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ESSAYS ON ASSET PRICING WITH FINANCIAL FRICTIONS

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Sven Klingler

ESSAYS ON ASSET PRICING WITH FINANCIAL FRICTIONS

The PhD School in Economics and Management

PhD Series 28.2017

CBS  COPENHAGEN BUSINESS SCHOOL
HANDELSHØJSKOLEN

Essays on Asset Pricing with Financial Frictions

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Foreword

This thesis is the product of my PhD studies at the Department of Finance and Center for Financial Frictions (FRIC) at Copenhagen Business School. The thesis consists of three self-contained essays, which can be read independently. The common theme throughout the thesis is the effect of funding frictions on asset prices.

The first essay (co-authored with David Lando) shows how financial regulation creates a demand for credit default swap (CDS) contracts on safe sovereigns. Derivatives-dealing banks either face tighter funding conditions or purchase CDS contracts on safe sovereigns to free regulatory capital. The second essay (co-authored with Suresh Sundaresan) provides evidence that underfunded pension plans, which face funding constraints because they are restricted from using direct leverage, have a demand for long-dated interest rate swaps. This demand by underfunded pension plans can explain the persistent negative 30-year swap spread. The third essay uses deviations from the covered interest rate parity (CIP) as a proxy for market-wide funding conditions and shows that hedge funds with higher exposure to that risk underperform funds with lower exposure. That is, hedge funds with higher exposure to funding frictions, generate lower risk-adjusted returns than hedge funds with a lower exposure to that risk.

This thesis has benefited from the advice, helpful comments, and support of many people over the years. I would like to take this opportunity and thank some of them. First and foremost, I am grateful to my two advisors David Lando and Lasse Pedersen for their support. David Lando's help, advice, and guidance, as well as his positivity, is what enabled me to write this thesis. Many fruitful discussions with Lasse Pedersen helped me to develop and sharpen my research ideas. I also want to thank Suresh Sundaresan for hosting my visits at Columbia University and for our joint efforts in explaining negative swap spreads. I am also grateful to Bob Hodrick for his valuable feedback.

I would also like to thank my friends and colleagues for their support. Here, I am especially grateful to Simon Rottke and Valeri Sokolovsky, for many great discussions and good times. I am also indebted to Nigel Baradalle and Christian Wagner for their valuable feedback; as well as to Davide Tomio, Desi Volker, and Aleksandra Rzenik for their support.

Additionally to that, I want to thank my colleagues and fellow PhD students for making work at CBS so enjoyable. Finally, I am grateful to Frank Fabozzi and Marliese Uhrig-Homburg, who got me interested in an academic career in Finance in the first place, and who helped me getting a position at Copenhagen Business School.

Most importantly, I am grateful to my parents for their support and for always believing in me, even in times when I did not.

Sven Klingler

Copenhagen, April 2017

Summary

Summary in English

Essay 1: Safe-Haven CDS Premiums (co-authored with David Lando)

The first essay focuses on Credit default swap (CDS) premiums of safe sovereigns, that is, the insurance against the default of countries with a low credit risk, like Germany, Japan, or the United States. We motivate the essay by establishing the following two stylized facts. First, we document that there is a large market for insurance against the default of safe sovereigns and that the CDS premiums for these sovereigns are substantial, sometimes exceeding 100 basis points. Second, we show that there is virtually no relationship between CDS premiums and bond yield spreads, which are measured as the spread between bond yield and risk-free rate, for safe sovereigns. This finding is in opposition to the no-arbitrage theory that CDS premiums and yield spreads should move in lockstep. Motivated by these stylized facts, we investigate the following two questions: First, what are the motives behind purchasing insurance against the default of safe sovereigns? Second, what drives safe-haven CDS premiums if not credit risk?

Our answer to the first question is that financial regulation provides an incentive for derivatives-dealing banks to purchase sovereign CDS. Basel III, the new financial regulation, introduces a capital charge against uncollateralized OTC derivatives (even if the counterparty is a safe sovereign) and gives derivatives dealing banks a choice between facing an addition to regulatory capital or purchasing CDS. We analyze this regulatory friction in an equilibrium model where a derivatives-dealing bank demands CDS to free regulatory capital. Due to a lack of natural CDS sellers, an end-user of derivatives provides the CDS to the bank. Because selling CDS is costly for end users due to an associated margin requirement, the end user demands a positive CDS premium, even for a risk-free sovereign. The model enables us to derive an equilibrium CDS premium that depends on the demand for freeing regulatory capital by banks and the opportunity cost of selling CDS. Hence, the answer to our second question is that regulatory frictions can be a major driver of sovereign CDS

premiums.

We provide five pieces of evidence in favor of our hypothesis. First, we document that uncollateralized derivatives positions are subject to the Basel III capital charge and that banks use sovereign CDS to avoid this capital charge. Second, we provide sample calculations to show that the orders of magnitude of CDS outstanding are consistent with the derivatives hedging motive for Germany. Third, we document that derivatives dealers are net *buyers* of sovereign CDS, as opposed to being net sellers of CDS, which is common in most other markets. Fourth, we show that there is no link between CDS premium and bond yield for safe sovereigns, but that the link between the two becomes stronger for more risky sovereigns. Finally, we find that regulatory proxies for the exposure towards sovereigns and for banks' financial constraints are significant drivers of sovereign CDS.

Essay 2: An Explanation of Negative Swap Spreads (co-authored with Suresh Sundaresan)

The second essay offers an explanation for the persistent negative 30-year swap spreads that have been observed in the U.S. after the default of Lehman Brothers. Swap spreads are the difference between the fixed rate in an interest rate swap (IRS) and the yield of a Treasury bond with the same maturity and should be positive according to the following argument. In an IRS, a fixed rate is exchanged against LIBOR payments, which contain a credit-risk component. Hence, to compensate for this credit risk, the swap rate should be above the risk-free rate. Furthermore, the Treasury yield should be below the risk-free rate because investors value the safety and liquidity of U.S. government bonds and therefore accept a rate below the risk-free rate for the convenience of holding such an asset. Despite this intuition, the 30-year swap spread turned negative in 2008 and is still negative as of today.

We provide an explanation for this pricing anomaly by linking swap spreads to the demand for duration hedging from underfunded pension plans. Pension funds have liabilities with a long duration, namely pension obligations to their clients. To hedge this duration risk, they can either purchase long-dated bonds or receive fixed in a long-dated IRS. The advantage of using IRS instead of investing in long-dated bonds is that an IRS has an initial value of zero and, hence, no initial investment is required. Because pension funds are not allowed to use direct leverage, this advantage is relevant. Underfunding is important because, if pension funds are underfunded, they want to shift toward more risky asset holdings in an attempt to generate higher expected returns, which increase the probability of becoming fully funded again. To do so, pension funds shift their asset allocation from safe long-term bonds to stocks. This shift toward stocks causes a mismatch in the fund's asset duration

and liability duration and one way of hedging this duration mismatch is to receive fixed in a long-dated IRS, which, in contrast to holding bonds, requires no initial investment.

To shed light on this channel, we use data from the financial accounts of the U.S. to construct an aggregate measure of the underfunded ratio (UFR) for U.S. pension funds. We then show that UFR is a significant explanatory variable for 30-year swap spreads, even after controlling for other commonly-used drivers of swap spreads, such as market volatility and the term premium. Additionally to that, we find that UFR is more significant in regimes when pension funds are underfunded compared to periods when pension funds are fully funded. Furthermore, we document that the measure is statistically and economically significant for 30-year swap spreads, but not for swap spreads with shorter maturities.

Essay 3: High Funding Risk, Low Return

The third essay shows that hedge funds with a higher exposure to a simple risk measure generate lower returns than hedge funds with a lower exposure to the same measure. The risk measure is based on deviations from the covered interest rate parity (CIP), and I show that the measure spikes when market-wide funding conditions deteriorate. This “high funding risk, low return” finding raises two central questions that I investigate in the essay: Why does a hedge fund manager choose a higher exposure to this risk without getting compensated for it? Why do investors put their money in funds with greater risk without getting compensated for the additional risk?

To answer these questions, I start by developing a simple model in which hedge fund managers with access to less profitable strategies invest more aggressively in their strategies in an attempt to generate competitive returns. By investing more aggressively, these fund managers hold a lower cash buffer against deteriorating funding conditions, thereby having a higher exposure to funding shocks. This increased risk lowers the funds’ expected returns but enhances the probability that the fund generates returns that are competitive with those of better managers, that is managers with access to more profitable strategies. In the model, investors are initially unaware of the managers’ risk-taking and withdraw from funds with access to less profitable strategies if they do not generate competitive returns.

Empirically, I find that hedge funds with a high loading on the CIP risk measure severely underperform hedge funds with a low loading on that measure. The average difference in risk-adjusted returns between high-risk and low-risk funds is 0.54% (t -statistic of 2.46). Moreover, in line with the model’s predictions, I document that hedge funds with a high loading on the funding risk measure experience more equity withdrawals than funds with a low loading on that measure. Additionally to that, the returns of high-risk funds are more

sensitive to equity withdrawals, which supports my hypothesis that high-risk funds hold a lower cash buffer against unexpected funding shocks. Furthermore, the link between a high loading on the CIP risk measure and lower expected returns is less significant for funds which face a lower risk of being exposed to funding shocks, that is funds that impose stricter redemption terms on their investors or funds that have multiple prime brokers.

