanto CDS spread over the full sample. Panel C and D report the same quantities but for the debt crisis period (August 2010 to December 2012). The model is estimated using maximum likelihood estimation in cojunction with the unscented Kalman filter based on USD CDS data, quanto CDS data (both from Markit), and currency options data from Bloomberg. The sample consists of 281 weekly observations from August 2010 to April 2016.

Panel A: Full sample (August 2010 - April 2016)										
	Mean covariance component (bps)					Max covariance component (bps)				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
IRE	8.09	18.95	23.49	25.55	27.25	31.26	59.04	60.60	57.23	54.62
ITA	3.98	12.08	16.35	16.84	15.40	14.37	42.19	55.23	55.15	48.39
POR	5.98	13.02	15.16	15.54	15.51	32.24	64.36	70.97	69.32	64.56
SPA	5.66	9.86	9.24	8.03	6.77	24.06	43.61	38.51	30.23	21.89
	Share of spread from covariance risk					Max share of spread from covariance risk				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
IRE	0.54	0.72	0.75	0.77	0.79	0.82	0.96	0.98	0.98	0.99
ITA	0.20	0.34	0.38	0.38	0.37	0.39	0.60	0.65	0.65	0.62
POR	0.21	0.31	0.35	0.37	0.40	0.56	0.72	0.76	0.78	0.78
SPA	0.13	0.17	0.20	0.21	0.21	0.35	0.46	0.40	0.31	0.21
Panel B: Debt Crisis Period (August 2010 - December 2012)										
	Mean covariance component (bps)					Max covariance component (bps)				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
IRE	15.78	32.27	35.81	36.04	36.02	31.26	59.04	60.60	57.23	54.62
ITA	7.77	23.14	30.70	30.92	27.34	14.37	42.19	55.23	55.15	48.39
POR	11.95	24.62	27.18	26.57	25.11	32.24	64.36	70.97	69.32	64.56
SPA	12.49	22.00	18.35	13.23	8.18	24.06	43.61	38.51	30.23	21.89
	Share of spread from covariance risk					Max share of spread from covariance risk				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
IRE	0.37	0.54	0.58	0.60	0.62	0.61	0.74	0.76	0.77	0.78
ITA	0.27	0.48	0.53	0.53	0.50	0.39	0.60	0.65	0.65	0.62
POR	0.23	0.35	0.38	0.39	0.40	0.55	0.72	0.76	0.78	0.78
SPA	0.22	0.30	0.25	0.18	0.11	0.35	0.46	0.40	0.31	0.21

Table 1.10: **Summary statistics for risk premiums of USD CDS and quanto CDS.** This table shows risk premiums associated with holding USD CDS and quanto CDS for Ireland, Italy, Portugal, and Spain. Panel A reports the mean risk premiums for holding USD CDS and quanto CDS in basis points at maturities of 1-10 years. Panel B reports the average risk premiums for USD CDS and quanto CDS as a fraction of total spreads. The model is estimated using maximum likelihood estimation in conjunction with the unscented Kalman filter based on USD CDS data, quanto CDS data (both from Markit), and currency options data from Bloomberg. The sample consists of 281 weekly observations from August 2010 to April 2016.

Panel A: Mean risk premium in bps

	USD CDS					Quanto CDS				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
IRE	0.66	1.90	2.34	1.68	-3.18	0.92	2.36	3.14	3.44	3.33
ITA	36.60	96.01	134.81	156.92	171.41	3.76	11.46	15.94	16.88	15.81
POR	55.96	146.97	211.29	251.03	278.33	3.00	7.96	11.18	12.83	13.76
SPA	28.84	76.59	114.31	144.67	177.44	2.85	6.64	8.48	9.25	8.56

Panel B: Mean risk premium as a fraction of spread

	USD CDS					Quanto CDS				
	1 yr	3 yrs	5 yrs	7 yrs	10 yrs	1 yr	3 yrs	5 yrs	7 yrs	10 yrs
IRE	0.03	0.05	0.06	0.06	0.05	0.25	0.17	0.15	0.14	0.12
ITA	0.38	0.58	0.66	0.69	0.72	0.63	0.69	0.73	0.75	0.77
POR	0.42	0.57	0.64	0.69	0.75	0.25	0.42	0.40	0.39	0.38
SPA	0.28	0.48	0.59	0.65	0.71	0.49	0.51	0.61	0.45	0.67

Table 1.11: **Summary statistics for observed quanto yield spreads.** This table reports summary statistics for observed quanto yield spreads. The synthetic quanto yield spread is the difference in yields between a USD bond and a synthetic USD bond, which is constructed based on EUR bond credit spreads. The synthetic USD bond is constructed such that it matches the coupon scheme, notional value, and time to maturity of the USD bond. The quanto bond yield spread is computed as the difference in yields on coupon bonds denominated in USD and EUR with similar maturities corrected for the riskless interest rate differential. Newey-West t-statistics of the means are reported in square brackets. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1%, respectively. The sample period is from August 2010 to April 2016 (281 observations).

Panel A: Debt Crisis (August 2010 – March 2013)						
	IT (synth)	IT (bond)	ES (synth)	ES (bond)	PT (synth)	PT (bond)
Mean (bps)	40.82***	59.67***	62.65***	98.98***	4.31	28.61
	[5.42]	[9.45]	[6.31]	[7.85]	[0.35]	[1.02]
Std (%)	0.39	0.39	0.62	0.81	0.67	1.91
Skew	0.44	0.46	0.65	0.61	0.25	1.00
Q5 (bps)	-16.35	1.76	-26.61	-3.44	-91.89	-198.73
Q95 (bps)	123.06	140.06	178.21	240.36	112.53	360.88
Fraction > 0	0.86	0.95	0.89	0.91	0.50	0.40
Panel B: Post Debt Crisis (March 2013 – April 2016)						
	IT (synth)	IT (bond)	ES (synth)	ES (bond)		
Mean (bps)	14.04	25.81***	33.31***	22.15***		
	[0.84]	[3.21]	[4.47]	[3.47]		
Std (%)	0.27	0.22	0.19	0.20		
Skew	-0.34	-0.17	-0.41	0.09		
Q5 (bps)	-34.66	-8.84	-1.12	-8.30		
Q95 (bps)	55.74	60.70	60.35	53.22		
Fraction > 0	0.74	0.88	0.94	0.85		

Table 1.12: **Regressions of observed quanto yield spreads on model-implied quanto yield spreads.** This table shows the results from regressing observed quanto yield spreads on model-implied quanto yield spreads (Model QY) and observed 5-year quanto CDS spreads (5Y QCDS). The observed synthetic quanto yield spread is the difference in yields on a USD bond and a synthetic USD bond, which is constructed from EUR credit spreads. The synthetic USD bond is constructed such that it matches the coupon scheme, notional value, and time to maturity of the USD bond. The quanto bond yield spread is computed as the difference in yields on comparable coupon bonds denominated in USD and EUR corrected for the riskless interest rate differential. Newey-West t-statistics are reported in square brackets. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1%, respectively. The sample period is from August 2010 to April 2016 (281 observations).

Panel A: Debt Crisis (August 2010 – March 2013)						
	IT (synth)	IT (bond)	ES (synth)	ES (bond)	PT (synth)	PT (bond)
Model QY	0.99	1.03**	1.37***	1.60***	1.16*	2.85
	[1.43]	[2.28]	[2.99]	[6.60]	[1.70]	[1.03]
Intercept	0.00	-0.00	0.00	0.01	-0.01	−0.02
	[-0.29]	[-0.03]	[0.32]	[-0.99]	[-1.57]	[-0.60]
R^2 (%)	8.66	15.90	20.75	35.54	5.83	3.29
5Y QCDS	0.92**	1.19***	1.24***	1.93***	0.15	0.16
	[2.07]	[3.22]	[4.19]	[4.30]	[0.21]	[0.06]
Intercept	0.00	0.00	0.00	-0.01	0.00	0.01
	[-0.68]	[-0.71]	[-1.69]	[-1.30]	[-0.19]	[0.27]
R^2 (%)	17.06	29.19	25.28	35.44	0.19	0.00
Panel B: Post Debt Crisis (March 2013 – April 2016)						
	IT (synth)	IT (bond)	ES (synth)	ES (bond)		
Model QY	-1.37	-0.54	0.27	0.38*		
	[-1.60]	[-0.51]	[-0.93]	[1.66]		
Intercept	0.00***	0.00**	0.00*	0.00		
	[2.66]	[2.85]	[1.88]	[0.54]		
R^2 (%)	17.60	5.71	3.38	6.81		
5Y QCDS	-0.49	-0.08	0.24	0.32*		
	[-1.13]	[-0.21]	[1.44]	[1.86]		
Intercept	0.00***	0.00***	0.00***	0.00***		
	[2.66]	[3.17]	[4.23]	[2.35]		
R^2 (%)	7.72	0.26	3.86	6.50		

1.11 Appendix: Discrete-Time Model

1.11.1 Crash Risk Consistent with No-Arbitrage

Define the time t risk-neutral expectation of the exchange rate, i.e., the time t forward price:

$$E_t^Q(X_{t+1}) = F \tag{1.39}$$

Crash risk in the exchange rate upon default is modeled as follows. If default occurs between t and $t+1$, the exchange rate takes a hit of δ compared to the time t forward price

$$E_t^Q(X_{t+1}|\tau = t+1) = \delta E_t^Q(X_{t+1}) = \delta \cdot F \tag{1.40}$$

We refer to δ as the expected depreciation upon default or the crash risk parameter. Combining equations (1.39) and (1.40) gives

$$F = E_t^Q(X_{t+1}|\tau > t)Q(\tau > t+1|\tau > t) + E_t^Q(X_{t+1}|\tau = t)Q(\tau = t+1|\tau > t) \tag{1.41}$$

Rearranging,

$$E_t^Q(X_{t+1}|\tau > t+1) = \frac{1 - \delta Q(\tau = t+1|\tau > t)}{1 - Q(\tau = t+1|\tau > t)} F = \frac{1 - \delta\lambda_t}{1 - \lambda_t} F \tag{1.42}$$

Assume the exchange rate appreciates unconditionally with u with probability q and depreciates u^{-1} with probability $1-q$. Then we obtain an arbitrage-free model in each node by scaling the states of the exchange rate conditional on default with δ and the states of the exchange rate conditional on survival with $C(\lambda_t)$:

$$\frac{F}{X_t} = qu + (1-q)u^{-1} = \lambda_t \delta \Big(qu + (1-q)u^{-1} \Big) + (1-\lambda_t)C(\lambda_t)\Big(qu + (1-q)u^{-1} \Big)$$

Since each node is free of arbitrage, the entire model is free of arbitrage. Furthermore, for a given forward price, we see that

$$q = \frac{\frac{F}{X_t} - u}{u - u^{-1}}$$

1.11.2 Proofs in the Discrete-Time Model

Domestic CDS Premium

Define the unconditional mean default probability that prevails in the next period as $\bar{\lambda} = q^{\lambda}\lambda^{U} + (1-q^{\lambda})\lambda^{D}$. Then we can express the CDS premium as:

$$S^{d}(0,T) = (1-R)\frac{P_d^2\bar{\lambda}\left(1-\bar{\lambda}\right)\left(\frac{1-\left(\left(1-\bar{\lambda}\right)P_d\right)^{T-1}}{1-\left(1-\bar{\lambda}\right)P_d}\right)+P_d\lambda_0}{\left(1-\lambda_0\right)P_d\left(\frac{\left(\left(1-\bar{\lambda}\right)P_d\right)^{T}}{1-\left(1-\bar{\lambda}\right)P_d}\right)} \tag{1.43}$$

Proof. In general, the discrete-time CDS premium in domestic currency with maturity T, $S^{d}(0,T)$, is given by:

$$S^{d}(0,T) = (1-R)\frac{\sum_{t=1}^{N} P_d(0,t)E_0^{Q}\left(1_{\tau=t}\right)}{\sum_{i=1}^{N} P_d(0,t)E_0^{Q}\left(1_{\tau>t}\right)} = \frac{\sum_{t=1}^{N} P_d^{t}E_0^{Q}\left(1_{\tau=t}\right)}{\sum_{i=1}^{N} P_d^{t}E_0^{Q}\left(1_{\tau>t}\right)} \tag{1.44}$$

The last equal sign follows from the assumption of a flat interest rate term structure such that $P_d(0,t) = P_d^{t}$, where P_d is a one-period domestic discount bond.

The survival probability up and until time t is straightforward to compute in the model, since the one-period survival probabilities are independent across time:

$$\begin{aligned}
E_0^{Q}\left(1_{\tau>t}\right) = Q_0\left(\tau>t\right) &= (1-\lambda_0)E_0^{Q}\left(\prod_{i=1}^{t-1}1-\lambda_i\right) = (1-\lambda_0)\prod_{i=1}^{t-1}E_0^{Q}\left(1-\lambda_i\right)\\
&= (1-\lambda_0)\prod_{i=1}^{t-1}1-\left(q^{\lambda}\lambda^{U}+(1-q^{\lambda})\lambda^{D}\right) = (1-\lambda_0)\prod_{i=1}^{t-1}\left(1-\bar{\lambda}\right)\\
&= (1-\lambda_0)\left(1-\bar{\lambda}\right)^{t-1}
\end{aligned}$$

For two periods or longer, we can express the default probability in terms of the difference between the survival probability up and until time $t-1$ and survival probability up to time

t:

$$E_0^Q\left(1_{\tau=t}\right) = Q_0\left(\tau > t-1\right) - Q_0\left(\tau > t\right) = \left(1-\lambda_0\right)\left(\left(1-\bar{\lambda}\right)^{t-2} - \left(1-\bar{\lambda}\right)^{t-1}\right)$$
$$= \bar{\lambda}\cdot\left(1-\lambda_0\right)\left(1-\bar{\lambda}\right)^{t-2} \quad \text{for} \quad t \geq 2$$

Plugging the premium and protection leg payments into (1.44) and by using the expression of a geometric series, we get

$$S^d(0,T) = (1-R)\frac{P_d^2\bar{\lambda}\left(1-\lambda_0\right)\sum_{t=0}^{T-2} P_d^t\left(1-\bar{\lambda}\right)^t + P_d\lambda_0}{\left(1-\lambda_0\right)P_d\sum_{t=0}^{T-1} P_d^t\left(1-\bar{\lambda}\right)^t}$$
$$= (1-R)\frac{P_d^2\bar{\lambda}\left(1-\bar{\lambda}\right)\left(\frac{1-\left(\left(1-\bar{\lambda}\right)P_d\right)^{T-1}}{1-\left(1-\bar{\lambda}\right)P_d}\right) + P_d\lambda_0}{\left(1-\lambda_0\right)P_d\left(\frac{\left(\left(1-\bar{\lambda}\right)P_d\right)^T}{1-\left(1-\bar{\lambda}\right)P_d}\right)}$$

□

In the specific case when $\lambda_0 = \bar{\lambda}$, we have:

$$S^d(0,T) = (1-R)\frac{P_d^2\bar{\lambda}\left(1-\lambda_0\right)\sum_{t=0}^{T-2} P_d^t\left(1-\bar{\lambda}\right)^t + P_d\lambda_0}{\left(1-\bar{\lambda}\right)P_d\sum_{t=0}^{T-1} P_d^t\left(1-\bar{\lambda}\right)^t}$$
$$= (1-R)\frac{P_d^2\bar{\lambda}\left(1-\bar{\lambda}\right)\left(\frac{1-\left(\left(1-\bar{\lambda}\right)P_d\right)^{T-1}}{1-\left(1-\bar{\lambda}\right)P_d}\right) + P_d\lambda_0}{\left(1-\lambda_0\right)P_d\left(\frac{\left(\left(1-\bar{\lambda}\right)P_d\right)^T}{1-\left(1-\bar{\lambda}\right)P_d}\right)}$$
$$= (1-R)\frac{P_d^2\bar{\lambda}(1-\bar{\lambda})\left(1-\left(\left(1-\bar{\lambda}\right)P_d\right)^{T-1}\right) + P_d\bar{\lambda}\left(1-(1-\bar{\lambda})P_d\right)}{(1-\bar{\lambda})P_d\left(1-((1-\bar{\lambda})P_d)^T\right)}$$
$$= (1-R)\frac{P_d\bar{\lambda}\left(P_d(1-\bar{\lambda}) - \left(\left(1-\bar{\lambda}\right)P_d\right)^T + (1-(1-\bar{\lambda})P_d)\right)}{P_d(1-\bar{\lambda})\left(1-\left(\left(1-\bar{\lambda}\right)P_d\right)^T\right)}$$
$$= (1-R)\frac{\bar{\lambda}}{1-\bar{\lambda}} \tag{1.45}$$

Derivation of Foreign CDS Premium

In this section, we show that the expression for the discrete-time foreign CDS premium is:

$$S^f(0,t) = (1-R)\frac{P_d^2\,(F-L)\,L_0\frac{1-(LP_d)^{T-1}}{1-LP_d} + P_d(F-L_0)}{P_d L_0\frac{1-(LP_d)^T}{1-LP_d}} \tag{1.46}$$

where $L_0 = F\,(1-\delta\lambda_0)$, $L = F(1-\delta\bar{\lambda}) - K\delta\rho\,(u-u^{-1})\left(\lambda^U - \lambda^D\right)$ and $K = \sqrt{qq^\lambda(1-q)(1-q^\lambda)}$.

Proof. When determining the foreign CDS premium, $S^f(0,T)$, we exchange the payment stream of the premium leg and the protection leg into units of domestic currency using M_t defined in (1.8):

$$S^f(0,T) = (1-R)\frac{\sum_{t=1}^N P_f^t E_0^{Q^f}(1_{\tau=t})}{\sum_{t=1}^N P_f^t E_0^{Q^f}(1_{\tau>t})} = (1-R)\frac{\sum_{t=1}^N P_d^t E_0^Q\left(\frac{X_t}{X_0}1_{\tau=i}\right)}{\sum_{t=1}^N P_d^t E_0^Q\left(\frac{X_t}{X_0}1_{\tau>t}\right)}$$

At each point in time, there are 4 possible states for the default probability, λ_t and the one-period relate changes in the exchange rate $\frac{X_{t+1}}{X_t}$: $((u,\tilde{\lambda}^1),(u,\tilde{\lambda}^0),(u^{-1},\tilde{\lambda}^1),(u^{-1},\tilde{\lambda}^0))$, which are reached with respective probabilities $(Q_{11},Q_{10},Q_{01},Q_{00})$, where we have used the notation $\tilde{\lambda}^1 = \lambda^U$ and $\tilde{\lambda}^0 = \lambda^D$.

For each survival step, the exchange rate needs to be adjusted for the compensating factor defined as: $C(\lambda) = \frac{1-\delta\lambda}{1-\lambda}$ in order to preclude arbitrage opportunities. Important to mention is that the levels of the one-step survival probabilities are independent of one another, and so are the relative changes in the exchange rate. In summary, only the one-step changes in the exchange rate from t to $t+1$ and the default probability at time t are correlated (this is what gives us the FX/default covariance risk effect). These assumptions give us the following expression for the price of a defaultable foreign bond in terms of domestic currency:

$$\begin{aligned}
P_f^t E_0^{Q^f}\left(1_{\tau>t}\right) &= P_d^t E_0^{Q}\left(X_t 1_{\tau>t}\right) \\
&= P_d^t E_0^{Q}\left(\prod_{k=0}^{t-1}(1-\lambda_k)C(\lambda_k)\frac{X_{k+1}}{X_k}\right) \\
&= P_d^t E_0^{Q}\left(\prod_{k=0}^{t-1}(1-\delta\lambda_k)\frac{X_{k+1}}{X_k}\right) \\
&= P_d^t\left(1-\delta\lambda_0\right)E_0^{Q}\left(\frac{X_1}{X_0}\right)\prod_{i=1}^{t-1}E_0^{Q}\left((1-\delta\lambda_k)\frac{X_{k+1}}{X_k}\right) \\
&= P_d^t\left(1-\delta\lambda_0\right)F\prod_{i=1}^{t-1}\left(\sum_{i,j=0,1}Q_{ij}(1-\delta\tilde{\lambda^j})u^{2i-1}\right) \\
&= P_d^t\left(1-\delta\lambda_0\right)F\left(\sum_{i,j=0,1}Q_{ij}(1-\delta\tilde{\lambda^j})u^{2i-1}\right)^{t-1} \qquad (1.47)
\end{aligned}$$

Next, we calculate an expression for the last term in equation (1.47) by plugging in the Q_{ij}s:

$$\begin{aligned}
\left(\sum_{i,j=0,1}Q_{ij}(1-\delta\tilde{\lambda^j})u^{2i-1}\right) &= qu\Big(q^\lambda\left(1-\delta\lambda^U\right)+\left(1-q^\lambda\right)\left(1-\delta\lambda^D\right)\Big) \\
&+(1-q)u^{-1}\Big(q^\lambda\left(1-\delta\lambda^U\right)+\left(1-q^\lambda\right)\left(1-\delta\lambda^D\right)\Big) \\
&+quA_1\Big(\left(1-\delta\lambda^U\right)-\left(1-\delta\lambda^D\right)\Big) \\
&+(1-q)u^{-1}A_0\Big(\left(1-\delta\lambda^D\right)-\left(1-\delta\lambda^U\right)\Big) \\
&= \Big(qu+(1-q)u^{-1}\Big)\Big(q^\lambda\left(1-\delta\lambda^U\right)+\left(1-q^\lambda\right)\left(1-\delta\lambda^D\right)\Big) \\
&+quA_1\delta(\lambda^D-\lambda^U)-(1-q)u^{-1}A_0\delta(\lambda^D-\lambda^U) \\
&= F(1-\delta\bar{\lambda})-K\delta\rho\left(u-u^{-1}\right)\left(\lambda^U-\lambda^D\right)\equiv L \qquad (1.48)
\end{aligned}$$

where $K=\sqrt{qq^\lambda(1-q)(1-q^\lambda)}$. In the last equal sign, we use the no-arbitrage condition of a one-period forward contract, $F=qu+(1-q)u^{-1}$, and the fact that $qA_1=(1-q)A_0=\rho\sqrt{qq^\lambda(1-q)(1-q^\lambda)}$.

Next step is to express $E_0^{Q^f}\left(1_{\tau=t}\right)$ in terms of $E_0^{Q^f}\left(1_{\tau>t}\right)$. First, from the derivations above, we can express the premium payments on the compact form:

$$P_f^t E_0^{Q^f}\left(1_{\tau>t}\right) = \begin{cases} P_d^t L_0 & \text{if } t=1 \\ P_d^t L_0 L^{t-1} & \text{if } t \geq 2 \end{cases}$$

where $L_0 = F\left(1-\delta\lambda_0\right)$ and L and K are defined above. In order to compute $E_0^{Q^f}\left(1_{\tau=t}\right)$ for $t \geq 2$, in terms of domestic currency, we express it in terms of differences between defaultable zero-coupon bonds in foreign currency:

$$\begin{aligned} P_f^t E_0^{Q^f}\left(1_{\tau=t}\right) &= P_d^t E_0^{Q}\left(\frac{X_t}{X_0} 1_{\tau=t}\right) \\ &= P_d^t E_0^{Q}\left(\frac{X_t}{X_0} 1_{\tau>t-1}\right) - P_d^t E_0^{Q}\left(\frac{X_t}{X_0} 1_{\tau>t}\right) \\ &= P_d^t L_0 \big(F \cdot L^{t-2} - L^{t-1}\big) = P_d^t L_0 L^{t-2}\big(F-L\big) \end{aligned}$$

Above, we have used $E_0^Q\left(\frac{X_t}{X_0} 1_{\tau>t-1}\right) = F E_0^Q\left(\frac{X_{t-1}}{X_0} 1_{\tau>t-1}\right)$. Thus, we can express the protection leg payments on the following compact form:

$$P_f^t E_0^{Q^f}\left(1_{\tau=t}\right) = \begin{cases} P_d^t \delta\lambda_0 F & \text{if } t=1 \\ P_d^t L_0 L^{t-2}\left(F-L\right) & \text{if } t \geq 2 \end{cases} \tag{1.49}$$

We then obtain the expression for the foreign CDS premium in (1.46) by plugging in the compact form expressions for the premium and protection leg payments, and make use of the expression for a geometric series:

$$\begin{aligned} S^f(0,t) &= (1-R)\frac{\sum_{t=2}^T P_d^t L_0 L^{t-2}\left(F-L\right) + P_d \delta\lambda_0 F}{\sum_{t=1}^T P_d^t L_0 L^{t-1}} \\ &= (1-R)\frac{P_d^2\left(F-L\right)L_0 \sum_{t=2}^T (P_d L)^{t-2} + P_d\left(F-L_0\right)}{L_0 P_d \sum_{t=1}^T (P_d L)^{t-1}} \\ &= (1-R)\frac{P_d^2\left(F-L\right)L_0 \sum_{t=2}^T (P_d L)^{t-2} + P_d\left(F-L_0\right)}{L_0 P_d \sum_{t=1}^T (P_d L)^{t-1}} \\ &= (1-R)\frac{P_d^2\left(F-L\right)L_0 \frac{1-(LP_d)^{T-1}}{1-LP_d} + P_d(F-L_0)}{P_d L_0 \frac{1-(LP_d)^T}{1-LP_d}} \end{aligned}$$

$\square$

Proof of Proposition 1 and 2

The domestic CDS premium is unaffected by changes in the severity in foreign currency at default, δ, hence all we need to show is that the foreign CDS premium in (1.46) is **increasing** in δ such that the quanto spread, $QS(0,T) = S^d(0,T) - S^f(0,T)$, is **decreasing** in δ.

Evidently both L_0 and L are decreasing functions in δ (holding any other parameters fixed), so if we can show that the CDS premium is decreasing in L_0 and L, we are done. First, we split the CDS premium up in two expressions:

$$
\begin{aligned}
S^f(0,T) &= (1-R)\frac{P_d^2\,(F-L)\,L_0\frac{1-(LP_d)^{T-1}}{1-LP_d} + P_d(F-L_0)}{L_0 P_d \frac{1-(LP_d)^T}{1-LP_d}} \\
&= \frac{(1-R)}{P_d}\left(\underbrace{(F-L)\,\frac{P_d^2\left(1-(LP_d)^{T-1}\right)}{1-(LP_d)^T}}_{A} + \underbrace{\frac{P_d\left(\frac{F}{L_0}-1\right)}{\frac{1-(LP_d)^T}{1-LP_d}}}_{B}\right)
\end{aligned}
$$

Next, we show that both A and B are decreasing in δ. Consider the expression A. Since the riskless bond is assumed to be more expensive than a risky bond, we have $F - L > 0$ and $1 - LP_d > 0$. This implies that A is decreasing in L if and only if $\frac{\left(1-(LP_d)^{T-1}\right)}{1-(LP_d)^T}$ is decreasing in L, since $F - L$ obviously is decreasing in L. We show that $\frac{\left(1-(LP_d)^{T-1}\right)}{1-(LP_d)^T}$ is indeed decreasing in L by defining the function:

$$
f(m) = \frac{\left(1-(mP_d)^{T-1}\right)}{1-(mP_d)^T}
$$

Differentiating f with respect to m yields

$$
f'(m) = -\frac{(P_d m)^t\left((P_d m)^t - tP_d m + t - 1\right)}{P_d m^2\left((P_d m)^t - 1\right)^2}
$$

From this expression, we see that f' is negative if and only if $(P_d m)^t - tP_d m + t - 1$ is positive, which is indeed the case, since this function is strictly convex with a minimum of 0 at $m = \frac{1}{P_d}$. Hence, we showed that the expression A is decreasing in M and hereby in δ

as well.

An analogue argument can be used to show that $\left(\frac{1-(LP_d)^T}{1-LP_d}\right)^{-1} > 0$ is decreasing in L and hence in δ. Likewise is $\frac{F}{L_0} - 1 > 0$ and decreasing in L_0 and therefore in δ. Hence, the expression B is decreasing in δ as a product of two positive monotonically decreasing functions in δ.

The proof for Proposition 2 is conducted in an analogous manner to the proof of Proposition 1. In Proposition 1, we show that the quanto spread is decreasing in L, and since $L = F(1-\delta\lambda) - K\rho\,(u - u^{-1})\,\delta\left(\lambda^U - \lambda^D\right)$ is decreasing in ρ, then the foreign CDS premium increases in ρ. Evidently from the expression of L, L is increasing in $\lambda^U - \lambda^D$ if $\rho < 0$. The foreign CDS premium is therefore decreasing (increasing) in $\lambda^U - \lambda^D$ when $\rho < 0$ ($\rho > 0$).

Derivation of the Expressions (1.9)-(1.10)

First, the expression (1.9) follows immediately from (1.45) with $\lambda^U = \lambda^D = \lambda$ and $\bar{\lambda} = \lambda$, where λ is the fixed probability of default. Hence it follows that for any maturity T:

$$S^d(0,T) = (1-R)\frac{\lambda}{1-\lambda}$$

In order to derive the foreign-denominated CDS premium in the presence of crash risk and fixed default risk, we first notice that $L_0 = L = (1-\delta\lambda)\,F$. Inserting this into (1.46) gives

$$S^f(0,T) = (1-R)\frac{P_d^2 F\left(1-\left(1-\delta\lambda\right)\right)F\left(1-\delta\lambda\right)\frac{1-\left(F(1-\delta\lambda)P_d\right)^{T-1}}{1-\left(F(1-\delta\lambda)P_d\right)} + P_d\left(F - F(1-\delta\lambda)\right)}{F\left(1-\delta\lambda\right)\frac{1-\left(F(1-\delta\lambda)P_d\right)^{T-1}}{1-\left(F(1-\delta\lambda)P_d\right)}}$$

$$= (1-R)\frac{P_f^2\delta\lambda\left(1-\delta\lambda\right)\left(\frac{1-\left((1-\delta\lambda)P_f\right)^{T-1}}{1-(1-\delta\lambda)P_f}\right) + P_f\delta\lambda}{\left(1-\delta\lambda\right)P_f\left(\frac{\left((1-\delta\lambda)P_f\right)^{T}}{1-(1-\delta\lambda)P_f}\right)} = (1-R)\frac{\delta\lambda}{1-\delta\lambda}$$

The last equal sign follows from repeating the exact same calculations that led us to equation (1.45), with P_d replaced with P_f and $\bar{\lambda}$ replaced with $\delta\lambda$.

1.12 Appendix: Affine Model

1.12.1 Market Price of Risk

In this section we provide a proposition which help us specify the pricing kernel between the data-generating measure, the domestic measure, and the foreign measure. Cheridito et al. (2007) show that if an affine diffusion exists under a measure M^0, and it does not hit the boundary of the state space, then there also exists an affine diffusion under a measure M^1 which does not hit the boundary of the state space. More formally, they show (in the case of affine models without jumps) that if the drift and diffusion functions under M^0 and M^1 both satisfy the boundary non-attainment condition and the existence condition[9], then a true martingale exists defining the measure change from M^0 to M^1.

Lemma 1. *Assume that* $\left(\mu^{M^0}, \sigma\right)$ *and* $\left(\mu^{M^1}, \sigma\right)$ *satisfy the boundary non-attainment condition and the existence condition. Define the Radon-Nikodym derivative from* M^0 *to* M^1*:*

$$L_t = -L_{t-}\gamma_t dW_t^{M^0} + L_{t-}\sum_{i=1}^{K}(dZ_{i,t}^{M^0} + \lambda_{i,t}^{M^0}\zeta_i dt)$$

Where $dZ_{i,t}^{M^0}$ *is a pure jump process with intensity* λ_i *and the jump size distribution with mean jump size* ζ_i*. The jump times for* Z_i *are serially and cross-sectionally independent. Define for* $j = 0, 1$*:*

$$\mu^{M^j} : D \to \mathbb{R}^n,\ \mu^{M^j}(y) = a^{M^j} + b^{M^j}y,\ \sigma : D \to \mathbb{R}^{n\times n},\ \sigma(y)\sigma^T(y) = a_{ij} + b_{ij}y$$
$$L^{M^j} : D \to \mathbb{R}^n,\ L^{M^j}(y) = l_0^{M^j} + l_1^{M^j}y$$

Then the following three statements hold:

[9] The existence criterium is a necessary restriction on μ, σ, λ and D in order for an SDE to have a solution. Essentially the matrix $\sigma(Y_t)\sigma^T(Y_t)$ has to be positive definite on the interior of the state space and positive semi-definite on the closure of state space. In order for the latter to be fulfilled the drift term has to be positive on the closure of D and $\sigma(Y_t)\sigma^T(Y_t)$ has to approach the 0-matrix. These two requirements make sure that $\sigma(Y_t)\sigma^T(Y_t)$ is positive definite on D and does not fail to be positive semi-definite on the closure of D. The boundary non attainment condition makes sure that the volatility for each coordinate in Y_t remains strictly positive. For a detailed discussion of the existence of a solution to SDEs, see Duffie and Kan (1996) and Cheridito et al. (2007).

1. There exists a stochastic process Y_t that solves the SDE:

$$Y_t = Y_0 + \mu^{M^0}(Y_t)dt + \sigma(Y_t)dW_t^{M^0}$$

2. There exists a measure M^1 equivalent to M^0 such that:

$$Y_t = Y_0 + \mu^{M^1}(Y_t)dt + \sigma(Y_t)dW_t^{M^1}$$

3. The jump intensities and drifts under M^0 and M^1 are related as:

$$\lambda_i^{M^1}(Y_t) = (1+\zeta_i)\lambda_i^{M^0}(Y_t), \quad \mu^{M^1}(Y_t) = \mu^{M^0}(Y_t) - \sigma(Y_t)\gamma_t$$

Proof. Cheridito et al. (2007) show that continuous process: $dL_t^C = -\gamma_t dW_t^{M^0}$ is indeed a true martingale with $E_t^{M^0}(L_T^C) = 1$, provided that the existence and boundary non-attainment condition holds under both M^0 and M^1. The compensated jump process $Z_{i,t} + \lambda_{i,t}\zeta_{i,t}$ is also a true $\mathbb{M}^0$- martingale, since the mean jump size for each $Z_{i,t}$ is bounded and only exhibits a finite number of jumps. Hence the process $\sum_{i=1}^{K}(dZ_{t,i}^{M^0} + \lambda_{i,t}^{M^0}\zeta_i dt)$ is a true M^0-martingale since it is a finite sum of martingales. The process L_t is therefore a true martingale and hence 1.-3. follows from Girsanov's theorem for jump processes. □

1.12.2 Pricing of CDS in Affine Framework

Pricing of Domestic CDS

All the state-variables that are used to price the domestic CDS premium are independent. This makes the expressions for the ordinary differential much more simple, since the variance-covariance structure of the state-variables is a diagonal matrix. Therefore, we can represent the system of ordinary differential equations used for computing (1.17)-(1.18) for the domestic denominated CDS as

$$\frac{\partial\beta(t,T)}{\partial t} = \omega - K_1^T\beta(t,T) - \frac{1}{2}H\beta(t,T)\circ\beta(t,T), \qquad \frac{\alpha(t,T)}{\partial t} = -K_0^T\beta(t,T) \tag{1.50}$$

$$\frac{\partial B(t,T)}{\partial t} = -K_1^T B(t,T) - \frac{1}{2}H\beta(t,T)\circ B(t,T), \qquad \frac{A(t,T)}{\partial t} = -K_0^T B(t,T) \tag{1.51}$$

Where $\circ$ is the Hadamard product, and

$$\omega = \begin{bmatrix} 1 \\ 1 \\ 0 \end{bmatrix}, \ K_0 = \begin{bmatrix} \kappa_l \theta_l \\ 0 \\ \kappa_m \theta_m \end{bmatrix}, \ K_1 = \begin{bmatrix} -\kappa_l & 0 & 0 \\ 0 & -\kappa_z & \kappa_z \\ 0 & 0 & -\kappa_m \end{bmatrix}, \ H = \begin{bmatrix} \sigma_l^2 & 0 & 0 \\ 0 & \sigma_z^2 & 0 \\ 0 & 0 & \sigma_m^2 \end{bmatrix}$$

The boundary conditions are $\alpha(T,T) = 0$, $\beta(T,T) = [0,0,0]$, $A(T,T) = 0$ and $B(T,T) = [1,1,0]$

Pricing of Foreign CDS

The foreign CDS premium is a bit more involved than the domestic CDS premiums but also fits into the affine framework. Define the vector $\beta_j(t,T) = [\beta_v(t,T), \beta_l(t,T), \beta_z(t,T), \beta_m(t,T)]$, where $\beta_j(t,T)$ corresponds to the beta for state variable j, then the ordinary differential equation for state variable j is given by:

$$\frac{\partial \beta_j(t,T)}{\partial t} = \omega - K_1^T \beta_j(t,T) - \frac{1}{2}\beta_j(t,T) H_j \beta_j(t,T), \quad \frac{\alpha_j(t,T)}{\partial t} = -K_0^T \beta_j(t,T) \tag{1.52}$$

$$\frac{\partial B_j(t,T)}{\partial t} = -K_1^T B_j(t,T) - \frac{1}{2}\beta_j(t,T) H_j B_j(t,T), \quad \frac{A_j(t,T)}{\partial t} = -K_0^T B_j(t,T) \tag{1.53}$$

where:

$$\omega = \begin{bmatrix} 0 \\ (1+\zeta) \\ (1+\zeta) \\ 0 \end{bmatrix}, \ K_0 = \begin{bmatrix} \kappa_v^f \theta_v^f \\ \kappa_l \theta_l \\ 0 \\ \kappa_m \theta_m \end{bmatrix}, \ K_1 = \begin{bmatrix} -\kappa_v^f & 0 & 0 & 0 \\ \frac{1}{2}\sigma_l \rho \left(\frac{\theta_l}{\theta_v}\right)^{\frac{1}{2}} & \frac{1}{2}\sigma_l \rho \left(\frac{\theta_v}{\theta_l}\right)^{\frac{1}{2}} - \kappa_l & 0 & 0 \\ 0 & 0 & -\kappa_z & \kappa_z \\ 0 & 0 & 0 & -\kappa_m \end{bmatrix}$$

$$H_v = \begin{bmatrix} \sigma_v^2 & \frac{1}{2}\sigma_l \sigma_v \left(\frac{\theta_l}{\theta_v}\right)^{\frac{1}{2}} & 0 & 0 \\ \frac{1}{2}\sigma_l \sigma_v \left(\frac{\theta_l}{\theta_v}\right)^{\frac{1}{2}} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}, \ H_l = \begin{bmatrix} 0 & \frac{1}{2}\sigma_l \sigma_v \left(\frac{\theta_v}{\theta_l}\right)^{\frac{1}{2}} & 0 & 0 \\ \frac{1}{2}\sigma_l \sigma_v \left(\frac{\theta_v}{\theta_l}\right)^{\frac{1}{2}} & \sigma_l^2 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}$$

$$H_z = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & \sigma_z^2 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}, \ H_m = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & \sigma_m^2 \end{bmatrix}$$

The boundary conditions are $\alpha(T,T) = 0$, $\beta(T,T) = [0,0,0,0]$, $A(T,T) = 0$, and $B(T,T) = [0,(1+\zeta),(1+\zeta),0,0]$

1.13 Appendix: Estimation Approach

We estimate the model in two steps. In the first step, we apply maximum likelihood estimation (MLE) in conjunction with the unscented Kalman filter to infer a time-series of the instantaneous currency volatility process v_t and estimates of its risk-neutral and objective parameters ($[\kappa_v, \theta_v, \sigma_v, \kappa_v^P, \theta_v^P]$). We refer to section 1.13.1 for details on the Unscented Kalman filter and why we use this estimation approach. In this step, we only have one state variable, and the measurements consist of currency implied volatilities. We use a stochastic volatility model a la Heston (1993) as the currency options model, i.e, we assume that instantaneous currency volatility dynamics are unaffected by the jump components in the exchange rate arising from sovereign defaults specified in (1.15). Importantly, this does not mean that we ignore the correlation between sovereign credit and currency risk or the jump risk when pricing the sovereign CDS contracts, which is the focus of the analysis.

For pricing the currency options and CDS premiums, the discount factors in Euro and U.S. dollar are needed, which we bootstrap from their respective overnight index swap rates. The model-implied option prices are derived using the Fast Fourier Transform of Carr and Madan (1999) which we then transform into implied volatilities using the Garman and Kohlhagen (1983) formula such that they are comparable to the observables. We use implied volatilities rather than option prices since these are more stable than option prices along the moneyness and maturity dimension (see e.g., Schwartz and Trolle (2009)). Denoting x_t the time t state variable vector, then the measurement equation in the Kalman filter is given by

$$y_t = h(x_t) + e_t \tag{1.54}$$

where y_t is the vector of observables, $h(x_t)$ is the pricing function at state x_t, and e_t is the vector of measurement errors. In this particular case: $x_t = v_t$, y_t is the vector of observed implied volatilities, $h(x_t)$ is the vector of corresponding Heston (1993) implied volatilities, and e_t is a vector of IID Gaussian measurement errors with covariance matrix R. To reduce the number of parameters, we make the common assumption that the measurement errors are cross-sectionally uncorrelated (i.e., R is a diagonal matrix), and furthermore, we assume that the standard deviations of the measurement errors are identical for all options, σ_O.