Summary in Danish

Essay 1: Safe-Haven CDS-præmier (medforfattet af David Lando)

Det første afsnit omhandler Credit default swap (CDS) præmier skrevet på sikre stater, dvs. en forsikring mod fallit af lande med lav kreditrisiko, såsom Tyskland, Japan, eller USA. Vi motiverer dette afsnit med to stiliserede fakta. For det første, sådokumenterer vi, at der er et stort marked for forsikring mod fallit af sikre stater og, at CDS-præmier er betydelige—til tider overskrider de 100 basis point. For det andet, såviser vi, at der stort set ikke er nogen sammenhæng mellem CDS-præmier og obligationsspænd, der er målt som forskellen mellem den effektive obligationsrente og den risikofrie rente for sikre stater. Dette empiriske faktum er i modsætning til teorien om ingen arbitrage, hvor CDS-præmier og obligationsspænd skal bevæge sig perfekt med hinanden. Motiveret af disse to stiliserede fakta undersøger vi de følgende to spørgsmål. Først, hvad er motiverne bag køb af forsikring mod fallit af sikre stater? For det andet, hvad driver CDS-præmier på sikre stater, hvis det ikke er kreditrisiko?

Svaret på det første spørgsmål er, at finansiell regulering giver et incitament for derivathandlende banker til at købe CDS påstater. Basel III, den nye finansielle regulering, introducerer, at der er kapitalomkostninger forbundet ved handel med OTC-derivater uden sikkerhedsstillelse (selv hvis modparten er en sikker stat), og efterlader derved derivathandlende banker i valget mellem at øge den regulatoriske kapital, eller at erhverve CDS. Vi analyserer denne regulatoriske friktion i en generel ligevægtsmodel, hvor derivathandlende banker efterspørger CDS for at frigøre regulatorisk kapital. På grund af mangel på naturlige sælgere af CDS, såudbyder slutbrugere af derivater CDS til banken. Eftersom det er bekosteligt at sælge CDS for slutbrugerne pga. et associeret margin krav, såkræves af slutbrugerne en positiv CDS-præmie, selv hvis staten er risikofri. Modellen muliggør at beregne en ligevægts CDS-præmie, der afhænger af efterspørgslen for at frigøre regulatorisk kapital og offeromkostningerne ved at sælge CDS. Derved er svaret til vores andet spørgsmål, at regulatoriske friktioner kan være en afgørende faktor for CDS-præmier påstater.

Vi fremlægger fem stykker af beviser, der understøtter vores hypotese. Først, sådokumenterer vi, at positioner i derivater uden sikkerhedsstillelse er genstand for en Basel III kapi-

talomkostning, og at banker benytter CDS til at undgå denne kapitalomkostning. For det andet, så laver vi stikprøveberegninger, der viser at størrelsen på de stående CDS er konsistent med et hedging motiv i tilfældet af Tyskland. For det tredje, så viser vi, at derivathandlere er nettokøbere af CDS påstater, i modsætning til nettokøbere af CDS, der er typisk for de fleste andre markeder. For det fjerde, så viser vi, at der ikke er en forbindelse mellem CDS-præmier og obligationsspænd for sikre stater, men at denne sammenhæng er mere markant for mere risikable stater. Afslutningsvis, så finder vi, at regulatoriske proxier for eksponering imod stater og for bankers finansielle begrænsninger er væsentlige drivkræfter for CDS påstater.

Essay 2: En forklaring på negative swap spread (medforfattet af Suresh Sundaresan)

Essay nummer to kommer med en forklaring på det negative 30-års swap spread, der konsistent har været observeret i USA efter Lehman Brothers fallit. Swap spreads er forskellen på den faste rente i et rentebyt og yieldet på en statsobligation med samme løbetid. Denne faste rente, bytterenten, burde være positivt jf. følgende argument. I et rentebyt bytter man en fast rente mod LIBOR renten, som inderholder kreditrisiko. For at kompensere for denne kreditrisiko bør bytterenten altså være højere end den risikofri rente. Derudover bør renten på statsobligationer være lavere end den risikofri rente, da investorer sætter pris på den sikkerhed og likviditet som statsobligationer tilbyder. På trods af denne intuition har det 30-årige swap spread dog været negativt siden 2008 og er det stadig i dag.

Vi kommer med en forklaring på denne anomali ved at linke swap spreads til efterspørgslen for løbetidshedging fra underfinansierede pensionsplaner. Pensionskasser har passiver med lang løbetid jf. deres pensionsforpligtelser til deres kunder. For at hedge denne løbetidsrisiko kan pensionskasserne enten købe obligationer med lang løbetid eller modtage en bytterente med lang løbetid. Fordelen ved at modtage en bytterente frem for at købe obligationer er at et rentebyt ikke kræver nogen investering. Da pensionskasser ikke må bruge gearing direkte, er denne fordel relevant. Underfinansiering er vigtigt, fordi det gør at pensionskasserne vil holde mere risikable aktiver for at forsøge at generere højere afkast, da dette øger sandsynligheden for at de bliver fuldt finansierede igen. For at opnå dette skifter pensionskasserne fra sikre obligationer med lang løbetid til aktier. Dette skift mod aktier skaber en skævvridning mellem løbetiden på fondens aktiver og løbetiden på fondens passiver, og en måde hvorpå man kan rette op på denne skævvridning, er ved at modtage bytterenten i et rentebyt, hvilket i modsætning til at holde obligationer, ikke kræver nogen investering.

Vi bruger data for regnskaber fra USA til at skabe et aggregeret mål for underfinansier-

ing af pensionskasser (UFR), og vi bruger dette mål til at kaste lys på underfinansierings effekt på swap spreads. Vi viser, at UFR signifikant forklarer det 30-årige swap spread, selv efter vi kontrollerer for andre variable der normalt bliver brugt til at forklare swap spreads, såsom markedsvolatilitet og rentestrukturspræmien. Derudover finder vi at UFR er mere signifikant i regimer hvor pensionskasser er underfinansieret relativt til regimer hvor pensionskasser er fuldt finansierede. Derudover dokumenterer vi, at målet er statistisk og økonomisk signifikant for 30-års swap spreads men ikke for swap spreads med kort løbetid.

Essay 3: Høj finansieringsrisiko, lave afkast

Det tredje essay viser at hedge fonde som er stærkt eksponerede mod et simpelt risikomål genererer lavere afkast end hedge fonde med en lavere eksponering mod det samme risikomål. Risikomålet er baseret på afvigelser i den dækkede renteparitet (CIP), og jeg viser at målet slår ud når mulighederne for at opnå finansiering forværres. Denne observation af "høj finansieringsrisiko, lavt afkast" leder til to centrale spørgsmål som jeg vil undersøge i dette essay. Først, hvorfor vælger en bestyrer af en hedge fond at påtage sig denne eksponering uden at blive kompenseret for det? For det andet, hvorfor vælger investorer at placere penge hos fonde som er eksponeret mod denne risiko uden at blive kompenseret for det?

For at besvare disse spørgsmål udvikler jeg først en simpel model hvori hedge fond bestyrere med relativt mindre profitable strategier, investerer mere aggressivt i et forsøg på at generere konkurrencedygtige afkast. Ved at investere mere aggressivt holder disse investorer en lavere kontantbeholdning som beskyttelse mod højere finansieringsomkostninger og er derved mere eksponerede mod et potentielt chok til finansieringsomkostningerne. Denne øgede risiko mindsker fondens forventede afkast, men øger sandsynligheden for at fondens afkast er af samme størrelse som afkast fra bedre fonde, dvs. fonde med mere profitable handelsstrategier. I modellen er investorerne som udgangspunkt ikke klar over fondens risikoeksponering, og trækker deres investering ud af fonden såfremt fonden ikke leverer et konkurrencedygtigt afkast.

Empirisk finder jeg at hedge fonde med høj eksponering mod denne finansieringsrisiko underpræsterer i forhold til hedge fonde med relativt lavere eksponering. Den gennemsnitlige forskel i risikojusterede afkast mellem høj- og lav-risiko fonde er 0,54% (t-teststørrelse på 2.46). Ydermere kan jeg i overensstemmelse med modellens prædiktioner, dokumentere at fonde med høj eksponering mod finansieringsrisikomålet oplever større udstrømning af egenkapital i forhold til fonde med relativt lavere eksponering mod dette mål. Det vises også at afkastene for disse højrisiko fonde er mere følsomme overfor udstrømninger i egenkapitalen, hvilket styrker min hypotese om at højrisiko fonde har lavere kontantbeholdninger i

beskyttelse mod chok til fondens finansieringsmuligheder. Til sidst vises det at sammenhængen mellem eksponering mod CIP risikomålet og forventede afkast, er mindre signifikant for fonde som er mindre eksponerede overfor chok til deres finansieringsmuligheder, det være sig fonde med strengere regler vedrørende udbetaling til investorer eller fonde med flere prime brokers.

Introduction

This thesis consists of three self-contained essays, which can be read independently. All three essays investigate the influence of funding frictions on asset prices and financial markets. The first essay (co-authored with David Lando) shows how financial regulation creates a demand for credit default swap (CDS) contracts on safe sovereigns. Derivatives-dealing banks either face tighter funding conditions or purchase CDS contracts on safe sovereigns to free regulatory capital and alleviate their funding constraint. The second essay (co-authored with Suresh Sundaresan) shows that underfunded pension plans, which face funding constraints because they are restricted from using direct leverage, have a demand for long-dated interest rate swaps. This demand by underfunded pension plans is an explanation for the persistent negative 30-year swap spread. The third essay uses deviations from the covered interest rate parity (CIP) as a proxy for market-wide funding conditions and shows that hedge funds with higher exposure to funding risk underperform funds with lower exposure to that risk.