We approximate the distribution of v_t with a Gaussian distribution such that the moments of the Gaussian distribution match the first two moments of v_t. All moments are computed by means of an Euler discretization, and we then cast the model into state space form

$$x_t = A + \phi x_{t-1} + \sqrt{Q_{t-1}}\varepsilon_t, \qquad \varepsilon_t \sim \mathrm{N(0,I)} \tag{1.55}$$

where in this particular case

$$A = \kappa_v^P \theta_v^P \cdot dt, \qquad \phi = e^{-\kappa^P dt}, \qquad Q_t = \sigma_v^2 v_t \cdot dt \tag{1.56}$$

Through the UKF iterations, we obtain $t-1$ predictions of the observables at time t, $\bar{y}_t$, and the corresponding prediction error covariance matrix $\bar{\Sigma}_{yy,t}$. With those at hand, we can then express the log-likelihood function using the prediction error decomposition

$$l(\Theta) = \sum_{t=1}^{N} -\frac{1}{2}\log|\bar{\Sigma}_{yy,t}| - \frac{1}{2}(y_t - \bar{y}_t)^T \bar{\Sigma}_{yy,t}^{-1}(y_t - \bar{y}_t) \tag{1.57}$$

where N is the number of observations, using weekly sampling we have $N = 281$ observations. We then find the maximum likelihood estimate of the parameters by maximizing (1.57).

In the second step, we estimate the parameters of the default intensities for one sovereign at the time using CDS premiums denominated in EUR and USD, now treating v_t as observable and its parameters as given. In this step, we use MLE in conjunction with the UKF to filter out the default intensity state variables, $[l_t, z_t, m_t]$, and to estimate their objective and risk-neutral parameters.

The measurements are the CDS premiums denominated in USD and the quanto CDS spread. In the pricing model, the USD contract is taken to be the domestic CDS contract, and the EUR contract is considered to be the foreign-denominated CDS contract. Their respective model-implied CDS premiums are henceforth derived according to (1.20), with the relevant transforms reported in Appendix, equations (1.51) and (1.53), respectively.

We assume that the measurement errors are the same for all maturities and for each type of contract, and we denote them σ_U and σ_{UE} for the USD-denominated CDS and the quanto CDS spread, respectively. The state space form of the discretized state variable dynamics, i.e., equation (1.55), is represented by the transition matrices

$$A = \begin{bmatrix} \kappa_l^P \theta_l^P \\ 0 \\ \kappa_m^P \theta_m^P \end{bmatrix} dt, \quad \phi = \begin{bmatrix} e^{-\kappa_l^P dt} & 0 & 0 \\ 0 & e^{-\kappa_z^P dt} & -\kappa_z^P dt \\ 0 & 0 & e^{-\kappa_m^P dt} \end{bmatrix}, \quad Q_t = \begin{bmatrix} \sigma_l^2 l_t & \sigma_v \sigma_l \sqrt{l_t v_t} & 0 \\ \sigma_v \sigma_l \sqrt{l_t v_t} & \sigma_z^2 z_t & 0 \\ 0 & 0 & \sigma_m^2 m_t \end{bmatrix} dt \tag{1.58}$$

With the model represented on state space form, we can then compute the maximum likelihood estimates by maximizing the log-likelihood function in (1.57).

Various specifications of the model above have been implemented, and our estimations reveal that it is important that the model allows for a drift adjustment for currency/default covariance risk which depends on the level of the default intensity. For instance, we implemented a simple affine model capturing default/currency covariance risk in which the systematic default intensity is a fixed fraction of the currency volatility: $\beta_i v_t$. This model has a closed form solution for the foreign CDS premium, without using any approximations. The problem with this specification, however, is that is not well-suited for handling differences in time trends in the credit spreads and the currency volatility. In the sample, the EURUSD currency volatility is persistent and exhibits strong mean-reversion, while sovereign eurozone default risk unambiguously trends downward during the latter period of the sample period.

1.13.1 The Unscented Kalman Filter

In the standard Kalman filter both the state vector equation and the measurement equation are linear in the state variables and both have Gaussian noise. To be specific, the (Gaussian) state space representation of such a system is:

$$x_t = A + \phi x_{t-1} + \sqrt{Q_{t-1}}\varepsilon_t, \qquad \varepsilon_t \sim \mathrm{N(0,I)} \tag{1.59}$$

$$y_t = Hx_t + e_t, \qquad e_t \sim \mathrm{N(0,R)} \tag{1.60}$$

x_t is the state vector and y_t are the measurements (in our case CDS premiums and option implied volatilities). We denote the forecasts at time $t-1$ of the state variables at time t and their covariance matrix as $\bar{x}_t$ and $\bar{\Sigma}_{xx,t}$, and $\hat{x}_t$ and $\hat{\Sigma}_{xx,t}$ are their updates at time t (updated based on new information inherit in y_t). $\bar{y}_t$ and $\bar{\Sigma}_{yy,t}$ represent the $t-1$ model forecast errors of the measurements at time t and their covariance matrix. The forecasts of the state variables and their covariance matrix are given by

$$\bar{x}_t = A + \phi\hat{x}_{t-1},\ \bar{\Sigma}_{xx,t} = \phi\hat{\Sigma}_{xx,t-1}\phi^T + Q_{t-1} \tag{1.61}$$

and the forecasts of the measurements and their covariance matrix, and their covariance with the state variables are given by:

$$\bar{y}_t = H\bar{x}_t,\ \bar{\Sigma}_{yy,t} = H\bar{\Sigma}_{xx,t}H^T + R,\ \bar{\Sigma}_{xy,t} = \bar{\Sigma}_{xx,t}H^T \tag{1.62}$$

The updated state variables and their covariance are calculated as

$$\hat{x}_t = \bar{x}_t + K_t(y_t - \bar{y}_t),\ \hat{\Sigma}_{xx,t} = \bar{\Sigma}_{xx,t} - K_t\bar{\Sigma}_{yy,t}K_t^T \tag{1.63}$$

where $K_t = \bar{\Sigma}_{xy,t}\bar{\Sigma}_{yy,t}^{-1}$. Given the (exponential) affine structure of the dynamics of the state vector, we can represent the discretized dynamics of the state variables as in (1.64) below with system matrices as specified in (1.58). Since neither the CDS premiums or options are linear in the state variables, the measurement equation (1.65) is governed by a non-linear function h:

$$x_t = A + \phi x_{t-1} + \sqrt{Q_{t-1}}\varepsilon_t, \qquad \varepsilon_t \sim \mathrm{N(0,I)} \tag{1.64}$$

$$y_t = h(x_t) + e_t, \qquad e_t \sim \mathrm{N(0,R)} \tag{1.65}$$

The UKF is one method for handling this non-linearity. In the UKF, the mean and covariance matrix of the forecasts of the measurement series and its covariance with the state variables are derived using a set of deterministic sampling points denoted sigma points, $\mathcal{X}_{t,i}$. The sigma points are chosen such that their mean and covariance match $\bar{x}_t$ and $\bar{\Sigma}_{xx,t}$, respectively. Based on the sigma points, new measurements, $\mathcal{Y}_{t,i}$, are generated $h(\mathcal{X}_{t,i}) = \mathcal{Y}_{t,i}$. From $\mathcal{Y}_{t,i}$, we then estimate the moments of the forecasts of the measurements as:

$$\bar{y}_t = \sum_{i=0}^{2p} w_i \mathcal{Y}_{t,i},\ \bar{\Sigma}_{yy,t} = \sum_{i=0}^{2p} w_i \left[\mathcal{Y}_{t,i} - \bar{y}_t\right]\left[\mathcal{Y}_{t,i} - \bar{y}_t\right]^T + R,\ \bar{\Sigma}_{xy,t} = \sum_{i=0}^{2p} w_i[\mathcal{X}_{t,i} - \bar{x}_t]\left[\mathcal{Y}_{t,i} - \bar{y}_t\right]^T$$

where the sigma points and the weights are defined as

$$\begin{aligned}
\mathcal{X}_{t,0} &= \bar{x}_t, \quad \mathcal{X}_{t,i} = \bar{x}_t \pm \sqrt{(p+\delta)\left(\Sigma_{xx,t}\right)_j} \;\; j = 1, \cdots, p,\ i = 1, \cdots, 2p \\
w_0 &= \frac{\delta}{p+\delta}, \quad w_i = \frac{1}{2(p+\delta)}, \;\; j = 1, \cdots, 2p
\end{aligned}$$

where p is the dimension of the state vector and $\delta > 0$. We then use the Kalman filter as described above to obtain forecasts and updates of the state variables. Assuming normality of the forecast errors, we can use the forecast error decomposition of the log-likelihood function for the sample:

$$l(\Theta) = \sum_{t=1}^{N} -\frac{1}{2}\log|\bar{\Sigma}_{yy,t}| - \frac{1}{2}(y_t - \bar{y}_t)^T \bar{\Sigma}_{yy,t}^{-1}(y_t - \bar{y}_t)$$

Essay 2

Forward-Looking Currency Betas

Forward-Looking Currency Betas

Andreas Bang Nielsen*

Abstract

I propose a model-free method to derive forward-looking betas to currency portfolios from cross-pair currency options. Using the dollar factor—an equal-weighted basket of all foreign currencies against the U.S. dollar—as the systematic factor, I find that these option-implied betas are significantly better predictors of realized betas and currency excess returns compared to traditional rolling window betas. Constructing portfolios based on option-implied betas leads to a significantly positive relation between ex-ante betas and ex-post portfolio returns, whereas there is an insignificant relation when rolling window betas are used.

Keywords: Exchange rate risk premiums, factor models, currency options, option-implied betas, foreign exchange volatility

JEL Codes: G12, G15, F31

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2.1 Introduction

This paper proposes a novel method to compute currency factor exposures (betas) that are purely forward-looking and adjust immediately to new information. The currency market provides a unique opportunity to calculate forward-looking betas because covariances can be retrieved from cross-pair currency option prices without assuming any parametric structure on variances and correlations.

In particular, the covariance between any pair of currencies against, say, the U.S. dollar, can be expressed in terms of their U.S. dollar variances and their cross-pair variance. In this manner, the exchange rate covariance structure can be constructed from option-based variances, and as a result, forward-looking betas of currency portfolios can be calculated. Forward-looking betas are unique to currencies because covariances in other major asset classes, such as stocks, cannot be derived from options, since there is no (liquid) market for options for which the payoff depends on the price evolution of two securities.

Factor models have most commonly been used for stocks and bonds, but a growing literature, pioneered by Lustig, Roussanov, and Verdelhan (2011), has emerged, which explains currency risk premiums as exposures to factors built from currencies. In this literature, betas are estimated by means of rolling window regressions of realized currency returns on the realized systematic factors. The betas estimated using this method, however, suffer from a number of caveats. They are backward-looking, adjust slowly to new information, and the econometrician has to decide on which particular subset of the data to use for the estimation. In contrast, since the option-implied betas are inferred from the latest cross-section of option prices, they require neither historical data nor choices of estimation window and frequency.

In order to compare the empirical properties of the option-implied betas to the rolling window betas, I use the dollar factor of Lustig, Roussanov, and Verdelhan (2011)—which is an equally weighted portfolio of all foreign currencies taking the perspective as a U.S. investor—as the systematic factor in currency excess returns. However, the methodology that I propose can be used for any currency risk factor. The excess return on the dollar factor is the excess return a U.S. investor receives from borrowing money at home and investing in all (developed) foreign currencies equally weighted, and it carries a significant risk premium and explains a large share of the time-series variation in exchange rates (Lustig, Roussanov, and Verdelhan, 2011; Verdelhan, 2017). Lustig, Roussanov, and Verdelhan (2014) show that the dollar factor tends to appreciate (depreciate) whenever the average short-term foreign interest rates is above (below) the short-term U.S. interest rate. As a result, a conditional dollar factor, which is long the dollar factor whenever the average forward discount (U.S. minus average foreign interest rates) is negative and short otherwise, has collected a larger excess return than the (unconditional) dollar factor (Lustig, Roussanov, and Verdelhan, 2014; Verdelhan, 2017).

Conditional on the average foreign discount, I find a significantly positive relation between ex-ante option-implied dollar factor betas and ex-post portfolio excess returns, while there is an insignificant relation when using rolling window betas. Interestingly, the option-implied betas are strong predictors of portfolio excess returns, because they predict spot exchange rate changes of the portfolios, while rolling window betas exhibit no predictability of spot exchange rate changes. I provide evidence that this is because option-implied betas are more powerful and less biased predictors of realized betas than rolling window betas.

Specifically, conditional on the average foreign discount, I sort currencies into portfolios on the basis of dollar factor betas, for each type of beta separately, and construct an HML dollar factor which dynamically buys high-beta and shorts low-beta currencies. At a 1-month holding period, when using the option-implied betas, the HML dollar factor has a significant mean annualized excess return of 3.35 percent (Sharpe ratio of 0.41), where 2.35% stems from the spot change component. On the other hand, constructing the HML dollar factor based on 252-day rolling window betas leads to an insignificant mean excess return of 0.95 percent (Sharpe ratio of 0.11), with a spot change component of -0.18%. Thus, the difference in the HML dollar excess returns is entirely due to the fact that option-implied betas are stronger predictors of currency spot changes.

The results for the beta-sorted portfolios do not necessarily imply that the option-implied betas are more accurate forecasters of realized currency returns. Betas could be inaccurately measured that would cause large model prediction errors in the time series and still properly rank currencies on betas. The option-implied betas, however, are not only better at ranking currencies on betas, they are also significantly better predictors of portfolio returns in the time series, with smaller mean squared model prediction errors across all portfolios and forecast horizons. Furthermore, using the rolling window betas to forecast portfolio returns delivers biased predictions, whereas the option-implied beta predictions are virtually unbiased. The expected low-beta portfolios, based on rolling window betas, tend to exhibit larger realized returns than the expectation, and vice versa, the expected high-beta portfolios exhibit lower returns than expected.

I show that this superior model performance when using option-implied betas is because they are stronger and less biased predictors of realized betas. At any forecast horizon, for both portfolios and individual currencies, the option-implied betas provide significantly smaller prediction errors than the rolling window betas. Furthermore, consistent with the prediction bias for portfolio excess returns, when using rolling window betas, the expected beta of (high) low-beta portfolios tends to be (smaller) larger ex-post than the expectation, while the option-implied betas deliver virtually unbiased predictions.

2.2 Related Literature

There are, to the best of my knowledge, no papers that have studied option-implied betas in currencies, while there are several papers that use options to estimate betas in the equity literature, for example: French, Groth, and Kolari (1983); Siegel (1995); and more recently, Buss and Vilkov (2012); Chang, Christoffersen, Jacobs, and Vainberg (2011); Christoffersen, Fournier, and Jacobs (2017). Since there is not (yet) a liquid market for options that depend on the price evolution of two stocks, stock correlations cannot be implied out from options without assumptions. French, Groth, and Kolari (1983) suggest computing betas using a mixture of option-implied volatilities and correlations estimated from historical data. Chang, Christoffersen, Jacobs, and Vainberg (2011) compute purely forward-looking betas under the assumption that stock returns follow a linear factor model where idiosyncratic shocks have no skew, and Buss and Vilkov (2012) derive option-implied betas by parametrically linking risk-neutral and objective correlations (estimated from past returns). I add to this literature by computing betas in currency markets which only use option market information and require no distributional assumptions.

This paper is related to the literature that uses option-based information to predict realized returns and moments. Busch, Christensen, and Nielsen (2011) find that implied volatilities are stronger predictors of realized volatilities than historical volatility estimates in fixed income, equity, and foreign exchange. Jorion (1995) finds consistent results in currency markets for a different sample period. Buss and Vilkov (2012) show that option-implied betas are significantly better at explaining the cross-section of stock returns and predicting realized CAPM betas compared to rolling window betas. I contribute to this literature by showing that purely option-implied betas in currency markets are strong predictors of future realized currency returns and betas.

The option-implied betas proposed in this paper can be used to estimate risk exposures in any currency factor model, for instance, in the models suggested by Lustig, Roussanov, and Verdelhan (2011), Lustig, Roussanov, and Verdelhan (2014), and Verdelhan (2017), who focus on the carry factor (long high interest rate currencies and short low interest rate currencies), and the dollar factor (long equal-weighted basket of all foreign currencies). Other notable currency factors include the momentum factor of Menkhoff, Sarno, Schmeling, and Schrimpf (2012b) and Asness, Moskowitz, and Pedersen (2013), the global volatility factor of Menkhoff, Sarno, Schmeling, and Schrimpf (2012a), and the international correlation dispersion factor of Mueller, Stathopoulos, and Vedolin (2017).

The work of Verdelhan (2017) is perhaps closest to this paper. He documents that time-varying exposure to the dollar factor is of key importance in explaining the cross-section of currency returns and the time-series variation in currencies. I contribute to this paper by documenting that for the G10 currencies, option-implied betas better explain the cross-section of currency returns and exhibit smaller time-series predictions errors for realized returns and betas than the historical rolling window betas.

More generally, this paper is related to the literature on time-varying currency risk premiums (Lustig, Roussanov, and Verdelhan, 2014; Mueller, Stathopoulos, and Vedolin, 2017; Sarno, Schneider, and Wagner, 2012). For instance, Lustig, Roussanov, and Verdelhan (2014) show that a static carry trade, which is long currencies with the highest average interest rates and short those with the lowest, only explains about one third of the returns to a dynamic carry trade, which is long-short based on time-varying betas to the carry factor. Therefore, a central theme in this literature is time-varying betas, which I show can be measured in real time using currency options for any given currency factor.

This paper is related to the relatively scarce literature which uses cross-pair currency options to study currency risk premiums. Two notable papers are Mueller, Stathopoulos, and Vedolin (2017) and Jurek and Xu (2014). Mueller, Stathopoulos, and Vedolin (2017) find that currency correlations are counter-cyclical and they construct a correlation dispersion measure that explains the cross-section of currency excess returns. More importantly for this paper, they show how to compute risk-neutral covariances between exchange rates by using model-free cross-pair variances derived from options using, e.g., using Britten-Jones and Neuberger (2000). However, rather than constructing a factor from the covariances between currencies, I use them to measure forward-looking risk exposures to currency portfolios.

Jurek and Xu (2014) estimate risk premiums using currency options in a latent factor model in which the common factor follows a sufficiently rich dynamic structure that captures the most salient features of currency returns. In contrast, I specify exactly what the systematic factor is and estimate risk premiums without imposing specific distributional assumptions on the common factor. One of the key strengths of the option-implied betas suggested in this paper is that no parameters have to be estimated, which makes them easy to implement and computationally efficient.

2.3 Option-Implied Risk Exposures

In the equity literature, there has been a long tradition of modeling expected returns to individual stocks and portfolios as their covariation with a set of systematic factors, but the popular factors used in the equity literature, e.g., the three factors of Fama and French (1993), have little explanatory power for currency returns (Burnside, Eichenbaum, and Rebelo, 2011). Likewise, macro-based models have failed to "beat" the random walk in predicting currency returns (Cheung, Chinn, and Pascual, 2005; Meese and Rogoff, 1983).

Recently, a new stream of literature has emerged which has found that the cross-section and time-variation of currency returns appear to be well-explained by exposure to portfolios of currencies. Arguably, the most notable currency factors are the carry and dollar factors introduced in Lustig, Roussanov, and Verdelhan (2011) [1].

A central element in this research is time-varying betas, especially for cross-sectional analysis, which critically relies on accurate measurements of betas. Common to this liter-

[1] Other notable examples of papers that use factor models to model currency returns: Menkhoff, Sarno, Schmeling, and Schrimpf, 2012a,b; Mueller, Stathopoulos, and Vedolin, 2017; Ready, Roussanov, and Ward, 2017.

ature is that the betas are estimated using rolling window regressions of currency returns on the proposed factors, which implicitly assumes that historical realizations reflect future outcomes. There are a number of caveats with historical betas. First, they do not adjust immediately to structural changes in the currency market conditions, for instance due to unforeseen changes in a country's monetary policy, that is, they are slow-moving. Second, the econometrician has to decide on a particular time frame and data frequency used for the estimation, both of which are subjective decisions.

The betas derived from currency option markets do not suffer from any of these issues; the option market provides the betas in real time, and only numerical implementation errors affect their measurements, which tend to be of minor impact (Della Corte, Ramadorai, and Sarno, 2016; Mueller, Stathopoulos, and Vedolin, 2017). Arguably, the forward-looking nature of the option-implied betas is especially valuable in periods in which future expectations of exchange rates deviate substantially from the past, as was the case, for example, during the financial crisis, the European debt crisis, or the Asian crisis.

2.3.1 Model Setup

Define the exchange rate S^{ji} as units of currency i per 1 unit of currency j, that is, an appreciation in the exchange rate corresponds to an increase in currency j relative to currency i. Moreover, define the log change over $[t, t+m]$ as $\Delta s^{ji}_{t,t+m} \equiv \log S^{ji}_{t+m} - \log S^{ji}_{t}$. I assume that the log currency dynamics is governed by a single-factor model:

$$\Delta s^{ji}_{t,t+m} = i^{i}_{t,t+m} - i^{j}_{t,t+m} + \beta^{ji}_{t,t+m} G_{t,t+m} + \varepsilon^{ji}_{t+m} \tag{2.1}$$

where $i^{i}_{t,t+m} - i^{j}_{t,t+m}$ is the interest rate differential between currency i and j over a horizon of length m, $G_{t,t+m}$ are shocks in the systematic factor, $\beta^{ji}_{t,t+m}$ measures the sensitivity of currency j to shocks in the systematic factor, and ε^{ji}_{t+m} is idiosyncratic risk (non-priced risk).

The conditional expected excess return for holding currency j is then given by

$$E_t\left(rx^{ji}_{t,t+m}\right) = \beta^{ji}_{t,t+m} \cdot \lambda^{G}_{t,t+m} \tag{2.2}$$

where $rx^{ji}_{t,t+m} = \Delta s^{ji}_{t,t+m} - \left(i^{i}_{t,t+m} - i^{j}_{t,t+m}\right)$, and $\lambda^{G}_{t,t+m}$ is the price of risk for the systematic factor. $rx^{ji}_{t,t+m}$ is the excess return from borrowing money in currency i and investing in currency j, measured in terms of currency i. The risk premium defined in (2.2) is based on a log approximation, which has been used in the majority of papers studying currency risk premiums, dating back to Bilson (1981) and Fama (1984). In the actual empirical implementation of currency excess returns, I use discrete returns, rather than log returns, but in general, the difference is miniscule and does not alter the main conclusions of the paper.

The model is able to capture time-varying risk premiums through time-dependent risk exposures (and prices of risk), which has been documented in the literature as an important salient feature of exchange rates, e.g., in order to match the failure of the uncovered interest rate parity (Sarno, Schneider, and Wagner, 2012).

Formally, a factor structure in the log exchange rates can be constructed in an international complete market model by imposing a factor structure in the law of motion of each country's log pricing kernel. This is because a standard no-arbitrage argument shows that the difference in log changes of each country's pricing kernels governs the law of motion of their bilateral log exchange rate. In this manner, heterogeneity in exposures to shocks in the systematic factors drives the cross-section of currency excess returns. This modeling approach has been taken by several papers (e.g., Lustig, Roussanov, and Verdelhan, 2014; Menkhoff, Sarno, Schmeling, and Schrimpf, 2012b; Mueller, Stathopoulos, and Vedolin, 2017; Verdelhan, 2017). However, the approach that I use does not necessarily assume complete markets but only that there is a factor structure in currency excess returns.

2.3.2 Option-Implied Currency Betas

In the following, I will suppress the base currency index whenever it is clear from the context that the method applies to any base currency. The time t conditional beta for currency j, $\beta^j_{t,t+m}$ over $[t, t+m]$ in the model (2.1), is given by:

$$\beta^j_{t,t+m} = \frac{Cov_t\left(\Delta s^j_{t,t+m}, G_{t,t+m}\right)}{V_t\left(G_{t,t+m}\right)} \tag{2.3}$$

If $G_{t,t+m}$ is the m-period innovation in a portfolio consisting of N currencies with weights w_k, that is, $G_{t,t+m} = \sum_{k=1}^{N} w_k \Delta s_{t,t+m}^{k}$ [2], then the beta in (2.3) can be expressed as

$$\beta_{t,t+m}^{j} = \frac{\sum_{k=1}^{N} w_k Cov_t \left(\Delta s_{t,t+m}^{j}, \Delta s_{t,t+m}^{k} \right)}{\sum_{k=1}^{N} \sum_{l=1}^{N} w_k w_l Cov_t \left(\Delta s_{t,t+m}^{k}, \Delta s_{t,t+m}^{l} \right)} \tag{2.4}$$

The numerator in (2.4) contains the covariances between currency j and all the constituents of the systematic factor portfolio. The denominator contains all the variances and covariances of the constituents of the factor portfolio. These moments can be computed using traditional rolling window estimates, or they can be implied out from options. In the next section, I show how to compute risk-neutral covariances and variances of exchange rates by means of currency options, and hence how to compute option-implied model-free betas.

The model-free measure of the covariance between two currencies, against a given base currency, can be constructed from their respective exchange rates versus the base currency and their cross-pair exchange rate. Denote the cross-pair exchange rate between two foreign currencies k and j, S^{kj}, and the respective base exchange rates S^{k} and S^{j}. Let the base currency be USD, then one unit of j equals S^{j} units of USD, which can be converted into $S^{j} \cdot (S^{k})^{-1}$ units of k. Hence, in the absence of triangular arbitrage, then $S^{j} \cdot (S^{k})^{-1} = (S^{kj})^{-1}$ [3]. Assuming the absence of triangular arbitrage at time t and $t+m$ and taking logs then gives

$$\Delta s_{t,t+m}^{kj} = \Delta s_{t,t+m}^{k} - \Delta s_{t,t+m}^{j} \tag{2.5}$$

Taking risk-neutral variance on both sides of (2.5), and rearranging, gives:

$$Cov_t^{Q} \left(\Delta s_{t,t+m}^{k}, \Delta s_{t,t+m}^{j} \right) = \frac{1}{2} \Bigg(V_t^{Q} \left(\Delta s_{t,t+m}^{k} \right) + V_t^{Q} \left(\Delta s_{t,t+m}^{j} \right) - V_t^{Q} \left(\Delta s_{t,t+m}^{kj} \right) \Bigg) \tag{2.6}$$

Equation (2.6) expresses the risk-neutral covariance between two exchange rates against the same base currency, say the U.S dollar, in terms of the risk-neutral variances of the

[2]Since interest rates over $[t, t+m]$ are known at time t, they do not impact variances and covariances, therefore we may equivalently think of G_{t+m} as the log excess return on the dollar factor.

[3]The data used in Mueller, Stathopoulos, and Vedolin (2017) show that triangular arbitrage spreads on average are below 1 basis point and last for less than a second (Fenn, Howison, McDonald, Williams, and Johnson (2009) report similar quantities). Consequently, the spreads from triangular arbitrage are so small that they do not significantly affect the option-implied moments.

respective exchange rates against the U.S. dollar and their risk-neutral cross-pair variance. By using the expression of Britten-Jones and Neuberger (2000), I compute each risk-neutral variances in (2.6) by integrating over a continuum of put and call prices:

$$V_t^Q(\Delta s_{t,t+m}) = 2e^{i_{t,t+m}} \left(\int_0^{S_t} \frac{1}{K^2} P(K,t,t+m)dK + \int_{S_t}^{\infty} \frac{1}{K^2} C(K,t,t+m)dK \right) \quad (2.7)$$

where $i_{t,t+m}$ is the riskless interest rate of the base currency over horizon m, and $P(K,t,t+m)$ and $C(K,t,t+m)$ are put and call prices, respectively, with maturity m and strike K. From the expressions (2.6)-(2.7), the entire covariance matrix for all exchange rates can be constructed, and hence option-implied betas to currency portfolios.

Deriving the model-free variance from expression (2.7) requires a continuum of put and call prices at different strikes. Options in currency markets are, in general, quoted in terms of Garman and Kohlhagen (1983) implied volatilities at five different strikes, spread evenly across moneyness (see section 2.4 for details on the options data). I interpolate between those available strike/implied volatility pairs using a cubic spline, as in Della Corte, Ramadorai, and Sarno (2016)[4] and use the Garman and Kohlhagen (1983) formula to convert each strike/volatility pair into put and call prices, which are then used to calculate the integral in (2.7).

Jiang and Tian (2005) point out that discretization errors arise from performing the numerical integration of the integral in (2.7), however, Mueller, Stathopoulos, and Vedolin (2017) report that the discretization errors do not exceed 0.5 percentage points of the implied volatilities in currency markets. In the literature, there are different variations on how to compute the model-free moments. As a robustness check, I derived the variances and covariances using the expressions of Martin (2017) and Bakshi, Kapadia, and Madan (2003) and the differences were negligible [5]. The historical moments are computed using daily log changes in the exchange rates. Specifically, the annualized realized variance of exchange

[4]Della Corte, Ramadorai, and Sarno (2016) analyze different interpolation schemes, including the no-arbitrage vanna-volga method of Castagna and Mercurio (2007), and find virtually no differences in the derived variances.

[5]Mueller, Stathopoulos, and Vedolin (2017) find a statistically insignificant difference between option implied currency correlations using the model-free variances of Martin (2017) and Britten-Jones and Neuberger (2000).

rate i and its covariance with currency k, using a window of length L, are calculated as

$$RV^{i}_{t,t-L} = \frac{252}{L}\sum_{j=0}^{L-1}\left(\Delta s^{i}_{t-j}\right)^{2} \tag{2.8}$$

$$RCOV^{ik}_{t,t-L} = \frac{252}{L}\sum_{j=0}^{L-1}\Delta s^{i}_{t-j}\Delta s^{k}_{t-j} \tag{2.9}$$

where Δs_t denote daily log changes in the exchange rate.

2.3.3 Dollar Factor Betas

Let the m-period innovation to the dollar factor, defined as an equal-weighted portfolio of foreign currencies against the U.S. dollar, be denoted:

$$\Delta Dol_{t,t+m} \equiv \frac{1}{N}\sum_{i=1}^{N}\Delta s^{i}_{t,t+m} \tag{2.10}$$

The time t variance over $[t, t+m]$ of the dollar factor and its covariance with exchange rate j under measure M, which may either be the objective measure P or the risk-neutral measure Q, are given by

$$V^{M}_{t}\left(\Delta Dol_{t,t+m}\right) = \frac{1}{N^2}\sum_{i=1}^{N}V^{M}_{t}\left(\Delta s^{i}_{t,t+m}\right) + \frac{1}{N^2}\sum_{i=1}^{N}\sum_{k\neq i}^{N}Cov^{M}_{t}(\Delta s^{i}_{t,t+m}, \Delta s^{k}_{t,t+m}) \tag{2.11}$$

$$Cov^{M}_{t}\left(\Delta s^{j}_{t,t+m}, \Delta Dol_{t,t+m}\right) = \frac{1}{N}\sum_{i=1}^{N}Cov^{M}_{t}\left(\Delta s^{j}_{t,t+m}, \Delta s^{i}_{t,t+m},\right) \tag{2.12}$$

Following Verdelhan (2017), I exclude the relevant currency from the dollar factor when computing betas to avoid a mechanical relation, i.e., the time t dollar beta for currency j under measure M is defined as

$$\beta^{M}_{jt,t+m} = \frac{Cov^{M}_{t}\left(\Delta s^{j}_{t,t+m}, \Delta Dol_{t,t+m|j}\right)}{V^{M}_{t}\left(\Delta Dol_{t,t+m|j}\right)} = \frac{\frac{1}{N-1}\sum_{i\in N|j}Cov^{M}_{t}\left(\Delta s^{j}_{t,t+m}, \Delta s^{i}_{t,t+m},\right)}{V^{M}_{t}\left(\Delta Dol_{t,t+m|j}\right)} \tag{2.13}$$

where $Dol_{t,t+m|j}$ denotes the dollar factor excluding currency j. The m-month risk-neutral dollar factor beta is then derived from (2.11)-(2.12), excluding currency j from the dollar factor, by plugging into (2.13). Variances/covariances are computed from the expressions

(2.6)-(2.7) with m-month maturity options. The historical dollar factor betas are computed similarly from the expression (2.13) by plugging in realized variances/covariances derived from (2.8)-(2.9), or equivalently from regressions of log currency changes onto changes in the dollar factor.

From a theoretical point of view, the moments used in the calculation of beta must be under the objective measure, rather than the risk-neutral measure; that is, the risk-neutral moments must be corrected for risk premiums for bearing variance/covariance risk. Even though this is indeed true, I take a pragmatic view: if there is empirical evidence for the usefulness of option-implied betas in forecasting realized currency returns and betas, it seems worthwhile using them for, e.g., constructing currency trading strategies and for risk assessment of portfolios.

In the equity literature a similar view has been taken, and there is a vast literature documenting that option-based information is useful for forecasting volatilities, betas, and returns, without correcting for risk premiums. For instance, there is evidence that option-implied volatilities are more powerful predictors of realized stock volatilities than volatility estimates based on historical data (e.g. Busch, Christensen, and Nielsen, 2011; Christensen and Hansen, 2002; Christensen and Prabhala, 1998). Even state of the art methodologies using high-frequency historical data underperform raw option-implied volatilities in forecasting volatility for currencies, stocks, and bonds (Busch, Christensen, and Nielsen, 2011). Buss and Vilkov (2012) find that option-based CAPM betas are better at identifying a monotonic relation between ex-ante betas and ex-post portfolio returns and exhibit less biased estimates of realized betas compared to traditional rolling window CAPM betas.

2.4 The Data

In this section, the datasets for spot and forward exchange rates as well as the options data are presented.

2.4.1 Currency Spot and Forward Data

I collect data on spot exchange rates and forward exchange rates for all G10 currencies (AUD, CAD, CHF, EUR, GBP, JPY, NOK, NZD, and SEK) against the U.S. dollar from

January 1996 to August 2016 at a daily frequency. The spot and forward contract data are obtained from Reuters via Datastream, and the maturities of the forward contracts are one, two, three, six, nine, and 12 months. Exchange rate j is denoted S_t^j and defined as units of U.S. dollar per unit of currency j, i.e., if the exchange rate rises then currency j appreciates against the U.S. dollar. The time t forward price of currency j at maturity m is defined analogously and denoted $F_{t,t+m}^j$.

From the spot and forward data, I construct the m-period excess return of a long position in currency j as:

$$RX_{t,t+m}^j = \frac{S_{t+m}^j - F_{t,t+m}^j}{S_t^j} \tag{2.14}$$

Thus, the excess return defined in (2.14) is the return of buying $1/S_t^j$ units of currency j forward and selling it at the prevailing spot rate at time $t+m$.

Akram, Rime, and Sarno (2008) show that the covered interest rate parity (CIP) holds closely at daily and lower frequencies, which implies that $F_{t,t+m}^j = S_t^j e^{i_{t,t+m} - i_{t,t+m}^j}$, where $i_{t,t+m}$ and $i_{t,t+m}^j$ are the m-month riskless interest rates in U.S. dollar and currency j, respectively. Hence, provided that the CIP holds, then

$$RX_{t,t+m}^j \approx \frac{S_{t+m}^j - S_t^j}{S_t^j} + i_{t,t+m}^j - i_{t,t+m} \tag{2.15}$$

So the currency excess return can be decomposed into a spot change component and an interest rate component. Since the financial crisis, however, a persistent cross-currency basis has emerged, reflecting that the CIP does not hold in the traditional sense. Rime and Syrstad (2016) attribute the cross-currency basis to market segmentation between the interbank money market in different currencies (e.g., the overnight and LIBOR markets), which is primarily confined to top-tier banks, and the market for currency funding for the typical arbitrageurs (e.g., the currency forward and cross-currency swap market).

Therefore, the currency excess returns considered in this paper ought to be considered from the perspective of an investor who implements currency trading strategies and examines currency risk premiums through the lens of the forward market, rather than through the interbank money markets. The expression in (2.15) is, however, still useful for decomposing excess currency returns into a spot and carry component.

2.4.2 Currency Options Data

The currency options data consist of quotes from J.P. Morgan collected daily from January 1998 to August 2016 for all G10 currencies, including cross-pairs, which gives rise to a total of 45 quoted currency pairs (9 pairs against the U.S. dollar and 36 cross-pairs).

The currency options are quoted in terms of Garman and Kohlhagen (1983) implied volatilities (IVs) of delta-neutral straddles, 10-delta and 25-delta risk-reversals and 10-delta and 25-delta butterflies. All options are traded at fixed maturities of 1 month, 2 months, 3 months, 6 months, 9 months, and 12 months.

The delta-neutral straddle is defined as a long position in a put and call option (with the same strike) such that the straddle has a delta of 0—often referred to as the at-the money (ATM) straddle. The risk reversal consists of a long position in an out-of-the money (OTM) call and a short position in an OTM put with equal absolute deltas. The butterfly is defined as the difference between the average IV of an OTM call and an OTM put and the IV of the ATM straddle.

Using this data, it is possible to recover five strikes from the implied volatility data (at each maturity): two strikes below the prevailing forward price, one at the forward price, and two strikes above the prevailing forward price (more precisely, strikes for -10-delta puts, -25-delta puts, -50-delta puts, 25-delta, and 10-delta calls). For more details on how to compute strikes from IVs at different deltas, see Jurek (2014) and Della Corte, Sarno, Schmeling, and Wagner (2016).

2.5 Empirical Results

In this section, I compare the historical rolling window betas with the option-implied betas using the dollar factor model.

2.5.1 The Dollar Carry Trade

The dollar factor has been documented to carry a risk premium (Lustig, Roussanov, and Verdelhan, 2011), and Lustig, Roussanov, and Verdelhan (2014) provide evidence for this risk premium is counter-cyclical. More specifically, define the average forward discount as $AFD_{t,t+m} = \frac{1}{N}\sum_j^N f_{t,t+m}^j - s_t^j \approx i_{t,t+m} - \frac{1}{N}\sum_j^N i_{t,t+m}^j$, then buying the dollar factor when the AFD is negative and shorting it when it is positive, significantly enhances the excess to holding the dollar factor. Buying the dollar factor conditional on the sign of the AFD is referred to as the dollar carry trade, and the portfolio it is holding as the conditional dollar factor.

The high excess returns of the dollar carry trade is regarded as compensation for risk, rather than a pricing anomaly (Lustig, Roussanov, and Verdelhan, 2014; Verdelhan, 2017). The risk-based explanation is that when U.S. short-term interest rates are low relative to other developed economies, the U.S. economy tends to be in recession, thereby exposing U.S. investors who are long foreign currencies (and short the U.S. dollar) to the risk that the U.S. dollar appreciates when their marginal utility is high. As a result, a U.S. investor demands a risk premium for holding the conditional dollar factor.

Table 2.1 reports annualized mean excess returns and Sharpe ratios for the dollar carry trade and the HML carry trade of Lustig, Roussanov, and Verdelhan (2011). The portfolios are monthly rebalanced and the holding periods range from 1-12 months. The HML carry trade buys the upper tertile interest rate currencies and shorts the lower tertile interest rate currencies [6]. The brackets below the mean excess returns report Newey and West (1987) t-statistics, with automatic lag selection according to Newey and West (1994). Figure 2.1 shows the cumulative returns of the 1-month dollar carry trade and the HML carry trade, and in the panel below, the annualized 1-month AFD is plotted.

For the G10 currencies from 1998-2016, the dollar carry strategy has on average been profitable. For instance, at the 1-month horizon, it delivers a mean excess return of 3.45% and a Sharpe ratio of 0.41. It is important to note that the excess returns to the dollar carry trade are driven by the spot component, in contrast to the HML carry trade where the excess returns stem from the interest rate component. At the 1-month month horizon, for instance, the HML carry trade and the dollar carry trade returns have an insignificant correlation of 19%, and only 0.55% per annum of the mean excess return on the dollar carry trade is explained by exposure the HML carry factor. This suggests that the dollar carry trade and the HML carry trade are driven by different factors, consistent with the results

[6] More precisely, it buys the upper tertile forward contracts trading at a discount, and shorts the lower tertile forward contracts trading at a premium.

reported in Lustig, Roussanov, and Verdelhan (2014) and Verdelhan (2017).

As the maturity gets longer, the mean excess return on the dollar carry trade declines from 3.45% at the 1-month horizon to 1.80% at the 12-month horizon. The interest rate component of the excess returns is stable across maturities, but the spot component declines, i.e., the AFD's predictability of future spot changes is weaker at longer horizons. The HML carry trade, on the other hand, is almost exclusively driven by the interest rate component and delivers virtually the same mean excess returns at horizons between 1-12 months, ranging from 3.27% to 2.93%.

2.5.2 Measuring Dollar Factor Betas

Following Verdelhan (2017), I compute historical dollar factor betas by regressing changes in exchange rates on changes in the dollar factor (excluding the exchange rate under consideration). Verdelhan (2017) uses 60-month past currency returns in his benchmark estimation of betas. This choice is likely because it has been the standard approach in the estimation of betas in the equity literature, at least since Jensen, Black, and Scholes (1972). Given the relatively small sample used in this paper, and because I want to obtain the most powerful estimates of betas, I use daily data for the beta estimation. In the benchmark estimation, I use a 252-day rolling window, but the results for historical betas are virtually unaffected when using 126-day or 504-day rolling windows.