Safe-Haven CDS Premiums

In the first essay, we argue that credit risk plays a limited role in explaining the Credit Default Swap (CDS) premiums on safe sovereigns, and that the level of the sovereign CDS premiums for safe sovereigns is more likely due to financial regulation. Derivatives-dealing banks engage in OTC derivatives, such as interest rate swaps, with sovereigns. Most sovereigns do not post collateral in these transactions and this leaves the dealer banks exposed to counterparty-credit risk. This risk adds to the dealer banks' risk-weighted assets (RWAs) even when the sovereign is safe, because counterparty risk for regulatory purposes is quantified using CDS premiums. As long as there is some credit risk and hence a positive CDS premium, however small, this creates an incentive for dealer banks to buy CDS protection with the purpose of obtaining capital relief. But selling CDS, even on an almost risk-free entity, is not cost-free. The seller of the CDS must use a share of his own capital to provide the initial margin, and the opportunity cost of providing this margin causes the seller to require a positive CDS premium. This creates a type of self-fulfilling prophecy in which the

CDS premium settles at a significantly higher level than what can be explained by credit risk alone.

We build a one-period model in which two agents face different margin and capital constraints. The first agent is a derivatives-dealing bank who is engaged in a derivatives transaction with a sovereign. Due to regulatory requirements, this derivatives transaction adds to the dealer banks's RWA, thereby lowering its capital available for other investments. To free up capital, the bank can buy CDS on the sovereign. The second agent is an end user of derivatives, who has no exposure to the sovereign and acts as seller of credit protection. The end user weighs the benefit of receiving the CDS premium against the cost of having to invest less in the risky asset. For simplicity and in order to characterize a premium that is independent of credit risk, we assume that there is no default risk associated with the sovereign. In our model, the only reason for buying the CDS is regulatory requirements. We also provide an extension of the model that incorporates credit risk. In both cases, the CDS premium can be viewed as an addition to the CDS premium of a credit-risky sovereign that can be attributed to dealer bank's demand for capital relief arising from uncollateralized derivatives transaction with dealer banks.

We provide five pieces of evidence in favor of our hypotheses. First, we document that uncollateralized derivatives positions are subject to the Basel III capital charge and that banks use sovereign CDS to avoid this capital charge. Second, we provide sample calculations to show that the orders of magnitude of CDS outstanding are consistent with the derivatives hedging motive for Germany. Third, we find that derivatives dealers are net *buyers* of sovereign CDS, as opposed to being net sellers of CDS, which is common in most other markets. Fourth, we show that there is no link between CDS premium and bond yield for safe sovereigns, but that the link between the two becomes stronger for more risky sovereigns. This finding suggests that safe-haven CDS premiums are not driven by credit risk. Finally, we find that regulatory proxies for the exposure towards sovereigns and for banks' financial constraints are significant drivers of sovereign CDS.

An Explanation of Negative Swap Spreads

While the first essay investigates CDS premiums and the influence of funding frictions that financial regulation imposes on banks, the second essay provides a link between interest rate swap rates and funding frictions faced by pension funds. In the second essay, we examine the persistent negative 30-year swap spread, which is defined as the difference between the swap rate (which is the fixed-rate in the swap) of a 30-year interest rate swap (IRS) and the yield of a Treasury bond with the same maturity. Negative swap spreads are a pricing anomaly and present a challenge to views that have been held prior to the financial crisis

that suggested that swap spreads are indicators of market uncertainty, which *increase* in times of financial distress. This is because the fixed payment in an IRS is exchanged against a credit-risky floating payment, which should cause swap rates to be above the risk-free rate and increase in times of financial distress. Additionally to that, treasuries have a status as “safe haven,” i.e., assets that investors value for their safety and liquidity. In times of financial distress, investors value the convenience of holding safe and liquid assets even more, which decreases the treasury yield below the risk-free rate. In summary, these arguments show that the 30-year swap spread should have increased around the default of Lehman Brothers.

We offer a new perspective on the possible reasons behind this pricing anomaly. Our hypothesis is that demand for duration hedging by underfunded pension plans coupled with balance sheet constraints faced by swap dealers puts pressure on long-term swap fixed rates and ultimately turned the 30-year swap spread negative. We first develop a model where underfunded pension plans’ demand for duration hedging leads them to optimally receive the fixed rate in IRS with long maturities. Pension funds have long-term liabilities in the form of unfunded pension claims and invest in a portfolio of assets, such as stocks, as well as in other long-term assets, like government bonds. They can balance their asset-liability duration by investing in long-term bonds or by receiving fixed in an IRS with long maturity. Our theory predicts that, if pension funds are underfunded, they prefer to hedge their duration risk with IRS rather than buying Treasuries, which may be not feasible given their funding status. The preference for IRS to hedge duration risk arises because the swap requires only modest investment to cover margins, whereas buying a government bond to match duration requires outright investment. Thus, the use of IRS allows the underfunded pension funds to invest their scarce funds in assets (such as stocks) with higher expected return.

Using data from the financial accounts of the United States (former flow of funds table) from the Federal Reserve, we construct a measure of the aggregate under-funded status of DB plans (both private and public) in the United States. We then use this measure to test the relationship between the underfunded ratio (UFR) of DB plans and long-term swap spreads in a regression setting. Even after controlling for other common drivers of swap spreads, recognized in the literature, such as the spread between LIBOR and repo rates, Debt-to-GDP ratio, dealer-broker leverage, market volatility, and level as well as the slope of the yield curve, we show that the UFR is a significant variable in explaining 30-year swap spreads. In line with our narrative, we also show that swap spreads of shorter maturities are *not* affected by changes in UFR. We use stock prices as an instrumental variable in a two-stage least square setting to address possible endogeneity concerns and to further show the robustness of our conclusions.

High Funding Risk, Low Return

While the first two essays illustrate how a deviation from the law of one price can be caused by funding frictions, the third essay uses a deviation from the law of one price to measure market-wide funding conditions and studies the effect of these frictions on hedge fund returns. The risk measure that I examine in this essay is an index of deviations from the covered interest rate parity (CIP) across several different currencies and maturities. Using a large cross-section of hedge fund returns, I then form decile portfolios based on the funds' loading on the CIP measure over the past three years and rebalance the portfolios on a monthly basis. I find that hedge funds with a low loading on the CIP measure, that is, hedge funds with a low funding risk, outperform hedge funds with a high loading on the CIP measure. This result demonstrates that a high loading on funding risk predicts poor fund performance. Instead of being a "priced risk factor," funding risk, as measured by the CIP measure, has the opposite effect: a *higher* loading on the CIP measure predicts *lower* risk-adjusted returns.

To rationalize this finding, I develop a simple model, in which two hedge funds differ with respect to the return that they can generate from investing in an alpha-generating strategy. Funding risk arises because both funds face an exogenous risk of outflows which can force them to unwind their strategies early at a cost. Investors are initially unaware of the difference in the funds' alpha-generating strategies and withdraw from the bad fund, which is the fund with the lower alpha-generating strategy, once they can identify it. The bad fund, therefore, invests more aggressively in its funding-risky strategy to avoid being revealed as bad. Hence, if the funding shock is small, investors are unable to identify the bad fund. It is only if the funding shock is large enough that the bad fund generates losses. These losses due to the funding shock predict lower returns in the next period and enable the investors to identify the fund as the bad fund.

I test the model predictions and obtain three main findings. First, I find that hedge funds with a high loading on the funding risk measure experience more equity withdrawals than funds with a low loading on that measure. Second, the returns of high-risk funds are more sensitive equity withdrawals than the returns of low-risk funds, indicating that high-risk funds hold a lower cash buffer against deteriorating funding conditions. Finally, the link between a high loading on the funding risk measure and low expected returns is less significant for funds that impose stricter redemption terms on their investors and for funds that have multiple prime brokers.

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Essay 1

Safe Haven CDS Premiums¹

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Abstract

Credit Default Swaps can be used to lower capital requirements of derivatives dealing banks who enter into uncollateralized derivatives positions with sovereigns. This makes CDS contracts valuable to dealer banks and contributes to a disconnect between bond yield spreads and CDS premiums, which is particularly pronounced for safe sovereigns. We describe part of the regulation that gives banks the incentive to obtain capital relief using CDS and incorporate the basic features into a simple model. A variety of empirical tests related to volumes of contracts outstanding, yield spreads and CDS premiums, regulatory proxies, and corporate bond markets support our explanation.

1.1 Introduction

We argue in this paper that regulatory incentives to buy Credit Default Swap (CDS) contracts affect CDS premiums and notional amounts outstanding. For safe sovereigns, changes in CDS premiums are virtually unrelated to changes in yield spreads, and we attribute this disconnect to regulatory uses of CDS contracts. In short, derivatives-dealing banks engage in OTC derivatives, such as interest rate swaps, with sovereigns. Most sovereigns do not post collateral in these transactions and this leaves the dealer banks exposed to counterparty-credit risk. This risk adds to the dealer banks' risk-weighted assets (RWAs), and hence to their capital requirements. This is true even when the sovereign is safe, because counterparty risk for regulatory purposes is quantified using CDS premiums. As long as there is some credit risk and hence a non-zero CDS premium, however small, dealer banks have an incentive to buy CDS protection with the purpose of obtaining capital relief. The value of capital relief may dominate the value of the default protection, especially for safe sovereigns. It also requires a higher CDS premium to induce sellers to offer default protection, even on an almost risk-free entity, because the seller of the CDS must provide initial margin, and there is an opportunity cost of providing this margin. The end result is an equilibrium in which the CDS premium is significantly higher than what can be explained by credit risk alone.

We explain in how variation in the so-called Credit Value Adjustments of uncollateralized derivatives positions contributes to the regulatory capital requirements of dealer banks, and we build a one-period model incorporating the essential features. The model has two agents: The first agent is a derivatives-dealing bank who is engaged in a derivatives transaction with a sovereign. Due to regulatory requirements, this derivatives transaction adds to the dealer banks' RWA, thereby lowering its capital available for investment in a risky asset. To free up capital, the bank can buy CDS on the sovereign. The second agent is an end user of derivatives, who has no exposure to the sovereign and acts as seller of credit protection. The end user weighs the benefit of receiving the CDS premium against the cost of having to invest less in the risky asset. In our model, the only reason for buying the CDS is regulatory requirements. Our model offers quantitative guidance as to how CDS premiums depend on margin requirements for the seller and the buyer of CDS protection, capital requirements of the dealer bank and limits on leveraged investment in the risky asset.

We present a variety of empirical tests of our explanation. First, we look at connections between derivatives positions of banks with sovereign counterparties and the net notional amounts of sovereign CDS outstanding. As a first reality check, we confirm that derivatives dealers are net buyers of sovereign CDS, and that the level and volatility of CDS premiums

can justify purchasing protection on safe sovereigns for regulatory purposes. Our ballpark estimates of the amount of CDS notional that can potentially be explained by the outstanding amounts of derivatives with sovereign counterparties show that the CDS demand due to Basel III's CVA capital charges can account for more than 50% of the sovereign CDS volume outstanding, a number that is in line with estimates found in industry research letters. We also look at information on bank derivative exposures toward sovereigns from EBA bank stress tests and find that a significant relationship between these exposures and CDS amounts outstanding.

Second, changes in bond yield spreads and in CDS premiums are almost unrelated for safe sovereigns. A central prediction of our model is that the regulatory component of CDS premiums is larger for safe sovereigns than for less safe sovereigns. Figure 1.1 shows that regressing bond yields on a proxy for the riskless rate and CDS premiums reveals a clear pattern in which the CDS premium explains a larger part of bond yields the riskier the sovereign becomes. For Germany, Japan, and the United States CDS premiums are not a significant explanatory variable for bond yields. For Great Britain the CDS premium is significant, but only at a 10% level. For the three risky European sovereigns in our sample (Italy, Portugal, and Spain), the regression coefficient on the CDS premium is close to one. We perform robustness checks to rule out explanations based on convenience benefits of safe assets.

Third, we test whether proxies for the constraints imposed by regulatory capital help explaining CDS premiums. We find that for the risky sovereigns, Italy, Portugal, and Spain, CDS premiums are mainly driven by credit risk. For the low-risk sovereigns Austria, Finland, and France, both credit and our regulatory capital proxies, have strong explanatory power for CDS premiums. Therefore, our theory does not only apply to safe-haven sovereigns but extends to entities with a low credit risk. For the safe havens Germany, UK, Japan, and the US, we find that regulatory proxies are significant and can explain up to 33% of the variation in CDS premiums.

Finally, evidence from corporate bonds suggests that the disconnect also carries over to safe corporate issuers. Using data for corporates offer two advantages over sovereigns. First, corporate CDS contracts have been actively traded prior to the financial crisis. Second, we can distinguish between financial firms and non-financial firms. Non-financial counterparties typically do not post collateral in their derivatives transactions and we would therefore expect to see a similar pattern of falling correlation between CDS premiums and bond yield spreads as credit quality increases. Financial firms are more likely to collateralize their derivatives positions and we would therefore expect a stronger relationship between CDS premiums and bond yield spreads for these issuers.

1.2 Related Literature

Figure 1.2 illustrates the disconnect between CDS premiums and bond spreads for Germany and the much closer connection for France and Italy. The observed patterns could not occur in a frictionless market where an increase in the CDS premium would also increase the corresponding bond yield. More precisely, the CDS premium and bond yield spread should be equal due to an arbitrage relationship. Hence, our work is related to the growing literature on the limits of arbitrage, as introduced by Shleifer and Vishny (1997) and studied by Gromb and Vayanos (2002) for the case when arbitrageurs need to collateralize their positions. Gromb and Vayanos (2010) survey the literature on limits of arbitrage and summarize the basic idea in these models. An exogenous demand shock for a certain asset occurs to outside investors and arbitrageurs, who both are utility-maximizing and constrained, and take advantage of the shock by providing the asset. We contribute to this literature by providing a parsimonious model in the spirit of Gârleanu and Pedersen (2011), which incorporates the supply and demand side, as well as the explicit financial frictions that drive the potential mispricing.

Yorulmazer (2013) is an early contribution arguing that capital relief is an important motive for banks to buy CDS protection. His main concern is how this may lead to increased systemic risk in the banking system. We prove solutions for CDS premiums that incorporate the exact institutional features of CDS trading and capital relief, and we provide empirical support in several dimensions.

The difference between the CDS premium and the yield spread is commonly referred to as the CDS-bond basis and there is a large strand of literature aiming to explain this basis. Augustin, Subrahmanyam, Tang, and Wang (2014) provide an extensive survey on CDS premiums. Empirically, the CDS-bond basis has been studied for corporate issuers by, among others, Blanco, Brennan, and Marsh (2005), Longstaff, Mithal, and Neis (2005), and Bai and Collin-Dufresne (2013). O’Kane (2012), Gyntelberg, Hørdahl, Ters, and Urban (2013), and Fontana and Scheicher (2014) analyze the CDS-bond basis for European sovereigns. Our empirical analysis complements this strand of literature by showing that there is not only a CDS-bond basis for safe government bonds, rather CDS premiums and yield spreads are completely unrelated.

The drivers of sovereign CDS premiums have been widely studied. Pan and Singleton (2008) and Longstaff, Pan, Pedersen, and Singleton (2011) explain them by global investors’ risk appetite, Ang and Longstaff (2013) suggest systemic risk as one potential driver, and Antón, Mayordomo, and Rodríguez-Moreno (2015) suggest that buying pressure of CDS dealers plays a role. While investors’ risk appetite explains why risky sovereign CDS premi-

ums increase in times of market distress this explanation fails to explain why safe sovereign CDS increase at the same time. Our model gives an alternative explanation for why safe sovereign CDS premiums increase in times of market distress. In addition, our theory helps explaining changes in the amounts of CDS outstanding, which have been studied by Oehmke and Zawadowski (2016) for corporate CDS and by Augustin, Sokolovski, Subrahmanyam, and Tomio (2016) for sovereigns.

Chernov, Schmid, and Schneider (2015) model default risk premiums of the US government, and CDS premiums on US government debt are also touched upon in Brown and Pennacchi (2015), who argue that there may well be a credit risk element in US Treasuries arising from underfunding of pension plans, and that US CDS premiums reflect default risk. We agree that there may well be default-risk premiums for safe-sovereign CDS contracts, but we argue that the regulatory incentive to hold these contracts dominates in their pricing.

Illiquidity premiums in CDS have been studied in Bongaerts, de Jong, and Driessen (2011) and Junge and Trolle (2014), but these papers do not deal with sovereign CDS which judging from volumes outstanding and trading activity are by far the most liquid contracts.

1.3 Regulation and Sovereign CDS Demand

We first highlight the essential features of regulation of uncollateralized derivatives positions that motivate our model and our empirical findings. A significant part of large dealer banks' exposure to sovereign entities comes from interest rate swaps and other over-the-counter (OTC) derivative positions. Unlike financial entities, most sovereigns do not post collateral in OTC derivatives positions and this leaves dealer banks exposed to counterparty credit risk. The current regulatory regime, referred to in short as Basel III (see Basel Committee on Banking Supervision, 2011), contains a charge related to this counterparty credit risk. While the risk of losses related to outright default of a derivatives counterparty had been dealt with earlier, this new capital charge was motivated the significant losses in values of derivatives positions that arose from deteriorating credit quality (but not outright default) of counterparties during the financial crisis.

A bank will suffer marked-to-market losses if an OTC exposure has positive value to the bank and the credit quality of the counterparty deteriorates. In technical terms, a deteriorating credit quality will lead to an adjustment in the so-called Credit Value Adjustment (CVA) of the bank's position. The CVA measures the difference between the value of the OTC exposure if held against a default-free counterparty versus a risky counterparty. When this difference increases, it implies a loss to the bank. Basel III imposes an addition to the bank's Risk Weighted Assets (RWAs), and therefore ultimately to its capital requirement,

related to the risk of changes in the CVA. Importantly, the default risk of the counterparty that goes into the CVA calculation is measured using CDS premiums. This means that regardless of how safe the counterparty is, there is a capital charge as long as the CDS premium on the counterparty is strictly positive and variable.

Basel III gives banks the option to avoid this addition to RWAs by purchasing CDS on the counterparty. Hence, this regulatory framework gives dealer banks an incentive to buy sovereign CDS instead of merely acting as net sellers of CDS contracts, which is common in most other markets. In line with this argument, Figure 1.3 shows that from 2010 on, after the announcement of Basel III, derivatives dealers are indeed net buyers of sovereign CDS.² The notional amount of CDS that the bank will have to buy to obtain full capital relief is equal to so-called expected exposure (EE) arising from the OTC position. If the position is left unhedged, it will lead to an increase in RWAs of EE and therefore a corresponding increase in the bank's capital requirement equal to a fraction of EE . It is the trade-off between the cost of buying protection and the benefit of obtaining capital relief that is fundamental to our model in the next section. More details on the computation of expected exposures and CVAs can be found in Appendix 1.7.1.

1.4 The Model

We set up a simple one-period model that focuses on determining the CDS premium. In this model, a bank has an incentive to purchase CDS protection on a (riskless) entity to obtain capital relief. An end user can earn the CDS premium by selling CDS to the bank, but needs trading capital to do so.