Buss and Vilkov (2012) compare historical and option-implied equity betas in the context of the CAPM and find that historical betas estimated from 252-day data exhibit better performance than 60-month betas. Frazzini and Pedersen (2014) argue that correlations tend to move more slowly than volatilities, and thus, the best beta estimates are achieved by estimating volatilities using 128-252 daily observations and correlations using a 3-year rolling window of three-day overlapping data, which resembles the daily data approach taken here the most.

The various different methods used in the literature to estimate betas underline some of the complications that arise when estimating betas from historical data. Which data frequency should be used? What is the appropriate time frame? Should correlations and volatilities be estimated over different horizons/frequencies? On the one hand, a sufficient amount of data is needed to get statistically reliable estimates of betas, and on the other hand, time-variation in betas should also be captured. In contrast, very much like option-implied volatilities, which are observable in the market place, the option-implied betas reflect the market's (risk-neutral) view on betas.

Table 2.2 reports the summary statistics for the 1-12-month option-implied dollar factor betas (Q-betas) and the 252-day rolling window dollar factor betas (P-betas). The Q-betas are computed from the expression (2.13) using only options data available on a given day. Starting with the Q-betas, we note that the time-series averages and standard deviations are similar in magnitude at different maturities. There is, however, substantial time-variation in the term structure of Q-betas as seen from Figure 2.4, which plots the spread between the 12-month Q-beta and the 1-month Q-beta. Table 2.3 reports correlations between the Q-betas at different maturities, and it corroborates that the spread between Q-betas across different maturities varies over time. For example, the correlation between the 1-month and 12-month betas ranges from 67% for the AUD to 94% for the CAD, which suggests that there is potentially additional information embedded in the term structure of betas. The availability of a term structure of betas is a special feature of the Q-betas and something which cannot be achieved through time-series regressions. In the empirical section, I explore further if the information content inherent in the term structure of betas can help produce better forecasts of realized betas and currency returns.

As a first step in assessing the relationship between Q-betas and P-betas, I plot in Figure 2.2 their respective time-averaged means for each currency against each another. The Q-betas are computed using 1-month daily option prices, and the P-betas are computed using daily historical realized currency returns (252-day rolling windows). The plot shows that the time-averaged P-betas against the time-averaged Q-betas line up almost perfectly around the linear regression line (R^2 = 97.7% and the regression coefficient is 1.086). This is perhaps a bit surprising, since the betas are derived using vastly different methodologies—the first one being based on a single cross-section of option prices and the latter on slow-moving regressions.

Figure 2.3 illustrates the time series of P-betas and the 1-month Q-betas for the G10 currencies. For any exchange rate, the betas appear to be positively related, and they tend to move together over longer horizons. This is encouraging in the sense that for the Q-betas to be useful predictors of realized betas, the two cannot be entirely disconnected, but on the other hand, it may also raise the concern that they are substitutes for one another. The Q and P-betas, however, are far from perfectly correlated, as shown in Table 2.3, which reports their contemporaneous time-series correlations. The correlation between the 1-month Q-beta and the P-beta is lowest for the NOK at 34%, and it is largest for the CAD at 83%. The average correlation across all currencies is 44%. The empirical results presented below corroborate that the two types of betas are not interchangeable, as they produce vastly different predictions of realized betas and returns.

2.5.3 Dollar Factor Beta-Sorted Portfolios

If the conditional dollar factor model is an appropriate model, the expected excess return of a portfolio should increase monotonically in the portfolio's expected conditional dollar factor exposure, entailing that a high minus low conditional dollar factor portfolio delivers a positive expected excess return. In practice, identifying such a risk-return relation relies critically on accurate measurements of dollar factor exposures. The slow-moving nature of the rolling window betas may not be very informative of the realized risk exposures over the course of, say, the next month, and especially not if there are rapid changes in the factors that drive exchange rates. Naturally, we may then ask if the ex-ante nature of option-implied betas, and their ability to instantaneously incorporate new information, make them better at anticipating future returns than historical betas.

As a first step in the comparative analysis of the betas, I construct beta-sorted portfolios using both methodologies. I follow the portfolio construction procedure of Verdelhan (2017). Specifically, each month, for each type of beta separately, I allocate the currencies into three equal-weighted portfolios from low to high based on their dollar factor betas. I then construct three portfolios P_1, P_2, and P_3 which are long the respective beta-sorted portfolios whenever the average foreign discount is negative (i.e., average foreign interest rate is larger than the U.S. dollar interest rate) and short otherwise. In other words, the portfolios are constructed based on their exposure to the conditional dollar factor. Table 2.4 shows mean excess returns, standard deviations, and Sharpe ratios for each of the portfolios at horizons of 1-12 months.

The brackets below the mean excess returns are t-statistics based on Newey and West (1987), with the automatic lag selection of Newey and West (1994). In the construction of the Q-beta-sorted portfolios, I use options with the same time to expiry as the holding period of the forward contracts. The P-beta used for portfolio construction is calculated on the basis of overlapping daily rolling window regressions with a length of 252, i.e., the same beta is used for each holding period. Both shorter (126 days) and longer (504 days) rolling windows produce similar results, therefore I only report results for the 252-day P-betas.

Table 2.4 reveals a clear pattern: at any holding period, the mean excess portfolio returns increase in the ex-ante Q-beta, while a more dispersed pattern is to be found when P-betas are used for portfolio construction. E.g., for the 1-month holding period, a high minus low (HML) factor based on Q-betas, which buys portfolio P_3 and shorts portfolio P_1, gives a significant (t-statistic of 2.58) mean annualized excess return of 3.35 percent (Sharpe ratio 0.41), while the HML factor based on P-betas has an insignificant (t-statistic of 0.57) mean excess return of 0.95 percent (Sharpe ratio 0.11). The mean excess returns to the Q-beta HML factors are positive at longer horizons as well (albeit only significant at the 2-month holding period), and larger than for the corresponding P-beta HML factors.

Figure 2.5 illustrates the cumulative returns to monthly rebalanced HML dollar factors, for both types of betas, along with the annualized 1-month Q-volatility and the 252-day rolling volatility of the dollar factor. We see that implementing an HML dollar strategy based on Q-betas, rather than P-betas, gives larger returns throughout the sample period. Interestingly, while the dollar carry trade performs poorly from 2010-2016 (Figure 2.1), the HML dollar strategy continues to deliver high positive excess returns. In this period, the AFD is negative, therefore, the dollar carry trade is short the U.S. dollar and long the equal-weighted basket of foreign currencies. Thus, the dollar carry trade is exposed to an upwards shift in the level of the U.S. dollar relative to all foreign currencies, which in fact occurred over this period. The HML dollar factor, on the other hand, is immune to level shifts in the U.S. dollar, since the long and short side of the portfolio are affected equally.

Notably, there is no obvious link between the volatility of the dollar factor and the returns to the HML dollar strategy. For instance, the HML factor does not crash during the financial turmoil in 2008, as the HML carry trade (Jurek, 2014; Menkhoff, Sarno, Schmeling, and Schrimpf, 2012a). This highlights that the HML factor and the HML carry trade appear to be driven by different risk factors (in this sample, their correlation is $\sim 19\%$).

Table 2.5 shows the mean excess returns of each beta-sorted portfolio decomposed into a spot and interest rate component. Interestingly, at any holding period, the larger mean excess returns on the Q-beta HML factors compared to the P-beta HML factors stem entirely from the spot component. For example, at the 1-month holding period, the spot components are 2.35% and -0.18%, and the interest rate components are 1.00% and 1.12%, for the Q-beta and P-beta HML factors, respectively. Thus, for the Q-beta HML factor the largest proportion of the excess return is due to spot changes, which is in contrast to the HML carry trade, where the return is primarily driven by the interest rate differential (see Table 2.1). For both types of beta, the annualized interest rate components are virtually the same for the HML factor portfolios across different holding periods, whereas the spot components decrease, i.e., the Q-beta predictability of currency spot changes is confined to shorter horizons.

One potential explanation for why the Q-betas are better at explaining the cross-section of currency returns, relative to the P-betas, is that they are better at predicting realized volatility and to a lesser extent because they more accurately forecast correlations with the dollar factor. For instance, Jorion (1995) and Busch, Christensen, and Nielsen (2011) provide evidence that implied volatilities from currency options are better predictors of realized currency volatility than historical volatility measures.

I examine if this is the case by constructing portfolios on the basis of betas which are built from a mixture of Q and P-moments. Specifically, I follow the method suggested by French, Groth, and Kolari (1983), in which betas are constructed from historical correlations and option-based variances. Supposedly, if the Q-correlations are good predictors of realized correlations, this beta method will be less successful at identifying high and low-beta currencies ex-ante. In the same spirit, I also construct mixed betas based on Q-correlations in conjunction with P-variances.

Following the exact same procedure as previously, portfolios are constructed based on both types of mixed betas. Table 2.6 shows the results. The HML factor constructed from betas combining P-correlations and Q-variances gives a lower mean excess return, at any horizon, compared to the Q-beta HML factor. For instance, at the 1-month horizon, the mean annualized excess return is 2.53% (t-statistic 1.53), compared to 3.35% when Q-correlations are used (Table 2.4). Furthermore, the Sharpe ratio declines to 0.30 (11 percentage points), suggesting that the Q-correlations are more effective at constructing betas for cross-sectional analysis than P-correlations.

Using betas built from Q-correlations and P-variances to construct portfolios reaffirms that Q-correlations are useful for computing betas. Using this mixed dollar factor beta, the HML factors have larger mean excess returns and Sharpe ratios relative to the corresponding P-beta HML factors at all horizons. There is a monotonic relation between ex-ante portfolio betas and ex-post portfolio excess returns when sorting on the mixed beta, albeit the HML factor excess returns are insignificant.

Among all types of betas, the pure Q-betas perform best for portfolio construction, which corroborates that both Q-correlations and Q-variances contain useful information for the computation of betas.

2.5.4 Evaluation of Model Predictions

Another important aspect of betas is how strong predictors they are of realized returns when used as inputs in the factor model. We may erroneously reject an accurate model due to poorly measured betas. Although the Q-betas identify a monotonic relation between ex-ante portfolio betas and ex-post returns, it is not certain that they perform well in the context of the model. Betas could be flawed and still capable of properly ranking currencies on ex-ante betas, and thus performing well in a portfolio sorting exercise. In the following, I examine in greater detail how the two types of betas perform in the context of the model.

Using the conditional dollar factor model implies a linear relation between expected portfolios excess returns and betas. As a first step to examine this, I plot in Figure 2.6, at the 1-month horizon, the average realized excess returns against the average model predicted excess returns for the dollar beta-sorted portfolios, for each type of beta. Table 2.7 reports the results at horizons of 1-12 months. The model predicted returns are computed assuming a fixed price of conditional dollar risk equal to its unconditional mean over the entire sample (this follows from the Euler equation since the beta of a tradable systematic factor is 1). Since the price of risk is fixed, the model predicted excess return for each portfolio is the average portfolio beta times the conditional dollar price of risk.

From Figure 2.6, we see that using the Q-betas leads to a relation between portfolio betas and returns that is too steep. The (low) high-beta portfolio has a (smaller) larger mean excess return than predicted by the model, whereas, when using P-betas, the opposite is the case. As a result, the realized risk-return relation is too flat for the P-betas, whereas for Q-betas it is too steep—but the model predictions and realized returns appear to be better aligned. These results should be considered as indicative and must be interpreted with caution. The exercise assumes a fixed price of risk for the conditional dollar factor (estimated over the entire sample period), and the prediction errors are uninformative of the model's performance in the time series. The takeaway from Figure 2.6, however, is consistent with the results obtained via dynamic time-series predictions (which I will discuss further below): when using Q-betas rather than the P-betas, the factor model exhibits less biased predictions of returns and the mean time-series prediction errors are smaller.

The results for time-series prediction errors are reported in Table 2.8 for both types of betas. For each portfolio, and beta type, the prediction error is computed as the realized excess return less the model predicted excess return. This is done monthly, and thus a time series of prediction errors is generated. The model excess return is computed as the portfolio's ex-ante beta times the realization of the conditional dollar factor.

Columns 1-4, Panels A and B, show the mean squared prediction errors (MSEs) in basis points for portfolios sorted on Q-betas and P-betas, respectively. In general, the MSEs are larger, for both types of betas, for high and low-beta portfolios. But at any horizon, the mean MSEs across all portfolios are smaller when using the Q-betas compared to the P-betas. For instance, at the 1-month horizon, the mean P-beta MSE is 15.41 bps compared to a mean Q-beta MSE of 13.42 bps, and their difference is statistically significant with a Newey and West (1987) t-statistic of 2.18 (reported in the ninth column).

For comparison, I report in Panels C and D the MSEs for a random walk forecast, which predicts that the future spot exchange rates are equal to their current values. For all portfolios, at all horizons, the random walk forecasts have larger mean MSEs relative to both the P and Q-beta predictions. The t-statistics for the difference in mean MSEs between the random walk forecast and the beta forecasts are reported in the column furthest to the right, and are at all horizons statistically significant at the 1% level. This is consistent with Lustig, Roussanov, and Verdelhan (2014), who document that the average forward discount (U.S. minus the average foreign interest rate) is a strong predictor of aggregate foreign currency changes versus the U.S. dollar—foreign currencies tend to appreciate (depreciate) versus the U.S. dollar when the AFD is negative (positive).

Columns 5-8 report the mean prediction errors (ME), in percentages, for the beta-sorted portfolios, and the difference in MEs between the high and low-beta portfolios. The model forecasts based on P-betas appear to have a systematic bias, while there is no notable bias when using Q-betas. At the 1-month horizon, the P-betas tend to underestimate returns to the low-beta portfolio, with an average of 1.17%, and on the other hand, they tend to overestimate returns to the high-beta portfolio, with a mean prediction error of -1.20%. Consequently, the difference in mean prediction errors between the high-beta and the low-beta portfolio is -2.38%, whereas the bias is slightly positive, 0.44%, for the Q-beta-sorted portfolios. This difference in the bias between the P and Q-betas is consistent with that the HML dollar factor portfolio based on Q-betas delivers substantially larger excess returns than the HML dollar factor portfolio constructed based on P-betas. To summarize, the model performs better when using the Q-betas—the mean squared prediction errors are smaller at all horizons, and there are no notable prediction biases across the beta-sorted portfolios.

Interestingly, a similar prediction bias for the rolling window betas also appears in the CAPM in which (high) low-beta stocks tend to have (smaller) larger returns than predicted by the CAPM—the low-risk anomaly (Frazzini and Pedersen, 2014). Consistent with my findings, Buss and Vilkov (2012) find that when betas are computed based on options, there are no notable biases between low and high-beta portfolios, whereas rolling window betas tend to underestimate returns (ex-ante) of low-beta stocks, and vice versa for high-beta stocks. Thus, this suggests that rolling window betas generally induce biased model predictions.

One potential explanation for the better performance of the model when using Q-betas is that they are more powerful and less biased predictors of realized betas. In the following, I investigate this for both portfolios and individual currencies.

2.5.5 Predicting Dollar Factor Betas for Portfolios

In a factor pricing model, the expected excess return on a security is given by the *expected* beta times the price of risk of the factor. Therefore, accurate predictions of *future* betas are crucial for empirically identifying a monotonic relation between ex-post returns and betas. A true model may erroneously be rejected if the betas are ranked in the wrong order, and furthermore, if the level of betas is inaccurately measured, this may cause large model prediction errors. The superior model performance when using the option-implied betas may thus be due to them being better predictors of ex-post betas. In the following, I investigate this hypothesis by comparing the predictive power of the two types of betas for portfolio betas.

Realized beta is not observable and needs to be proxied by a measurable quantity. Following Andersen, Bollerslev, Diebold, and Wu (2006), Buss and Vilkov (2012), and Chang, Christoffersen, Jacobs, and Vainberg (2011), I measure realized betas using daily rolling window regressions, where the length of the window matches the forecast horizon. That is, for forecasts at time t at horizon τ, the realized beta at time $t+\tau$ is estimated using daily data from $(t, t+\tau]$. I calculate the P-beta predictors of realized betas at horizon τ using a window of length τ, i.e., the time t P-beta forecasts are computed based on daily currency returns from $(t-\tau, t]$. At forecast horizons of up to one year, the maturity that I use to derive the Q-beta predictors matches the forecast horizon, and for forecast horizons exceeding one year, the 1-year options are used.

I calculate the portfolio beta prediction errors as follows for each forecast horizon: every day, for each type of beta, three portfolios are constructed based on their expected betas. For each portfolio, I calculate the daily model prediction error as the difference between the realized and the expected portfolio beta. It is important to note that the ex-ante portfolio of currencies is compared to the exact same set of currencies ex-post.

I conduct daily forecasts to increase the power of the predictive tests, as in, e.g., Della Corte, Sarno, and Tsiakas (2011), Chang, Christoffersen, Jacobs, and Vainberg (2011), and Jorion (1995). Due to the overlapping data used to derive the rolling window betas, which causes autocorrelation in the prediction errors, I use t-statistics based on Newey and West (1987), with number of lags that matches the forecast horizon, as in, e.g., Della Corte, Sarno, and Tsiakas (2011) and Chang, Christoffersen, Jacobs, and Vainberg (2011). I run the forecasts at horizons that span 6-18 months; i.e., the 1-3 month horizons are not included in the forecasts as was the case for the portfolio return predictions. This mismatch is because the forecast horizon has to be sufficiently long in order to get reliable beta estimates while at the same time avoiding overlapping the data used to estimate the ex-ante and ex-post rolling window betas.

Table 2.9 shows the prediction error results for the daily beta-sorted portfolios. First, I compare the biases for the forecasts. The left panel (ME) reports the mean of the time-series prediction errors for each portfolio, for both types of beta. At all horizons, the P-beta forecasts have a propensity to (overestimate) underestimate betas of the (high) low-beta portfolios. The mean prediction bias between the high and low-beta portfolios is highly significant at all horizons, with t-statistics ranging from -4.02 to -2.92.

At the 6-month horizon, as an example, the P-beta forecasts of the low-beta portfolios underestimate realized betas with an average of 7.44% (in beta units). On the other hand, high-beta portfolios are on average overestimated by 8.29%, resulting in a highly significant (t-statistic -2.92) high minus low beta bias of -15.73% (-33.47% relative bias). In comparison, the 6-month mean prediction errors for low and high Q-beta-sorted portfolios are -0.26% and -3.46%, respectively, resulting in an insignificant (t-statistic -1.42) high minus low mean prediction error of -3.20% (-5.92% relative bias).

This significant P-beta prediction bias makes them prone to misidentifying low-beta currencies as high-beta (and vice versa), which strikes as being a plausible explanation for why there is no monotonic relation between ex-ante betas and ex-post portfolio excess returns. In contrast, since there is no notable prediction bias for the Q-betas—with a relative bias that is nearly six times smaller—they are to a lesser extent subject to this issue. Furthermore, the difference in the bias between the Q-beta and P-beta forecasts of realized betas is consistent with that the P-betas have a tendency to (overestimate) underestimate the (high) low-beta portfolio excess returns, while there is no noticeable bias for the Q-betas.

The high minus low-beta bias increases in the length of the forecast horizon, albeit it is much smaller for the Q-betas at any horizon. The growing high minus low beta bias at longer horizons appears to be a likely explanation for the diminishing returns to the HML dollar portfolios for longer holding periods (see Table 2.4).

The panel on the right shows the mean squared prediction errors (MSEs) in percentages. We see that for all portfolios, at all horizons, the Q-beta forecasts have the smallest MSE. Notably, the Q-betas have much stronger predictive power for the low and high-beta portfolios. The average MSE across all three portfolios is smallest for the Q-beta forecasts at any forecast horizon. For example, at the 6-month horizon, the average MSE is 1.34% for the

Q-beta forecasts, while it is 1.90% based on the P-beta forecasts. The t-statistics for the difference in the average MSEs are reported for each forecast horizon in the column furthest to the right. We see that the Q-beta average MSEs are statistically significantly smaller at all forecast horizons, with t-statistics ranging from $2.30 - 3.56$. The Q-betas therefore deliver not only less biased predictions of betas, but also smaller prediction errors.

To conclude, the findings that I provide in this section suggest that the stronger predictive power of the Q-betas compared to the P-betas explain why the Q-betas correctly identify a monotonic relation between ex-ante betas and ex-post returns and why they exhibit the smallest model prediction errors of portfolio excess returns in the time series.

2.5.6 Predicting Dollar Factor Betas for Individual Currencies

As a final comparison between the predictive power of the two beta types, I conduct predictive regressions of realized betas for individual currencies on ex-ante Q-betas and P-betas:

$$\beta^P_{it+\tau} = \gamma^Q_{i0} + \gamma^Q_{i1}\beta^Q_{it} + \varepsilon^Q_{it+\tau} \tag{2.16}$$

$$\beta^P_{it+\tau} = \gamma^P_{i0} + \gamma^P_{i1}\beta^P_{it} + \varepsilon^P_{it+\tau} \tag{2.17}$$

Ideally, the intercept is zero and the slope coefficient unity. As for the portfolio forecasts, the ex-post betas are computed using daily data over the forecast horizon $(t, t+\tau]$, and the P-beta predictors are computed using daily data from $(t-\tau, t]$. The Q-beta predictors are based on option prices at time t for which the time to expiry matches the length of the forecast horizon as closely as possible.

The results for the Q-beta and P-beta predictive regressions are reported in Tables 2.10 and 2.11, respectively. As for the portfolio beta predictions, the predictive regressions are conducted on daily data to increase the power of the tests, and t-statistics (reported in brackets under the relevant coefficients) are adjusted for autocorrelation using Newey and West (1987) with number of lags that match the forecast horizon, as in, e.g., Della Corte, Sarno, and Tsiakas (2011) and Chang, Christoffersen, Jacobs, and Vainberg (2011).

The results for the individual predictive regressions corroborate that Q-betas are better predictors of ex-post betas than P-betas. At the 5% significance level, 30 out of 45 predictive regressions (nine exchange rates at five horizons) have a significant slope coefficient, while it is only significant in four cases when using the P-betas. In all the predictive regressions, for both beta types, the slope coefficient is less than unity and has a positive intercept, i.e., the beta predictors tend to be biased predictors of ex-post betas. The slope coefficient, however, is closer to unity when using the Q-betas in 30 out of 45 cases, and in the same regressions the intercept is closer to 0. Likewise, Q-betas have a higher explanatory power in 29 of the cases relative to the P-betas when using the R^2-metric.

It is important to note that there are a few currencies for which the P-betas have notably stronger explanatory power, namely, the GBP, NOK, and SEK. According to a central bank survey conducted by Bank for International Settlements (Bank of International Settlements, 2016), the latter two currencies are among the most illiquid currencies of the G10, e.g., they both account for less than 1% of the overall currency market turnover (other liquidity metrics, such as the volume in OTC currency interest rate derivatives, are consistent with this picture).

Thus, one plausible explanation for the weak forecast performance for the SEK and NOK is that the cross-pair options are illiquid, and perhaps especially so for longer maturities. In fact, using 1-month option-implied betas improves the explanatory power for these currencies, most notably at longer forecast horizons where longer-dated option maturities were used to construct predictors (similar improvements are obtained by using 2-month and 3-month maturities). Chang, Christoffersen, Jacobs, and Vainberg (2011) find a similar result in predictions of realized CAPM betas using option-implied CAPM betas, where option-implied betas based on short-term liquid options have a stronger predictive ability than betas based on longer-dated illiquid options, even at longer horizons. If the long maturity options are in fact less liquid, there appears to be an important trade-off between applying the more liquid short-term maturities vs. longer-dated maturities that match the forecast horizon better.

The overall predictive performance is better for Q-betas compared to the P-betas according to any metric considered in the analysis, both for portfolios and individual currencies. However, there are a few exceptions, as mentioned above, in which the forecasts would likely benefit from incorporating historical information. As a small step in this direction, I performed multivariate regressions using both types of beta as regressors, and I indeed found that the explanatory power tends to increase the most for currencies in which the Q-beta is a weak predictor. In the interest of space, only the adjusted R^2s of those regressions are reported in Table 2.12. It appears to be an interesting topic for future research to further explore if betas based on information inherent in the two types of betas are beneficial for portfolio construction and currency return predictions.

2.6 Conclusion

In recent currency literature, factor models have shown promising results toward explaining the time-variation and cross-section of currency returns. For these models to be applied successfully, accurate measurement of betas with respect to the factors is crucial. I develop a method for calculating purely forward-looking betas to portfolios of currencies using options on currency cross-pairs. Specifically, I compare the option-implied betas to the historical betas by using the dollar factor—an equal-weighted portfolio of foreign currencies against the U.S. dollar—as the common factor driving currency excess returns.

By allocating currencies into portfolios based on their dollar factor betas, I document that a portfolio which buys high dollar factor beta and shorts low dollar factor beta currencies delivers significant mean excess returns when the option-implied betas are used, while the mean excess returns are insignificant when historical betas are used. Furthermore, I find that the dollar factor model delivers smaller prediction errors and less biased predictions when using the option-implied betas compared to when using the historical betas. The historical betas tend to underestimate low-beta portfolio returns and overestimate high-beta portfolio returns. In comparison, the option-implied betas deliver virtually unbiased predictions.

I provide evidence that the superior model performance when using option-implied betas is because they are more powerful and less biased predictors of realized betas compared to historical betas. At all forecast horizons, the prediction errors of portfolio betas are statistically significantly smaller when using option-implied betas compared to the historical betas. The option-implied betas deliver unbiased forecasts of realized betas, i.e., they exhibit no propensity to underestimate or overestimate betas for low and high-beta currencies. In contrast, the historical betas tend to (overestimate) underestimate (high) betas for low-beta currencies. Furthermore, the results for the predictive regressions for individual currency betas suggest that option-implied betas are more powerful predictors of realized betas, both in terms of significance of the slope coefficients and explanatory power (R^2s).

2.7 Figures

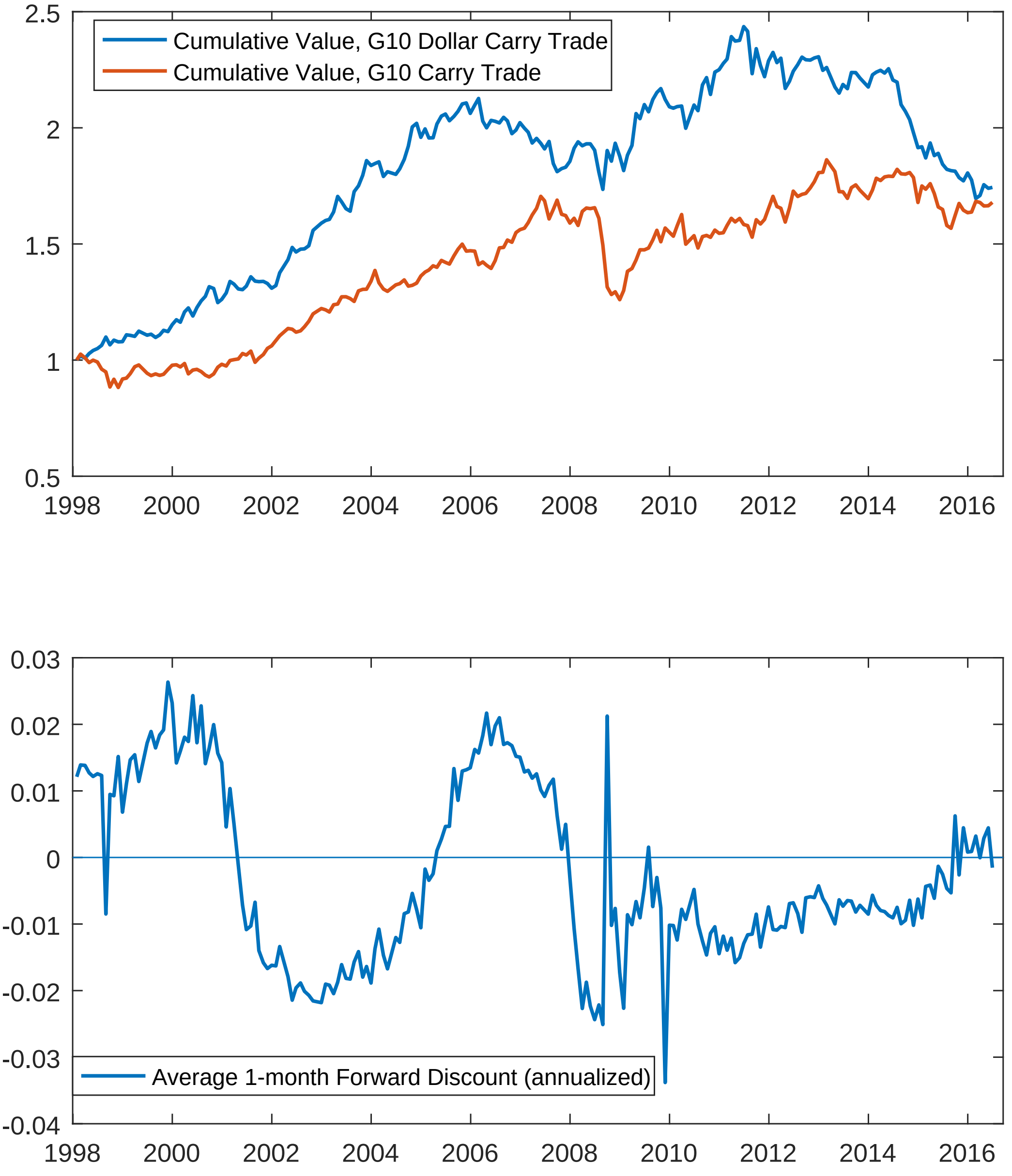


Figure 2.1: **Cumulative returns of the dollar carry trade and the HML carry trade.** The upper figure shows the 1-month annualized cumulative value of investing one dollar in the dollar carry trade and the HML carry trade. The dollar carry trade is long an equally weighted basket of all 1-month forward contracts when the average 1-month forward discount is negative and short the same basket of currencies when it is positive. The HML carry trade is monthly rebalanced and is long (short) the upper (lower) tertile interest rate currencies. The lower figure shows the annualized average 1-month forward discount for the G10 currencies. The sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

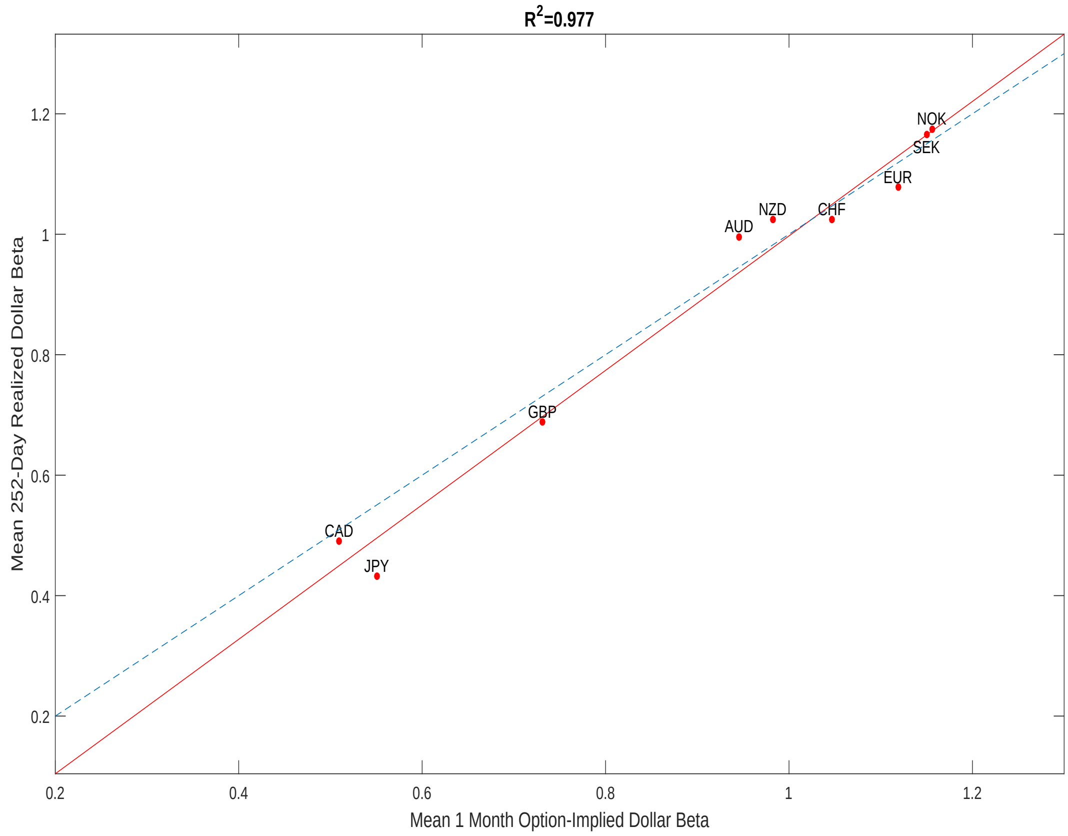


Figure 2.2: **Scatterplot of time-averaged means of P-betas against Q-betas.** This figure illustrates the scatterplot of the average historical dollar factor beta plotted against the average 1-month option-implied dollar factor beta. The dollar factor beta for a currency i is computed as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, which is defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed using 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the sample comprises 4659 daily observations.

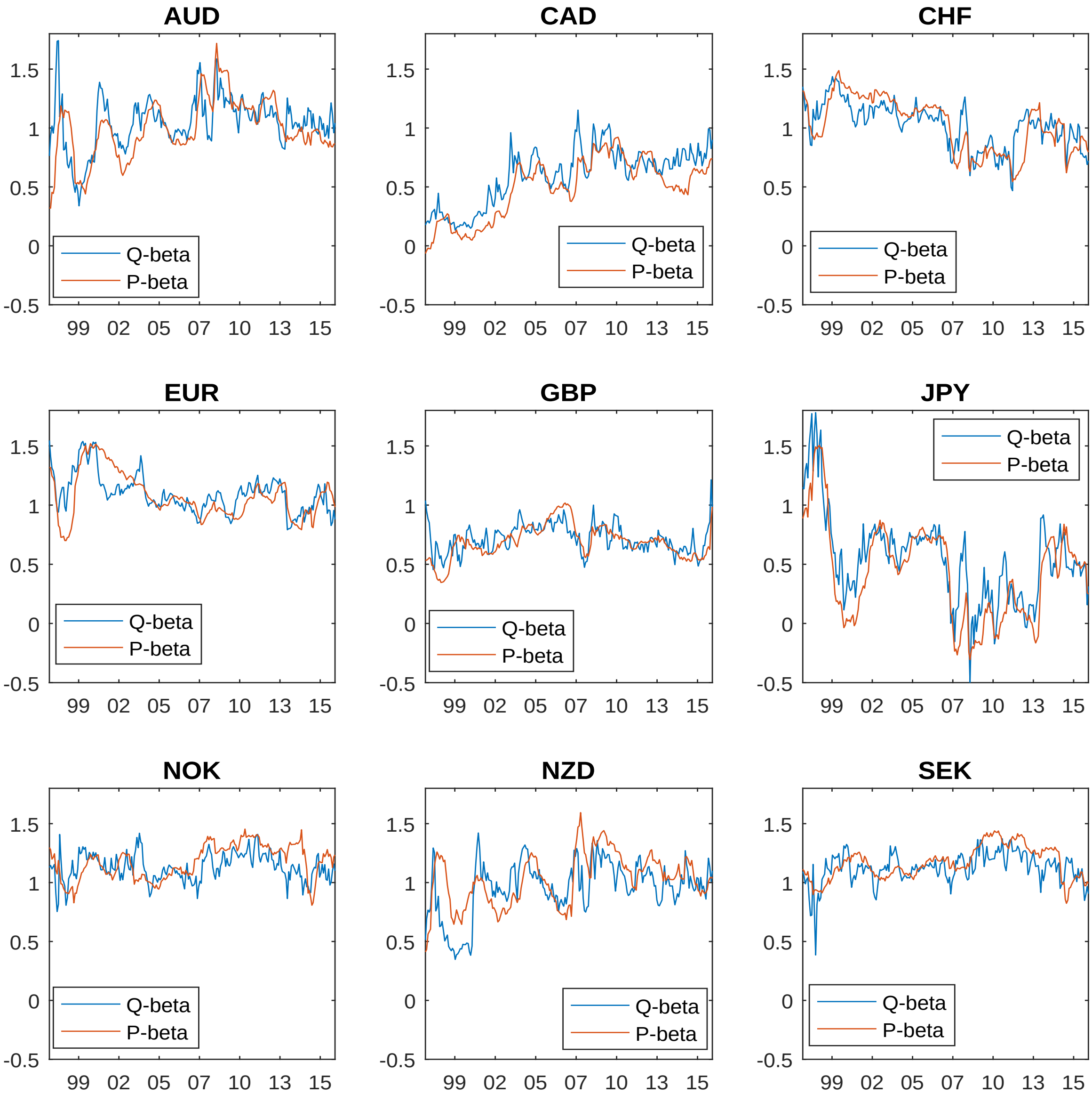


Figure 2.3: **Time series of 1-month Q-betas and P-betas for the G10 currencies.** This figure illustrates the time series of the 1-month option-implied dollar factor beta and the 252-day rolling window dollar factor beta for each currency. The dollar factor beta for currency i is computed as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, which is defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed from 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

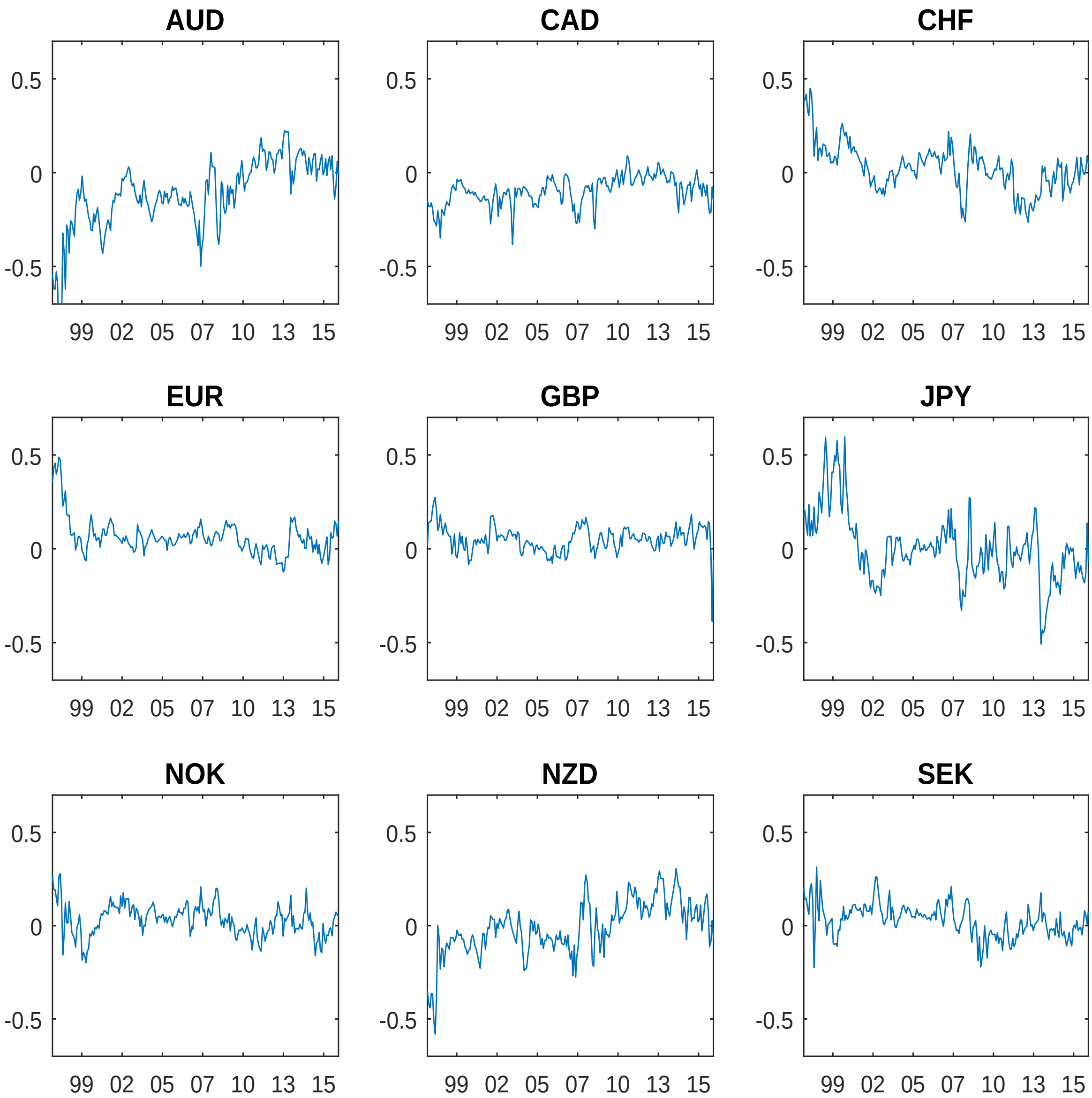
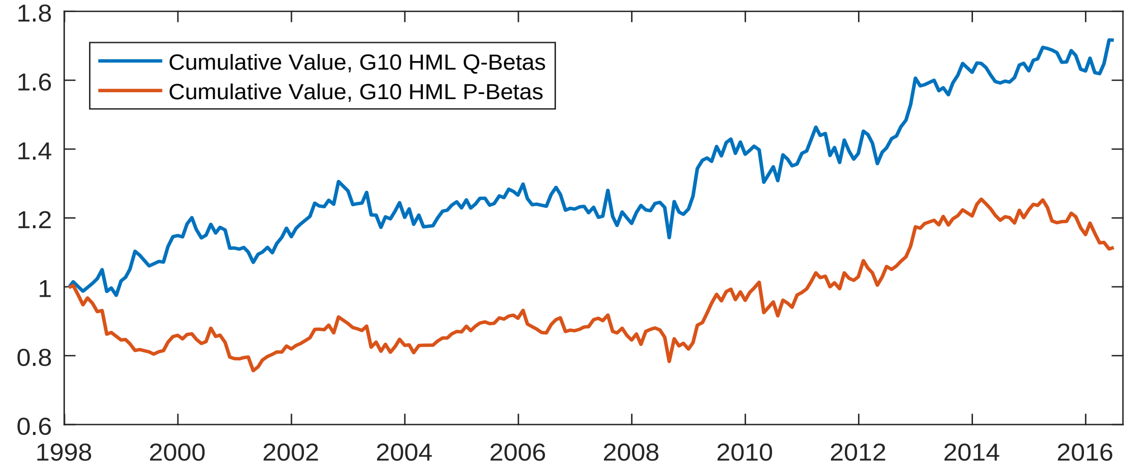


Figure 2.4: **Time series of the 12-1 month Q-beta spread for the G10 currencies.** This figure illustrates the time series of the difference between the 12-month and 1-month option-implied dollar factor betas for each currency. The dollar factor beta for a currency i is computed as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, which is defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed from 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.