1.4.1 The Assets

There are three different assets in the economy. First, there is a risky asset with price normalized to one, and normally-distributed time-1 payoff $\tilde{r} \sim \mathcal{N}(1 + \mu, \sigma^2)$. To streamline our expositions, we focus on the CDS premium, taking μ and σ^2 as exogenously given constants. The risky asset has a margin requirement m for both buying and short-selling the asset. Hence, one unit of wealth can at most support a long or short position of $1/m$ in the risky asset. From a regulatory perspective, the risky asset contributes to risk weighted assets of the bank. We choose for simplicity to let m also denote the contribution to the

²Unfortunately, there is no information for the buyers and sellers of individual sovereigns available. Hence, we cannot claim that the variation of the notional amount of sovereign CDS bought by dealers can only be traced to financial regulation. It is also possible that, especially during the European debt crisis, the end users' demand for CDS on risky sovereigns increased.

capital requirement for the bank associated with holding one unit of the risky asset. Second, a risk-free asset which pays off $1 + r$ for each unit invested in it at time 0. We assume that the risk-free asset is in perfectly elastic supply and that r is an exogenously given constant. Third, a CDS contract on an entity which is not part of the model and can be thought of as a safe sovereign. The CDS premium s is the main focus of our model and will be determined in equilibrium. We denote by $\tilde{s}$ the random payoff on the CDS as seen from the protection buyer:

$$\tilde{s} := \begin{cases} -s, & \text{with probability } 1 - p \\ LGD, & \text{with probability } p \end{cases}$$

and hence the expected pay-off as seen from the protection buyer is

$$\bar{s} := pLGD - (1 - p)s.$$

The initial margin for buying and selling the CDS is n^+ and n^- respectively. The notional amount of CDS outstanding is determined in equilibrium. s, n^+ , and n^- are all per unit of insured notional, so the relevant dollar amounts are obtained by multiplying the numbers with the notional amount on the CDS contract. We refer to a long position in the CDS as representing a purchase of insurance. If, for example, $s = 45$ bps, a purchase of insurance of 1 dollars of notional, requires a payment of 0.0045 dollars at the end of the period if there is no default, and leads to a positive cash flow equal to $LGD = 0.6$ if there is a default.

1.4.2 The Agents and Their Constraints

There are two different agents, a derivatives-dealing bank B and an end user of derivatives E . Agent i 's wealth at time 1 then given as:

$$W_1^i = W_0^i(1 + r) + g(\tilde{r} - r) + \bar{g}\tilde{s},$$

where $g \in \{b, e\}$ denotes the dollar amount of wealth invested in the risky asset for each agent type, and $\bar{g} \in \{\bar{b}, \bar{e}\}$ denotes the notional amount insured by the CDS for each agent type. So, for example, $\bar{b}$ refers to the dollar amount on which the bank has bought protection (if $\bar{b}$ is positive) or sold protection (if $\bar{b}$ is negative). We assume that agents solve a mean-variance problem in which the optimization objective takes the form

$$\max_{g, \bar{g}} \left[g(\mu - r) + \bar{g}\bar{s} - \frac{1}{2}(\sigma g)^2 \right].$$

We have chosen a risk aversion parameter for both agents to be the same and set it to $\gamma = 1$. There will only be a supply of CDS from the end user when expected return on buying CDS protection is negative, i.e. $\bar{s} < 0$, so that there is a compensation for risk for selling protection, and this will be the case in equilibrium. We disregard, however, the variance contribution for this contract to simplify the analysis and focus on the trade-off between the risky asset and the CDS, and not the size of the risk premium in the CDS contract.

The agents' constraints involve capital requirements of the bank and funding requirements of the end user. Recall that the amount of wealth required to establish a position g in the risky asset is the same for long and short positions and given by $m|g|$. We refer to $m|g|$ as the margin requirement and to the wealth constraint due to margin requirements as the margin constraint. The margin requirement for establishing a long position $\bar{g} > 0$ in the CDS (buying protection) is given by $n^+\bar{g}$ and $n^-|\bar{g}|$ for establishing a short position $\bar{g} < 0$ (selling protection). We think of the agent as having to deposit the amount of cash in a margin account where it earns the risk-free rate r .

The bank and the end user differ in their margin constraints. The end user's constraint is given as:

$$me + n^-|\bar{e}| \leq W_0^E. \quad (1.1)$$

Equation (1.1) can be interpreted as follows. The end user can invest a maximum amount of $\frac{W_0^E}{m}$ in the risky asset. This would rule out taking a position in the CDS contract because any non-zero position in the CDS contract reduces the degree to which the agent can make a levered investment in the risky asset. In equilibrium, the end user will take only long positions in the risky asset. Further, the end user will only consider selling the CDS in order to earn the CDS premium if it offers a positive expected return to do so.

The bank faces a different constraint arising from regulatory capital requirements. We assume that the bank has an interest rate swap with the reference entity of the CDS outstanding. This position adds to the risk-weighted assets of the bank and reduces the bank's ability to lever its risky asset or take positions in the CDS market. As explained in Section 1.3, the contribution to risk-weighted assets is proportional to the expected exposure EE of the interest rate swap. The proportionality factor x depends on the risk that the credit quality of the counterparty deteriorates over the lifetime of the interest rate swap. This risk is measured through the level and the volatility of the CDS premium. The bank can free up capital by purchasing CDS, and a CDS with notional amount equal to EE removes the capital charge entirely. Obtaining capital relief by removing the capital charge from the margin constraint is the reason why the bank considers buying CDS contracts and is willing

to pay a premium which implies a negative excess return on the position. The bank does not gain any capital relief from buying protection on a larger notional than EE . Rather than representing this as a kink in the margin constraint, we add the constraint $\bar{b} \leq EE$ to our optimization problem. Therefore, the bank's margin constraint can be written as:

$$\begin{aligned} mb + n^+ \bar{b} + x(EE - \bar{b}) &\leq W_0^B \\ \bar{b} &\leq EE. \end{aligned} \tag{1.2}$$

In equilibrium, the bank takes a long position in the risky asset and has a non-negative position in the CDS. This is because the only other agent involved in the CDS market is the end user who, in equilibrium, sells CDS.

1.4.3 Equilibrium

Equilibrium in our model is defined as follows:

Definition 1. *In the market described above, equilibrium is defined by a premium s on the CDS contract and positions in the CDS contracts such that*

(i) *Agents maximize the mean-variance utility*

$$\left[g(\mu - r) + \bar{g}\bar{s} - \frac{1}{2}(\sigma g)^2 \right]$$

subject to the constraints (1.1) and (1.2) respectively.

(ii) *The CDS market clears:*

$$\bar{b} + \bar{e} = 0. \tag{1.3}$$

Before stating our main result, we introduce the following three parameter restrictions that we label 'regularity conditions:'

$$\frac{\mu - r}{\sigma^2} > \frac{1}{m} \max(W_0^E, W_0^B - n^+ EE) \tag{1.4}$$

$$W_0^B - xEE > 0 \tag{1.5}$$

$$x > n^+ \tag{1.6}$$

Condition (1.4) ensures that the agents are margin-constrained and conditions (1.5) and (1.6) ensure that the bank has capital for investing in the risky asset and can potentially benefit from purchasing the CDS. Under these regularity conditions we can now state our main result.

Proposition 1. *Assume that the regularity conditions are satisfied. Let*

$$s^b = \frac{1}{1-p} \left(\frac{x-n^+}{m} \left((\mu-r) - \frac{\sigma^2}{m} (W_0^E - n^- EE) + pLGD \right) \right) \quad (1.7)$$

$$s^B := \frac{1}{1-p} \left(\frac{x-n^+}{m} \left((\mu-r) - \frac{\sigma^2}{m} (W_0^B - xEE) \right) + pLGD \right) \quad (1.8)$$

(i) *Define*

$$s_f^e = \frac{1}{1-p} \left(\frac{n^-}{m} \left((\mu-r) - \frac{\sigma^2}{m} (W_0^E - n^- EE) + pLGD \right) \right). \quad (1.9)$$

If $s_f^e \leq s^b$, then s_f^e is the unique equilibrium CDS premium and in this equilibrium, the bank buys full protection on its entire expected exposure $\bar{b} = EE$ from the end user.

(ii) *Define*

$$s_p^e := \frac{1}{1-p} \left[\frac{1}{m \left(\frac{1}{(x-n^+)^2} + \frac{1}{(n^-)^2} \right)} \times \left((\mu-r) \left(\frac{1}{x-n^+} + \frac{1}{n^-} \right) - \frac{\sigma^2}{m} \left(\frac{W_0^B - xEE}{x-n^+} + \frac{W_0^E}{n^-} \right) \right) - pLGD \right]. \quad (1.10)$$

If $s^b < s_p^e \leq s^B$, then s_p^e is the unique equilibrium CDS premium and in this equilibrium, the bank buys partial CDS protection equal to the amount:

$$\bar{b} = \frac{1}{x-n^+} \left(\frac{m}{\sigma^2} \left(\mu-r - \frac{m[(1-p)s - pLGD]}{x-n^+} \right) - (W_0^B - xEE) \right).$$

The proof of Proposition 1 can be found in Appendix 1.7.6.

Numerical Example

In Figure 1.5 we illustrate the model by plotting, for a set of parameters, the supply $-\bar{e}$ and demand $\bar{b}$ for CDS as a function of the CDS premium. With our choice of parameters, described below, the end user starts selling CDS for $s > 84$ basis points and would in fact be buying CDS for $s < 9$ basis points. The bank is willing to buy CDS up to a value of the premium equal to 192 bps. The CDS market clears for a CDS premium of $s = 93$ basis points.

Our motivation for the choice of parameters is as follows: We set the expected excess return to $\mu - r = 0.055$. The standard deviation of the risky asset is given as $\sigma = 0.2$, which is approximately the long-term mean of the S&P 500 implied volatility index VIX. The

initial wealth of bank and end user are set to $W_0^B = W_0^E = 0.2$ to obtain binding margin constraints for both agents. Trading the risky asset requires an initial margin of $m = 0.2$ and this is also the addition to the capital requirement of the bank per unit of additional risky asset. We follow Gârleanu and Pedersen (2011) and assume a margin requirement of 5% for low risk CDS entities. Fourth, the default probability of the sovereign is $p = 0.75\%$ with $LGD = 0.6$, which in a risk neutral world would correspond to a CDS premium of 45 basis points. The bank either faces an addition to its risk-weighted assets of $xEE = 0.06$ with $x = 0.15$ and $EE = 0.4$ or buys CDS to free regulatory capital. Our choice of x is justified in Section 1.5.1, where we perform sample CVA VaR calculations for different sovereigns. EE is chosen as a large number relative to the bank's and end user's wealth for illustrative purposes.

Model Implications

Focusing first on the case where the bank buys full protection, the solution for the CDS premium given in Equation (1.9) has the following implications. First, an increasing expected exposure (EE) on the bank's swap position, which, in equilibrium, increases the demand for CDS protection, increases the premium. Second, a higher margin requirement for selling the CDS (i.e. a higher n^-), increases the CDS premium. However, it is important to keep in mind that the expression for the equilibrium CDS premium only holds if $s^e < s^b$. Therefore, if margin requirements become too high, this may cause a decreasing demand for CDS protection by the bank and therefore a lower CDS premium. Third, a capital-constrained bank is willing to pay an additional premium for CDS protection. Fourth, a higher excess return implies a higher CDS premium. Therefore, our theory provides an alternative explanation for why stock returns are important in explaining changes in CDS premiums, even if the stock returns do not affect credit risk. Finally, assuming that the expected excess return is fixed, Equation (1.9) implies that a higher volatility of the risky asset decreases the CDS premium. This is because investments in the risky asset become less attractive as the volatility increases when expected excess return is fixed.

1.5 Empirical Evidence

We now turn to data and divide our empirical analysis into four broad categories: First, we investigate whether the regulatory relief per unit of CDS protection bought gives institutions an incentive to buy protection, and we investigate the volumes of CDS outstanding compared to the aggregate derivatives exposures of banks to sovereigns. Next, we investigate the covariation between CDS premiums and sovereign bond spreads. The regulatory incentive to

buy CDS protection should lead to smaller correlation between CDS premiums and corporate bond yields for safe sovereigns, where the regulatory component can be large compared to the credit risk component. Third, we test whether different proxies for bank's incentives to hedge (capital constraints, increases in the size and risk of expected exposures) have an effect on CDS premiums. And finally, we see if the pattern of smaller correlation between CDS premiums and yield spreads for safe entities can also be found in US corporate bond markets, and whether the pattern is different for financial firms and non-financial firms. Our data are described in each part separately.

1.5.1 Linking CDS Volume to CVA Hedging

According to several industry research notes, a large fraction of the outstanding sovereign CDS volume can be a consequence of financial regulation. For example, the fraction is estimated to be 25% in Carver (2011) and up to 50% in ICMA (2011). In Appendix 1.7.3, we provide more anecdotal evidence to support our claim that derivatives dealers use sovereign CDS to hedge CVA risk as well as more detailed sample calculations. In this section, we focus on sample calculations and statistical tests.

To justify the use of sovereign CDS for CVA hedging, we need to make sure that the amount of capital relief per unit of CDS notional bought, $x(s_t)$ as defined in (1.19) in Appendix 1.7.1, is large enough to outweigh the margin costs associated with buying CDS contracts. Note that $x(s_t)$ can be computed from historical CDS data. We use CDS premiums for 10 different sovereigns, and our calculations of $x(s_t)$ show that it is typically optimal for banks to hedge their entire CVA VaR using CDS contracts. This then suggests, that there should be a connection between the volume of bank derivatives positions with sovereign counterparties and the amount of CDS contracts outstanding.

Data

We collect data on OTC derivatives outstanding for 28 different sovereigns from the 2013 EBA stress tests and 28 countries from the 2015 stress tests. The data refer to all OTC derivatives that a sovereign, or a government-sponsored entity, has with derivatives dealing banks which were part of the EBA stress test.³

³Stress tests were conducted on banks in all European countries, including Great Britain. However, volumes for derivatives-dealing banks in Switzerland and the United States are not included in the notional amounts. Hence, the amounts from the EBA stress tests underestimate the exposure of all derivatives dealers toward these sovereigns.

CVA and Risk Charges Associated with Derivatives

We initially focus on the 10 sovereigns for which we later do regressions of bond yields on credit spreads. In column 1 and 2 of Table 1.1, we report the notional value and the fair value of all derivatives for these 10 sovereigns that have positive fair value for banks. DTCC provides data on the aggregate dealer holdings of sovereign CDS. The fair value of all derivatives with positive value gives an indication of how deep the derivatives are in-the-money, without accounting for the option-like feature of Expected Exposure discussed in the appendix and without taking netting possibilities into account. While netting of a bank's exposure with a sovereign might imply a smaller expected exposure than the amount indicated by the fair value, there are other reasons why the expected exposure may be larger. First, the current fair value of a derivative nets out positive and negative values that the derivative may have in the future, whereas expected exposure takes into account only values in future states in which the derivative has positive value. Second, the EBA data do not account for OTC exposures that non-European banks have with these sovereigns.

Because banks would need to buy CDS protection on a notional amount equal to the expected exposure to hedge their OTC derivatives exposure towards sovereigns, the fair value of derivatives give an indication of whether the order of magnitude of such positions is comparable to the amounts of CDS outstanding. Column 4 of Table 1.1 reports the amount of sovereign CDS outstanding for the respective countries, and we note that in all cases except for the US, the notional amounts of CDS outstanding are of the same order of magnitude as the fair value of derivatives positions with positive value. We will test the relationship between CDS volume and derivatives positions on a larger sample below.

Table 1.1 also shows in column 9 (furthest to the right) the amount of capital relief $x(s_t)$ that 1 unit of sovereign CDS purchase will provide. Columns 6-9 provide the necessary input to make this calculation, as explained in Appendix 1.7.1. As we can see, the value ranges from lowest value of $x(s_t) = 0.052$ for the US to the highest value of $x(s_t) = 0.821$ for Portugal. In Proposition 1 $x(s_t)$ is written as x , and we note that the regularity condition $x > n^+$ is satisfied for all countries since we assume $n^+ = 0.05$. It is likely that the margin requirement for buying CDS - especially on safe sovereigns, is in fact smaller than 0.05 because the margin would easily exceed the present value of the CDS contract even if the premium dropped to zero, and that therefore we in all cases can justify the purchase of a CDS as providing capital relief.

Testing the Link Between CDS Volumes and CVA Risk

After having established that the orders of magnitude for CVA hedging are large compared to the size of the sovereign CDS market for our sample of 10 sovereigns, we next test whether there is a link between CDS volumes outstanding and sovereign's derivatives exposures to banks on a larger sample. To that end, we expand the sample to include all sovereigns that have derivatives positions with a positive fair value for European and UK banks. We also add the results from the December 2015 stress tests. Panel (a) of Figure 1.6 shows a scatter plot of CDS volumes outstanding (measured as the net notional outstanding) against the fair value of all derivatives with positive value for reporting banks. As we can see from the figure, there is a strong relationship between the two numbers. One exception is China, where the CDS net notional outstanding is significantly larger than the fair value of banks' derivatives positions.

To test the significance of the relationship between sovereign CDS outstanding and banks' derivatives exposures, we next run cross-sectional regressions of the the following form:

$$\log(CDS_{i,t}) = \alpha + \beta \log(Derivatives_{i,t}) + Controls_{i,t} + \varepsilon_{i,t}, \quad (1.11)$$

where $Derivatives_{i,t}$ is the fair value of all derivatives with positive fair value to banks. Table 1.2 shows the results of this test. In Panel (1), we run regression (1.11) without additional controls. We add a dummy variable for the level and the slope coefficient in Panel (2). The dummy variable is equal to one if the data is from the 2015 stress test and zero otherwise, As we can see from the table, the fair value of all derivatives outstanding is a significant explanatory variable for the total amount of CDS outstanding. Overall 45% of the cross-sectional variation in CDS net notional outstanding can be explained by derivatives. Moreover, neither the level or the slope of the main regression change significantly from 2013 to 2015.

To rule out that the link between sovereign CDS outstanding and dealer banks' sovereign derivatives positions is purely driven by the amount of sovereign debt outstanding, we add the total debt outstanding for each of the sovereigns as a control variable to our regression in Panel (3) of Table 1.2. As we can see from the table, controlling for sovereign debt outstanding lowers the statistical and economic significance of our variable. However, even after controlling for the sovereigns' debt outstanding, cf. Augustin et al. (2016), the fair value of banks' derivatives positions with sovereigns is still statistically significant at a 1% level. Moreover, adding a dummy variable for the level and the two slope coefficients shows that the effect of debt outstanding does not change significantly from 2013 to 2015.