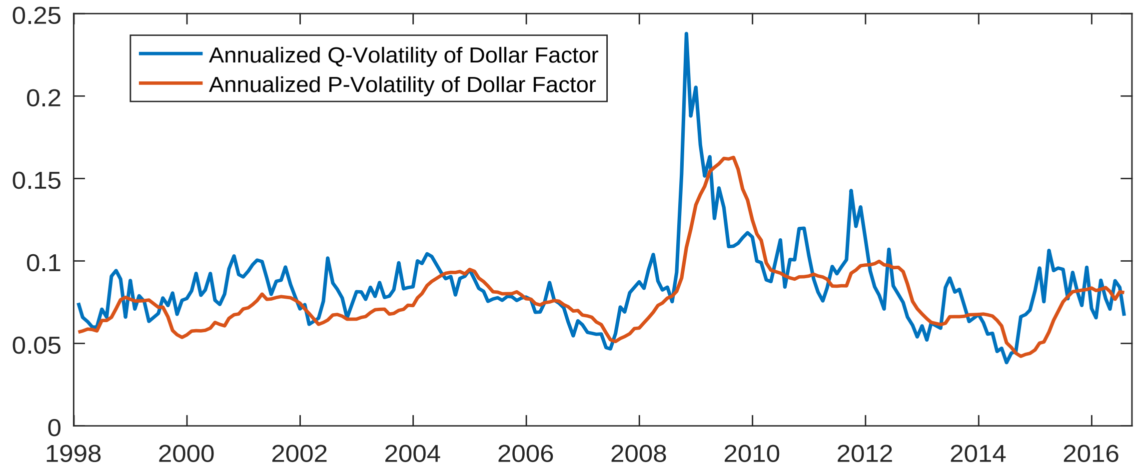


Figure 2.5: **Cumulative returns of HML dollar factor and dollar factor volatility.** The upper panel shows the cumulative return, for P and Q-betas for the HML dollar factor. Each month, and for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios: P_1, P_2, and P_3. Each month, the HML dollar factor buys (sells) P_3 and sells (buys) P_1 when the average forward discount is negative (positive). The Q-betas are derived from 1-month maturity options using the expression in (2.13)—i.e. as the covariance between the relevant currency and a equal-weighted portfolio of all foreign currencies vs. U.S. dollar (the dollar factor) normalized by the risk-neutral variance of the dollar factor. The P-betas are computed using 252-day rolling windows and updated every month. The lower panel shows the annualized 252-day rolling window volatility of the dollar factor along with its risk-neutral annualized 1-month volatility. The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

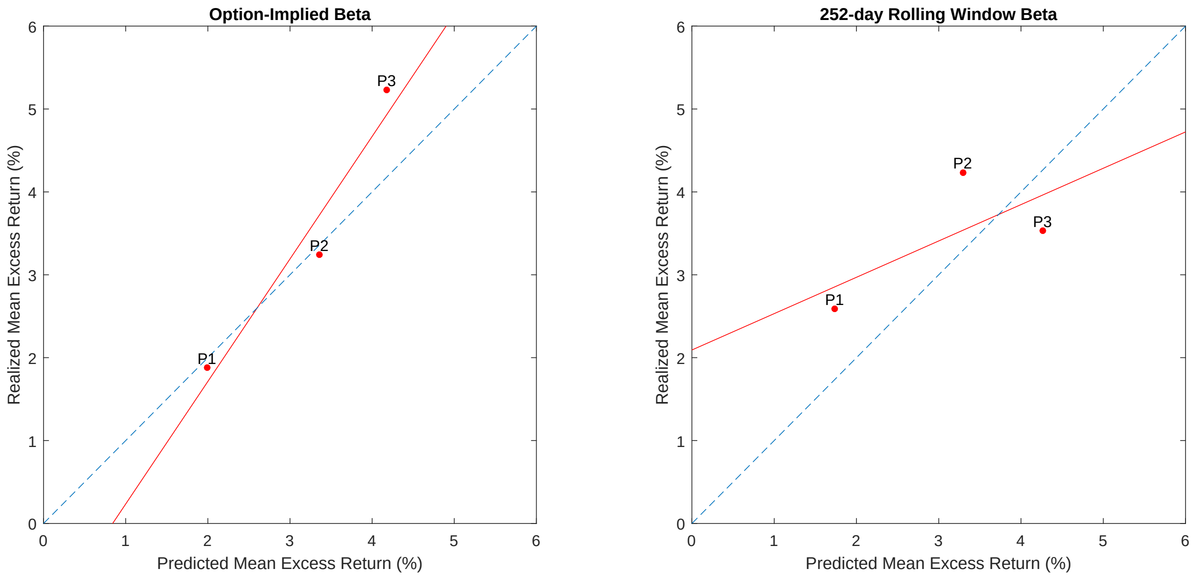


Figure 2.6: **Mean realized vs. mean predicted excess returns for Q and P-betas.** This figure shows the scatterplot of mean realized portfolio excess returns plotted against model predicted mean excess returns at a holding period of one month. For each type of beta, the model predicted excess return is computed as mean beta of each portfolio times the unconditional mean excess return of the conditional dollar factor. The conditional dollar factor is long the dollar factor—equal-weighted basket of all foreign currencies vs. U.S. dollar—if the average foreign discount is negative and short this portfolio otherwise. The dollar factor beta for currency i is computed as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor. The option-implied beta portfolios are constructed based on the 1-month option-implied betas, using expression (2.13). The P-betas are computed from 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

2.8 Tables

Table 2.1: **Excess returns on the dollar carry trade and HML carry trade.** This table shows annualized excess mean returns, the spot exchange rate component, forward discounts and standard deviations of the dollar carry trade and the HML carry trade. The dollar carry trade is long an equal-weighted basket of all foreign currencies if the AFD is negative and short the same set of currencies otherwise. The HML carry trade is long the upper tertile interest rate currencies and short the lower decile interest rate currencies. The excess returns for each strategy are reported at six different maturities: 1, 2, 3, 6, 9 and 12 months, and the portfolios are rebalanced at the end of each month. Newey and West (1987) t-statistics are reported in brackets and ***, **, and * indicate significance at a 1%, 5%, and 10% level, respectively. The exchange rate data are from Reuters through Datastream and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

Panel A: Dollar Carry Trade					
Horizon	Mean	$\frac{\Delta S_{t+m}}{S_t}$	Δi_{t+m}	Std	Sharpe Ratio
1 mo	3.45*	2.32	1.32	8.32	0.41
	[1.89]				
2 mo	2.71	1.58	1.31	8.57	0.32
	[1.24]				
3 mo	2.16	1.04	1.31	8.79	0.25
	[0.87]				
6 mo	1.72	0.63	1.33	9.51	0.18
	[0.42]				
9 mo	1.80	0.73	1.34	9.67	0.19
	[0.49]				
12 mo	1.80	0.74	1.35	9.76	0.18
	[0.54]				
Panel B: HML Carry Trade					
Horizon	Mean	$\frac{\Delta S_{t+m}}{S_t}$	Δi_{t+m}	Std	Sharpe Ratio
1 mo	3.27	-0.67	3.93	8.56	0.38
	[1.58]				
2 mo	3.35	-0.53	3.88	8.65	0.39
	[1.24]				
3 mo	3.13	-0.70	3.84	9.01	0.35
	[1.28]				
6 mo	2.98	-0.82	3.80	8.91	0.33
	[0.82]				
9 mo	2.93	-0.85	3.77	8.51	0.34
	[0.98]				
12 mo	2.93	-0.84	3.76	7.95	0.37
	[1.09]				

Table 2.2: **Descriptive statistics for dollar factor betas.** This table shows descriptive statistics for the dollar factor betas. The dollar factor beta for currency i is computed as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, which is defined as an equally weighted basket of all currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed from 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

Panel A: Q-betas									
1-month:	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
Mean	0.95	0.51	1.05	1.12	0.73	0.55	1.16	0.98	1.15
Std	0.23	0.27	0.20	0.17	0.13	0.39	0.13	0.25	0.14
2-month:									
Mean	0.94	0.50	1.06	1.13	0.75	0.54	1.16	0.97	1.15
Std	0.22	0.26	0.21	0.17	0.12	0.38	0.12	0.24	0.12
3-month:									
Mean	0.93	0.50	1.05	1.14	0.75	0.55	1.16	0.97	1.16
Std	0.22	0.26	0.20	0.17	0.11	0.39	0.11	0.25	0.11
6-month									
Mean	0.93	0.50	1.06	1.15	0.76	0.55	1.16	0.96	1.16
Std	0.23	0.26	0.22	0.18	0.10	0.40	0.10	0.26	0.11
9-month:									
Mean	0.92	0.50	1.05	1.15	0.76	0.54	1.16	0.96	1.16
Std	0.24	0.26	0.22	0.18	0.10	0.41	0.10	0.26	0.10
12-month:									
Mean	0.92	0.50	1.05	1.15	0.77	0.54	1.17	0.96	1.16
Std	0.24	0.26	0.23	0.18	0.09	0.42	0.11	0.27	0.10
Panel B: P**-betas**									
	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
Mean	1.00	0.49	1.02	1.08	0.69	0.43	1.17	1.02	1.17
Std	0.25	0.25	0.23	0.19	0.14	0.41	0.15	0.22	0.14

Table 2.3: **Contemporaneous correlations between dollar factor betas.** This table reports the contemporaneous correlations between dollar factor betas. Panel A reports the contemporaneous correlations between the 252-day rolling window beta and the Q-betas at maturities from 1-12 months. Panel B reports the contemporaneous correlations between the 1-month beta and 2-12 month betas. The dollar factor beta for a currency i is computed as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, which is defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed using 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

Panel A: Correlations between P and Q-betas									
Maturity	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	0.54	0.83	0.70	0.60	0.54	0.81	0.34	0.45	0.48
2 mo	0.66	0.89	0.77	0.60	0.53	0.83	0.40	0.52	0.45
3 mo	0.67	0.90	0.77	0.61	0.50	0.83	0.37	0.54	0.43
6 mo	0.67	0.91	0.80	0.59	0.46	0.83	0.46	0.57	0.43
9 mo	0.67	0.92	0.80	0.57	0.45	0.83	0.46	0.59	0.40
12 mo	0.65	0.91	0.77	0.55	0.40	0.81	0.43	0.58	0.40
Panel B: Correlations between Q-betas									
Maturity	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1/2 mo	0.92	0.97	0.98	0.98	0.96	0.99	0.96	0.97	0.97
1/3 mo	0.86	0.97	0.95	0.95	0.93	0.97	0.92	0.95	0.93
1/6 mo	0.76	0.95	0.89	0.90	0.87	0.93	0.83	0.88	0.85
1/9 mo	0.70	0.94	0.86	0.86	0.83	0.90	0.77	0.84	0.78
1/12 mo	0.67	0.94	0.82	0.83	0.80	0.88	0.72	0.82	0.73

Table 2.4: **Descriptive statistics of portfolios sorted on dollar factor betas.** This table reports the means, standard deviations and Sharpe ratios of excess returns on monthly rebalanced portfolios sorted on Q and P dollar factor betas at horizons of 1-12 months. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios P_1, P_2, and P_3. The investor buys (sells) each portfolio when the average forward discount is negative (positive). For Q-beta-sorted portfolios, the length of the holding period and the maturity of the options are the same. The dollar factor beta for a currency i is computed according to (2.13)—i.e. as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i).The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-beta is computed using a 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

Panel A: Q-betas												
	Mean				Std				Sharpe Ratio			
Horizon	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1
1 mo	1.88	3.24	5.23**	3.35**	6.68	9.53	10.96	8.20	0.28	0.34	0.48	0.41
	[1.08]	[1.41]	[2.31]	[2.58]								
2 mo	1.57	2.03	4.53**	2.96*	7.08	9.86	11.03	8.02	0.22	0.21	0.41	0.37
	[0.86]	[0.87]	[1.97]	[1.93]								
3 mo	1.13	2.02	3.32	2.18	7.28	10.10	11.24	7.82	0.16	0.20	0.30	0.28
	[0.62]	[0.75]	1.3]	[1.26]								
6 mo	0.89	2.22	2.05	1.16	7.91	11.02	11.39	7.54	0.11	0.20	0.18	0.15
	[0.49]	[0.84]	[0.84]	[0.70]								
9 mo	1.00	2.42	1.99	0.99	8.05	10.80	11.67	7.03	0.12	0.22	0.17	0.14
	[0.61]	[1.06]	[0.88]	[0.70]								
12 mo	1.08	2.34	1.97	0.89	8.28	10.63	11.81	6.96	0.13	0.22	0.17	0.13
	[0.70]	[1.08]	[0.93]	[0.72]								
Panel B: P-betas												
	Mean				Std				Sharpe Ratio			
Horizon	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1
1 mo	2.59	4.23*	3.53	0.95	6.89	9.74	10.80	8.27	0.38	0.44	0.33	0.11
	[1.40]	[1.93]	[1.52]	[0.57]								
2 mo	1.76	3.65*	2.72	0.97	7.00	9.91	10.98	7.91	0.25	0.37	0.25	0.12
	[0.95]	[1.71]	[1.07]	[0.53]								
3 mo	1.44	3.09	1.93	0.49	7.19	9.94	11.47	8.29	0.20	0.31	0.17	0.06
	[0.72]	[1.29]	[0.71]	[0.30]								
6 mo	1.14	2.89	1.14	0.00	7.96	10.33	12.14	8.27	0.14	0.28	0.09	0.00
	[0.58]	[1.21]	[0.42	[-0.01]								
9 mo	1.27	2.81	1.33	0.06	8.17	10.63	12.02	7.88	0.16	0.26	0.11	0.00
	[0.71]	[1.26]	[0.58]	[0.07]								
12 mo	1.39	2.46	1.54	0.15	8.45	10.94	11.78	7.63	0.16	0.23	0.13	0.02
	[0.79]	[1.17]	[0.74]	[0.20]								

Table 2.5: **Spot and interest rate components for portfolios sorted on dollar factor betas.** This table reports spot and interest rate components of monthly rebalanced portfolios sorted on Q and P dollar factor betas at horizons of 1-12 months. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios P_1, P_2, and P_3. The investor buys (sells) each portfolio when the average forward discount is negative (positive). For Q-beta sorted portfolios, the length of the holding period and the maturity of the options are the same. The dollar factor beta for a currency i is computed according to (2.13)—i.e. as the covariance between innovations in currency i and the dollar portfolio normalized by the variance of the dollar portfolio, defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-beta is computed using a 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

Panel A: Q-betas								
	$\frac{\Delta S_{t+m}}{S_t}$				Δi_{t+m}			
Horizon	P_1	P_2	P_3	$P_3 - P_1$	P_1	P_2	P_3	$P_3 - P_1$
1 mo	1.27	2.08	3.62	2.35*	0.61	1.17	1.61	1.00
	[0.78]	[0.94]	[1.65]	[1.78]				
2 mo	1.01	0.79	2.94	1.93	0.56	1.24	1.59	1.03
	[0.58]	[0.37]	[1.33]	[1.19]				
3 mo	0.60	0.86	1.66	1.06	0.53	1.16	1.65	1.12
	[0.34]	[0.36]	[0.69]	[0.58]				
6 mo	0.47	1.13	0.29	-0.18	0.43	1.09	1.77	1.34
	[0.26]	[0.47]	[0.12]	[0.10]				
9 mo	0.62	1.46	0.10	-0.51	0.38	0.96	1.89	1.50
	[0.36]	[0.66]	[0.04]	[-0.34]				
12 mo	0.76	1.48	-0.03	-0.79	0.32	0.86	2.00	1.67
	[0.46]	[0.73]	[-0.01]	[-0.67]				
Panel B: P-betas								
	$\frac{\Delta S_{t+m}}{S_t}$				Δi_{t+m}			
Horizon	P_1	P_2	P_3	$P_3 - P_1$	P_1	P_2	P_3	$P_3 - P_1$
1 mo	1.97	3.21	1.79	-0.18	0.62	1.03	1.74	1.12
	[1.15]	[1.48]	[0.79]	[-0.11]				
2 mo	1.14	2.62	0.98	-0.16	0.62	1.03	1.74	1.12
	[0.70]	[1.28]	[0.41]	[-0.08]				
3 mo	0.83	2.07	0.23	-0.60	0.61	1.02	1.71	1.09
	[0.49]	[0.91]	[0.09]	[-0.29]				
6 mo	0.57	1.87	-0.56	-1.14	0.57	1.01	1.71	1.14
	[0.32]	[0.85]	[-0.22]	[-0.44]				
9 mo	0.73	1.81	-0.36	-1.08	0.54	1.00	1.68	1.14
	[0.42]	[0.84]	[-0.15]	[-0.48]				
12 mo	0.86	1.47	-0.13	-0.99	0.52	0.99	1.67	1.14
	[0.52]	[0.71]	[-0.05]	[-0.48]				

Table 2.6: **Portfolio sorts on dollar factor exposure with mixed betas.** This table reports annualized means, standard deviations and Sharpe ratios of excess returns of monthly rebalanced portfolios sorted on mixed dollar factor betas for horizons of 1-12 months. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios P_1, P_2, and P_3. The investor buys (sells) each portfolio when the average forward discount is negative (positive). Panel A reports the results for betas based on correlations estimated using a 252-day rolling window and Q-variances derived via expressions (2.6)-(2.7). Panel B reports results for betas based on Q-correlations and P-variances (252-day rolling window variances). The maturity of the options used to compute the Q-variances matches the holding period of the forward contracts. The dollar factor beta for a currency i is computed according to (2.13)—i.e. as the covariance between innovations in currency i and the dollar factor normalized by the variance of the dollar factor, which is defined as an equally weighted basket of all foreign currencies against the U.S. dollar (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the comprise 216 monthly observations.

Panel A: Betas using P-correlations and Q-variances												
	Mean				Std				Sharpe Ratio			
Horizon	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1
1 mo	1.97	3.87*	4.50**	2.53	6.88	9.99	10.70	8.33	0.29	0.39	0.42	0.30
	[1.08]	[1.71]	[2.13]	[1.54]								
2 mo	1.53	2.56	4.04*	2.51	7.16	10.40	10.47	7.43	0.21	0.25	0.39	0.34
	[0.84]	[1.17]	[1.90]	[1.63]								
3 mo	1.29	2.20	2.98	1.69	7.33	10.55	10.71	7.45	0.18	0.21	0.28	0.23
	[0.70]	[0.89]	[1.29]	[1.07]								
6 mo	0.99	2.37	1.81	0.82	7.98	10.99	11.40	7.60	0.12	0.22	0.16	0.11
	[0.53]	[0.97]	[0.75]	[0.55]								
9 mo	1.17	2.58	1.66	0.49	8.19	10.98	11.52	7.35	0.14	0.24	0.14	0.07
	[0.65]	[1.12]	[0.72]	[0.37]								
12 mo	1.31	2.45	1.63	0.32	8.45	11.02	11.62	7.54	0.16	0.22	0.14	0.04
	[0.77]	[1.08]	[0.73]	[0.25]								
Panel B: Betas using Q-correlations and P-variances												
	Mean				Std				Sharpe Ratio			
Horizon	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1	P_1	P_2	P_3	P_3-P_1
1 mo	2.62	3.54	4.12*	1.50	6.76	9.77	10.82	8.11	0.39	0.36	0.38	0.19
	[1.59]	[1.59]	[1.81]	[1.05]								
2 mo	1.52	3.11	3.33	1.81	7.11	9.67	11.03	7.75	0.21	0.32	0.30	0.23
	[0.91]	[1.49]	[1.42]	[1.10]								
3 mo	1.04	2.66	2.62	1.57	7.32	9.69	11.45	8.02	0.14	0.27	0.23	0.20
	[0.57]	[1.14]	[1.02]	[0.92]								
6 mo	0.89	2.68	1.57	0.68	7.82	10.41	12.26	8.12	0.11	0.26	0.13	0.08
	[0.49]	[1.11]	[0.60]	[0.43]								
9 mo	0.79	3.20	1.70	0.91	7.94	10.35	12.32	7.48	0.10	0.31	0.14	0.12
	[0.47]	[1.43]	[0.68]	[0.68]								
12 mo	0.89	2.99	1.86	0.97	7.95	10.70	11.98	6.90	0.11	0.28	0.16	0.14
	[0.56]	[1.38]	[0.81]	[0.85]								

Table 2.7: **Mean prediction errors with constant price of conditional dollar risk.** This table shows model prediction errors for Q and P-beta-sorted portfolios based on a fixed price of conditional dollar risk. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios P_1, P_2, and P_3. The investor buys (sells) each portfolio when the average forward discount is negative (positive). For each portfolio, the mean model predicted excess return and the mean realized excess return are reported. The model predicted portfolio excess returns are computed as the portfolio's beta times the mean excess return on the conditional dollar factor (the price of conditional dollar risk). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed from 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

Panel A: Mean Excess Returns based on Q-betas (%)						
	Model Prediction			Realized		
Horizon	P_1	P_2	P_3	P_1	P_2	P_3
1 mo	1.99	3.36	4.18	1.88	3.24	5.23
2 mo	0.95	1.69	2.10	1.57	2.03	4.53
3 mo	0.63	1.13	1.40	1.13	2.02	3.32
6 mo	0.31	0.57	0.70	0.89	2.22	2.05
9 mo	0.21	0.38	0.47	1.00	2.42	1.99
12 mo	0.15	0.28	0.35	1.08	2.34	1.97
Panel B: Mean Excess Returns based on P-betas (%)						
	Model Prediction			Realized		
Horizon	P_1	P_2	P_3	P_1	P_2	P_3
1 mo	1.74	3.29	4.26	2.59	4.23	3.53
2 mo	0.87	1.65	2.13	1.76	3.65	2.72
3 mo	0.58	1.10	1.42	1.44	3.09	1.93
6 mo	0.29	0.55	0.71	1.14	2.89	1.14
9 mo	0.19	0.37	0.48	1.27	2.81	1.33
12 mo	0.14	0.28	0.36	1.39	2.46	1.54

Table 2.8: **Model prediction errors.** This table reports average model prediction errors for the Q and P-beta-sorted portfolios. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios P_1, P_2, and P_3. The investor buys (sells) each portfolio when the average forward discount is negative (positive). Each month, the prediction error is computed as the difference between the realized and the model predicted return. The model predicted return is computed as ex-ante portfolio beta times the realized excess return on the conditional dollar factor. Columns 1-4 report the mean squared prediction errors (MSEs) in basis points for each portfolio and the mean MSE across all portfolios. Columns 5-8 report the mean error (ME) in % for each portfolio as well as the difference in MEs between the high and low-beta portfolio. Column 9 reports the t-statistic for the difference in mean MSEs for the P-beta and Q-beta. The column 10 reports the t-statistic for the difference in mean MSEs for the random walk forecast and the beta forecasts. Panels C and D report the MSEs for the random walk forecast for the Q and P-betas, respectively. The options data are from JP Morgan Dataquery and the exchange rate spot and forward data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

Panel A: Model forecast Q-beta

	MSE (bps)				ME (%)				$\overline{MSE}^P$-$\overline{MSE}^Q$ T-statistic	$\overline{MSE}^B$-$\overline{MSE}^Q$ T-statistic
Horizon	P_1	P_2	P_3	$\overline{P}$	P_1	P_2	P_3	$P_3 - P_1$		
1 mo	13.63	10.13	16.49	13.42	0.14	-0.25	0.58	0.44	2.18	8.02
2 mo	16.43	11.69	15.92	14.68	0.43	-0.78	0.73	0.30	0.02	6.08
3 mo	17.02	12.98	14.80	14.93	0.32	-0.18	0.18	-0.14	0.10	5.44
6 mo	18.21	10.63	14.41	14.42	0.36	0.46	-0.45	-0.81	0.70	4.55
9 mo	17.78	8.98	11.23	12.66	0.34	0.54	-0.56	-0.90	1.46	4.73
12 mo	19.14	8.07	11.54	12.92	0.46	0.46	-0.63	-1.09	1.92	5.48

Panel B: Model forecast P-beta

	MSE (bps)				ME (%)				$\overline{MSE}^B$-$\overline{MSE}^Q$ T-statistic
Horizon	P_1	P_2	P_3	$\overline{P}$	P_1	P_2	P_3	$P_3 - P_1$	
1 mo	17.74	12.05	16.45	15.41	1.17	1.02	-1.20	-2.38	7.67
2 mo	16.41	12.09	15.59	14.70	0.75	1.17	-1.13	-1.87	5.77
3 mo	17.30	11.66	16.23	15.06	0.66	1.14	-1.18	-1.84	4.81
6 mo	20.37	10.37	15.14	15.29	0.55	1.34	-1.37	-1.92	3.69
9 mo	20.24	10.79	13.12	14.72	0.63	1.15	-1.16	-1.79	3.73
12 mo	22.24	12.09	14.05	16.13	0.83	0.79	-0.91	-1.75	4.57

Panel C: Random walk forecast Q-beta

	MSE (bps)			
Horizon	P_1	P_2	P_3	$\overline{P}$
1 mo	44.51	90.95	120.65	85.37
2 mo	50.20	96.52	122.86	89.86
3 mo	52.91	101.74	125.76	93.47
6 mo	62.47	118.24	129.79	103.50
9 mo	64.77	111.51	137.88	104.72
12 mo	68.45	107.79	139.49	105.24

Panel D: Random walk forecast P-beta

	MSE (bps)			
Horizon	P_1	P_2	P_3	$\overline{P}$
1 mo	47.23	95.45	116.67	86.45
2 mo	48.34	98.99	120.90	89.41
3 mo	50.44	99.06	132.04	93.85
6 mo	60.50	105.51	150.07	105.36
9 mo	62.74	109.24	150.90	107.63
12 mo	66.80	111.86	146.01	108.22

Table 2.9: **Beta prediction errors for portfolios.** This table shows beta prediction errors for portfolios sorted on Q and P-betas. Every day, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into three equal-weighted portfolios P_1, P_2, and P_3. For each type of beta, the daily prediction error is computed as the difference between realized and expected portfolio beta at horizons from 6-18 months. The Q-betas used in the portfolio construction match the forecast horizon as closely as possible. For P-betas, the length of the rolling window matches the forecast horizon. In Panels A and B, columns 1-3 report mean prediction errors in % (ME) for each portfolio. Column 4 reports the difference in MEs between P_3 and P_1, and column 5 reports the t-statistic for their significance. In the right panel, Columns 1-3 report the mean squared prediction errors (MSE) in % for each portfolio, column 4 reports the mean MSE across all portfolios, and the final column reports the t-statistic for the difference between mean MSEs based on P-betas and Q-betas. All t-statistics are based on Newey and West (1987) with lags equal to the forecast horizon. Panel C reports the mean expected betas for each portfolio, for each type of beta. The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

	ME (%)				$\overline{ME}^{P_3}$-$\overline{ME}^{P_1}$ T-statistic	MSE (%)				$\overline{MSE}^{P}$-$\overline{MSE}^{Q}$ T-statistic
Panel A: Q-beta Prediction Errors										
Horizon	P_1	P_2	P_3	$P_3 - P_1$		P_1	P_2	P_3	$\bar{P}$	
6 mo	-0.26	0.55	-3.46	-3.20	-1.42	1.03	1.06	1.94	1.34	3.55
9 mo	0.52	0.94	-4.76	-5.28	-2.42	0.89	1.01	1.91	1.27	3.05
12 mo	0.96	1.61	-5.67	-6.62	-2.82	0.78	1.00	1.89	1.22	3.26
15 mo	1.35	1.85	-5.89	-7.24	-2.67	0.76	1.00	1.93	1.23	3.56
18 mo	1.91	1.87	-6.29	-8.21	-2.64	0.84	1.05	2.04	1.31	2.30
Panel B: P-beta Prediction Errors										
6 mo	7.44	1.67	-8.29	-15.73	-2.92	2.15	1.04	2.51	1.90	—
9 mo	8.09	2.34	-9.22	-17.31	-4.02	1.93	1.01	2.49	1.81	—
12 mo	8.48	1.93	-9.18	-17.66	-3.71	1.77	1.19	2.55	1.84	—
15 mo	9.20	1.76	-9.67	-18.88	-3.73	1.91	1.27	2.61	1.93	—
18 mo	9.02	1.64	-9.75	-18.78	-3.78	1.84	1.10	2.33	1.75	—
Panel C: Mean Predicted Betas										
	Mean Q-Beta					Mean P-Beta				
Horizon	P_1	P_2	P_3	$P_3 - P_1$		P_1	P_2	P_3	$P_3 - P_1$	
6 mo	0.65	0.99	1.18	0.54		0.64	0.99	1.16	0.47	
9 mo	0.62	1.00	1.20	0.57		0.66	0.99	1.14	0.48	
12 mo	0.62	1.01	1.20	0.57		0.66	1.00	1.13	0.47	
15 mo	0.62	1.01	1.20	0.59		0.67	1.00	1.13	0.45	
18 mo	0.61	1.01	1.21	0.60		0.68	1.00	1.12	0.44	

Table 2.10: **Q-beta predictions for individual dollar factor betas.** This table reports the results of daily predictive regressions of realized betas, β^P_{t+m}, using Q-betas as predictors for each G10 currency:

$$\beta^P_{t+m} = \gamma^Q_{0,m} + \gamma^Q_{1,m}\beta^Q_t + \varepsilon^Q_{t+m}$$

The forecast horizons are 6-18 months. The realized betas for a forecast horizon of length m are computed from daily rolling window regressions over the time interval $[t, t+m]$. For the 6-12-month forecast horizons, the Q-betas are derived from options with the same maturity as the forecast horizon. At the 15-month and 18-month horizons, the 1-year Q-beta is used as predictor. The options data are from JP Morgan Dataquery and the exchange rate data is obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

		Forecast Horizon				
		6 mo	9 mo	12 mo	15 mo	18 mo
AUD	$\gamma^Q_{1,m}$	0.46***	0.43**	0.41**	0.32*	0.29*
		[3.58]	[2.56]	[2.27]	[1.78]	[1.90]
	$\gamma^Q_{0,m}$	0.52***	0.60***	0.63***	0.72***	0.76***
		[3.02]	[2.84]	[2.78]	[2.73]	[3.05]
	R^2	0.16	0.15	0.16	0.13	0.12
CAD	$\gamma^Q_{1,m}$	0.79***	0.72***	0.75***	0.75***	0.73***
		[10.54]	[8.16]	[8.78]	[8.28]	[7.02]
	$\gamma^Q_{0,m}$	0.04	0.16**	0.15**	0.16**	0.17**
		[0.71]	[2.00]	[2.40]	[2.54]	[2.30]
	R^2	0.54	0.58	0.66	0.70	0.69
CHF	$\gamma^Q_{1,m}$	0.81***	0.79***	0.77***	0.67***	0.67***
		[7.14]	[7.70]	[6.97]	[5.74]	[5.92]
	$\gamma^Q_{0,m}$	0.19	0.18	0.20*	0.31**	0.31**
		[1.62]	[1.57]	[1.68]	[2.24]	[2.25]
	R^2	0.38	0.44	0.44	0.44	0.46
EUR	$\gamma^Q_{1,m}$	0.69***	0.54***	0.49**	0.38*	0.40**
		[5.15]	[2.60]	[2.27]	[1.73]	[1.98]
	$\gamma^Q_{0,m}$	0.32**	0.47**	0.52**	0.63***	0.62***
		[2.35]	[2.22]	[2.30]	[2.70]	[2.87]
	R^2	0.32	0.21	0.20	0.15	0.18
GBP	$\gamma^Q_{1,m}$	0.40**	0.22	0.20	0.13	0.16
		[2.53]	[1.08]	[0.84]	[0.35]	[0.52]
	$\gamma^Q_{0,m}$	0.41***	0.54***	0.55***	0.61***	0.58***
		[3.97]	[3.77]	[4.00]	[2.68]	[2.93]
	R^2	0.10	0.03	0.03	0.01	0.02
JPY	$\gamma^Q_{1,m}$	0.63***	0.52***	0.44**	0.28	0.22
		[4.57]	[2.89]	[2.16]	[1.37]	[1.01]
	$\gamma^Q_{0,m}$	0.07	0.12	0.15	0.22	0.25
		[0.58]	[0.80]	[1.00]	[1.05]	[1.12]
	R^2	0.32	0.25	0.21	0.12	0.08
NOK	$\gamma^Q_{1,m}$	0.26*	0.10	0.01	−0.02	−0.04
		[1.81]	[0.66]	[0.03]	[−0.10]	[−0.14]
	$\gamma^Q_{0,m}$	0.88***	1.05***	1.17***	1.21***	1.23***
		[4.61]	[4.99]	[6.72]	[5.48]	[4.80]
	R^2	0.04	0.01	0.00	0.00	0.00
NZD	$\gamma^Q_{1,m}$	0.29***	0.21*	0.21*	0.20**	0.19***
		[2.69]	[1.83]	[1.82]	[2.04]	[2.92]
	$\gamma^Q_{0,m}$	0.76***	0.83***	0.84***	0.85***	0.86***
		[5.49]	[5.42]	[4.75]	[4.70]	[6.17]
	R^2	0.08	0.05	0.06	0.09	0.09
SEK	$\gamma^Q_{1,m}$	0.45***	0.43***	0.39***	0.30**	0.27*
		[5.26]	[4.37]	[3.46]	[2.15]	[1.78]
	$\gamma^Q_{0,m}$	0.66***	0.67***	0.72***	0.83***	0.87***
		[6.77]	[6.14]	[5.95]	[4.95]	[4.60]
	R^2	0.16	0.13	0.10	0.05	0.04

Table 2.11: ***P*-beta predictions for individual dollar factor betas.** This table reports the results of daily predictive regressions of realized dollar factor betas, β^P_{t+m}, using P-betas as predictors for each G10 currency:

$$\beta^P_{t+m} = \gamma^P_{0,m} + \gamma^P_{1,m}\beta^P_t + \varepsilon^P_{t+m}$$

The forecast horizons are 6-18 months. The realized betas for a forecast horizon of length m are computed from daily rolling window regressions over the time interval $[t, t+m]$, and the P-beta predictions are computed from daily rolling window regressions over the time interval $[t-m, t)$. The exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

		Forecast Horizon				
		6 mo	9 mo	12 mo	15 mo	18 mo
AUD	$\gamma^P_{1,m}$	0.32*	0.29	0.24	0.19	0.21
		[1.68]	[1.17]	[0.55]	[0.46]	[0.97]
	$\gamma^P_{0,m}$	0.68**	0.72**	0.77**	0.83***	0.81***
		[2.36]	[2.28]	[2.56]	[3.26]	[5.63]
	R^2	0.11	0.09	0.07	0.04	0.06
CAD	$\gamma^P_{1,m}$	0.62**	0.67**	0.72*	0.70	0.67
		[2.53]	[2.35]	[1.79]	[1.46]	[1.13]
	$\gamma^P_{0,m}$	0.21	0.18	0.17	0.19	0.21
		[0.98]	[0.73]	[0.76]	[0.66]	[0.52]
	R^2	0.40	0.52	0.62	0.60	0.57
CHF	$\gamma^P_{1,m}$	0.51	0.55	0.58	0.58	0.58
		[0.98]	[1.09]	[0.82]	[0.60]	[0.54]
	$\gamma^P_{0,m}$	0.49	0.45	0.43	0.42	0.42
		[0.68]	[0.64]	[0.45]	[0.32]	[0.28]
	R^2	0.26	0.31	0.33	0.33	0.33
EUR	$\gamma^P_{1,m}$	0.59*	0.43	0.33	0.24	0.17
		[1.95]	[0.77]	[0.58]	[0.39]	[0.24]
	$\gamma^P_{0,m}$	0.44	0.61	0.72	0.82	0.90
		[1.15]	[0.89]	[1.00]	[1.06]	[0.95]
	R^2	0.35	0.19	0.12	0.06	0.03
GBP	$\gamma^P_{1,m}$	0.49*	0.46	0.41	0.36	0.35
		[1.82]	[0.48]	[0.28]	[0.38]	[0.47]
	$\gamma^P_{0,m}$	0.36*	0.38	0.42	0.46	0.46
		[1.73]	[0.47]	[0.40]	[0.81]	[1.09]
	R^2	0.22	0.21	0.18	0.15	0.16
JPY	$\gamma^P_{1,m}$	0.52	0.45	0.38	0.28	0.17
		[1.47]	[0.96]	[0.48]	[0.27]	[0.13]
	$\gamma^P_{0,m}$	0.19	0.21	0.23	0.26	0.30
		[0.48]	[0.47]	[0.41]	[0.41]	[0.31]
	R^2	0.28	0.23	0.18	0.10	0.04
NOK	$\gamma^P_{1,m}$	0.31	0.30	0.32	0.35	0.40
		[1.48]	[0.95]	[1.08]	[1.33]	[0.85]
	$\gamma^P_{0,m}$	0.81***	0.82**	0.81**	0.77**	0.71
		[3.01]	[2.43]	[2.34]	[2.24]	[1.09]
	R^2	0.10	0.09	0.10	0.13	0.17
NZD	$\gamma^P_{1,m}$	0.13	0.11	0.13	0.11	0.11
		[0.48]	[0.29]	[0.29]	[0.22]	[0.27]
	$\gamma^P_{0,m}$	0.91***	0.92*	0.91*	0.93**	0.93***
		[3.05]	[1.76]	[1.73]	[2.29]	[3.05]
	R^2	0.02	0.01	0.02	0.01	0.02
SEK	$\gamma^P_{1,m}$	0.57***	0.56***	0.51*	0.46	0.40
		[3.78]	[2.69]	[1.70]	[1.20]	[0.92]
	$\gamma^P_{0,m}$	0.50***	0.51**	0.58*	0.64	0.71
		[2.81]	[2.46]	[1.95]	[1.64]	[1.50]
	R^2	0.31	0.31	0.26	0.21	0.16

Table 2.12: **Adjusted R^2s for forecasts of realized betas using P-betas and Q-betas as predictors.** This table reports the results for regressions of realized dollar factor betas, β^P_{t+m}, regressed on Q-beta and P-beta forecasts for each G10 currency:

$$\beta^P_{t+m} = \gamma^P_{0,m} + \gamma^P_{1,m}\beta^P_t + \gamma^Q_{1,m}\beta^Q_t + \varepsilon^{PQ}_{t+m}$$

The forecast horizons are $6, 9, 12, 15$ and 18 months. For the 6-12 month forecast horizons, the Q-beta predictors are derived from options with the same maturity as the forecast horizon. At the 15-month and 18-month horizons, the 1-year Q-beta is used as predictor. The P-betas are based on daily rolling window regressions with the same length as the forecast horizon. The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 4659 daily observations.

	Adjusted R^2				
	6 mo	9 mo	12 mo	15 mo	18 mo
AUD	0.16	0.13	0.11	0.09	0.08
CAD	0.54	0.58	0.68	0.69	0.68
CHF	0.38	0.41	0.42	0.44	0.45
EUR	0.39	0.28	0.24	0.24	0.23
GBP	0.22	0.21	0.18	0.15	0.16
JPY	0.33	0.27	0.23	0.20	0.17
NOK	0.10	0.09	0.10	0.13	0.17
NZD	0.08	0.05	0.05	0.05	0.05
SEK	0.32	0.34	0.30	0.26	0.22

2.9 Additional Tables

Table 2.13: **Descriptive statistics for G10 currencies.** This table reports descriptive statistics for the G10 currencies. For each foreign currency the mean, standard deviation and Sharpe ratio are reported for the U.S. dollar excess return at horizons of 1-12 months. All quantities are reported in percentage terms. Panel A reports the mean annualized currency excess return. Panel B reports annualized forward discounts (interest rate differentials provided CIP holds). Panel C reports annualized standard deviations of currency excess returns. Panel D reports annualized Sharpe ratios for the currency excess returns. Before January 1999, the Euro is replaced by the Deutsche Mark. The data are obtained from Reuters via Datastream and the sample period is from 1998-2016 and the data comprise 4659 daily observations.

Panel A: Mean Excess Returns

Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	2.91	0.82	0.63	-0.23	-0.36	-1.31	0.61	3.84	-0.43
2 mo	2.86	0.75	0.65	-0.20	-0.36	-1.22	0.61	3.85	-0.42
3 mo	2.80	0.72	0.71	-0.19	-0.25	-1.26	0.61	3.82	-0.40
6 mo	2.97	0.77	0.79	-0.21	-0.16	-1.27	0.55	3.93	-0.39
9 mo	3.09	0.80	0.69	-0.36	-0.10	-1.41	0.43	4.05	-0.41
12 mo	3.04	0.88	0.63	-0.46	-0.02	-1.77	0.38	4.01	-0.38

Panel B: Forward Discounts ($f_{t,t+m} - s_t$)

Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	-2.27	-0.18	1.71	0.46	-0.81	2.42	-1.19	-2.72	0.14
2 mo	-2.25	-0.18	1.67	0.43	-0.81	2.40	-1.18	-2.70	0.13
3 mo	-2.22	-0.18	1.65	0.42	-0.80	2.40	-1.17	-2.67	0.12
6 mo	-2.20	-0.18	1.64	0.40	-0.80	2.44	-1.12	-2.66	0.11
9 mo	-2.18	-0.19	1.63	0.40	-0.79	2.48	-1.09	-2.65	0.07
12 mo	-2.17	-0.20	1.62	0.41	-0.78	2.54	-1.07	-2.64	0.04

Panel C: Standard Deviations

Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	12.96	9.23	10.63	10.30	8.61	10.93	11.54	13.48	11.42
2 mo	13.46	8.98	10.32	10.40	8.85	10.76	11.56	13.47	11.60
3 mo	13.51	9.02	10.13	10.57	9.13	11.27	12.04	13.45	11.82
6 mo	14.21	9.25	10.04	10.85	9.85	11.32	12.56	14.30	12.70
9 mo	14.25	9.19	9.79	11.00	9.64	10.70	12.74	14.60	13.11
12 mo	14.11	9.12	10.00	11.26	9.36	11.22	12.72	14.61	13.15

Panel D: Sharpe Ratios

Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	22.46	8.86	5.97	-2.25	-4.23	-11.97	5.30	28.46	-3.78
2 mo	21.26	8.32	6.28	-2.22	-4.10	-11.32	5.29	28.61	-3.65
3 mo	20.72	7.93	6.99	-1.79	-2.71	-11.20	5.08	28.43	-3.40
6 mo	20.90	8.37	7.86	-1.96	-1.58	-11.26	4.35	27.49	-3.09
9 mo	21.66	8.71	7.01	-3.28	-1.07	-13.22	3.37	27.75	-3.12
12 mo	21.57	9.60	6.28	-4.07	-0.27	-15.80	2.98	27.43	-2.91

Table 2.14: **Descriptive statistics of portfolios sorted on dollar factor betas.** This table shows the means, standard deviations and Sharpe ratios of excess returns of monthly rebalanced portfolios sorted on Q and P-betas to the conditional dollar factor at horizons of 1-12 months. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into two equal-weighted portfolios P_1 and P_2. The investor buys (sells) each portfolio when the average forward discount is negative (positive). For Q-beta-sorted portfolios, the length of the holding period and the maturity of the options are the same. The dollar factor beta for a currency i is computed according to (2.13).The m-month Q-betas are computed using the model-free measures of covariance/variance implied from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed using a 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the comprise 216 monthly observations.

Panel A: Q-betas									
	Mean			Std			Sharpe Ratio		
Horizon	P_1	P_2	$P_2 - P_1$	P_1	P_2	$P_2 - P_1$	P_1	P_2	$P_2 - P_1$
	2.12	5.11	2.99	7.40	10.54	6.53	0.29	0.48	0.46
	[1.19]	[2.40]	[2.41]						
	1.33	4.43	3.10	7.70	10.68	6.44	0.17	0.41	0.48
	[0.67]	[1.78]	[2.27]						
	1.20	3.35	2.14	7.91	10.87	6.36	0.15	0.31	0.34
	[0.52]	[1.23]	[1.57]						
	1.19	2.39	1.21	8.92	10.96	5.53	0.13	0.22	0.22
	[0.30]	[0.62]	[1.01]						
	1.52	2.15	0.63	8.86	11.40	5.60	0.17	0.19	0.11
	[0.43]	[0.60]	[0.60]						
	1.67	1.95	0.27	8.77	11.71	5.83	0.19	0.17	0.05
	[0.57]	[0.58]	[0.26]						
Panel B: P-betas									
	Mean			Std			Sharpe Ratio		
Horizon	P_1	P_2	$P_2 - P_1$	P_1	P_2	$P_2 - P_1$	P_1	P_2	$P_2 - P_1$
	2.86	4.19	1.32	7.40	10.60	6.68	0.39	0.39	0.20
	[1.57]	[1.96]	[1.04]						
	2.09	3.48	1.39	7.65	10.76	6.56	0.27	0.32	0.21
	[1.04]	[1.36]	[0.93]						
	1.71	2.71	1.00	7.74	11.18	6.82	0.22	0.24	0.15
	[0.75]	[0.92]	[0.62]						
	1.51	1.99	0.48	8.31	11.95	6.90	0.18	0.17	0.07
	[0.45]	0.40]	0.22]						
	1.62	2.03	0.41	8.48	11.96	6.34	0.19	0.17	0.06
	[0.54]	0.46]	0.23]						
	1.59	2.05	0.46	8.83	11.72	6.04	0.18	0.18	0.08
	[0.56]	[0.53]	[0.30]						

Table 2.15: **Spot and forward components of portfolios sorted on dollar factor betas.** This table shows means, standard deviations and Sharpe ratios of excess returns of monthly rebalanced portfolios sorted on the Q and P-beta to the conditional dollar factor at horizons of 1-12 months. Each month, for each type of beta separately, the currencies are ranked in ascending order based on their dollar factor betas and allocated into two equal-weighted portfolios P_1 and P_2. The investor buys (sells) each portfolio when the average forward discount is negative (positive). The dollar factor beta for a currency i is computed according to (2.13)—i.e. as the covariance between innovations in currency i and the dollar portfolio normalized by the variance of the dollar portfolio, defined defined as an equally weighted basket of all foreign currencies vs. U.S. dollar (excluding currency i). The m-month Q-betas are computed using the model-free measures of covariance and variance implied out from currency options with m-month maturity (expressions: (2.6)-(2.7)). The P-betas are computed using 252-day rolling window regressions of daily innovations in the exchange i against daily innovations in the dollar factor (excluding currency i). The options data are from JP Morgan Dataquery and the exchange rate data are obtained from Reuters through Datastream. The sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

Panel A: Q-betas						
	$\frac{\Delta S_{t+m}}{S_t}$			Δi_{t+m}		
Horizon	P_1	P_2	$P_2 - P_1$	P_1	P_2	$P_2 - P_1$
	1.38	3.49	2.11	0.74	1.62	0.88
	[0.78]	[1.65]	[1.71]			
	0.56	2.86	2.30	0.77	1.57	0.80
	[0.28]	[1.16]	[1.74]			
	0.49	1.74	1.25	0.72	1.61	0.89
	[0.21]	[0.64]	[0.94]			
	0.57	0.70	0.14	0.62	1.69	1.07
	[0.15]	[0.18]	[0.12]			
	0.97	0.42	-0.54	0.56	1.73	1.17
	[0.30]	[0.12]	[-0.52]			
	1.67	1.95	0.27	0.48	1.79	1.31
	[1.20]	[0.16]	[-1.04]			
Panel B: P-betas						
	$\frac{\Delta S_{t+m}}{S_t}$			Δi_{t+m}		
Horizon	P_1	P_2	$P_2 - P_1$	P_1	P_2	$P_2 - P_1$
	2.20	2.47	0.27	0.66	1.71	1.05
	[1.24]	[1.16]	[0.21]			
	1.43	1.77	0.34	0.66	1.71	1.05
	[0.73]	[0.70]	[0.23]			
	1.05	1.03	-0.02	0.66	1.68	1.02
	[0.48]	[0.35]	[-0.01]			
	0.88	0.31	-0.57	0.63	1.68	1.05
	[0.28]	[0.061]	[-0.26]			
	1.01	0.37	-0.64	0.61	1.66	1.05
	[0.36]	[0.08]	[-0.36]			
	0.99	0.41	-0.58	0.60	1.64	1.04
	[0.37]	[0.11]	[-0.39]			

Essay 3

Systematic Currency Volatility Risk Premia

Systematic Currency Volatility Risk Premia

Andreas Bang Nielsen *

Abstract

I show that volatility risk of the dollar factor—an equally weighted basket of developed U.S. dollar exchange rates—carries a significant risk premium and that it is priced in the cross-section of currency volatility excess returns. The dollar factor volatility risk premium is negative on average with an upward sloping and concave term structure. Consistent with this pattern, I find that dollar factor volatility risk is most significantly priced in the cross-section of volatility excess returns at shorter maturities. A trading strategy that sells (buys) volatility insurance on currencies with high (low) exposure to dollar factor volatility risk delivers high mean excess returns and Sharpe ratios. At shorter maturities, the profitability of this strategy cannot be explained by exposure to traditional currency factors, equity factors, or currency volatility carry factors.

Keywords: Volatility risk premia, factor models, foreign exchange volatility, currency options, option-implied betas

JEL Codes: G12, G15, F31

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3.1 Introduction

It is well-documented that currency volatility risk premia—differences between realized and expected risk-neutral volatilities—are, on average, significantly negative. This indicates that currency volatility risk is priced by investors and that it has been profitable to sell insurance on realized volatility. Currency volatility risk premia are strong predictors of currency excess returns, but very little is known about the underlying risk sources that drive currency volatility risk premia themselves. In a standard asset pricing model—for instance a linear factor model—the negative currency volatility risk premia must reflect compensation exclusively for taking exposure to systematic risk factors, or in other words, idiosyncratic volatility risk is diversifiable and thus irrelevant for the pricing of volatility insurance.

In this paper, I develop a method that separates the contribution of idiosyncratic and systematic variance risk to the total variance of exchange rates which I use to investigate if the share of systematic variance risk accounts for the cross-section of currency volatility risk premia. The main hypothesis is that the large volatility excess returns for individual currencies documented in previous research are primarily attributable to the systematic component, since the idiosyncratic component is diversifiable.

More specifically, I use a linear factor model where the dollar factor—an equally weighted portfolio of foreign currencies versus the U.S. dollar—is the systematic factor driving currency excess returns. Based on this model, I decompose variances of exchange rates into a systematic component, i.e., variance induced by exposure to the dollar factor, and an idiosyncratic variance component (the residual part). The systematic variance component is calculated for each exchange rate based on a model-free technique that uses currency options to measure the dollar factor variance and risk exposures (betas). The systematic variance components are thus forward-looking and cover the same time horizons as the option maturities.

The main finding of the paper is that I identify a highly significant negative relation between volatility excess returns and the share of systematic variance risk. As a result, a long-short portfolio that buys (sells) volatility protection on currencies with the smallest (largest) shares of systematic variance has delivered significant mean excess returns and high Sharpe ratios. This risk premium is largest at shorter maturities and is not subsumed by traditional currency factors, equity factors, or the volatility carry factor proposed by Della Corte et al. (2017).

I use the dollar factor to study volatility risk premia in currencies because Verdelhan (2017) shows that the dollar factor accounts for a substantial part of the variation of bilateral exchange rates. Other natural candidate factors, i.e., the conditional and unconditional carry factors of Lustig et al. (2011), have a relatively small contribution in comparison for most exchange rates. The findings of Verdelhan (2017) therefore suggest that the carry factors, although priced in the cross-section of currency excess returns, play a smaller role as a source of exchange rate volatility. In contrast, the dollar factor appears intimately connected to foreign exchange volatility risk, even for exchange rates that do not involve the U.S. dollar. The methodology that I propose, however, can be used for any currency factor model including multi-factor models.

While most studies on volatility risk premia focus only on the difference between one-month realized and risk-neutral volatility, I also investigate if systematic variance risk is priced at longer horizons. Specifically, I empirically test if exposure to dollar factor variance risk is priced in the cross-section of excess returns on volatility swaps and forward volatility agreements (FVAs). The volatility swap pays at maturity the difference between realized volatility over the life of the contract and the spot implied volatility fixed at the inception of the contract (the swap rate). An FVA is a forward contract on spot implied volatility, i.e., it allows the holder to enter into a volatility swap at future point in time at a swap rate known today. The FVAs therefore contain information about the term structure of volatility risk premia, while volatility swaps are informative about the short-term volatility risk premia.

The market for currency volatility derivatives is enormous. For example, according to Della Corte, Kozhan, and Neuberger (2017), the market for FVAs has a daily average turnover of 254 billion USD and notional amounts outstanding of 11.7 trillion USD, as of April 2016, underlining a large demand from market participants to hedge against future currency volatility shocks.

Since my main objective is to study if dollar factor variance risk is an important source of volatility excess returns, I begin the empirical analysis by documenting some key properties of the volatility risk premia of the dollar factor. I show that the term structure of dollar factor volatility is, on average, upward sloping and concave, i.e., it is particularly steep at the short end and flattens out at longer maturities. As a consequence, the average dollar factor volatility risk premia are negative with an upward sloping term structure. Intuitively, this implies that investors are willing to pay a higher price for insuring against short-term systematic variance risk compared to long-term systematic variance risk.

In order to test if dollar factor variance risk is priced in the cross-section volatility excess returns, I define a measure for the share of systematic variance risk (SYS) for each exchange rate as its systematic variance component divided by its total objective variance. The systematic variance component is the product of the dollar factor variance and the squared dollar factor beta, both of which are inferred from cross-pair currency options as in Nielsen (2017). I justify that the SYS measure identifies systematic variance risk by showing that the cross-section of expected variance excess returns is solely explained by SYS under two assumptions. First, there is a negative variance risk premium on the systematic factor. Second, idiosyncratic variance risk premia are equal across exchange rates per unit volatility (e.g., they are all zero).

Each month, I allocate volatility swaps into portfolios based on their SYS measures and document a strong negative relation between SYS and the portfolio excess returns. For example, at the 1-month horizon, a long-short portfolio that buys (sells) volatility swaps with low (high) SYS measures delivers a significant monthly mean excess return of 4.47% and an annualized Sharpe ratio of 0.71. The mean excess returns of the long-short portfolios, however, decrease in maturity and are only significant for maturities of up to six months. Furthermore, I find that the share of systematic variance is priced in the cross-section of FVA excess returns, in particular at shorter maturities. For example, when the forward contract and its underlying volatility both have a 1-month maturity, a long-short portfolio that buys (sells) FVAs with a low (high) share of systematic variance has a monthly mean excess return of 2.73% and an annualized Sharpe ratio of 0.95. At shorter maturities, the risk premia on the long-short portfolios based on systematic variance risk cannot be explained by exposure to the conditional dollar factor of Lustig et al. (2014), the G10 HML carry factor of Lustig et al. (2011), the G10 FX momentum factor of Asness et al. (2013), or the volatility carry factor of Della Corte et al. (2017). At longer maturities of the FVAs, however, the risk premia of the systematic variance risk portfolios are subsumed by the volatility carry factor. However, this is partly explained by the fact that the volatility risk premia decrease in maturity.

There is a growing literature that studies volatility risk premia in currency markets (Della Corte, Kozhan, and Neuberger, 2017; Della Corte, Sarno, and Tsiakas, 2011; Londono and Zhou, 2017). But, to the best of my knowledge, there are no previous papers that directly estimate systematic versus idiosyncratic volatility risk premia in currencies, which is the main objective of this paper. Della Corte, Sarno, and Tsiakas (2011) show that forward volatility prices are biased predictors of future spot implied volatility. Building on this idea, Della Corte, Kozhan, and Neuberger (2017) document high mean excess returns on a long-short portfolio that buys (sells) FVAs that trade at a high (low) forward volatility discount, i.e., a volatility analogue to the traditional carry trade.

While Della Corte, Kozhan, and Neuberger (2017) focus on constructing portfolios of longer-term FVAs based on the slope of the volatility term structure, the systematic variance risk that I identify is most significantly priced for FVAs at shorter maturities and for volatility swaps. In general, at shorter maturities of the forward contract, the volatility carry factor explains only a small part of the systematic forward volatility risk premia that I identify from FVAs. Moreover, the volatility carry factor based on volatility swaps does not explain the excess returns of the long-short portfolio of volatility swaps that I construct using the share of systematic variance measure.

The finding that systematic variance risk is priced in currency volatility excess returns appears consistent with a general phenomenon of investors requiring significantly larger risk premia for holding systematic variance vis-à-vis diversifiable variance risk (Bollerslev, Tauchen, and Zhou, 2009; Carr and Wu, 2008; Duan and Wei, 2008). For instance, Carr and Wu (2008) find that variance risk premia on individual stocks are mostly explained by their exposure to the S&P 500 index variance, and Duan and Wei (2008) identify a positive relationship between the share of systematic variance for individual stocks and the level and slope of their implied volatility curve. Christoffersen, Fournier, and Jacobs (2017) construct a no-arbitrage option pricing model with a systematic factor in equity returns, and they uncover a positive relation between a stock's beta with respect to the S&P 500 index and the level and slope of its implied volatility curve.

3.2 Systematic Variance Risk Premia

The central hypothesis of this paper is that exposure to systematic currency variance risk explains the cross-section of currency variance risk premiums. To this end, I study two types of derivatives that allow investors to gain direct exposure to the future direction of variance, variance swaps and forward variance agreements. I assume that there is a factor structure in the exchange rates in which the dollar factor of Lustig et al. (2011, 2014) is the systematic factor driving currency excess returns, and I then deduce a simple theoretical relation between excess returns for providing variance insurance and exposure to dollar factor variance risk (systematic variance risk).

In what follows, I present the theory and its testable implications based upon variance swaps and forward variance agreements, and not volatility derivatives, because decompositions of variances into systematic and idiosyncratic components are exact, while similar decompositions for volatilities are convexity biased due to Jensen's inequality. But in the empirical analysis, I study volatility excess returns to be consistent with common market conventions and the academic and practitioner literature. However, whether I use variance excess returns or volatility excess returns are not important for the main results of the paper.

3.2.1 Variance Swaps

A variance swap entered into at time t with maturity τ is a contractual agreement in which the buyer pays at time τ a time t fixed price for annualized variance over $[t, t+\tau]$, $SVAR_t^\tau$ (the variance swap rate), and receives the realized annualized variance over the same interval at time τ, $RVAR_t^\tau$. A variance swap with a notional value of N thus delivers the following payoff at maturity:

$$N \cdot (RVAR_t^\tau - SVAR_t^\tau) \tag{3.1}$$

The variance swap rate is determined such that the variance swap at inception has a zero net market value. As a result, the variance swap rate equals the time t risk-neutral expectation of the realized variance over the life of the contract (e.g., Carr and Wu (2008)):

$$SVAR_t^{\tau} = E_t^Q \left(RVAR_t^{\tau} \right) \tag{3.2}$$

The payoff at maturity specified in (3.1) does not adjust for the level of the variance. Thus, in order to make returns more easily comparable across currencies at different variance levels, I compute the variance swap excess returns by normalizing with the current spot variance:

$$RX_t^{\tau} = \frac{RVAR_t^{\tau} - SVAR_t^{\tau}}{RVAR_{t-\tau}^{t}} \tag{3.3}$$

This method for deriving excess returns is similar to the most commonly used definition of excess returns on currency forward contracts (Della Corte, Ramadorai, and Sarno, 2016; Fama, 1984; Menkhoff, Sarno, Schmeling, and Schrimpf, 2012a), with $RVAR_t^{\tau}$ as the spot price and $SVAR_t^{\tau}$ as the forward price of the underlying instrument. Normalizing with the current spot implied variance, rather than the realized variance, makes negligible differences for the analysis. Furthermore, the main results of paper also hold for payoffs, i.e., without normalizing. This will be discussed further in the empirical analysis.

3.2.2 Forward Variance Agreements

In a forward variance agreement (FVARA), the buyer agrees to purchase future implied spot variance, covering a given interval, at a fixed price set upon entry of the contract. An FVARA is therefore a bet on the direction of future implied variance and does not directly depend on realized variance of the underlying exchange rate. More precisely, the FVARA entered into at time t with maturity τ_1 written on spot variance covering $[t+\tau_1, t+\tau]$, where $\tau_1 + \tau_2 = \tau$, is defined as follows. At time t, the holder of the FVARA agrees to pay the annualized forward variance price $FVAR_{t,\tau_1}^{\tau_2}$ at time τ_1 against simultaneously receiving the annualized spot variance covering $[t+\tau_1, t+\tau]$, $SVAR_{t+\tau_1}^{\tau_2}$. Therefore, the time τ_1-payoff for the buyer of the FVARA with a notional of N is given by

$$N \cdot \left(SVAR_{t+\tau_1}^{\tau_2} - FVAR_{t,\tau_1}^{\tau_2} \right) \tag{3.4}$$

Since the FVA has a zero net market value at entry, the forward variance price must equal the risk-neutral expected spot implied variance (Della Corte, Kozhan, and Neuberger (2017)):

$$FVAR_{t,\tau_1}^{\tau_2} = E_t^Q\left(SVAR_{t+\tau_1}^{\tau_2}\right) \tag{3.5}$$

I define the τ_1-period excess return of the FVARA by normalizing its payoff with the current spot implied variance (with the same maturity as the underlying spot implied variance):

$$RX_{t,\tau_1}^{\tau_2} = \frac{SVAR_{t+\tau_1}^{\tau_2} - FVAR_{t,\tau_1}^{\tau_2}}{SVAR_t^{\tau_2}} \tag{3.6}$$

A useful way to think of the FVARA is that it gives the buyer the right to enter into a variance swap that covers a future time period at a variance swap rate known today. Hence, if investors are willing to pay a premium to hedge against a rise in the future variance swap rate, which is typically the case, then the mean FVARA excess return is negative.

3.2.3 Forward Variance Price

To derive the forward variance price, I express the spot variance that covers $[t, t+\tau]$ in terms of two variances that cover disjoint time intervals. The first time interval spans from the forward contract initiation and until its expiry ($[t, t+\tau_1]$), and the second time interval spans from the forward contract expiry to the end point of the volatility coverage period ($(t+\tau_1, t+\tau]$):

$$\begin{aligned} \overline{SVAR}_t^{\tau} &= E_t^Q\left(\overline{RVAR}_t^{\tau}\right) = E_t^Q\left(\overline{RVAR}_t^{t+\tau_1}\right) + E_t^Q\left(\overline{RVAR}_{t+\tau_1}^{t+\tau}\right) \\ &= \overline{SVAR}_t^{\tau_1} + \overline{FVAR}_{t,\tau_1}^{\tau_2} \end{aligned} \tag{3.7}$$

where $\overline{RVAR}$ and $\overline{SVAR}$ are non-annualized realized and implied variances, respectively. $\overline{FVAR}_{t,\tau_1}^{\tau_2}$ is the time t risk-neutral expectation of non-annualized realized variance over $[t+\tau_1, t+\tau]$, i.e., the t-forward price with maturity τ_1 on variance covering $[t+\tau_1, t+\tau]$. By annualizing the variances in (3.7), we get that the annualized forward variance price is given by:

$$FVAR_{t,\tau_1}^{\tau_2} = \frac{\tau}{\tau_2} SVAR_t^{\tau} - \frac{\tau_1}{\tau_2} SVAR_t^{\tau_1} \tag{3.8}$$

3.2.4 Dollar Factor Model

Define the exchange rate S_j as units of U.S. dollar per unit of currency j, and define the exchange rate log change over $[t, t+\tau]$ as $\Delta s_{t,j}^{\tau} \equiv \log S_{t+\tau,j} - \log S_{t,j}$. I assume that the log currency excess return dynamics, taking the perspective of a U.S. investor, follows a factor structure where the dollar factor is the systematic factor:

$$\Delta s_{t,j}^{\tau} + i_{t,j}^{\tau} - i_t^{\tau} = \beta_{t,j}^{\tau} \cdot \Delta dol_t^{\tau} + \varepsilon_{t,j}^{\tau} \tag{3.9}$$

where $i_{t,j}^{\tau} - i_t^{\tau}$ is the riskless interest rate differential for currency j over the interval $[t, t+\tau]$. Δdol_t^{τ} is the τ-period log excess return of an equal-weighted portfolio of all foreign currencies, i.e., $\Delta dol_t^{\tau} = \frac{1}{N} \sum_i \Delta s_{t,i}^{\tau} + \bar{i}_t^{\tau} - i_t^{\tau}$, where N is the number of currencies in the economy and $\bar{i}_t^{\tau}$ is average foreign interest rate. $\beta_{t,j}^{\tau}$ measures the sensitivity for currency j to shocks in the dollar factor, and $\varepsilon_{t,j}^{\tau}$ is an idiosyncratic zero-mean random variable realized at time τ. The factor structure in currency excess returns implies that the total variance of the log currency changes[1] consists of a systematic and idiosyncratic component:

$$E_t\left(RVAR_{t,j}^{\tau}\right) = \left(\beta_t^{\tau}\right)^2 \cdot E_t\left(RVAR_{t,dol}^{\tau}\right) + E_t\left(RVAR_{t,\varepsilon_j}^{\tau}\right) \tag{3.10}$$

where $RVAR_{t,j}^{\tau}$ is the total variance of currency j, $RVAR_{t,dol}^{\tau}$ is the variance of the dollar factor, and $RVAR_{t,\varepsilon_j}^{\tau}$ is the idiosyncratic variance component of currency j. Christoffersen, Fournier, and Jacobs, 2017 make a similar decomposition of equity variances into a market variance component—proxied by the variance of the S&P 500 index—and an idiosyncratic variance risk component. We may think of the variance decomposition (3.10) as the currency equivalent to such an equity variance decomposition using the dollar factor as the currency market factor.

[1]The interest rates are known in advance, implying that the variances of the log excess returns equal the variances of the log currency changes.

3.2.5 Forward-Looking Dollar Factor Betas

Identifying the systematic variance risk component requires the dollar factor beta as well as the variance of the dollar factor. As explained in Nielsen (2017), both of these components can be derived from the full panel of options on cross-pair exchange rates. Specifically, in the absence of triangular arbitrage opportunities—i.e., the same exchange rate applies whether an investor exchanges one currency into another via the U.S. dollar or directly via their cross-pair rate—the conditional covariance between two currencies k and j against a given base currency i (typically U.S. dollar) is given by (Mueller et al., 2017; Nielsen, 2017)

$$Cov_t^Q\left(\Delta s_{t,k}^{\tau}, \Delta s_{t,j}^{\tau}\right) = \frac{1}{2}\bigg(E_t^Q\left(RVAR_{t,k}^{\tau}\right) + E_t^Q\left(RVAR_{t,j}^{\tau}\right) - E_t^Q\left(RVAR_{t,kj}^{\tau}\right)\bigg) \tag{3.11}$$

Each of the variances on the right hand side can be derived from currency option prices on the relevant exchange rates, as discussed in more detail below. I calculate the covariances between all currencies using equation (3.11) from which I derive forward-looking dollar factor betas as in Nielsen (2017):

$$\beta_{t,j}^{\tau} = \frac{\frac{1}{N}\sum_{i\in N} Cov_t^Q\left(\Delta s_{t,j}^{\tau}, \Delta s_{t,i}^{\tau}\right)}{SVAR_{t,dol}^{\tau}} \tag{3.12}$$

Using the above dollar factor beta and the variance of the dollar factor, the variance of a given exchange rate can be decomposed into a systematic and an idiosyncratic component using equation (3.10).

3.2.6 The Share of Systematic Variance

The main hypothesis explored in this paper is that there is a monotonic relation between exposure to systematic variance risk and returns to providing variance risk protection. To this end, I construct a simple measure based on the assumption that idiosyncratic shocks carry no variance risk premia, or more generally, that they are the same across currencies per unit of spot variance. Besides this, I assume that the dollar factor betas carry no risk premia, or more precisely that expected betas under the risk-neutral and the objective measure are the same. This assumption is, for example, consistent with deterministic and time-varying loadings on the systematic factors. I thus relax the assumption of the underlying securities having constant loadings on the systematic factors, as often assumed in the literature that studies variance premia/option pricing in equity factor structure models, e.g., Christoffersen et al. (2017); Serban et al. (2008).

Under the assumption that expected betas are the same under the objective and the risk-neutral measure and that idiosyncratic shocks carry no variance risk premia, that is $E_t^P\left(RVAR_{t,\varepsilon}^{\tau}\right) = E_t^Q\left(RVAR_{t,\varepsilon}^{\tau}\right)$, the expected excess return on a variance swap is:

$$E_t^P\left(RX_t^{\tau}\right) = \frac{(\beta_t^{\tau})^2 \cdot \left(E_t^P\left(RVAR_{t,dol}^{\tau}\right) - E_t^Q\left(RVAR_{t,dol}^{\tau}\right)\right)}{RVAR_{t-\tau}^{t}} \tag{3.13}$$

The above expression is intuitive. The first term on the right hand side is the exposure to dollar factor variance captured by the squared dollar factor beta, and the second term is the variance risk premium on the dollar factor. Equation (3.13) stipulates that under the assumption of a negative variance risk premium on the dollar factor and zero variance risk premia on idiosyncratic shocks, the expected excess return on a variance swap decreases in $\frac{(\beta_t^{\tau})^2}{RVAR_{t-\tau}^{t}}$ or, equivalently, in the ratio of systematic variance to current total spot variance: $\left(\frac{(\beta_t^{\tau})^2 \cdot SVAR_{t,dol}^{\tau}}{RVAR_{t-\tau}^{t}}\right)$. Therefore, under the assumption that only the dollar factor carries a variance risk premium, we expect a monotonic negative relation between the share of systematic variance, defined as:

$$SYS_t^{\tau} = \frac{(\beta_t^{\tau})^2 \cdot SVAR_{t,dol}^{\tau}}{RVAR_{t-\tau}^{t}} \tag{3.14}$$

and average realized excess returns on variance swaps. The assumption of zero variance risk premia on idiosyncratic shocks can be relaxed when testing for whether systematic variance risk is priced in the cross-section of variance swap excess returns. It is, for example, sufficient to assume that idiosyncratic variance risk premia are equal across exchange rates. To see this, take two currencies A and B, with different betas but with the same current objective variance. Under these assumptions, the difference in expected variance swap excess returns of the two currencies is given by:

$$
\begin{aligned}
&E_t^P\left(RX_{t,A}^{\tau}\right) - E_t^P\left(RX_{t,B}^{\tau}\right) = \\
&\left(\frac{\left(\beta_{t,A}^{\tau}\right)^2}{RVAR_{t-\tau}^{t}} - \frac{\left(\beta_{t,B}^{\tau}\right)^2}{RVAR_{t-\tau}^{t}}\right) \cdot \left(E_t^P\left(RVAR_{t,dol}^{\tau}\right) - E_t^Q\left(RVAR_{t,dol}^{\tau}\right)\right) \\
&+ \underbrace{\frac{E_t^P\left(RVAR_{t,A}^{\tau}\right) - E_t^Q\left(RVAR_{t,A}^{\tau}\right) - \left(E_t^P\left(RVAR_{t,B}^{\tau}\right) - E_t^Q\left(RVAR_{t,B}^{\tau}\right)\right)}{RVAR_{t-\tau}^{t}}}_{=0}
\end{aligned}
\tag{3.15}
$$

The above expression shows that the expected variance swap excess return is smallest (i.e., most negative) for the currency with the largest share of systematic variance (SYS) if the idiosyncratic variance risk premia are equal for the two currencies (or their is difference sufficiently small compared to the systematic variance risk component).

3.2.7 Systematic Variance Risk Premia in Cross-Currencies

In the absence of triangular arbitrage opportunities, the factor structure in the log U.S. dollar exchange rates also induces a factor structure in the log cross-pair exchange rates. Specifically, in the absence of triangular arbitrage opportunities, the cross-pair exchange rate dynamics is:

$$
\Delta s_{t,ij}^{\tau} + i_{t,j}^{\tau} - i_{t,i}^{\tau} = \underbrace{\left(\beta_{t,i}^{\tau} - \beta_{t,j}^{\tau}\right)}_{\beta_{t,ij}^{\tau}} \Delta dol_{t,\tau}^{\tau} + \varepsilon_{t,ij}^{\tau} \tag{3.16}
$$

The beta of the cross-pair rate with respect to the dollar factor, $\beta_{t,ij}^{\tau}$, is thus the difference between the respective currencies' dollar factor betas. Therefore, the expected variance swap excess return on a cross-pair exchange rate decreases in the SYS measure under the previous stated assumptions, i.e., the dollar factor carries a negative variance risk premium and idiosyncratic shocks carry none (or they are the same across currencies). In the empirical section, I test this hypothesis by investigating if there is a monotonic relation between the level of the SYS measure and excess returns for SYS-sorted portfolios of cross-pair volatility swaps.

3.2.8 The Forward Share of Systematic Variance

The factor structure in variances, see equation (3.10), transmits into a factor structure in the forward implied variances too, since expectation is a linear operator. First, I express the future spot implied variance as:

$$\begin{aligned} SVAR_{t+\tau_1}^{\tau_2} &= \left(\beta_{t+\tau_1}^{\tau_2}\right)^2 \cdot E_{t+\tau_1}^{Q}\left(RVAR_{t+\tau_1,dol}^{\tau_2}\right) + E_{t+\tau_1}^{Q}\left(RVAR_{t+\tau_1,\varepsilon}^{\tau_2}\right) \\ &= \left(\beta_{t+\tau_1}^{\tau_2}\right)^2 \cdot SVAR_{t+\tau_1,dol}^{\tau_2} + SVAR_{t+\tau_1,\varepsilon}^{\tau_2} \end{aligned} \tag{3.17}$$

The currency spot variance is unknown prior to time $t+\tau_1$, but we can calculate its forward price by taking its time t risk-neutral expectation:

$$FVAR_{t,\tau_1}^{\tau_2} = \left(F\beta_{t,\tau_1}^{\tau_2}\right)^2 \cdot E_t^{Q}\left(SVAR_{t+\tau_1,dol}^{\tau_2}\right) + E_t^{Q}\left(SVAR_{t+\tau_1,\varepsilon}^{\tau_2}\right) \tag{3.18}$$

where $F\beta_{t,\tau_1}^{\tau_2} = E_t^{Q}\left(\left(\beta_{t+\tau_1}^{\tau_2}\right)^2\right)$ is the forward dollar factor beta. Assuming no forward variance risk premia on idiosyncratic shocks, then inserting equation (3.18) into equation (3.6) gives that the expected excess return on an FVARA is given by:

$$E_t^{P}\left(RX_{t,\tau_1}^{\tau_2}\right) = \frac{\left(F\beta_{t,\tau_1}^{\tau_2}\right)^2 \cdot \left(E_t^{P}\left(SVAR_{t+\tau_1,dol}^{\tau_2}\right) - E_t^{Q}\left(SVAR_{t+\tau_1,dol}^{\tau_2}\right)\right)}{SVAR_t^{\tau_2}} \tag{3.19}$$

In the empirical section, I show that the forward variance risk premium on the dollar factor, i.e., $E_t^{P}\left(SVAR_{t+\tau_1,dol}^{\tau_2}\right) - E_t^{Q}\left(SVAR_{t+\tau_1,dol}^{\tau_2}\right)$, is typically negative. This implies that investors are willing to pay a higher price today for hedging against systematic variance risk over a given future time interval compared to τ_1 periods later. In this case, the expected excess return on an FVA decreases monotonically in:

$$FSYS_{t,\tau_1}^{\tau_2} = \frac{\left(F\beta_{t,\tau_1}^{\tau_2}\right)^2 \cdot FVAR_{t,\tau_1,dol}^{\tau_2}}{SVAR_t^{\tau_2}} \tag{3.20}$$

This measure, the forward share of systematic variance, is essentially the same measure as used for variance swaps in (3.14), but based on the forward beta and forward implied variances.

To derive the forward betas, we need to calculate the forward covariances between exchange rates. These can be derived using the same technique as for spot covariances, but rather than using the spot variances, as in equation (3.11), the forward variances are used. That is, the time t conditional covariance between two exchange rates, k and j, over a future time interval, $[t+\tau_1, t+\tau]$, is given by:

$$Cov_t^Q\left(\Delta s_{t+\tau_1,k}^{\tau}, \Delta s_{t+\tau_1,j}^{\tau}\right) = \frac{1}{2}\Bigg(FVAR_{t,\tau_1,k}^{\tau_2} + FVAR_{t,\tau_1,j}^{\tau_2} - FVAR_{t,\tau_1,kj}^{\tau_2}\Bigg) \tag{3.21}$$

Based on the forward variances/covariances, I compute forward dollar factor betas for any given currency j—i.e., the time t expected beta at time τ_1 covering $[t+\tau_1, t+\tau]$—as:

$$F\beta_{t,\tau_1,j}^{\tau_2} = \frac{\frac{1}{N}\sum_{i\in N} Cov_t\left(\Delta s_{t+\tau_1,j}^{\tau}, \Delta s_{t+\tau_1,i}^{\tau}\right)}{FVAR_{t,\tau_1,dol}^{\tau_2}} \tag{3.22}$$

3.2.9 Calculating Spot and Forward Variances

Britten-Jones and Neuberger (2000) show that when the instantaneous variance process follows a diffusion process, then the spot implied variance is given by an integral of put and call prices:

$$SVAR_t^{\tau} = 2e^{i_t^{\tau}}\left(\int_0^{S_t}\frac{1}{K^2}P(K,t,\tau)dK + \int_{S_t}^{\infty}\frac{1}{K^2}C(K,t,\tau)dK\right) \tag{3.23}$$

where i_t^{τ} is the riskless interest rate of the base currency, and $P(K,t,\tau)$ and $C(K,t,\tau)$ are put and call prices, respectively, with maturity τ and strike K.

The spot implied variances are the building blocks for calculating betas, forward variances, forward betas, the SYS measures, and the excess returns on FVAs and volatility swaps. Specifically, I calculate the covariances between all currencies from equation (3.11) from which dollar factor betas are derived using equation (3.12). The same procedure is applied to derive the forward dollar factor betas from the forward variances.

Moreover, to obtain the volatility excess return series, I need a measure for realized volatility. I follow, e.g., Carr and Wu (2008); Della Corte et al. (2016); Londono and Zhou (2017), and calculate the annualized realized volatility using daily observations and a length of the estimation window that matches the maturity of the volatility swap:

$$RVAR_t^{\tau} = \frac{252}{\tau}\sum_{j=0}^{\tau-1}\left(\Delta s_{t-j}\right)^2, \; RV_t^{\tau} = \sqrt{RVAR_t^{\tau}} \tag{3.24}$$

where Δs_t are daily log changes in the exchange rate. I use volatility excess returns, rather than variance excess returns, to be consistent with the academic literature and the market conventions (see Della Corte, Sarno, and Tsiakas (2011) for an elaborate discussion). I derive the spot and forward volatilities from their respective variances following Della Corte, Sarno, and Tsiakas (2011) and Della Corte et al. (2017) and use $SV_t^\tau = \sqrt{SVAR_t^\tau}$ and $FV_{t,\tau_1}^{\tau_2} = \sqrt{FVAR_{t,\tau_1}^{\tau_2}}$. These measures of implied volatilities are convexity biased since expected volatility is less than the square-root of the expected variance according to Jensen's inequality. Della Corte, Sarno, and Tsiakas (2011) find, however, that the convexity biases are small for empirical purposes for currency implied volatilities [2].

I calculate the volatility swap excess returns using the same definition as for the variance swap excess returns. I.e., the τ-month holding period volatility swap excess return is

$$RX_t^\tau = \frac{RV_t^\tau - SV_t^\tau}{RV_{t-\tau}^t} \tag{3.25}$$

Rather than using buy-and-hold FVA excess returns, I follow Della Corte, Kozhan, and Neuberger (2017); Dew-Becker, Giglio, Le, and Rodriguez (2017); and Gorton, Hayashi, and Rouwenhorst (2012) and use 1-month FVA excess returns to avoid overlapping observations in the excess return series. The 1-month FVA excess return is defined as follows. At time t, the investor buys an FVA with maturity τ_1 covering $[t+\tau_1, t+\tau]$ and then closes this position at time $t+1$ by selling an FVA with maturity $\tau_1 - 1$ covering the same interval. The payoff of this strategy is simply:

$$FV_{t+1,\tau_1-1}^{\tau_2} - FV_{t,\tau_1}^{\tau_2} \tag{3.26}$$

I then calculate the 1-month excess return by normalizing the payoff with the current forward volatility price with maturity $\tau_1 - 1$ [3]:

$$RX_{t+1}^{FVA} = \frac{FV_{t+1,\tau_1-1}^{\tau_2} - FV_{t,\tau_1}^{\tau_2}}{FV_{t,\tau_1-1}^{\tau_2}} \tag{3.27}$$

[2] In order to address the convexity bias, I also carried through the entire analysis using variance as the underlying instrument and the main conclusions are unaltered.