1.5.2 Sovereign CDS Premiums and Bond Yield Spreads

We now explore the relationship between CDS premiums and bond yields. The time-series and scatter plots in Figure 1.2 indicate that there is a larger disconnect between bond yield spreads and CDS premiums for safer countries and we now do a regression analysis to investigate whether this pattern is borne out in the data. The disconnect would be consistent with the model's prediction that the regulatory contribution to the CDS premiums is of fixed size and therefore likely to play a significant role for safer sovereigns.

The Data

We study the relationship between CDS premiums and bond yield spreads for 10 different sovereigns, using 5-year data based on weekly observations sampled every Wednesday. We focus our analysis on the period from January 2010 to December 2014 and restrict our considerations to sovereigns that have one of the four major currencies, US Dollar, Euro, Japanese Yen, and British Pound.⁴ We further restrict our considerations to the 7 Eurozone countries with the most frequent quotes for both CDS premium and yield spread. The reason for starting our analysis in 2010 is that the new regulatory requirements were first announced in 2010, and CDS data on safe sovereigns (as opposed to corporates) are not sufficiently rich before then to study an effect of the regulatory change. The 5-year sovereign CDS data are obtained from Markit. The CDS premium for the United States is denominated in Euro, all other CDS premiums are denominated in US Dollar. We use the Bloomberg system to obtain 5-year bond yields and corresponding risk-free rate proxies. Bloomberg uses the most recent issue of the 5-year benchmark bond to compute the yield. If there is no benchmark bond with matching maturity available, no yield is reported. As a proxy for the risk-free rate, we use 5-year swap rates based on overnight lending. In these contracts one party pays a periodic floating rate based on the overnight lending rate and in return receives a fixed rate, denoted the swap rate. We describe these rates (as well as all other data in this article) in more detail in Appendix 1.7.5.

Credit Risk in Bond Yields

To test whether the credit risk in government bonds is reflected by CDS premiums we run a regressions of the following type:

$$\Delta Yield_t^i = \alpha + \beta^{CDS} \Delta CDS_t^i + \beta^{rf} \Delta r_t^i + \varepsilon_t, \quad (1.12)$$

⁴We focus on the four major safe-haven currencies because of data availability. For instance, CDS contracts on Switzerland and Singapore are not among the top 1,000 DTCC most actively traded contracts and quotes exist only infrequently.

where $\Delta Yield_t^i$, ΔCDS_t^i , and $\Delta r f_t^i$ denote changes in the bond yield, CDS premium, and risk-free rate for country i . If CDS premiums were a clean measure of credit risk, we would expect that an increase of one basis point in the CDS premium increases the corresponding bond yield by one basis point. If β^{CDS} is significantly different from 1 and possibly even close to 0 it supports our theory. Using this specification instead of directly comparing yield spreads and CDS premiums has the advantage that we can also check whether our proxy for the risk-free rate is reasonable and reflected in the bond yield.

To get an overview of the results, we first sort the 10 sovereigns by their estimate for β^{CDS} from small to large. We then plot the parameter estimates and the 95% confidence intervals for the estimates (corresponding to two standard deviations) in Figure 1.1. Panel A shows the estimates for β^{CDS} for the 10 sovereigns. As we can see from the figure, the sorting according to β^{CDS} also corresponds to our intuitive sorting. The relationship between bond yields and CDS premiums for the safe-haven sovereigns Japan, US, Germany, and UK is lowest. In particular, none of the parameter estimates is significantly different from zero at a 5% confidence level. Then, β^{CDS} for Finland, France, and Austria, which we refer to as 'low-risk' sovereigns, is significantly different from zero but still well below one and below the estimate for the risky sovereigns, Italy, Spain, and Portugal. On the other hand, the estimates for β^{rf} , reported in panel (b), are all significantly different from zero (at a 5% confidence level) and are close to one. Notably, with the exception of Japan, Germany, and Finland, none of the estimates is significantly different from one at the 95% confidence level. Overall, Figure 1.1 illustrates that there is a large disconnect between CDS premiums and bond yield spreads for safe sovereigns.

Controlling for Convenience Yields

An alternative explanation for why β^{CDS} is insignificant for safe sovereigns could be that safe-haven bonds typically carry a "convenience yield" or "liquidity premium," meaning that investors are willing to accept a lower yield on very safe and liquid assets, see for example Krishnamurthy and Vissing-Jorgensen (2012b).

We start by discussing the convenience yield argument for the case of German government bonds. On the one hand, due to implicit and explicit guarantees for German banks during the financial crisis and due to its responsibilities in the Eurozone, it is conceivable that German government bonds are not entirely free of credit risk. On the other hand, German government bonds are arguably the safest and most liquid Euro-denominated assets. Hence, investors might accept a lower bond yield for the convenience of holding such a safe and liquid asset. We use a variety of different proxies for the convenience yield of government bonds. Our main proxy, which is available for all four sovereigns, is the difference between

the 3-month Eonia swap rate and 3-month sovereign bond yield. We use this as a proxy for convenience yield because the credit risk for a bond issuer with high credit quality is smallest for short maturities. Hence, the 3-month German benchmark bond can be viewed as almost free of credit risk and the difference to the 3-month Eonia swap rate can be attributed to the convenience yield.⁵

As indicated by Figure 1.2, the CDS premium started in our sample period at a level close to 30 basis points and reached a peak around 110 basis points while at the same time the yield spread decreases from around 20 basis points to around -40 basis points, i.e., a decrease of 60 basis points. For CDS to be an accurate measure for credit risk, there would have to be a 140 basis points increase of the convenience yield in this period, but this is far from the peak of the estimated convenience yield at around 60 basis points that we report in Figure 1.4.⁶

In addition to this proxy, we add the spread between bonds issued by the Kreditanstalt für Wiederaufbau (KfW) and the German government bond yields as a proxy for convenience yield for Germany. The argument here is that KfW bonds are guaranteed by the German government and, hence, have the same credit risk as German government bonds but a different liquidity. Therefore, the spread between KfW bonds and German government bonds reflects the liquidity premium in German government bonds. For the U.S., we add the spread between on-the-run and off-the-run bonds as an additional proxy for convenience yield. An increase in this spread points to a situation where there is an elevated demand for the more liquid on-the-run treasury bonds which indicates more demand for highly liquid assets. Finally, we add the weekly government bond turnover as another proxy for flight to liquidity. This variable is available on a weekly basis for the UK and the U.S.⁷

The results of this analysis are exhibited in Table 1.4. As we can see from the table, adding the convenience yield proxies to the regression does not change our inference about β^{CDS} . Out of the four sovereigns, β^{CDS} is only significant for the UK and only at a 10% level. Moreover, increases in our convenience yield proxy, measured as the difference between 3-month overnight swap rates and 3-month bond yields, correspond to decreasing bond yields.

⁵We note that this proxy for convenience yield might be problematic for the U.S., where debates about the debt ceiling lead to elevated CDS premiums on the U.S. for short-term contracts (see Brown and Pennacchi (2015)). We therefore add several additional proxies for convenience yield for the U.S.

⁶In a different study, Krishnamurthy and Vissing-Jorgensen (2012b) determine the size of the convenience yield of US treasury bonds as, on average, 72 basis points. The difference between their study and our study is that we compare the bond yield to a proxy for the risk-free rate while they compare it the rate of the safest and most liquid corporate bonds. Since even the safest corporate bonds are not considered as risk-free, their convenience yield can be viewed as an upper bound.

⁷For Japan, turnovers are available on a monthly basis. We do not add turnovers for Japan in Table 1.4 to keep the number of observations comparable across countries. However, adding turnover for Japan leaves our inference about β^{CDS} unchanged. For Germany turnovers are only available on a semi-annual basis.

However, this proxy for convenience yield is only significant for Germany. Additionally to that, the KfW spread is significant at a 1% level for Germany and increases in that spread correspond to decreases in German bond yields. For the U.S., the on-the-run off-the-run spread is significant at a 10% and increases in that spread correspond to lower bond yields. Changes in bond turnover are insignificant for the U.S. and significant at a 10% level for the UK. Interestingly, increases in turnover correspond to increases in government bond yields. Finally, we note that the R^2 values for Germany, the UK, and the U.S. are all above 0.8 which mitigates omitted variable concerns because we are capable of explaining most of the variation in bond yields with our explanatory variables.

1.5.3 Regulatory Frictions as Drivers of CDS Premiums

In our model, dealer banks have an incentive to use CDS for hedging when their capital constraints are binding, and the demand for CDS should increase if the expected exposure of their derivatives positions with sovereigns increase. In this section we test whether proxies for dealer capital constraints and expected exposure are significant in explaining CDS premiums.

The Data

The Expected Default Frequency (EDF) is an estimate of a firm's default risk which is computed by Moody's Analytics. The estimate builds on a two-step procedure. In the first stage, information on a firm's market value of equity and its liability structure is used to infer the firm's asset value and asset volatility, and from this a 'distance-to-default' is computed which measures the distance, scaled by volatility, of a firm's assets to a default boundary. In the second stage, the distance-to default is converted into a default probability, the EDF, using the result of a non-parametric regression which links distance-to-default to default probabilities using a large historical sample. We denote by EDF_t is the average of the Moody's Expected Default Frequency (EDF) for the 16 largest derivatives dealing banks (G16 banks).⁸ Since there is a strong connection between sovereign credit risk and bank credit risk (see, for instance, Kallestrup, Lando, and Murgoci (2016)), we first regress the average EDF on the yield spread of the respective sovereign and use the residual of this regression as EDF_t .⁹ Table 1.5 reports the results of the regression specified in Equation

⁸These 16 banks are: Morgan Stanley, JP Morgan, Bank of America, Wells Fargo, Citigroup, Goldman Sachs, Deutsche Bank, Nomura, Societe Generale, Barclays, HSBC, Credit Agricole, BNP Paribas, Credit Suisse, Royal Bank of Scotland, and UBS.