[3] Whether I normalize with $FV_{t,\tau_1-1}^{\tau_2}$ or $FV_{t,\tau_1}^{\tau_2}$ does not affect any of the main results of the paper. Della Corte, Kozhan, and Neuberger (2017) normalize with the former and Gorton, Hayashi, and Rouwenhorst (2012) and Dew-Becker, Giglio, Le, and Rodriguez (2017) with the latter.

3.3 Data

This section describes the currency options data and the currency spot and forward data that I use in the empirical analysis.

3.3.1 Currency Options Data

The currency options data consist of OTC quotes from J.P. Morgan collected daily from January 1998 to August 2016 for all G10 currencies, including cross-pairs, which gives rise to a total of 45 quoted currency pairs (9 pairs against the U.S. dollar and 36 cross-pairs). The currency options are European options and are quoted in terms of Garman and Kohlhagen (1983) implied volatilities (IVs) of put options trading at fixed deltas of -0.10 and -0.25 and call options trading at fixed deltas of 0.10, 0.25, and 0.50. The options have fixed maturities of one month, two months, three months, six months, nine months, and 12 months.

From the Garman and Kohlhagen (1983) formula of option prices (the Black-Scholes formula adjusted for the foreign interest rate), the strikes can be retrieved for each IV at the different deltas and for each maturity (for more details on how to extract strikes from implied volatilities, see Della Corte, Sarno, Schmeling, and Wagner (2016); Jurek (2014)). As in, e.g., Della Corte, Sarno, Schmeling, and Wagner (2016) and Jiang and Tian (2005), I approximate the integral in (3.23) by performing two steps. First, I use a cubic spline between the available strike/IV points, and at strikes outside the observed range, I assume a constant IV that equals the end-point IV. I then transform the fitted strike/IV grid into a strike/option price grid using the Garman and Kohlhagen (1983) formula which I then use to approximate the integral with a trapezoidal numerical integration procedure.

3.3.2 Spot and Forward Data

Currency spot and forward data are needed to obtain the strikes from the currency options data. Moreover, I use the currency spot data to estimate realized volatilities in the calculation of volatility swap excess returns. I collect data on spot exchange rates and forward exchange rates for all G10 currencies (AUD, CAD, CHF, EUR, GBP, JPY, NOK, NZD, and SEK) against the U.S. dollar from January 1997 to August 2016 at a daily frequency [4].

[4] The Euro was first introduced in January 1999. Before this, I proxy the Euro with the Deutsche Mark.

The data start one year earlier than the options data because I use lagged realized volatility to derive volatility swap excess returns. The spot and forward contract data are obtained from Reuters via Datastream, and the maturities of the forward contracts match those of the currency options, i.e., one, two, three, six, nine, and 12 months. The exchange rates are defined as units of U.S. dollar per unit of currency j, therefore, if the exchange rate rises then currency j appreciates against the U.S. dollar.

3.4 Empirical Results

This section explores the empirical relation between excess returns on volatility swaps and FVAs and exposure to systematic variance risk.

3.4.1 Currency Volatility Risk Premia

Table 3.1 shows descriptive statistics for excess returns of volatility swaps on each G10 U.S. dollar exchange rate and the dollar factor. The time series of volatility swap excess returns at maturity τ is calculated as excess returns on a monthly rebalanced portfolio that each month buys a volatility swap with maturity τ and holds it until expiry, i.e., the holding period equals the maturity of the contract.

At each maturity, the mean excess returns are negative for all currencies and the dollar factor, and they have tendency to increase in maturity, even though they are not annualized. This pattern is quite striking. For example, for the dollar factor, the annualized mean excess return is almost 15 times smaller (more negative) at the 1-month maturity compared to the 12-month maturity. At the 1-6 month horizon, volatility swaps on the dollar factor carries a highly significant negative risk premium (e.g., Newey and West (1987) t-statistic of -4.72 and -2.32 at the 1 and 6-month horizon, respectively) which is consistent with investors demanding compensation for being exposed to systematic volatility risk. However, the mean excess returns are insignificant at the 9 and 12-month maturities. Moreover, the standard deviations are roughly similar across maturities, and as a result, the annualized Sharpe ratios of the dollar factor gradually increase across maturities from -1.48 at the 1-month to -0.32 at the 12-month maturity. These findings suggest that especially short-term systematic volatility risk is priced.

The upper panel of Figure 3.1 shows for the dollar factor, the average monthly realized volatility along with its average 1-12 month annualized implied volatilities. As a natural convention, the average realized volatility is depicted as the 0-month maturity. The term structure of implied volatilities is on average upward sloping and concave, i.e., it is especially steep at the short end and then becomes almost flat at longer maturities. The spread between average 1-month implied volatility and realized volatility is 76.70 bps, much larger than the average spread between the 2 and 1-month volatility of just 9.20 bps, and the 12-11 month spread which is 2.26 bps. For this reason, it has been highly profitable to pocket the difference between 1-month implied volatility and monthly realized volatility, and thus the highly significant mean excess returns for selling 1-month volatility swaps.

The lower panel of Figure 3.1 illustrates for the dollar factor the time series of monthly realized volatility and 1 and 12-month implied volatilities. The volatility term structure tends to briefly invert in times of distress, most notably in 2008-2009, but also during the Asian financial crisis in 1998. Those times are associated with large losses on the volatility swaps as indicated by their highly skewed and fat tailed excess return distribution (see Table 3.1). Thus, selling systematic volatility insurance delivers high average excess returns and Sharpe ratios but exposes the investor to extreme losses in bad times.

The average upward sloping term structure of dollar factor volatilities implies that investors are willing to pay a higher risk premium for hedging long-term systematic currency volatility shocks compared to short-term shocks. As a consequence, it has, on average, been profitable to sell long-term dollar factor volatility swaps and buy short-term dollar factor volatility swaps. This pattern suggests that expected excess returns on FVAs are negative, since an FVA can be replicated by buying a long-term volatility swap and rolling over short positions in short-term volatility swaps. More specifically, the payoff of a long FVA entered into at time 0 with maturity τ_1 on τ_2-month volatility has the same payoff as the following strategy:

At time 0 buy a volatility swap with maturity $\tau_1 + \tau_2$ (τ) and sell a volatility swap with maturity τ_1. Then at time τ_1, sell a volatility swap with maturity τ_2. At τ_1 the strategy receives $SV_{\tau_1}^{\tau}$ and pays $(SV_t^{\tau} - SV_t^{\tau_1})$, which is the forward volatility price $FV_{t,\tau_1}^{\tau_2}$, i.e., the payoff of this strategy is equivalent to the FVA payoff. The excess returns of this strategy are expected to be negative because long-term volatility swaps tend to be more expensive than short-term volatility swaps. Furthermore, the concave shape of the implied volatility term structure suggests that the mean FVA excess returns increase (become less negative) in maturity because the implied volatility term spreads are smaller at longer maturities.

As suggested above, Table 3.2 indeed shows that average excess returns of FVAs are negative both on the dollar factor and for individual currencies (with the exception being the Canadian dollar). Figure 3.2 shows for the dollar factor the FVA mean excess returns and Sharpe ratios for each maturity combination. Indeed, as suggested by the concave shape of its implied volatility curve, the mean FVA excess returns are especially negative at short-term forward maturities with short-term volatility as the underlying instrument. The mean excess return and Sharpe ratio of the dollar factor are most negative when the maturity is one month for both the forward contract and its underlying volatility, and both gradually increase in the maturity of the forward contract and the underlying volatility. Overall, the results of Table 3.2 and Figure 3.2 suggest that substantial systematic volatility risk premia are mostly confined to short maturities.

Interestingly, Dew-Becker et al. (2017) document a similar declining pattern in Sharpe ratios and mean excess returns for shorting forward variance contracts on the S&P 500 index. The S&P 500 index forward variance contracts only deliver negative mean excess returns at 1-2 month maturities, while they are insignificant and, in fact, positive at maturities beyond three months. As such, it appears to be a general phenomenon across asset classes that investors are particularly willing to pay for hedging short-term systematic variance risk compared to long-term systematic variance risk.

3.4.2 Share of Systematic Variance

In the previous section, I presented evidence suggesting that the dollar factor has a negative volatility risk premium. As a result, under the hypothesis that idiosyncratic volatility risk is not priced, selling volatility swaps and FVAs on currencies with the largest share of systematic variance risk are expected to deliver the highest excess returns. More specifically, in a linear factor model, the difference between variance swap excess returns of two currencies monotonically decreases in the difference between their SYS measures if their idiosyncratic variance risk premia are the same.

Intuitively, the SYS measure is a dynamic and forward-looking version of the R^2-metric of a linear regression and captures how much of the variation in an exchange rate that is explained by exposure to the systematic factor(s). Verdelhan (2017) documents that the dollar factor explains a far greater proportion of the systematic variation in exchange rates compared to the unconditional and conditional carry trade factors of Lustig et al. (2011). This evidence indicates that the dollar factor is the most prominent model for analyzing systematic variance risk in currencies. However, the methodology that I propose can be used in the context of any currency factor model.

Next, I focus on the spot SYS measure to keep my discussion on the share of systematic variance clear and brief, however, for portfolio construction of FVAs, I use the forward SYS measure. I calculate the spot SYS measure at time t with a horizon of τ as $\frac{(\beta_t^\tau)^2 \cdot SVAR_{t,dol}^\tau}{RVAR_t^{t-\tau}}$, where β_t^τ is the option-implied dollar factor beta based on options with maturity τ, $SVAR_{t,dol}^\tau$ is the τ-month option-implied variance of the dollar factor, and $RVAR_t^{t-\tau}$ is the (total) realized variance of the exchange rate over $[t-\tau, t]$. Table 3.3 reports the mean and quantiles of the SYS measure for each U.S. dollar exchange rate at 1-12 month maturities. The Euro and the Swiss franc/Norwegian krone bear the largest share of systematic variance, on average, while the Japanese yen and the Canadian dollar have the smallest.

Verdelhan (2017) runs time-series regressions of changes in bilateral exchange rates on the dollar factor and two carry trade factors. Interesting, the ranking of currencies that Verdelhan (2017) uncovers based on the R^2s from those time-series regressions are basically in line with the ranking that I find based on the average SYS measure. For example, Verdelhan (2017) also ranks the Euro and Swiss franc/Norwegian krone as those with the largest share of systematic variance and the Japanese yen and Canadian dollar as those with the lowest. This is, perhaps, not surprising since there is an almost perfectly linear relationship between individual average betas estimated from time-series regressions and option-implied betas, as shown in Nielsen (2017). It is important to mention, however, that the average measures do not capture the time variation in betas and variances which is of key importance for dynamic portfolio construction of volatility swap and FVA portfolios.

3.4.3 Portfolios Sorted by Share of Systematic Variance

In this section, I examine the empirical relation between systematic variance risk and excess returns on volatility swaps and FVAs, for both U.S. dollar and cross-pair exchange rates.

To this end, I first construct portfolios of volatility swaps based on the SYS measure. Each month, and for each maturity τ, the volatility swaps are allocated into three equal-weighted portfolios from low (P_1) to high (P_3) based on their SYS measures with maturity τ. The SYS factor is defined as the monthly rebalanced portfolio that each month sells P_3 (high SYS) and buys P_1 (low SYS). Table 3.4 shows, for maturities from 1-12 months, descriptive statistics for U.S. dollar volatility swap excess returns for portfolios sorted based on the SYS measure and volatility swap excess returns for the dollar factor. The brackets below the mean excess returns are t-statistics based on Newey and West (1987), with the automatic lag selection of Newey and West (1994).

For each maturity, the mean excess returns and Sharpe ratios monotonically decrease across the portfolios. As a result, there is a positive mean excess return on the SYS factor for each holding period, albeit only significant at 1-3 months. For example, at the 1-month maturity, the mean excess returns and annualized Sharpe ratios for the first and third portfolios are -8.77% and -1.31, and -13.24% and -1.80, respectively. Therefore, the monthly mean excess return on the 1-month SYS factor is 4.47%, which is significant at the 1% level (t-statistic of 3.47). The 1-month annualized Sharpe ratio of the SYS factor is 0.71, which over the same sample period exceeds the Sharpe ratios of the G10 HML carry factor of Lustig et al. (2011) and the G10 conditional dollar factor of Lustig et al. (2014). The mean excess returns and standard deviations for the SYS factor are roughly in the same order of magnitude at different maturities, and hence its annualized Sharpe ratios decline in maturity. Since the SYS factor is constructed based on exposure to dollar factor variance risk, this finding is consistent with that mean excess returns and Sharpe ratios decrease in maturity on a short dollar factor volatility swap.

If the dollar factor captures global risk, it should not only be priced in the cross-section U.S. dollar volatility swaps but also in the cross-section of volatility swaps on cross-pair exchange rates (i.e., exchange rates that do not have the U.S. dollar as base currency), as discussed in section 3.2.7. For the G10 currencies, there are 36 cross-pair exchange rates so the portfolios are more diversified than the U.S. dollar volatility swap portfolios. In this aspect, they therefore provide a cleaner test for whether systematic variance risk is priced. Following the same procedure as above, I construct SYS-sorted portfolios of volatility swaps on cross-pair exchange rates and test for a negative relation between the SYS measure and volatility swap excess returns. The descriptive statistics for the SYS-sorted portfolios are reported in Table 3.5. As was the case for the SYS-sorted U.S. dollar volatility swap portfolios, I find that for each maturity, the mean excess returns monotonically decrease across the portfolios. As a result, the SYS factor has a significant mean excess return at the 1-9 month maturities and nearly significant at the 12-month maturity. The mean excess returns tend to decline in maturity ranging from 4.80% at the 12-month to 6.73% at the 1-month horizon, and likewise annualized Sharpe ratios decrease in maturity. For example, at the 1-month horizon, the annualized Sharpe ratio is 0.65, while it is 0.32 at the 12-month horizon. Overall, the results corroborate that exposure to systematic variance risk is priced in the cross-section of cross-pair currency volatility excess returns, in particular at shorter horizons.

In the previous section, I documented that there is a negative forward volatility risk premium associated with the dollar factor, especially at shorter maturities. As a natural next step, I test if dollar factor forward variance risk is priced in the cross-section of FVAs. To this end, I sort U.S. dollar FVAs in ascending order based on their forward share of systematic variance measures (FSYS measures), and I then allocate them into three equal-weighted portfolios. The FSYS factor sells the high-FSYS portfolio and buys the low-FSYS portfolio. Table 3.6 reports descriptive statistics for excess returns on the FSYS-sorted portfolios and the FSYS factor at six maturity combinations spread out across the maturity spectrum (i.e., τ_1/τ_2 equals 1/1 mo, 1/4 mo, 6/1 mo , 6/4 mo, 8/1 mo, and 8/4 mo).

For each maturity combination, the mean excess returns decrease monotonically across the FSYS-sorted portfolios, and the FSYS factor's mean excess return is significant at each maturity combination except for the 6/4 maturity. Figure 3.3 shows the FSYS factor's mean excess returns and corresponding t-statistics for each maturity combination (66 portfolios). In 50% of the cases, the FSYS factor has a significant mean excess return at the 5% level (77% at the 10% significance level). Interestingly, at the 1-month maturity of the forward contract ($\tau_1 = 1$ mo), the FSYS factor's mean excess return is significant at the 5% level for each maturity of the underlying volatility ($\tau_2 = 1, \ldots, 11$ mo). The mean excess returns and t-statistics tend to decrease in the maturity of both the forward contract and the underlying implied volatility, and most of the insignificant outcomes occur when the maturity is long for both the forward contract and its underlying volatility. This finding is consistent with that the dollar factor's variance risk premium increases in maturity, i.e., it becomes less negative, as discussed previously.

Finally, Table 3.7 shows descriptive statistics for excess returns of the FSYS-sorted portfolios and the FSYS factor constructed from FVAs on cross-pair exchange rates. The results suggest that dollar factor variance risk is priced in the cross-section of short-term cross-pair FVAs and that FSYS factor mean excess returns are positive, but insignificant, at longer maturities. The results, however, are in general weaker compared to the results based on the U.S. dollar FVAs. Table 3.7 shows that the FSYS factor's mean excess returns are positive in five out six cases, but only significant at maturities of 6/1 mo and 8/1 mo. Across all maturity combinations, the FSYS factor's mean excess returns are positive in 91% of the cases, out of which 26% are significant at the 5% level (37% at the 10% level). The significant outcomes mostly occur at shorter maturities. For example, at the 1-month maturity of implied volatility, the FSYS factor's mean excess returns are significant in seven out of eleven cases.

3.5 Explaining SYS-Sorted Portfolio Returns

In this section, I investigate the underlying sources driving the excess returns on the portfolios sorted on systematic variance risk documented in the previous section.

3.5.1 Currency Factors

In the equity market, it is well-documented that spikes in market variance tends to be associated with contemporaneously negative market returns which induce negative correlation between variance swap returns and market returns, the so-called leverage effect (Carr and Wu, 2008; Heston, 1993). As a consequence, a significant part of the S&P 500 variance swap excess returns is accounted for by exposure to market risk. It seems plausible that there is a similar relation between SYS factor excess returns—which are constructed based on exposure to dollar factor variance risk—and dollar factor excess returns. For this reason, I empirically examine whether the risk premia of the SYS-sorted portfolios and the SYS factor are attributable to dollar factor exposure. Furthermore, I include a set of additional factors in the analysis that have been documented in the foreign exchange literature to carry significant risk premia.

More specifically, I run time-series regressions of the excess returns of the SYS-sorted portfolios and the SYS factor on the excess returns of the following three currency factors: the conditional dollar factor of Lustig et al. (2014) which is long (short) the dollar factor when the 1-month U.S. interest rate is below (above) the average 1-month G10 interest rate, the G10 HML carry factor of Lustig et al. (2011) which buys (shorts) the upper (lower) tertile interest rate currencies, and the G10 FX momentum factor of Asness et al. (2013) which buys (sells) 1-month FX forward contracts with the highest (lowest) past 12-month returns [5]. For the volatility swap portfolios, I focus solely on the monthly excess returns—i.e., the 1-month volatility swap portfolios—to match the return frequency of the factors.

Table 3.8 shows the time-series regression coefficients using this three-factor currency model. The model only explains a small share of the mean excess returns of the SYS-sorted portfolios and the SYS factor. The alpha (intercept) for each portfolio is statistically significantly different from 0 at the 1% level. Most importantly, the SYS factor's monthly alpha is 4.29% (t-statistic of 3.72) which is not materially different from its monthly mean excess return of 4.47%. Furthermore, the model explains a quite modest part of the variation of the SYS-sorted portfolio excess returns, with R^2s ranging from 0.12-0.26. More importantly, the R^2 is just about 5% for the SYS factor, i.e., the model neither explains the mean excess returns nor the variation of the SYS factor.

The loadings on the carry factor decrease across the SYS-sorted portfolios, and as a result, the SYS factor has a significant positive carry factor loading. Menkhoff et al. (2012a) show that the carry trade is negatively exposed to global volatility risk, i.e., the carry factor tends to depreciate in times of currency market turmoil. Since the SYS factor is constructed based on exposure to systematic variance risk, it seems reasonable that a part of its risk premium is accounted for by carry factor risk.

Perhaps more interestingly, each of the SYS-sorted portfolios have substantial and highly significant negative loadings on the conditional dollar factor, ranging from -4.32 to -3.02. This is not only a statistically significant effect, but also an economically important effect. For instance, a monthly decline of 1% in the conditional dollar factor tends to be associated with monthly gains on the SYS-sorted portfolios ranging from 3.02-4.32%. This suggests

[5] For further details on the construction of the G10 currency momentum factor see AQR's website: https://www.aqr.com/Insights/Datasets/Time-Series-Momentum-Factors-Monthly.

that when the conditional dollar factor falls then short volatility swap positions experience large losses, on average. The factor loadings, however, are of the same order of magnitude across the portfolios, i.e., it is a level effect. Therefore, the mean excess return on the SYS factor is not explained by exposure to the conditional dollar factor, in fact, the SYS factor has a slightly negative loading, albeit insignificant.

Table 3.9 shows time-series regression coefficients for the FSYS-sorted portfolios and the FSYS factor using the three-factor currency model described above. The FSYS factor has delivered significant risk-adjusted returns when accounting for its exposure to FX momentum, FX carry, and the conditional dollar factor, with five out of six maturity combinations delivering highly significant alphas. For instance, at the 1/1-month maturity, the FSYS factor's monthly alpha is 3.01% (t-statistic of 4.29) which is, in fact, slightly larger than its monthly mean excess return.

For all maturities, the FSYS factor has an insignificant exposure to each of the three currency factors, and the model has little explanatory power, with R^2s ranging from just about 0% up to 4%. The three-factor model does, however, explain a larger part of the variation in the excess returns of the FSYS-sorted portfolios, primarily attributable to the conditional dollar factor. The R^2s range between $0.30 - 0.39$, and the conditional dollar factor loadings are all negative and significant at the 1% level. As was the case for the SYS-sorted portfolios of volatility swaps, the dollar factor loadings are similar across the portfolios and thus cannot explain the excess returns on the FSYS factor. For instance, at the 1/1-month maturity, the conditional dollar factor loadings range between -3.21 to -3.40, and therefore, the FSYS factor's loading on the conditional dollar factor is just -0.18 and insignificant. Figure 3.4 illustrates the FSYS factor alphas and their corresponding t-statistics for each maturity combination. The results are virtually indistinguishable from those presented in Figure 3.3, which shows the same plot for mean excess returns (i.e., the model under the null), further reinforcing that the three-factor model does not explain the FSYS factor's excess returns at any horizon.

3.5.2 Equity Factors

Next, I test whether the excess returns on the SYS-sorted portfolios can be explained by risk exposures to well-established equity factors. In particular, I use the five-factor model of

Fama and French (2015) which consists of the following factors: the value-weighted equity market portfolio (MKT), small minus big (SMB), high minus low book-to-market equity ratio (HML), profitability (RMW), and investment (CMA). The data are retrieved from Kenneth French's website.

Table 3.10 shows the time-series regression coefficients for the SYS-sorted portfolios and the SYS factor using the five-factor equity model. The monthly alpha for the SYS factor is 5.02% and significant at the 1% level which is slightly larger than its monthly mean excess return of 4.47%. The SYS-sorted portfolios have substantial and highly significant negative exposures to the equity market portfolio, with equity market betas between -2.09 and -1.72, i.e., buying volatility swaps on U.S. dollar exchange rates, on average, hedges against equity market downturns. The SYS factor, on the other hand, has an insignificant equity market beta of just -0.25 because of the roughly similar equity market betas of the SYS-sorted portfolios. The negative equity betas for the SYS-sorted portfolios, however, are not large enough to account for their risk premia, and each of the portfolios have significant alphas at the 1% level.

Table 3.11 shows the results for the FSYS-sorted portfolios and the FSYS factor constructed from FVAs. The results are similar to those reported for the SYS-sorted portfolios, so I keep it brief. The FSYS factor alphas are significant at the 5% level for five out of six of the maturities—which are the exact same portfolios that have significant mean excess returns. The alphas are virtually no different than their corresponding mean excess returns. The FSYS-sorted portfolios have significant negative equity market betas, but as is the case for the SYS-sorted portfolios, the loadings are roughly the same across portfolios. As a result, the FSYS factors have economically and statistically insignificant equity market betas. Overall, exposures to equity risk factors do not rationalize the risk premia of the SYS/FSYS-sorted portfolios or the SYS/FSYS factors.

3.5.3 Currency Volatility Factors

Traditional currency factors explain only a minor part of the excess returns on the SYS factors. In what follows, I investigate if the SYS factor risk premia can be attributed to exposure to two factors that are constructed from volatility swaps/FVAs. Della Corte et al. (2017) show that a FX volatility carry factor (VCA) delivers significant mean excess returns.

The VCA buys (sells) FVAs for which the implied volatility term structures are downward sloping (upward sloping). The essence of the VCA is that it exploits that forward implied volatilities are biased predictors of future spot implied volatilities (Della Corte et al., 2017, 2011). In particular, if the current forward volatility price trades at a premium (discount) to current spot implied volatility, then the excess returns on FVAs tend be small (large), i.e., a volatility analogue to the forward premium puzzle (Bilson (1981); Fama (1984)). I construct the VCA factor at each maturity combination as follows. At the end of each month, I allocate volatility swaps/FVAs into three equal-weighted portfolios based on the following slope measure of their implied volatility term structure:

$$SLOPE_t = \frac{SV_t^{12m} - SV_t^{3m}}{SV_t^{3m}}$$

where SV_t^{12m} and SV_t^{3m} are the current 12-month and 3-month implied volatilities, respectively [6]. I then construct the VCA factor as the portfolio that buys (sells) the volatility swaps/FVAs with the smallest (largest) SLOPE measure. In the asset pricing tests, I use the VCA factor that matches the maturity of the volatility swaps/FVAs underlying the SYS/FSYS-sorted portfolios.

Furthermore, I build three portfolios of volatility swaps/FVAs sorted in ascending order based on their current spot volatilities. I then define the VLS factor as the factor that sells portfolio three and buys portfolio one. I include the VLS factor for two reasons. First, the SYS measure is constructed by normalizing the systematic variance component of a currency with its current (total) spot variance. Therefore, it could be the case that spot variance is the dominating factor for the cross-sectional variation in the SYS measure. If so, the SYS factor excess returns are mechanically explained by exposure to the VLS factor. Second, James and Marsh (2017) document that it has been profitable to hold a portfolio that buys (sells) delta-neutral at-the-money straddles with low (high) implied volatilities. This strategy resembles buying the VLS factor, since a straddle is a delta-neutral trade designed to profit from large changes in the underlying exchange rate.

Table 3.12 shows the time-series regression coefficients for the SYS-sorted portfolios and the SYS factor based on a currency five-factor model that consists of VCA, VLS, and

[6] Della Corte et al. (2017) use the 24-month contract rather than the 12-month contract, but this maturity is unavailable in my data for cross-pair options.

the three traditional currency factors used in the previous tests. The SYS factor delivers significant risk-adjusted excess returns also when accounting for its exposure to the VCA and VLS factors. The SYS factor has a 3.97% monthly alpha which is statistically significant at the 1% level and just 32 bps lower than the three-factor currency alpha. The SYS factor loadings on the VCA and VLS factors are insignificant, and the FX momentum, FX carry, and the conditional dollar factor loadings are essentially unchanged relative to the three-factor model. The unadjusted R^2 of the model is 6%, which is practically a negligible improvement compared to the three-factor R^2 of 5%, further supporting that exposures to VLS and VCA are not major driving forces behind the SYS factor excess returns.

Table 3.13 reports the corresponding estimates for the FSYS-sorted portfolios and the FSYS factor constructed from FVAs. For the 1/1 mo and 1/4 mo maturities, the FSYS factor's monthly alphas are 2.84% and 1.05%, respectively, and highly significant. This corresponds to that the VLS and VCA factors account for a modest share of 17 bps and 34 bps of the respective FSYS factor alphas. For these maturity combinations, the FSYS factor's exposure to VLS and VCA are insignificant, and the R^2s are 6% and 4%, which are virtually indistinguishable from the three-factor currency model R^2s. Figure 3.5 shows the FSYS factor's monthly alphas and corresponding t-statistics for each maturity combination. For the one-month forward maturity ($\tau_1 = 1$), the five-factor alphas are statistically significant at the 5% level for all implied volatility maturities shorter than 6 months and significant at the 10% level for all maturities, further supporting that the FSYS factor excess returns at shorter maturities are not explained by the five-factor currency model.

However, at longer maturities of the forward contract and its underlying volatility, the FSYS factor's excess returns covary strongly with the VCA excess returns. At maturities of 6/1 mo, 6/4 mo, 8/1 mo, and 8/4 mo, the FSYS factor has an insignificant alpha and a highly significant positive loading on the VCA factor, and the R^2s are substantially larger compared to the three-factor model R^2s. For these maturities, the FSYS factor loadings on VCA are quite stable and range between 0.31-0.37, with t-statistics of 3.62-5.08. Figure 3.5 shows that the alphas especially tend to decrease in the maturity of the forward contract (τ_1) which, however, is also partly explained by that the mean excess returns on the FSYS factor decrease in τ_1. Overall, the risk premia on the FSYS factor are most significant at shorter maturities of the forward contract and in this case cannot be explained by exposure to other risk factors.

At each maturity, the FSYS-sorted portfolios and the FSYS factor are insignificantly exposed to the VLS factor. Therefore, the variation in the FSYS measure, and thus the variation in the FSYS-sorted portfolio excess returns, is not mechanically linked to the cross-sectional variation in spot implied variance (the denominator of the FSYS measure).

3.6 Dollar Factor Beta-Sorted Portfolios

The empirical analysis above is based on excess returns calculated by normalizing the volatility swap/FVA payoffs with spot volatilities. Alternatively, we can simply use payoffs as excess returns since volatility swaps/FVAs are costless to enter. If we assume negative variance risk premia on the dollar factor and the same variance risk premia on idiosyncratic risk (i.e., as in sections 3.2.6 and 3.2.8), then there is a negative relation between expected volatility payoffs and dollar factor betas. As a consequence, if dollar factor variance risk is priced, we would expect to also identify a negative relation between portfolio payoffs and their dollar factor betas. For this reason, to test the robustness of my results, I conduct the above empirical analysis using payoffs and with the dollar factor betas as basis for constructing long-short systematic volatility risk portfolios. The figures and tables are reported in section 3.10, Appendix: Supplementary Tables and Figures.

Broadly speaking, the main findings presented in this section are consistent with the results based on normalized excess returns. Table 3.14 reports results for portfolios of U.S. dollar volatility swaps sorted based on dollar factor betas. We see that the mean excess returns decrease across the portfolios for each maturity. As a result, the mean excess returns on the SYS factors are positive at all maturities, albeit only significant at maturities between three and 12 months. Perhaps more surprising is that the results are stronger for portfolios of cross-pair volatility swaps. In this case, the mean excess returns on the SYS factors are positive at each maturity and significant at the 1% level at maturities between two and 12 months. The alpha of the SYS factor constructed from 1-month U.S. dollar volatility swaps is barely significant at the 5% level (t-statistic of 1.95) but is actually slightly larger than its unconditional mean excess return.

The results for FVAs provide strong support for that dollar factor variance risk is priced.

Table 3.16 shows descriptive statistics for FSYS factors constructed based on forward dollar factor betas at six maturity combinations. The FSYS factor's mean excess returns are significant at the 1% level at all maturity combinations, and the Sharpe ratios range between 0.92-1.20, i.e., they are substantially higher compared to the case of normalized excess returns. Figure 3.6 shows that this is a general pattern. In fact, the FSYS factor's mean excess returns are significant at the 1% level for any maturity combination. Table 3.17 shows results for portfolios of cross-pair FVAs, and they provide further support for a negative relation between forward dollar factor betas and FVA payoffs. At all maturity combinations, the FSYS factor's mean excess returns are positive, and they are significant at the 5% level in four out six cases and in all cases at the 10% level.

Table 3.18 reports the five-factor currency alphas for the FSYS factor, we see that the FSYS factor's alphas are highly significant for all maturity combinations, apart from the 8/4-month maturity. Figure 3.8 shows the five-factor alphas at each maturity combination. We see that in 65% of the cases the alphas are significant at the 5% level and that most of the insignificant outcomes occur at longer maturities of the forward contract and the underlying volatility. In general, at shorter maturities, the currency five-factor model does not explain the mean excess returns of the FSYS factors. Taken together, the results based on dollar factor beta-sorted portfolios provide additional support for the main result of the paper: Systematic variance risk is priced in the cross-section of volatility excess returns and is not accounted for by well-established currency volatility factors or traditional currency factors, especially not at shorter horizons.

3.7 Conclusion

In this paper, I propose a method for decomposing variances of exchange rates into a systematic part and an idiosyncratic part which I use to study the relation between dollar factor variance risk and volatility excess returns. The systematic component is derived from forward-looking measures of betas and variances and covers the same horizons as currency options. The methodology proposed in this paper can be used to derive the systematic variance component for any currency factor in real time, and at any horizon for which options exist. More importantly, the dollar factor variance components of exchange rates contain useful information for understanding currency volatility excess returns.

The main result of the paper is that the (expected) share of systematic variance is priced in the cross-section of volatility excess returns, primarily at shorter horizons. More specifically, long-short factors that sell (buy) volatility insurance on currencies with the largest (smallest) share of systematic variance deliver significant mean excess returns and high Sharpe ratios that exceed those of traditional currency strategies. The mean excess returns are especially significant at shorter maturities and in this case cannot be explained by exposure to well-established currency risk factors, equity risk factors, or the volatility carry factor. The results of this paper suggest that the dollar factor is not only an important driver of currency excess returns as documented in previous research, it also plays a vital role for understanding currency volatility risk premia.

3.8 Figures

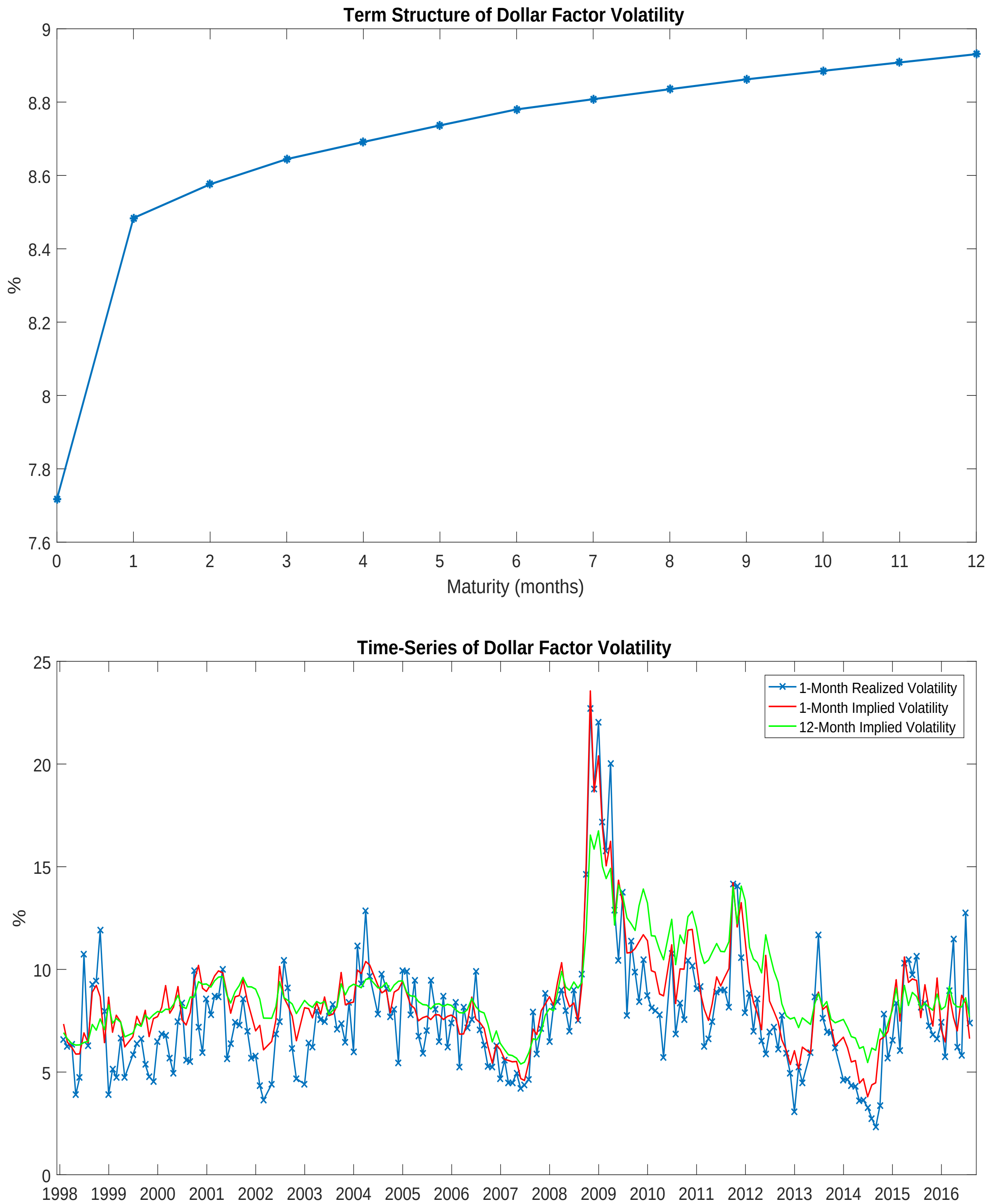

Figure 3.1: **Dollar factor implied volatility curve and volatility time series.** The top panel shows the time-series averages of monthly realized dollar factor volatility (maturity 0) and implied dollar factor volatility at each maturity from 1 to 12 months. The bottom panel plots the time series of monthly realized volatility, 1-month option-implied volatility, and the 12-month option implied volatility. Each time series is annualized by scaling with $\sqrt{12}$.

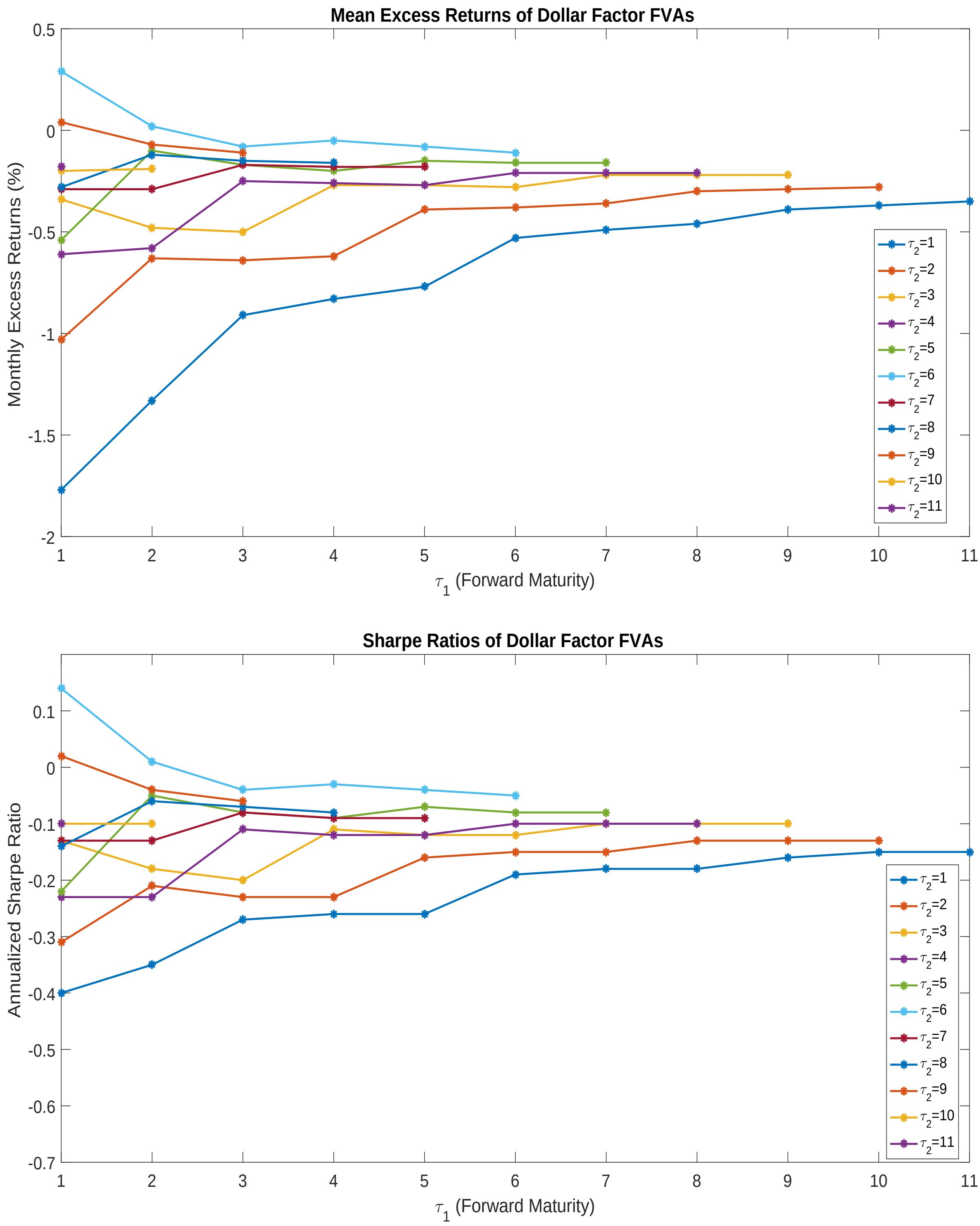

Figure 3.2: **Dollar factor FVA term structures and Sharpe ratios.** The top panel shows the time-series average of FVA excess returns on the dollar factor for each maturity combination (τ_1/τ_2). The bottom panel plots the FVA Sharpe ratios for the dollar factor for each maturity combination (τ_1/τ_2).

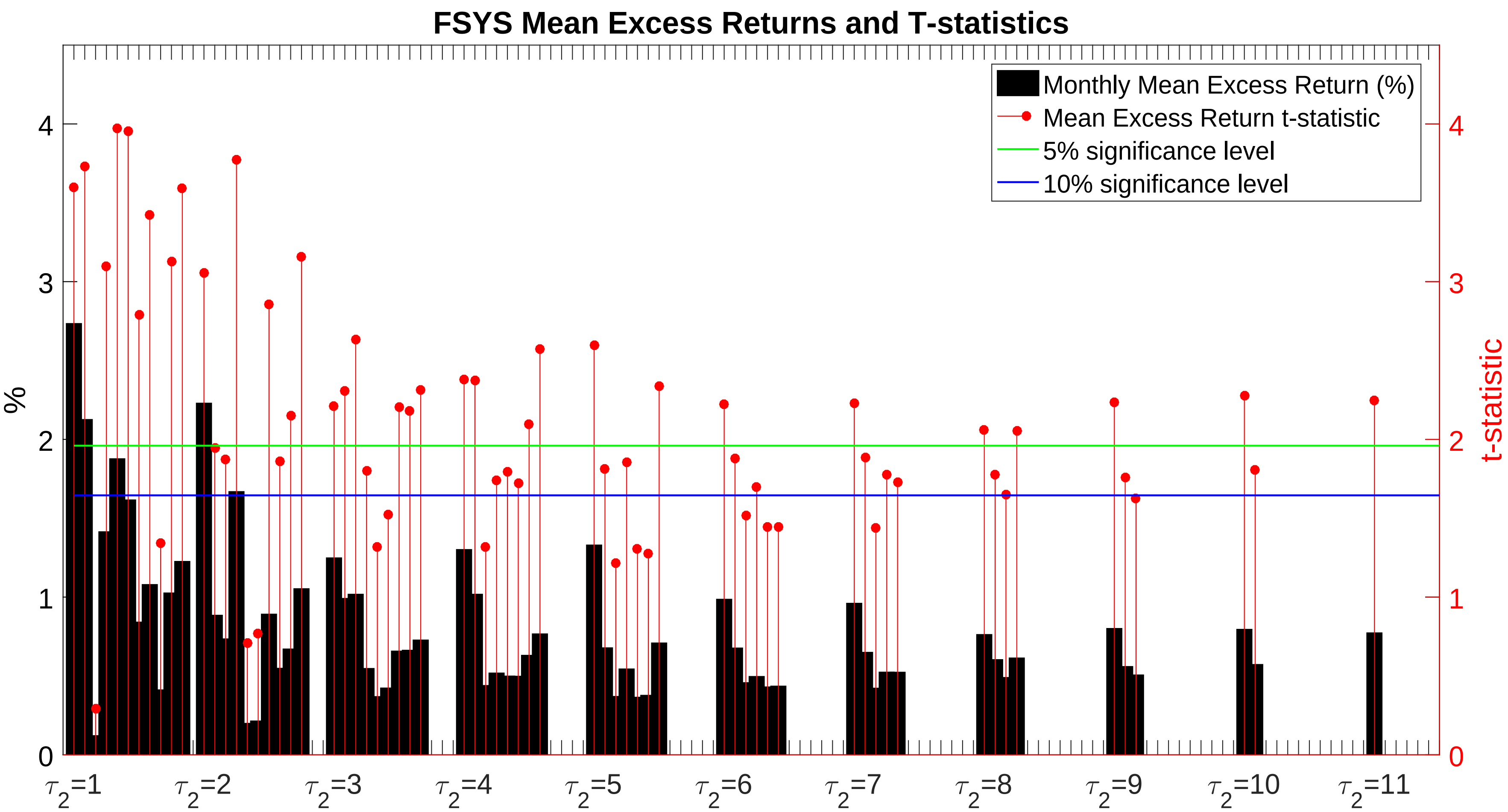

Figure 3.3: **Mean excess returns and t-statistics for the FSYS factor.** This figure illustrates mean excess returns and corresponding Newey and West (1987) t-statistics for the FSYS factor at each maturity/volatility coverage period combination τ_1/τ_2 (66 portfolios). The construction of the FSYS factors is described in the caption of Table 3.6.

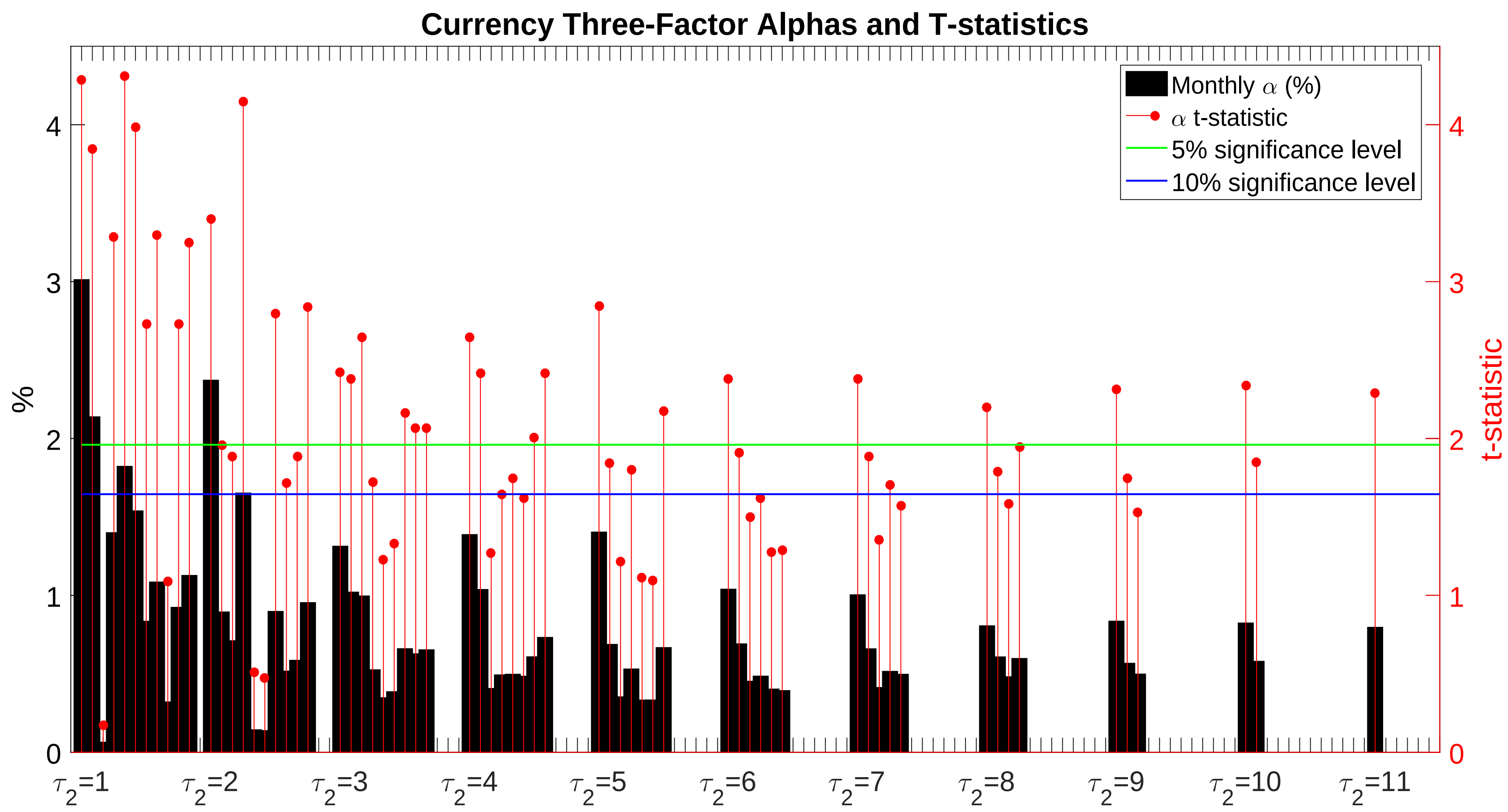

Figure 3.4: **Three-factor currency alphas for the FSYS factor.** This figure illustrates three-factor currency alphas and corresponding Newey and West (1987) t-statistics for the FSYS factor at each maturity/volatility coverage period combination τ_1/τ_2 (66 portfolios). The construction of the FSYS factors is described in the caption of Table 3.6. The currency factors are: the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL).

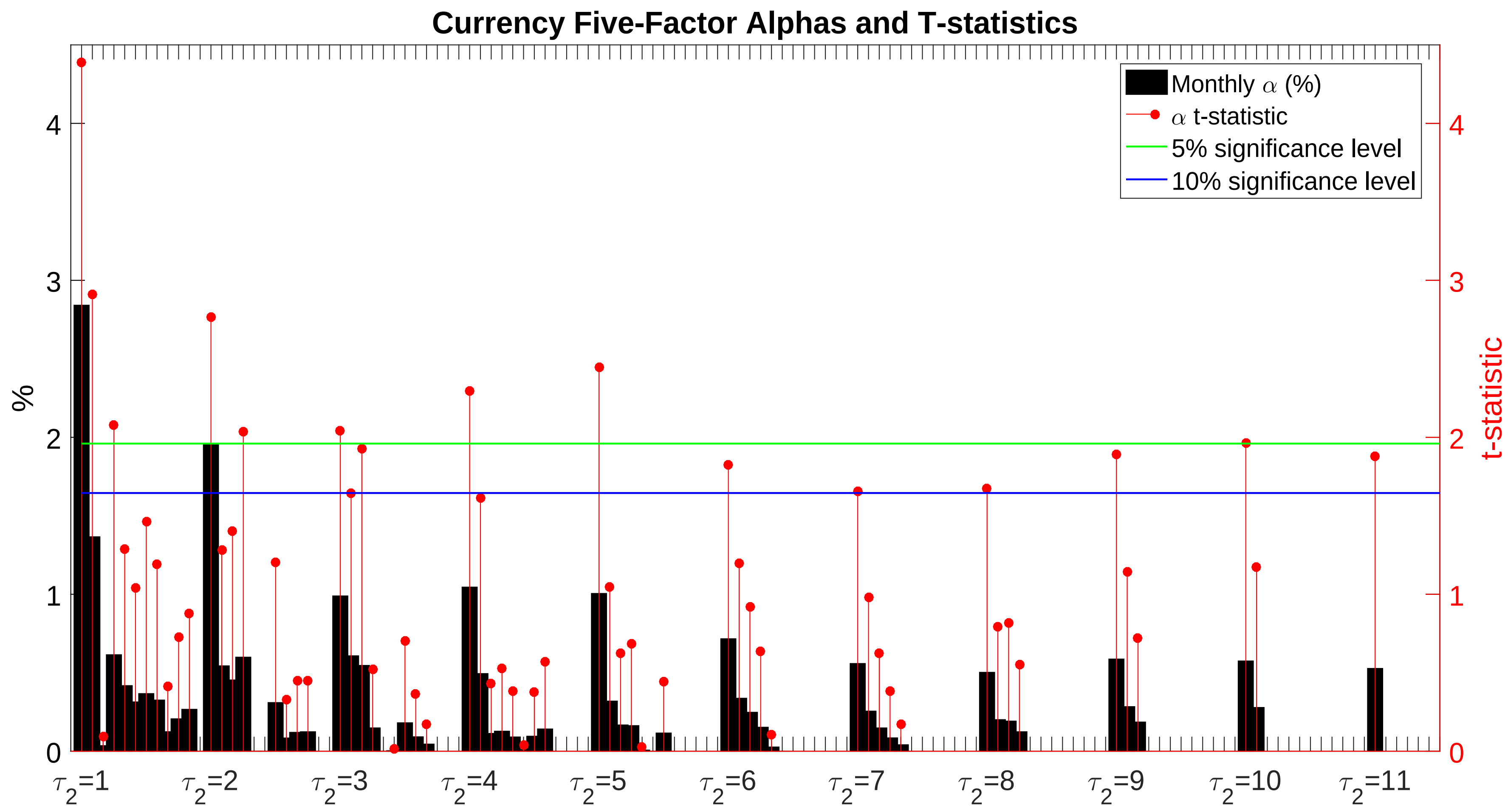

Figure 3.5: **Five-factor currency alphas for the FSYS factor**. This figure illustrates five-factor currency alphas and corresponding Newey and West (1987) t-statistics for the FSYS factor at each maturity/volatility coverage period combination τ_1/τ_2 (66 portfolios). The construction of the FSYS factors is described in the caption of Table 3.6. The currency factors are: the FX spot momentum factor (MOM), the high minus low carry factor (CAR), the conditional dollar factor (CDOL), the FVA volatility carry factor (VCA), and low minus high volatility FVA factor (VLS).

3.9 Tables

Table 3.1: **Descriptive statistics for volatility swap excess returns.** This table reports the means, standard deviations, skewness, and kurtosis for volatility swap excess returns of the G10 U.S. dollar exchange rates and the dollar factor. The volatility swap excess returns are calculated according to (3.3), with maturities of 1-12 months. The means and standard deviations are in percentages. The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	avg	dol
	Panel A: Mean Excess Returns										
1 mo	-7.73	-12.17	-7.51	-13.58	-10.77	-13.92	-11.31	-5.97	-12.27	-9.74	-11.35
2 mo	-5.35	-9.12	-6.40	-13.83	-11.00	-11.98	-10.06	-4.99	-11.24	-8.06	-10.15
3 mo	-3.53	-7.55	-5.78	-14.46	-12.14	-12.00	-9.19	-4.67	-11.10	-7.34	-10.04
6 mo	-2.80	-6.38	-3.44	-14.64	-14.06	-14.44	-8.35	-4.20	-10.40	-6.70	-9.69
9 mo	-2.80	-6.03	-3.11	-14.01	-14.73	-17.15	-7.87	-4.82	-9.51	-7.20	-9.43
12 mo	-2.70	-5.62	-3.48	-13.70	-14.83	-19.53	-8.04	-5.67	-8.97	-7.88	-9.34
	Panel B: Standard Deviation of Excess Returns										
1 mo	30.91	25.40	37.35	22.63	28.07	32.72	33.93	31.59	27.61	27.68	26.57
2 mo	38.32	28.41	38.22	19.65	26.53	30.40	29.70	31.09	23.35	30.83	24.66
3 mo	44.95	31.06	40.46	20.39	25.74	29.55	30.24	32.84	23.95	31.95	25.43
6 mo	42.79	31.53	49.28	24.66	29.74	29.16	31.07	33.15	26.83	31.51	27.78
9 mo	40.65	31.10	42.73	28.36	31.60	28.01	30.98	32.25	28.26	29.81	29.12
12 mo	40.79	31.21	38.39	32.09	34.34	28.06	31.69	31.03	30.69	29.40	31.22
	Panel C: Skewness of Excess Returns										
1 mo	1.84	0.79	5.94	0.15	1.57	0.88	2.20	1.90	0.64	2.48	0.63
2 mo	4.25	1.94	5.51	0.67	1.76	1.02	1.70	2.56	1.15	3.20	1.16
3 mo	4.71	2.31	4.80	0.79	1.92	0.82	1.55	2.54	1.87	2.84	1.33
6 mo	2.75	1.45	3.97	1.31	1.99	0.63	1.32	1.65	2.65	2.07	1.29
9 mo	1.91	1.05	2.68	1.36	1.86	0.62	1.35	1.10	2.57	1.60	1.36
12 mo	1.43	0.75	1.90	1.39	1.81	0.49	1.39	0.66	2.48	1.36	1.48
	Panel D: Kurtosis of Excess Returns										
1 mo	9.26	5.49	62.45	3.39	11.28	5.33	17.13	10.79	5.22	19.21	4.83
2 mo	32.17	10.62	43.69	4.78	9.78	4.73	8.99	14.25	7.69	20.91	6.07
3 mo	34.62	12.51	32.92	5.40	10.51	3.87	6.92	13.33	10.93	15.57	6.75
6 mo	13.72	6.51	21.95	6.42	9.35	3.29	5.23	7.28	14.56	8.43	6.55
9 mo	7.87	4.20	11.17	5.22	7.67	2.90	4.80	4.77	12.17	5.64	5.82
12 mo	5.30	3.28	6.65	4.98	6.84	2.75	4.66	3.17	9.92	4.45	5.70

Table 3.2: **Descriptive statistics for forward volatility agreement excess returns.** This table reports means, standard deviations, skewness, and kurtosis of forward volatility agreement (FVA) excess returns for the G10 U.S. dollar exchange rates and the dollar factor. The FVA excess returns are calculated according to the expression (3.27). The FVA maturities range from 1-8 months and the volatility coverage periods range from 1-4 months. The means and standard deviations are monthly and in percentages. The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

	Panel A: Mean Excess Returns (monthly)									
	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	dol
1/1 mo	-1.09	0.65	-2.55	-2.80	-3.84	-1.19	-1.52	-1.41	-1.59	-1.77
1/4 mo	-0.51	0.29	-1.33	-1.34	-1.83	-1.55	-0.61	-0.50	-0.64	-0.83
6/1 mo	0.03	0.50	-0.90	-0.84	-0.99	-1.34	-0.29	0.04	-0.33	0.29
6/4 mo	-0.22	0.12	-0.38	-0.34	-0.31	-1.13	-0.05	-0.28	-0.05	-0.05
8/2 mo	-0.14	0.22	-0.52	-0.47	-0.52	-1.12	-0.12	-0.17	-0.11	-0.28
8/4 mo	-0.02	0.28	-0.39	-0.33	-0.32	-0.92	0.03	-0.03	-0.09	-0.16
	Panel B: Std of Excess Returns (monthly)									
1/1 mo	18.85	16.32	15.36	15.16	16.31	16.93	16.35	17.18	15.39	15.26
1/4 mo	13.12	11.96	11.23	11.06	11.96	11.50	11.31	11.54	10.98	10.89
6/1 mo	9.96	9.89	9.13	8.76	10.68	8.66	8.96	8.75	8.67	7.22
6/4 mo	7.81	8.11	8.01	7.72	8.99	7.55	7.60	6.89	7.44	7.03
8/2 mo	7.75	8.22	7.72	7.53	8.46	7.49	7.41	6.88	7.42	7.20
8/4 mo	7.03	7.68	7.42	7.08	7.88	6.96	6.77	6.46	6.74	6.77
	Panel C: Skewness of Excess Returns									
1/1 mo	1.94	1.67	1.20	1.65	2.08	1.66	2.42	1.27	2.23	1.29
1/4 mo	2.27	1.24	1.10	1.73	2.27	1.64	2.06	1.44	2.15	1.27
6/1 mo	2.46	1.38	0.82	0.94	2.13	1.91	1.38	1.37	1.23	0.41
6/4 mo	1.87	0.99	0.71	0.76	1.72	1.32	1.04	1.02	1.05	0.72
8/2 mo	1.98	1.07	0.90	0.96	1.82	1.38	1.12	1.00	1.41	0.83
8/4 mo	1.99	1.05	0.90	0.93	1.45	1.30	0.90	0.95	1.30	0.82
	Panel D: Kurtosis of Excess Returns									
1/1 mo	10.50	10.10	5.58	9.70	13.64	8.17	14.92	6.13	14.25	6.91
1/4 mo	14.55	6.57	5.73	11.78	17.27	9.56	12.78	8.16	15.55	7.65
6/1 mo	18.07	6.94	4.92	6.37	13.92	13.03	7.35	8.25	7.72	3.42
6/4 mo	12.30	4.91	5.04	5.22	10.87	8.38	5.97	6.76	6.84	4.32
8/2 mo	12.85	5.09	5.73	6.25	12.48	8.43	6.19	6.21	8.29	5.14
8/4 mo	13.01	4.88	6.68	6.14	10.90	8.10	5.32	5.82	8.25	5.02

Table 3.3: **Descriptive statistics for the share of systematic variances.** This table reports means and quantiles for the share of systematic variance measure for the G10 U.S. dollar exchange rates. The share of systematic variance measure is defined for each exchange rate as: $\frac{(\beta_t^\tau)^2 \cdot SVAR_{t,dol}^\tau}{RVAR_t^{t-\tau}}$, where β_t^τ is the τ-month option-implied dollar factor beta, $SVAR_{t,dol}^\tau$ is the τ-month risk-neutral variance of the dollar factor, and $RVAR_t^{t-\tau}$ is the total realized exchange rate volatility over $[t-\tau, t]$. The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and the data comprise 216 monthly observations.

	Panel A: Mean								
Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	55.12	34.14	63.30	76.44	45.73	19.84	63.49	40.65	62.96
2 mo	45.59	25.79	65.41	78.90	46.73	19.42	65.22	41.49	65.06
3 mo	45.48	26.21	64.63	79.77	47.04	19.23	66.01	41.88	65.06
6 mo	45.68	25.68	65.43	80.41	47.26	18.16	65.96	41.49	65.43
9 mo	45.80	25.61	64.70	80.78	47.64	17.58	66.41	41.74	65.87
12 mo	46.27	26.13	64.03	80.95	48.19	17.46	67.16	41.93	65.72
	Panel B: 5% Quantile								
Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	16.77	4.22	34.27	56.81	23.25	0.30	36.37	7.40	39.28
2 mo	10.07	0.43	36.70	60.24	26.04	0.35	40.80	7.51	44.70
3 mo	9.01	0.41	37.29	62.17	27.32	0.23	42.83	7.01	45.89
6 mo	8.60	0.43	37.88	64.49	29.87	0.16	45.05	6.84	49.04
9 mo	8.45	0.39	37.55	65.98	31.19	0.15	45.94	7.63	49.80
12 mo	8.21	0.22	35.89	66.66	32.28	0.19	45.57	7.15	49.31
	Panel C: 95% Quantile								
Horizon	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
1 mo	77.73	63.39	84.00	89.88	67.10	46.42	88.23	67.70	82.05
2 mo	70.04	53.85	86.11	91.33	67.40	44.79	85.10	66.94	82.56
3 mo	68.13	53.99	85.28	93.14	66.22	45.03	86.57	67.57	81.59
6 mo	69.50	52.84	85.67	92.63	65.07	41.99	83.54	65.43	81.55
9 mo	70.48	52.32	85.56	91.81	65.11	41.63	83.21	65.88	82.25
12 mo	71.12	52.88	85.29	93.51	65.51	42.45	90.12	66.65	83.32

Table 3.4: **Excess returns of U.S. dollar volatility swap portfolios sorted by share of systematic variances.** This table shows the means, standard deviations, skewness, kurtosis, and annualized Sharpe ratios for U.S. dollar volatility swap portfolios sorted by their share of systematic variances. The means and standard deviations are in percentages. The SYS measure at maturity τ is $\frac{(\beta_t^\tau)^2 \cdot SVAR_{t,dol}^\tau}{RVAR_t^{t-\tau}}$, where β_t^τ is the option-implied dollar factor beta, $SVAR_{t,dol}^\tau$ is the option-implied variance of the dollar factor, and $RVAR_t^{t-\tau}$ is the (total) realized exchange rate variance over $[t-\tau, t]$. Each month, the volatility swaps are allocated equally into three portfolios from low to high based on their SYS measures that match the maturity of the volatility swaps (1, 2, 3, 6, 9, and 12 months). The SYS factor sells the high-SYS portfolio and buys the low-SYS portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1 mo					Panel B: 2 mo				
	P_1	P_2	P_3	SYS	dol	P_1	P_2	P_3	SYS	dol
Mean	-8.77	-9.74	-13.24	4.47	-11.35	-7.33	-8.19	-12.46	5.13	-10.15
	[-4.58]	[-4.40]	[-6.28]	[3.47]	[-4.72]	[-2.98]	[-2.90]	[-5.60]	[4.08]	[-3.85]
Std	23.20	23.40	25.44	21.89	26.57	24.75	26.86	22.88	19.08	24.66
Skew	1.66	1.51	0.79	0.15	0.63	2.43	2.93	1.48	-0.23	1.16
Kurt	9.58	9.15	5.50	5.89	4.83	14.20	19.99	8.05	12.45	6.07
SR	-1.31	-1.44	-1.80	0.71	-1.48	-0.73	-0.75	-1.33	0.66	-1.01
	Panel C: 3 mo					Panel D: 6 mo				
	P_1	P_2	P_3	SYS	dol	P_1	P_2	P_3	SYS	dol
Mean	-7.01	-8.36	-11.44	4.43	-10.04	-6.10	-9.33	-10.80	4.71	-9.69
	[-2.43]	[-2.66]	[-3.72]	[2.78]	[-3.26]	[-1.62]	[-2.06]	[-2.42]	[1.65]	[-2.32]
Std	24.93	27.15	27.34	17.07	25.43	26.05	29.74	31.64	20.99	27.78
Skew	2.14	2.75	2.49	-1.97	1.33	1.41	2.01	2.23	-2.57	1.29
Kurt	11.16	17.35	14.11	17.41	6.75	6.71	9.67	9.65	15.82	6.55
SR	-0.56	-0.62	-0.84	0.52	-0.79	-0.33	-0.44	-0.48	0.32	-0.49
	Panel E: 9 mo					Panel F: 12 mo				
	P_1	P_2	P_3	SYS	dol	P_1	P_2	P_3	SYS	dol
Mean	-6.64	-9.81	-10.23	3.59	-9.43	-6.62	-10.14	-10.75	4.14	-9.34
	[-1.54]	[-1.77]	[-1.88]	[1.04]	[-1.84]	[-1.36]	[-1.58]	[-1.75]	[1.27]	[-1.54]
Std	25.83	30.67	32.10	21.04	29.12	26.39	32.48	32.58	18.79	31.22
Skew	1.17	1.59	1.87	-1.79	1.36	0.93	1.55	1.65	-0.90	1.48
Kurt	4.90	6.33	6.95	10.47	5.82	3.84	5.48	5.74	5.73	5.70
SR	-0.30	-0.37	-0.37	0.20	-0.37	-0.25	-0.31	-0.33	0.22	-0.30

Table 3.5: **Excess returns of cross-pair volatility swap portfolios sorted by share of systematic variances.** This table shows means, standard deviations, skewness, kurtosis, and annualized Sharpe ratios for cross-pair volatility swap portfolios sorted by their share of systematic variances (SYS). The means and standard deviations are monthly and in percentages. The SYS measure at maturity τ is $\frac{(\beta_t^\tau)^2 \cdot SVAR_{t,dol}^\tau}{RVAR_t^{t-\tau}}$, where β_t^τ is the option-implied dollar factor beta, $SVAR_{t,dol}^\tau$ is the option-implied variance of the dollar factor, and $RVAR_t^{t-\tau}$ is the (total) realized variance of the exchange rate over $[t-\tau, t]$. Each month, the volatility swaps are allocated equally into three portfolios from low to high based on their SYS measures that match the maturity of the volatility swaps (1, 2, 3, 6, 9, and 12 months). The SYS factor sells the high-SYS portfolio and buys the low-SYS portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1 mo				Panel B: 2 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-6.83	-8.82	-13.56	6.73	-4.96	-8.38	-10.84	5.88
	[-2.26]	[-3.79]	[-4.26]	[2.23]	[-1.38]	[-3.11]	[-2.76]	[2.56]
Std	37.87	25.03	32.53	36.07	37.05	25.74	38.37	29.66
Skew	7.30	1.98	-2.05	6.76	5.13	2.21	2.27	2.53
Kurt	80.14	12.50	24.01	65.63	38.67	13.64	25.54	33.97
SR	-0.62	-1.22	-1.44	0.65	-0.33	-0.80	-0.69	0.49
	Panel C: 3 mo				Panel D: 6 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-4.82	-6.53	-10.66	5.84	-4.24	-5.55	-10.32	6.08
	[-1.22]	[-1.65]	[-2.91]	[2.57]	[-0.82]	[-1.11]	[-2.53]	[2.42]
Std	35.00	35.13	32.76	22.99	35.91	35.45	28.41	19.41
Skew	3.78	3.62	0.59	4.36	2.72	2.47	1.62	2.97
Kurt	21.88	20.87	16.02	31.81	11.93	11.64	8.65	17.09
SR	-0.28	-0.37	-0.65	0.51	-0.17	-0.22	-0.51	0.44
	Panel E: 9 mo				Panel F: 12 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-5.37	-5.37	-10.86	5.49	-5.89	-7.06	-10.70	4.80
	[-1.02]	[-0.93]	[-2.30]	[2.50]	[-1.03]	[-1.21]	[-1.88]	[1.88]
Std	31.77	34.40	27.03	14.69	30.75	31.35	29.49	14.81
Skew	1.73	1.90	1.61	2.37	1.26	1.60	1.52	1.16
Kurt	6.32	7.26	6.43	13.50	4.24	5.70	5.31	6.40
SR	-0.20	-0.18	-0.46	0.43	-0.19	-0.23	-0.36	0.32

Table 3.6: **Excess returns of U.S. dollar FVA portfolios sorted by forward share of systematic variances.** This table shows means, standard deviations, skewness, kurtosis, and annualized Sharpe ratios for portfolios of U.S. dollar forward volatility agreements (FVAs) sorted by their forward share of systematic variances (FSYS). The means and standard deviations are monthly and in percentages. The FSYS measure at maturity/volatility coverage period (τ_1/τ_2) is $\frac{\left(F\beta_{t,\tau_1}^{\tau_2}\right)^2 \cdot FVAR_{t,\tau_1,dol}^{\tau_2}}{SVAR_t^{\tau_2}}$, where $F\beta_{t,\tau_1}^{\tau_2}$ is the forward dollar factor beta, $FVAR_{t,\tau_1,dol}^{\tau_2}$ is the forward implied variance of the dollar factor, and $SVAR_t^{\tau_2}$ is the spot implied variance of the exchange rate. Each month, the FVAs are allocated equally into three portfolios from low to high based on their FSYS measures that match the maturity/volatility coverage period of the FVAs. The FSYS factor sells the high-FSYS portfolio and buys the low-FSYS portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo					Panel B: 1/4 mo				
	P_1	P_2	P_3	FSYS	dol	P_1	P_2	P_3	FSYS	dol
Mean	0.13	-2.63	-2.61	2.73	-1.77	-0.10	-1.17	-1.40	1.30	-0.83
	[0.09]	[-2.20]	[-2.04]	[3.53]	[-1.39]	[-0.12]	[-1.28]	[-1.68]	[2.18]	[-0.99]
Std	14.49	14.72	15.43	10.00	15.26	10.34	10.94	10.67	6.63	10.89
Skew	2.24	1.48	1.79	0.28	1.29	2.01	1.82	1.84	0.44	1.27
Kurt	15.02	7.80	10.68	5.47	6.91	14.82	10.12	13.69	4.75	7.65
SR	0.03	-0.62	-0.59	0.95	-0.40	-0.03	-0.37	-0.46	0.68	-0.26
	Panel C: 6/1 mo					Panel D: 6/4 mo				
Mean	0.60	-0.97	-1.01	1.62	0.29	0.04	-0.46	-0.46	0.50	-0.05
	[0.73]	[-1.44]	[-1.43]	[3.54]	[0.65]	[0.08]	[-0.89]	[-0.85]	[1.68]	[-0.12]
Std	8.37	8.07	8.29	5.42	7.22	6.73	7.07	7.12	4.34	7.03
Skew	1.70	1.32	1.05	0.91	0.41	0.98	1.28	0.80	0.49	0.72
Kurt	10.95	8.58	6.82	5.30	3.42	5.68	8.21	5.95	4.50	4.32
SR	0.25	-0.42	-0.42	1.03	0.14	0.02	-0.23	-0.22	0.40	-0.03
	Panel E: 8/1 mo					Panel F: 8/4 mo				
Mean	0.38	-0.66	-0.70	1.08	-0.28	0.27	-0.37	-0.50	0.77	-0.16
	[0.65]	[-1.27]	[-1.29]	[3.28]	[-0.53]	[0.58]	[-0.81]	[-1.03]	[2.56]	[-0.34]
Std	6.77	6.74	6.90	4.37	7.20	6.16	6.25	6.44	4.23	6.77
Skew	1.36	1.07	1.11	0.68	0.83	1.01	1.35	1.05	0.52	0.82
Kurt	7.73	7.21	7.53	4.85	5.14	5.57	9.29	7.46	4.32	5.02
SR	0.19	-0.34	-0.35	0.85	-0.14	0.15	-0.20	-0.27	0.63	-0.08

Table 3.7: **Excess returns of cross-pair FVA portfolios sorted by forward share of systematic variances.** This table shows means, standard deviations, skewness, kurtosis, and annualized Sharpe ratios for portfolios of cross-pair exchange rate forward volatility agreements (FVAs) sorted by their forward share of systematic variances (FSYS). The means and standard deviations are monthly and in percentages. The FSYS measure at maturity/volatility coverage period (τ_1/τ_2) is $\frac{(F\beta_{t,\tau_1}^{\tau_2})^2 \cdot FVAR_{t,\tau_1,dol}^{\tau_2}}{SVAR_t^{\tau_2}}$, where $F\beta_{t,\tau_1}^{\tau_2}$ is the forward dollar factor beta, $FVAR_{t,\tau_1,dol}^{\tau_2}$ is the forward implied variance of the dollar factor, and $SVAR_t^{\tau_2}$ is the spot implied variance of the exchange rate. Each month, the FVAs are allocated equally into three portfolios from low to high based on their FSYS measures that match the maturity/volatility coverage period of the FVAs. The FSYS factor sells the high-FSYS portfolio and buys the low-FSYS portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo				Panel B: 1/4 mo			
	P_1	P_2	P_3	FSYS	P_1	P_2	P_3	FSYS
Mean	-1.36	-2.63	-2.00	0.64	-0.58	-1.50	-1.08	0.49
	[-1.02]	[-1.95]	[-1.67]	[1.20]	[-0.65]	[-1.60]	[-1.31]	[1.59]
Std	14.94	14.46	14.38	6.84	10.23	10.63	10.20	4.60
Skew	3.05	2.97	2.34	0.12	3.54	3.66	2.50	0.25
Kurt	22.86	24.11	15.65	3.56	29.61	32.66	18.31	3.76
SR	-0.31	-0.63	-0.48	0.33	-0.20	-0.49	-0.37	0.37
	Panel C: 6/1 mo				Panel D: 6/4 mo			
Mean	-0.20	-0.55	-0.96	0.77	-0.37	-0.37	-0.69	0.32
	[-0.26]	[-0.71]	[-1.39]	[2.95]	[-0.70]	[-0.76]	[-1.53]	[1.77]
Std	8.05	8.21	8.03	3.98	6.47	6.48	6.43	3.28
Skew	3.09	2.47	2.26	-0.88	2.49	1.86	1.75	0.49
Kurt	24.93	17.52	15.07	9.40	18.62	12.77	11.85	5.02
SR	-0.08	-0.23	-0.42	0.67	-0.20	-0.20	-0.37	0.34
	Panel E: 8/1 mo				Panel F: 8/4 mo			
Mean	-0.30	-0.30	-0.84	0.54	-0.15	-0.11	-0.06	-0.09
	[-0.54]	-0.52]	[-1.70]	[2.79]	[-0.32]	[-0.21]	[-0.13]	[-0.27]
Std	6.41	6.46	6.26	3.40	5.82	6.11	6.27	4.17
Skew	2.45	2.28	1.84	0.51	2.06	1.87	1.59	-0.69
Kurt	18.00	14.45	11.85	5.83	15.20	11.94	9.63	7.20
SR	-0.16	-0.16	-0.47	0.55	-0.09	-0.06	-0.03	-0.07

Table 3.8: **Currency three-factor risk-adjusted returns for volatility swap portfolios sorted by share of systematic variances.** This table shows currency three-factor risk-adjusted returns (αs) and factor loadings for portfolios of 1-month U.S. dollar exchange rate volatility swaps sorted by their share of systematic variances. The portfolio construction follows the same procedure as described in Table 3.4. Columns 1-4 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on three currency factors: the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL). The excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	α	MOM	CAR	CDOL	R^2
P_1	-7.61	0.26	0.41	-4.32	0.19
	[-4.69]	[0.85]	[0.53]	[-4.28]	
P_2	-8.26	0.50	-1.25	-4.51	0.26
	[-4.57]	[0.99]	[-1.82]	[-5.26]	
P_3	-11.90	0.17	-1.54	-3.02	0.12
	[-5.90]	[0.48]	[-2.13]	[-3.94]	
SYS	4.29	0.08	1.95	-1.30	0.05
	[3.72]	[0.22]	[2.99]	[-1.70]	

Table 3.9: **Currency three-factor risk-adjusted returns for FVA portfolios sorted by forward share of systematic variances.** This table shows currency three-factor risk-adjusted returns (αs) and factor loadings for portfolios of U.S. dollar exchange rate FVAs sorted by their forward share of systematic variances. The portfolio construction follows the same procedure as described in Table 3.6. In each panel, columns 1-4 show intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on three currency factors: the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL). The excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo					Panel B: 1/4 mo				
	α	MOM	CAR	CDOL	R^2	α	MOM	CAR	CDOL	R^2
P_1	0.82	0.39	0.46	-3.40	0.32	0.37	0.36	0.23	-2.49	0.35
	[0.73]	[2.60]	[0.90]	[-3.98]		[0.51]	[2.96]	[0.56]	[-4.32]	
P_2	-1.99	0.62	-0.31	-3.21	0.33	-0.65	0.50	-0.26	-2.59	0.39
	[-1.99]	[3.99]	[-0.59]	[-5.10]		[-0.91]	[4.15]	[-0.68]	[-5.04]	
P_3	-2.19	0.76	0.05	-3.22	0.31	-1.01	0.46	0.09	-2.40	0.33
	[-2.04]	[4.23]	[0.09]	[-4.07]		[-1.51]	[3.39]	[0.19]	[-4.22]	
FSYS	3.01	-0.36	0.41	-0.18	0.04	1.39	-0.10	0.14	-0.09	0.01
	[4.29]	[-1.91]	[1.31]	[-0.58]		[2.64]	[-0.90]	[0.64]	[-0.42]	
	Panel C: 6/1 mo					Panel D: 6/4 mo				
	α	MOM	CAR	CDOL	R^2	α	MOM	CAR	CDOL	R^2
P_1	0.89	0.39	0.11	-1.93	0.35	0.30	0.30	-0.03	-1.53	0.35
	[1.30]	[3.22]	[0.37]	[-4.57]		[0.68]	[3.20]	[-0.14]	[-5.68]	
P_2	-0.62	0.33	-0.11	-1.78	0.33	-0.17	0.33	-0.11	-1.61	0.36
	[-1.10]	[2.89]	[-0.38]	[-4.44]		[-0.41]	[3.15]	[-0.43]	[-5.27]	
P_3	-0.64	0.32	-0.01	-1.88	0.33	-0.18	0.29	-0.00	-1.55	0.31
	[-1.16]	[3.02]	[-0.04]	[-5.18]		[-0.41]	[3.33]	[-0.01]	[-5.61]	
FSYS	1.54	0.07	0.12	-0.06	0.01	0.48	0.02	-0.03	0.02	0.00
	[3.99]	[0.93]	[0.67]	[-0.32]		[1.62]	[0.25]	[-0.21]	[0.15]	
	Panel E: 8/1 mo					Panel F: 8/4 mo				
	α	MOM	CAR	CDOL	R^2	α	MOM	CAR	CDOL	R^2
P_1	0.66	0.27	0.09	-1.60	0.35	0.50	0.27	-0.05	-1.35	0.33
	[1.38]	[2.73]	[0.36]	[-5.15]		[1.20]	[3.16]	[-0.23]	[-5.34]	
P_2	-0.37	0.29	-0.13	-1.49	0.34	-0.09	0.29	-0.13	-1.44	0.36
	[-0.85]	[3.05]	[-0.55]	[-4.85]		[-0.24]	[3.20]	[-0.55]	[-4.83]	
P_3	-0.42	0.28	0.02	-1.54	0.32	-0.23	0.22	0.06	-1.40	0.30
	[-0.97]	[3.23]	[0.07]	[-5.33]		[-0.57]	[2.76]	[0.23]	[-5.03]	
FSYS	1.08	-0.00	0.07	-0.06	0.00	0.73	0.05	-0.10	0.06	0.01
	[3.30]	[-0.06]	[0.44]	[-0.37]		[2.42]	[0.83]	[-0.65]	[0.44]	

Table 3.10: **Fama-French five-factor risk-adjusted returns for volatility swap portfolios sorted by share of systematic variances.** This table shows Fama and French (2015) five-factor alphas and factor loadings for portfolios of U.S. dollar exchange rate volatility swaps sorted by their share of systematic variances. The portfolio construction follows the same procedure as described in Table 3.4. Columns 1-6 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on the five Fama and French (2015) equity factors: market (MKT), size (SMB), value (HML), profitability (RMW), and investment (CMA). The excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	α	MKT	SMB	HML	RMW	CMA	R^2
P_1	-6.67	-1.97	-1.16	0.51	-1.91	-1.14	0.12
	[-3.32]	[-3.10]	[-1.92]	[0.71]	[-2.24]	[-1.26]	
P_2	-8.12	-2.09	-0.04	-0.22	-0.97	-1.00	0.13
	[-3.89]	[-4.00]	[-0.09]	[-0.29]	[-1.17]	[-1.09]	
P_3	-11.75	-1.72	0.27	0.37	-0.84	-1.95	0.07
	[-5.54]	[-3.50]	[0.43]	[0.35]	[-0.84]	[-1.53]	
SYS	5.08	-0.25	-1.43	0.14	-1.07	0.81	0.04
	[3.63]	[-0.46]	[-2.97]	[0.21]	[-1.46]	[0.81]	

Table 3.11: **Fama-French five-factor risk-adjusted returns for FVA portfolios sorted by forward share of systematic variances.** This table shows Fama and French (2015) five-factor alphas and factor loadings for portfolios of U.S. dollar exchange rate FVAs sorted by their forward share of systematic variances. The portfolio construction follows the same procedure as described in Table 3.6. In each panel, columns 1-6 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on the five Fama and French (2015) equity factors: market (MKT), size (SMB), value (HML), profitability (RMW), and investment (CMA). The excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo							Panel B: 1/4 mo						
	α	MKT	SMB	HML	RMW	CMA	R^2	α	MKT	SMB	HML	RMW	CMA	R^2
P_1	1.55	-1.75	-0.51	-0.03	-1.01	-0.48	0.24	0.82	-1.17	-0.43	-0.23	-0.61	-0.09	0.24
	[1.14]	[-3.61]	[-1.90]	[-0.05]	[-2.61]	[-1.01]		[0.88]	[-3.19]	[-2.16]	[-0.72]	[-2.21]	[-0.28]	
P_2	-1.39	-1.69	-0.31	-0.02	-0.71	-0.51	0.22	-0.26	-1.24	-0.40	-0.22	-0.50	-0.15	0.24
	[-1.16]	[-3.92]	[-1.06]	[-0.03]	[-1.66]	[-0.82]		[-0.27]	[-3.64]	[-1.99]	[-0.55]	[-1.61]	[-0.35]	
P_3	-1.52	-1.77	-0.08	-0.46	-0.66	0.20	0.24	-0.73	-1.10	-0.18	-0.44	-0.22	0.08	0.22
	[-1.18]	[-3.77]	[-0.21]	[-0.82]	[-1.45]	[0.34]		[-0.83]	[-2.85]	[-0.79]	[-1.22]	[-0.67]	[0.20]	
FSYS	3.07	0.01	-0.44	0.43	-0.35	-0.68	0.03	1.55	-0.06	-0.25	0.21	-0.39	-0.17	0.02
	[3.99]	[0.09]	[-1.50]	[1.71]	[-1.41]	[-1.60]		[2.60]	[-0.61]	[-1.14]	[1.11]	[-2.01]	[-0.60]	
	Panel C: 6/1 mo							Panel D: 6/4 mo						
P_1	1.14	-0.76	-0.43	-0.50	-0.21	0.21	0.22	0.45	-0.65	-0.25	-0.35	-0.16	0.22	0.22
	[1.63]	[-3.01]	[-2.87]	[-1.67]	[-1.09]	[0.70]		[0.88]	[-3.49]	[-1.98]	[-1.52]	[-0.96]	[0.90]	
P_2	-0.46	-0.79	-0.25	-0.55	-0.02	0.01	0.25	-0.03	-0.69	-0.19	-0.42	-0.03	-0.02	0.24
	[-0.75]	[-3.61]	[-1.77]	[-1.73]	[-0.07]	[0.03]		[-0.05]	[-3.67]	[-1.44]	[-1.39]	[-0.12]	[-0.09]	
P_3	-0.47	-0.86	-0.37	-0.62	-0.11	0.23	0.29	-0.02	-0.72	-0.28	-0.49	-0.05	0.16	0.27
	[-0.76]	[-4.30]	[-2.25]	[-2.21]	[-0.47]	[0.78]		[-0.03]	[-4.08]	[-1.86]	[-1.93]	[-0.23]	[0.60]	
FSYS	1.61	0.10	-0.06	0.12	-0.10	-0.02	0.01	0.46	0.08	0.03	0.13	-0.11	0.06	0.02
	[3.72]	[1.02]	[-0.41]	[0.98]	[-0.76]	[-0.09]		[1.55]	[1.09]	[0.24]	[1.22]	[-0.92]	[0.29]	
	Panel E: 8/1 mo							Panel F: 8/4 mo						
P_1	0.78	-0.65	-0.24	-0.40	-0.13	0.22	0.23	0.64	-0.58	-0.23	-0.32	-0.16	0.21	0.21
	[1.46]	[-3.21]	[-1.87]	[-1.66]	[-0.81]	[0.88]		[1.32]	[-3.30]	[-1.93]	[-1.45]	[-0.98]	[0.90]	
P_2	-0.25	-0.63	-0.23	-0.48	0.01	0.00	0.24	0.06	-0.63	-0.21	-0.39	-0.07	-0.03	0.25
	[-0.53]	[-3.30]	[-1.72]	[-1.84]	[0.03]	[0.01]		[0.12]	[-3.35]	[-1.73]	[-1.56]	[-0.32]	[-0.11]	
P_3	-0.23	-0.72	-0.29	-0.42	-0.10	0.10	0.27	-0.07	-0.65	-0.29	-0.38	-0.09	0.08	0.26
	[-0.46]	[-3.84]	[-2.01]	[-1.91]	[-0.51]	[0.39]		[-0.15]	[-3.52]	[-2.07]	[-1.67]	[-0.45]	[0.33]	
FSYS	1.01	0.07	0.05	0.02	-0.03	0.12	0.01	0.71	0.07	0.05	0.06	-0.07	0.12	0.02
	[2.93]	[0.89]	[0.44]	[0.18]	[-0.31]	[0.62]		[2.25]	[0.96]	[0.47]	[0.50]	[-0.57]	[0.62]	

Table 3.12: **Currency-five factor risk-adjusted returns for volatility swap portfolios sorted by share of systematic variances.** This table shows currency five-factor risk-adjusted returns (αs) and factor loadings for portfolios of 1-month U.S. dollar exchange rate volatility swaps sorted by their share of systematic variances. The portfolio construction follows the same procedure as described in Table 3.4. Columns 1-6 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on five currency factors: the volatility carry factor constructed from volatility swaps (VCA), the low minus high volatility factor constructed from volatility swaps (VLS), the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL). Excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	α	VCA	VLS	MOM	CAR	CDOL	R^2
P_1	-6.50	0.29	0.18	0.20	0.31	-4.93	0.25
	[-3.81]	[3.20]	[1.30]	[0.81]	[0.48]	[-5.09]	
P_2	-8.52	0.09	-0.04	0.47	-1.26	-4.53	0.26
	[-4.59]	[0.75]	[-0.30]	[0.93]	[-1.96]	[-5.54]	
P_3	-10.47	0.39	0.23	0.10	-1.68	-3.81	0.20
	[-5.25]	[2.94]	[1.72]	[0.27]	[-2.64]	[-5.90]	
SYS	3.97	-0.10	-0.05	0.10	1.98	-1.12	0.06
	[3.14]	[-0.75]	[-0.43]	[0.27]	[2.83]	[-1.43]	

Table 3.13: **Currency five-factor risk-adjusted returns of FVAs sorted by forward share of systematic variances.** This table shows currency five-factor factor alphas (risk-adjusted returns) and loadings for portfolios of U.S. dollar exchange rate FVAs sorted by their forward share of systematic variances. The portfolio construction follows the same procedure as described in Table 3.6. In each panel, columns 1-6 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on five currency factors: the volatility carry factor constructed from FVAs (VCA), the low minus high volatility factor constructed from FVAs (VLS), the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL). The FVAs used to construct VCA and VLS match the maturity of the FVAs in the test portfolios. Excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo							Panel B: 1/4 mo						
	α	VCA	VLS	MOM	CAR	CDOL	R^2	α	VCA	VLS	MOM	CAR	CDOL	R^2
P_1	0.32	0.20	-0.01	0.39	0.35	-3.49	0.34	-0.40	0.39	0.05	0.33	0.08	-2.53	0.41
	[0.28]	[1.57]	[-0.07]	[2.60]	[0.70]	[-4.20]		[-0.58]	[4.20]	[0.43]	[2.82]	[0.21]	[-4.69]	
P_2	-1.79	-0.08	-0.21	0.63	-0.27	-3.17	0.34	-0.90	0.15	-0.06	0.49	-0.30	-2.61	0.40
	[-1.77]	[-0.63]	[-1.43]	[4.18]	[-0.52]	[-4.93]		[-1.21]	[1.46]	[-0.36]	[3.97]	[-0.79]	[-5.24]	
P_3	-2.52	0.13	0.11	0.75	-0.02	-3.28	0.32	-1.45	0.25	-0.07	0.44	0.01	-2.42	0.36
	[-2.31]	[1.13]	[0.61]	[4.09]	[-0.03]	[-4.15]		[-2.18]	[2.49]	[-0.60	[3.29]	[0.01]	[-4.33]	
FSYS	2.84	0.07	-0.12	-0.36	0.37	-0.21	0.06	1.05	0.14	0.12	-0.11	0.08	-0.11	0.03
	[4.39]	[0.81]	[-0.74]	[-1.90]	[1.14]	[-0.69]		[2.29]	[1.27]	[0.80]	[-1.00]	[0.38]	[-0.50]	
	Panel C: 6/1 mo							Panel D: 6/4 mo						
P_1	-1.32	0.72	0.18	0.27	-0.02	-1.71	0.55	-0.71	0.60	0.17	0.24	-0.12	-1.41	0.48
	[-2.60]	[8.75]	[1.78]	[3.24]	[-0.07]	[-5.61]		[-2.01]	[7.43]	[1.54]	[3.41]	[-0.67]	[-6.54]	
P_2	-1.59	0.32	0.14	0.29	-0.16	-1.68	0.37	-0.67	0.30	0.08	0.30	-0.16	-1.56	0.39
	[-2.83]	[2.23]	[1.11]	[2.86]	[-0.62]	[-4.72]		[-1.68]	[2.27]	[0.61]	[3.22]	[-0.62]	[-5.68]	
P_3	-1.64	0.34	0.27	0.28	-0.05	-1.77	0.38	-0.72	0.29	0.17	0.26	-0.05	-1.49	0.34
	[-3.03]	[3.05]	[2.76]	[2.90]	[-0.20]	[-5.75]		[-1.64]	[3.06]	[1.48]	[3.18]	[-0.19]	[-5.98]	
FSYS	0.31	0.37	-0.10	-0.01	0.03	0.06	0.20	0.01	0.31	-0.00	-0.02	-0.08	0.08	0.09
	[1.04]	[4.27]	[-1.03]	[-0.15]	[0.22]	[0.42]		[0.04]	[3.62]	[-0.02]	[-0.30]	[-0.50]	[0.57]	
	Panel E: 8/1 mo							Panel F: 8/4 mo						
P_1	-0.73	0.67	0.14	0.21	0.03	-1.44	0.52	-0.60	0.63	0.14	0.20	-0.09	-1.25	0.49
	[-2.05]	[8.76]	[1.49]	[2.91]	[0.15]	[-5.76]		[-1.87]	[8.12]	[1.54]	[3.21]	[-0.51]	[-6.12]	
P_2	-1.13	0.36	0.12	0.26	-0.16	-1.40	0.38	-0.59	0.29	0.00	0.25	-0.14	-1.39	0.40
	[-2.51]	[2.75]	[0.96]	[3.29]	[-0.70]	[-5.01]		[-1.63]	[2.86]	[0.02]	[3.24]	[-0.66]	[-4.99]	
P_3	-1.06	0.30	0.21	0.25	0.00	-1.47	0.36	-0.74	0.28	0.11	0.19	0.04	-1.36	0.33
	[-2.35]	[3.38]	[1.73]	[3.10]	[0.01]	[-5.34]		[-1.88]	[3.28]	[0.91]	[2.49]	[0.15]	[-5.23]	
FSYS	0.33	0.37	-0.07	-0.05	0.03	0.03	0.16	0.14	0.35	0.03	0.01	-0.12	0.11	0.11
	[1.19]	[5.08]	[-0.67]	[-0.79]	[0.19]	[0.21]		[0.57]	[4.15]	[0.27]	[0.13]	[-0.84]	[0.71]	

3.10 Appendix: Supplementary Tables and Figures

This appendix provides results for portfolios constructed based on dollar factor betas.

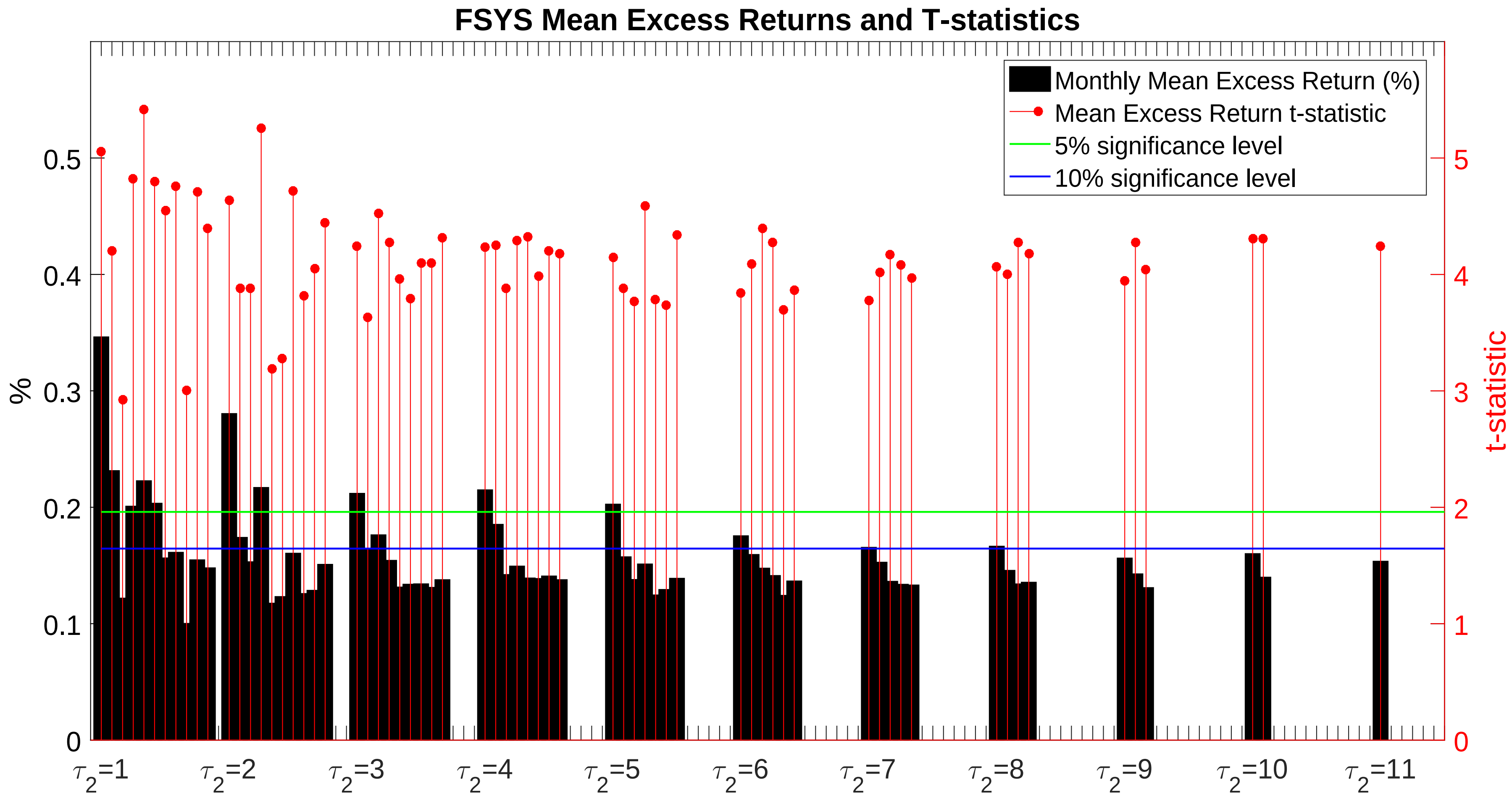

Figure 3.6: **Mean excess returns and t-statistics for FSYS factors constructed based on forward dollar factor betas.** This figure illustrates mean excess returns and corresponding Newey and West (1987) t-statistics for the FSYS factor constructed based on forward dollar factor betas at each maturity/volatility coverage period combination τ_1/τ_2 (66 portfolios). The construction of the FSYS factors based on forward dollar factor betas is described in the caption of Table 3.16.

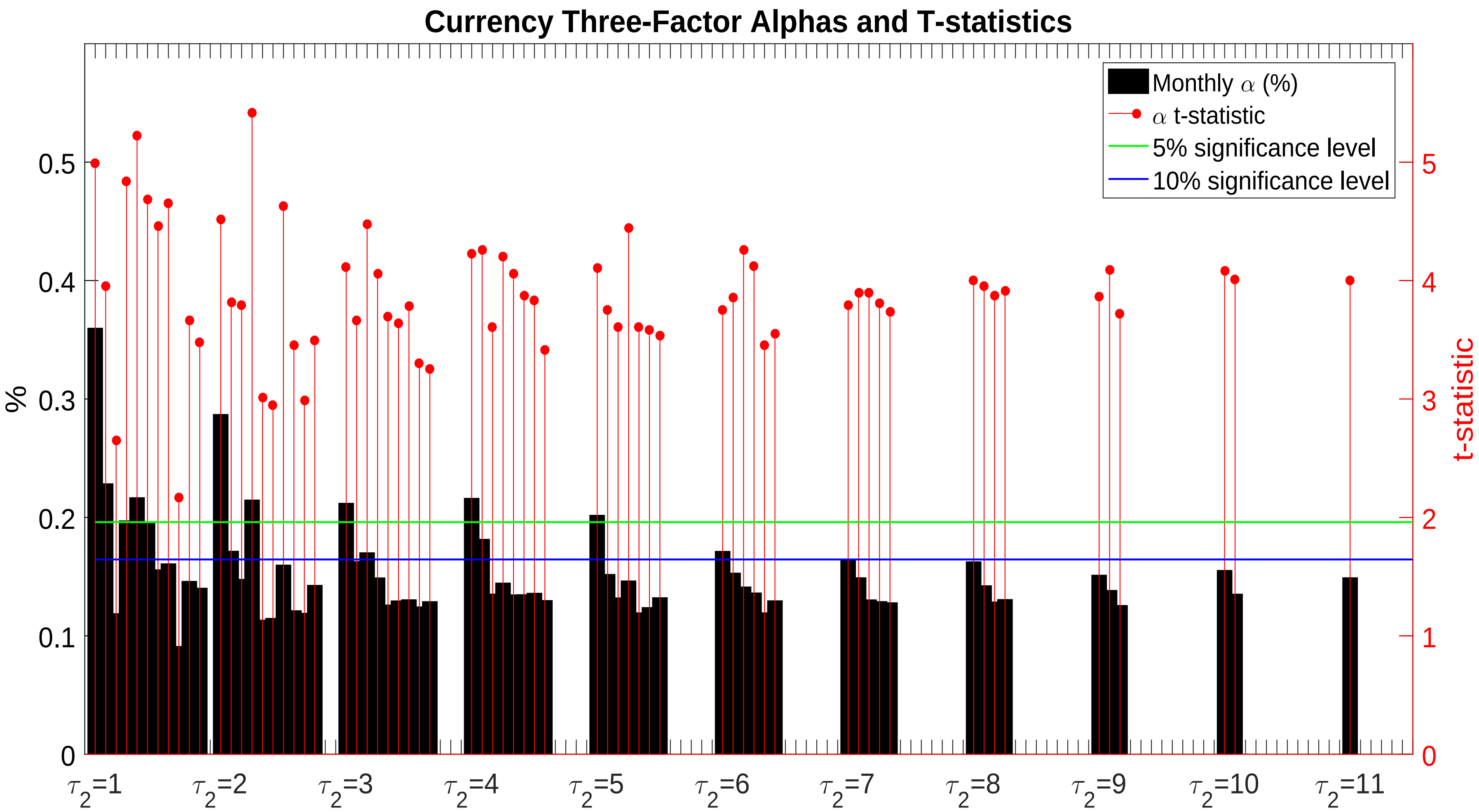

Figure 3.7: **Three-factor currency alphas and t-statistics for FSYS factors constructed based on forward dollar factor betas.** This figure illustrates three-factor currency alphas and corresponding Newey and West (1987) t-statistics for the FSYS factor constructed based on forward dollar factor betas at each maturity/volatility coverage period combination τ_1/τ_2 (66 portfolios). The construction of the FSYS factors based on forward dollar factor betas is described in the caption of Table 3.16. The three currency factors are: the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL).

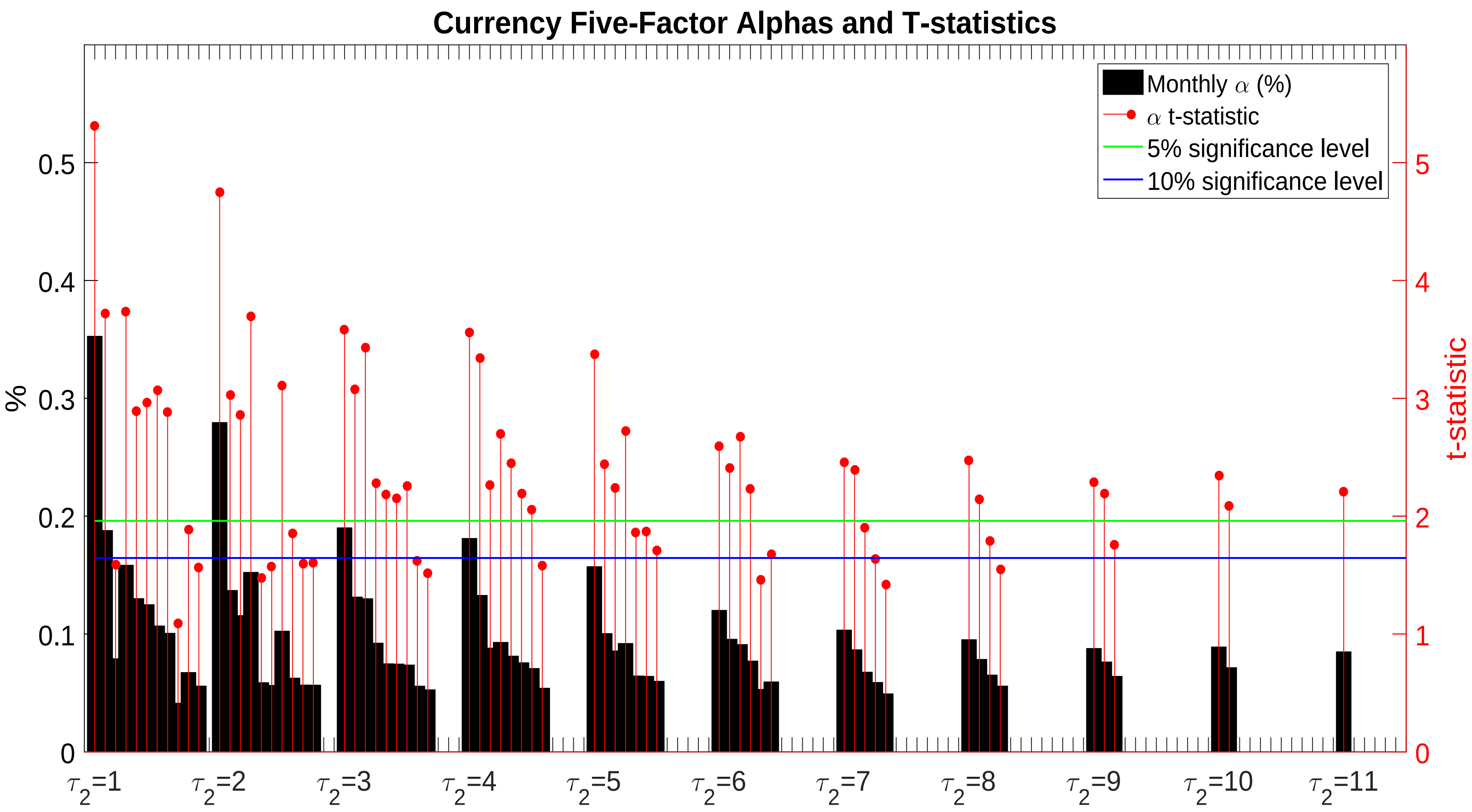

Figure 3.8: **Five-factor currency alphas and t-statistics for FSYS factors constructed based on forward dollar factor betas.**. This figure illustrates five-factor currency alphas and corresponding Newey and West (1987) t-statistics for the FSYS factor constructed based on forward betas at each maturity/volatility coverage period combination τ_1/τ_2 (66 portfolios). The construction of the FSYS factors based on forward dollar factor betas is described in the caption of Table 3.16. The five currency factors are: the FX spot momentum factor (MOM), the high minus low carry factor (CAR), the conditional dollar factor (CDOL), the FVA volatility carry factor (VCA), and the low minus high volatility FVA portfolio (VLS).

Table 3.14: **Excess returns of U.S. dollar volatility swap portfolios sorted by dollar factor betas.** This table shows means, standard deviations (in percentages), skewness, kurtosis, and annualized Sharpe ratios for U.S. dollar volatility swap portfolios sorted by their dollar factor betas. Each month, the volatility swaps are allocated equally into three portfolios from low to high based on their dollar factor betas that match the maturity of the volatility swaps (1, 2, 3, 6, 9 and 12 months). The SYS factor sells the high-beta portfolio and buys the low-beta portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1 mo				Panel B: 2 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-0.86	-1.00	-1.13	0.26	-0.77	-1.05	-1.04	0.27
	[-5.67]	[-4.98]	[-4.18]	[1.47]	[-3.83]	[-4.44	] [-3.18]	[1.43]
Std	2.06	2.48	3.31	2.20	2.14	2.37	3.39	2.06
Skew	1.58	1.28	3.91	-1.94	2.21	0.97	4.24	-2.74
Kurt	14.89	9.06	34.34	13.75	13.78	4.85	32.37	17.39
SR	-1.45	-1.40	-1.18	0.41	-0.88	-1.09	-0.75	0.32
	Panel C: 3 mo				Panel D: 6 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-0.78	-0.99	-1.19	0.41	-0.82	-0.96	-1.49	0.67
	[-3.22]	[-2.88]	[-3.58]	[2.32]	[-2.26]	[-1.82]	[-3.50]	[3.53]
Std	2.22	3.03	2.97	1.71	2.52	3.59	3.01	1.56
Skew	2.10	2.61	2.87	-1.68	1.40	2.21	1.75	-0.59
Kurt	12.08	16.50	20.19	11.55	7.17	10.99	10.43	5.78
SR	-0.70	-0.66	-0.80	0.48	-0.46	-0.38	-0.70	0.61
	Panel E: 9 mo				Panel F: 12 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-0.88	-1.07	-1.58	0.70	-0.89	-1.20	-1.67	0.78
	[-1.97]	[-1.71]	[-2.99]	[3.38]	[-1.81]	[-1.76]	[-2.65]	[3.00]
Std	2.63	3.67	3.16	1.48	2.70	3.63	3.39	1.56
Skew	0.91	1.66	1.24	-0.40	0.56	1.39	1.16	-0.56
Kurt	5.05	7.20	6.73	3.69	3.82	5.55	5.64	3.58
SR	-0.39	-0.34	-0.58	0.55	-0.33	-0.33	-0.49	0.50

Table 3.15: **Excess returns of cross-pair volatility swap portfolios sorted by dollar factor betas.** This table shows means, standard deviations (in percentages), skewness, kurtosis, and annualized Sharpe ratios for cross-pair volatility swap portfolios sorted by their dollar factor betas. Each month, the volatility swaps are allocated equally into three portfolios from low to high based on their dollar factor betas that match the maturity of the volatility swaps (1, 2, 3, 6, 9, and 12 months). The SYS factor sells the high-beta portfolio and buys the low-beta portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1 mo				Panel B: 2 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-0.77	-0.89	-1.03	0.26	-0.64	-0.86	-1.06	0.42
	[-4.16]	[-4.08]	[-3.72]	[1.40]	[-2.44]	[-3.38]	[-3.47]	[2.78]
Std	2.31	2.68	3.76	2.49	2.75	2.62	3.43	1.62
Skew	2.25	3.28	3.78	-3.07	3.64	2.53	2.68	-0.08
Kurt	13.47	30.10	35.33	27.59	25.91	17.32	20.13	6.57
SR	-1.15	-1.15	-0.95	0.35	-0.57	-0.80	-0.76	0.64
	Panel C: 3 mo				Panel D: 6 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-0.64	-0.81	-1.11	0.47	-0.67	-0.80	-1.40	0.73
	[-2.16]	[-2.43]	[-3.16]	[2.95]	[-1.66]	[-1.68]	[-3.01]	[3.98]
Std	2.75	3.07	3.32	1.48	2.87	3.37	3.38	1.40
Skew	2.80	3.60	2.39	-0.22	1.90	2.21	1.64	-0.05
Kurt	15.64	25.65	15.65	3.88	8.97	11.34	9.52	3.48
SR	-0.47	-0.53	-0.67	0.64	-0.33	-0.34	-0.59	0.74
	Panel E: 9 mo				Panel F: 12 mo			
	P_1	P_2	P_3	SYS	P_1	P_2	P_3	SYS
Mean	-0.69	-0.92	-1.57	0.87	-0.76	-1.07	-1.66	0.90
	[-1.44]	[-1.71]	[-2.71]	[4.00]	[-1.46]	[-1.86]	[-2.43]	[3.17]
Std	2.93	3.27	3.50	1.35	2.87	3.11	3.73	1.58
Skew	1.31	1.42	1.19	0.09	0.88	0.93	0.97	-0.36
Kurt	6.09	6.93	6.65	2.98	4.54	4.88	5.38	3.93
SR	-0.27	-0.33	-0.52	0.75	-0.26	-0.34	-0.44	0.57

Table 3.16: **Excess returns of U.S. dollar FVA portfolios sorted by forward dollar factor betas.** This table shows means, standard deviations (in percentages), skewness, kurtosis, and annualized Sharpe ratios for portfolios of U.S. dollar forward volatility agreements (FVAs) sorted by their forward dollar factor betas. Each month, the FVAs are allocated equally into three portfolios from low to high based on their forward dollar factor betas that match the maturity/volatility coverage period of the FVAs. The FSYS factor sells the high-beta portfolio and buys the low-beta portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo				Panel B: 1/4 mo			
	P_1	P_2	P_3	FSYS	P_1	P_2	P_3	FSYS
Mean	-0.02	-0.25	-0.36	0.35	-0.04	-0.11	-0.26	0.21
	[-0.12]	[-1.46]	[-2.34]	[4.92]	[-0.42]	[-0.99]	[-2.57]	[4.27]
Std	1.70	2.07	2.03	1.00	1.19	1.47	1.47	0.74
Skew	4.32	3.87	2.97	-0.17	3.08	3.23	2.28	-0.08
Kurt	40.89	34.22	25.77	3.39	28.06	27.60	20.81	3.24
SR	-0.04	-0.42	-0.62	1.20	-0.12	-0.26	-0.61	1.01
	Panel C: 6/1 mo				Panel D: 6/4 mo			
Mean	-0.00	-0.10	-0.21	0.20	-0.00	-0.06	-0.14	0.14
	[-0.04]	[-1.02]	[-2.16]	[4.39]	[-0.04]	[-0.92]	[-2.14]	[3.88]
Std	0.99	1.13	1.13	0.60	0.83	0.95	0.97	0.52
Skew	1.47	1.62	0.84	0.36	1.27	1.24	0.53	0.25
Kurt	11.62	12.27	8.62	3.84	9.92	8.93	7.78	3.48
SR	-0.01	-0.30	-0.63	1.17	-0.01	-0.22	-0.51	0.92
	Panel E: 8/1 mo				Panel F: 8/4 mo			
Mean	-0.00	-0.05	-0.16	0.16	0.02	-0.06	-0.12	0.14
	[-0.05]	[-0.74]	[-2.31]	[4.73]	[0.32]	[-0.95]	[-2.01]	[3.89]
Std	0.81	0.93	0.94	0.51	0.75	0.86	0.89	0.51
Skew	1.28	1.40	0.78	0.27	1.27	1.45	0.78	0.40
Kurt	9.98	10.04	9.15	3.60	10.43	10.14	9.36	4.32
SR	-0.01	-0.20	-0.61	1.09	0.08	-0.24	-0.47	0.94

Table 3.17: **Excess returns of cross-pair FVA portfolios sorted by forward dollar factor betas.** This table shows means, standard deviations (in percentages), skewness, kurtosis, and annualized Sharpe ratios for portfolios of cross-pair forward volatility agreements (FVAs) sorted by their forward dollar factor betas. Each month, the FVAs are allocated equally into three portfolios from low to high based on the forward dollar factor betas that match the maturity/volatility coverage period of the FVAs. The FSYS factor sells the high-beta portfolio and buys the low-beta portfolio. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo				Panel B: 1/4 mo			
	P_1	P_2	P_3	FSYS	P_1	P_2	P_3	FSYS
Mean	-0.14	-0.24	-0.27	0.13	-0.08	-0.17	-0.20	0.12
	[-0.99]	[-1.77]	[-1.53]	[1.95]	[-0.85]	[-1.53]	[-1.66]	[2.60]
Std	1.75	1.74	2.36	0.99	1.12	1.29	1.62	0.79
Skew	5.12	3.36	5.56	-1.18	4.06	4.33	4.04	-0.41
Kurt	51.38	29.39	58.40	15.90	37.51	41.59	40.08	12.22
SR	-0.29	-0.49	-0.40	0.44	-0.25	-0.44	-0.44	0.54
	Panel C: 6/1 mo				Panel D: 6/4 mo			
Mean	-0.08	-0.13	-0.16	0.08	-0.06	-0.08	-0.13	0.07
	[-0.94]	[-1.47]	[-1.45]	[2.05]	[-1.13]	[-1.53]	[-1.61]	[2.15]
Std	0.90	1.00	1.27	0.63	0.73	0.80	1.05	0.60
Skew	2.31	2.13	2.19	-0.01	1.83	1.88	1.53	0.56
Kurt	18.27	15.89	17.11	7.83	13.46	14.27	12.37	9.57
SR	-0.30	-0.47	-0.44	0.44	-0.30	-0.36	-0.42	0.38
	Panel E: 8/1 mo				Panel F: 8/4 mo			
Mean	-0.06	-0.09	-0.14	0.08	-0.03	-0.08	-0.09	0.06
	[-0.95]	[-1.27]	[-1.95]	[2.58]	[-0.68]	[-1.29]	[-1.24]	[1.77]
Std	0.73	0.79	1.04	0.61	0.67	0.73	1.02	0.67
Skew	1.91	2.07	1.80	0.49	1.73	1.52	1.60	0.40
Kurt	13.44	14.88	13.24	9.71	13.30	11.69	12.59	10.63
SR	-0.28	-0.38	-0.46	0.45	-0.18	-0.36	-0.31	0.30

Table 3.18: **Currency-five factor risk-adjusted returns for volatility swap portfolios sorted by dollar factor betas.** This table shows currency five-factor risk-adjusted returns (αs) and factor loadings for portfolios of 1-month U.S. dollar exchange rate volatility swaps sorted by their dollar factor betas. The portfolio construction follows the same procedure as described in Table 3.14. Columns 1-6 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on five currency factors: the volatility carry factor constructed from volatility swaps (VCA), the low minus high volatility factor constructed from volatility swaps (VLS), the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL). Excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	α	VCA	VLS	MOM	CAR	CDOL	R^2
P_1	-0.62	0.02	0.03	0.03	0.04	-0.42	0.24
	[-3.97]	[3.28]	[2.36]	[1.27]	[0.48]	[-4.07]	
P_2	-0.88	0.01	-0.00	0.06	-0.20	-0.39	0.22
	[-4.62]	[0.60]	[-0.16]	[1.05]	[-2.72]	[-5.73]	
P_3	-0.92	0.06	0.00	0.00	0.11	-0.72	0.32
	[-4.09]	[4.12]	[0.40]	[0.01]	[0.65]	[-3.08]	
SYS	0.30	-0.04	0.02	0.03	-0.07	0.30	0.26
	[1.95]	[-3.23]	[2.26]	[0.92]	[-0.68]	[1.99]	

Table 3.19: **Currency five-factor risk-adjusted returns of FVAs sorted by forward dollar factor betas.** This table shows currency five-factor alphas (risk-adjusted returns) and factor loadings for portfolios of U.S. dollar exchange rate FVAs sorted by their forward dollar factor betas. The portfolio construction follows the same procedure as described in Table 3.16. In each panel, columns 1-6 show the intercepts (αs) and loadings for time-series regressions of portfolio excess returns regressed on five currency factors: the volatility carry factor constructed from FVAs (VCA), the low minus high volatility factor constructed from FVAs (VLS), the FX spot momentum factor (MOM), the high minus low carry factor (CAR), and the conditional dollar factor (CDOL). The FVAs used to construct VCA and VLS match the maturity of the FVAs in the test portfolios. Excess returns are monthly and in percentages. The numbers in brackets are t-statistics based on Newey and West (1987). The G10 options data are from JP Morgan Dataquery and the sample period is from January 1998 to August 2016 and comprise 216 monthly observations.

	Panel A: 1/1 mo							Panel B: 1/4 mo						
	α	VCA	VLS	MOM	CAR	CDOL	R^2	α	VCA	VLS	MOM	CAR	CDOL	R^2
P_1	0.00	0.02	0.00	0.05	0.12	-0.42	0.37	-0.08	0.04	0.02	0.04	0.05	-0.30	0.41
	[0.03]	[1.56]	[0.16]	[2.22]	[1.27]	[-3.60]		[-1.10]	[4.51]	[1.22]	[2.74]	[0.77]	[-4.01]	
P_2	-0.19	0.00	-0.00	0.10	0.03	-0.49	0.38	-0.08	0.02	-0.02	0.07	-0.01	-0.37	0.42
	[-1.27]	[0.25]	[-0.13]	[4.47]	[0.30]	[-3.27]		[-0.85]	[2.25]	[-0.81]	[3.98]	[-0.12]	[-3.88]	
P_3	-0.35	0.01	-0.02	0.09	0.08	-0.45	0.33	-0.26	0.04	-0.03	0.06	0.02	-0.33	0.37
	[-2.66]	[1.09]	[-1.69]	[3.43]	[0.73]	[-3.46]		[-2.97]	[2.49]	[-2.02]	[3.20]	[0.25]	[-3.87]	
FSYS	0.35	0.00	0.02	-0.04	0.03	0.02	0.09	0.18	0.00	0.04	-0.02	0.03	0.03	0.16
	[5.31]	[0.28]	[1.65]	[-2.27]	[0.94]	[0.69]		[3.56]	[0.31]	[2.62]	[-2.08]	[1.00]	[1.10]	
	Panel C: 6/1 mo							Panel D: 6/4 mo						
P_1	-0.17	0.07	0.03	0.03	0.01	-0.19	0.41	-0.07	0.05	0.03	0.03	0.00	-0.17	0.40
	[-2.90]	[4.48]	[1.93]	[2.22]	[0.40]	[-4.70]		[-1.78]	[4.44]	[1.95]	[2.81]	[0.07]	[-5.16]	
P_2	-0.22	0.06	0.03	0.05	-0.04	-0.24	0.44	-0.11	0.05	0.02	0.04	-0.04	-0.21	0.44
	[-3.38]	[3.12]	[2.17]	[3.77]	[-0.89]	[-4.59]		[-2.26]	[3.57]	[1.56]	[4.02]	[-1.06]	[-5.19]	
P_3	-0.29	0.04	0.01	0.04	-0.01	-0.23	0.34	-0.15	0.03	-0.01	0.04	-0.01	-0.20	0.33
	[-3.80]	[1.99]	[0.51]	[2.97]	[-0.31]	[-5.33]		[-2.52]	[1.88]	[-0.86]	[3.56]	[-0.23]	[-5.55]	
FSYS	0.12	0.02	0.02	-0.01	0.02	0.03	0.09	0.08	0.02	0.05	-0.01	0.01	0.03	0.16
	[2.97]	[2.12]	[2.02]	[-1.28]	[0.97]	[1.86]		[2.19]	[2.18]	[3.77]	[-1.15]	[0.50]	[1.58]	
	Panel E: 8/1 mo							Panel F: 8/4 mo						
P_1	-0.09	0.06	0.03	0.02	0.01	-0.17	0.40	-0.06	0.06	0.03	0.02	0.00	-0.16	0.41
	[-2.17]	[5.00]	[1.75]	[2.28]	[0.34]	[-4.96]		[-1.64]	[5.70]	[2.12]	[2.58]	[0.12]	[-4.81]	
P_2	-0.14	0.06	0.03	0.04	-0.03	-0.20	0.44	-0.11	0.05	0.01	0.03	-0.03	-0.18	0.41
	[-2.50]	[3.76]	[1.89]	[3.92]	[-0.99]	[-4.43]		[-2.37]	[3.60]	[0.96]	[3.26]	[-0.87]	[-4.37]	
P_3	-0.19	0.03	-0.00	0.04	-0.00	-0.20	0.34	-0.12	0.02	-0.02	0.03	-0.01	-0.19	0.34
	[-3.21]	[2.24]	[-0.34]	[3.28]	[-0.02]	[-4.87]		[-2.10]	[1.96]	[-1.84]	[3.16]	[-0.21]	[-5.02]	
FSYS	0.10	0.03	0.03	-0.01	0.01	0.03	0.11	0.05	0.03	0.05	-0.01	0.01	0.03	0.20
	[2.88]	[3.34]	[2.88]	[-1.59]	[0.54]	[1.54]		[1.58]	[3.46]	[4.21]	[-1.17]	[0.55]	[2.23]	

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