⁹Our results are robust to several modifications of this specification. First, directly using the average EDF instead of the residual gives similar results regarding the statistical and economical significance of the regulatory proxies. Second, we modify the average EDF by dropping the EDFs of banks which are located in the respective country from the average EDF measure. For instance, if we ran a regression for Germany we computed the the average EDF without using Deutsche Bank. Again, we obtain similar results.

(1.13), where we group the sovereigns according to their β^{CDS} from Section 1.5.2.¹⁰

$Swptn_t$ is the (basis point) premium on an option to enter a 5-year swap position, as fixed payer or fixed receiver, in the respective currency, over the next 5 years. As discussed in Section 1.5.1, this variable captures the option-like feature of banks' expected exposure towards sovereigns and we therefore use it as a proxy for EE .

Regression Analysis

We now run the following regression:

$$\Delta CDS_t = \alpha + \beta^{YS} \Delta YS_t + \beta^{Swptn} \Delta Swptn_t + \beta^{EDF} \Delta EDF_t + \varepsilon_t. \quad (1.13)$$

YS_t is the difference between 5-year bond yield and 5-year overnight swap rate in the respective currency. We include this variable as a proxy for credit risk. The remaining two variables are independent of the sovereign's credit risk and we refer to them as regulatory proxies in the following.

Examining the results for the four safe-haven sovereigns in our sample, we find that the regulatory proxies are both statistically and economically significant. The R^2 of the regression ranges from 1% for the US to 33% for Germany. To confirm that the explanatory power comes from the regulatory proxies we run a separate regression of the CDS premium on the bond yield spread and report the ratio of the adjusted R^2 from this regression over the adjusted R^2 of the entire regression under 'Credit Ratio'. The credit ratio is zero for Japan, US, and Germany, indicating that the entire explanatory power comes from the two regulatory variables. Turning to the statistical significance, we can see that for Germany and Japan both regulatory proxies are statistically significant. For the UK and the US, ΔEDF_t is the only significant regulatory proxy. For the UK, the yield spread is statistically significant at a 1% level and the credit ratio is 0.19. As mentioned before, the UK started posting collateral in their OTC derivatives transactions in late 2012. The posting of collateral mitigates counterparty-credit risk and, therefore, lowers the CVA capital charge and the dealer banks' incentive to buy CDS protection. Therefore, it is in line with our theory that regulatory proxies are less significant for the UK. It is, however, unlikely that the effect is dramatic due to legacy positions that remain uncollateralized.

Turning to the results for the three low-risk sovereigns in our sample we find that our regulatory proxies have strong economical and statistical significance. With the exception of $\Delta Swptn_t$ for Austria, all regulatory proxies are statistically significant. The main difference

¹⁰Note that we do not include a proxy for the funding conditions of the protection seller. This is because, as explained in Section 1.7.3, there are no natural sellers of sovereign CDS.

between this group and the group of safe-haven sovereigns is that bond yield spreads are statistically significant at a 1% level and contribute to the explanatory power of our regression with a Credit Ratio ranging from 0.12 for Finland to 0.59 for Austria. Overall, the results for low-risk sovereigns confirm our model implications for credit-risky sovereigns that both credit risk and regulatory proxies help explaining the variation in CDS premiums. The finding is also in line with the anecdotal evidence provided in Section 1.7.3. An increased demand for sovereign CDS due to regulatory frictions, combined with a lack of natural sellers for these contracts can cause the CDS premium to increase, even if the fundamental credit risk remains constant.

Finally, turning to the three risky sovereigns in our sample, Italy, Portugal, and Spain, we first observe that yield spreads on bonds are clearly the major driver for CDS premiums. The parameter estimate for the yield spread is statistically significant at a 1% level and the credit ratio ranges from 0.76 for Italy to 0.94 for Spain. Interestingly, both regulatory proxies are statistically significant for Italy. This observation as well as the relatively low credit ratio for Italy can be explained by the fact that Italy is arguably the least risky of the three risky sovereigns and has a large notional amount of interest rate swaps outstanding.¹¹ Therefore, it supports our theory that regulatory proxies help explaining the variation in Italian CDS premiums.

1.5.4 Evidence from Corporate Bond Markets

Figure 1.1 illustrates the breakdown between CDS premium and bond yield for safe sovereigns. We argue that this breakdown is likely caused by regulatory incentives to buy CDS protection on sovereigns. Additionally to collateralized derivatives positions with sovereigns, banks also engage in uncollateralized derivatives positions with corporates, where they are also required to compute and report CVA for these positions. To the extent that banks hedge this CVA risk either for regulatory reasons or for accounting reasons (seeking to minimize earnings volatility arising from CVA volatility), we would expect to see a similar pattern of smaller correlation between CDS premiums and corporate yield spreads for corporations. Using data for corporates offer two advantages over sovereigns. First, corporate CDS contracts have been actively traded prior to the financial crisis. Second, we can distinguish between financial firms and non-financial firms. Non-financial counterparties typically do not post collateral in their derivatives transactions and we would therefore expect to see a similar pattern of falling correlation between CDS premiums and bond yield spreads as credit quality increases. Financial firms are more likely to collateralize their derivatives po-

¹¹See, for instance, <http://www.bloomberg.com/news/articles/2015-04-23/italy-is-euro-area-s-biggest-swap-loser-after-deals-backfired>.

sitions and we would therefore expect a stronger relationship between CDS premiums and bond yield spreads for these issuers.

Data

We obtain bond yields for corporate bonds with a credit rating, maturities between 3 years and 10 years, and a matching CDS premium with no restructuring (docclause XR) from TRACE. We use the last traded yield on each trading day and use a maturity-matched CDS premium, interpolated between the two CDS premiums with nearest maturity available. Similarly, we use a maturity-matched proxy for the risk-free rate, which are swap rates based on LIBOR (as in Bai and Collin-Dufresne (2013)). We clean the dataset for obvious outliers, that is, we remove firms where the average CDS-bond basis is above 1.000 basis points and individual observations where the CDS-bond basis is above 1.000 basis points. Table 1.6 provides summary statistics for our sample. We split the sample into five categories: Aaa-Aa-rated corporate bonds, A-rated corporate bonds, Baa-rated corporate bonds, and Ba-C-rated corporate bonds. As a control group, we also include Aaa-Aa-rated financials. The idea behind including financials is that, in contrast to corporates, many financials do post collateral in their derivatives transactions. Hence, we would expect a stronger relationship between CDS premiums and bond yields for this control group. We focus our analysis on individual bonds, that is, one firm could issue multiple bonds and we include all bonds that fulfill our criteria in the analysis.

Regression results

In line with our hypothesis that CDS premium on safe corporates is increased by demand for CVA hedging, Table 1.6 shows that the average and the median CDS-bond basis is only positive for Aaa-Aa corporates. We next investigate the relationship between bond yields and CDS premiums for our sample of corporate bonds. Table 1.7 shows the results of regressing changes in corporate bond yields on changes in CDS premiums, controlling changes in the risk-free rate, utilizing data from the entire sample period. As we can see from the table, β^{CDS} is 0.42 for Aaa-Aa corporates and significantly different from 1. For A and Baa corporates, β^{CDS} is close to one and not significantly different from one. Hence, for corporate bonds with low credit risk, the CDS premium seems to be driven by other factors than credit risk. Table 1.7 also shows that for non-investment grade corporates, β^{CDS} is also significantly different from one. In addition, β^{rf} is insignificant and close to zero for these bonds. One possible explanation for this observation could be a large illiquidity component in these bond yields (see, for instance, Longstaff et al., 2005).

We now investigate the breakdown of the relationship between bond yield and CDS premium for Aaa-Aa-rated corporate bonds further. To that end, we split the overall time series into three sub-periods: (i) July 2002 to June 2007, (ii) July 2007 to December 2009, and (iii) January 2010 to December 2014. The idea behind this split is that, according to our theory, there should be no breakdown between CDS premium and bond yield before the financial crisis because the new regulation was only announced afterward. During the financial crisis, the CDS-bond basis became massive (see, for instance, Duffie, 2010, Gârleanu and Pedersen, 2011, Bai and Collin-Dufresne, 2013, among many others) and therefore a breakdown of the relationship between CDS premium and bond yield is possible for other reasons than CVA hedging. Only in the third sub-period does our argument apply. We also analyze a sample of Aaa- Aa-rated financial bonds, where we expect a stronger link between CDS premiums and bond yields.

Table 1.8 shows the results of regressing changes in bond yields on changes in CDS premiums and risk-free rates, allowing for a different slope coefficient for corporate CDS, using Aaa-Aa-rated bonds from financial and non-financial issuers over the three different time intervals.

As we can see from the table, both non-financials and financials have a β^{CDS} that is not significantly different from one before the financial crisis. Moreover, there is no significant difference between β^{CDS} for financial and non-financial firms. During the financial crisis, β^{CDS} drops sharply and is significantly different from one for both samples. However, β^{CDS} is, again, not significantly different for financials than for non-financials. Only for the January 2010 to December 2014 sub-period do we observe a significant difference between β^{CDS} in the two samples. The β^{CDS} coefficient is only 0.50 for financials and -0.25 lower for corporates, indicating a massive disconnect between CDS premium and bond yield for non-financial firms after the financial crisis. In line with our hypothesis, this disconnect is less pronounced for financial firms.

1.6 Conclusion

Financial regulation requires derivatives-dealing banks to account for counterparty credit risk in their derivatives transactions with sovereigns. This counterparty risk adds to the capital requirements of dealer banks, unless it is hedged using CDS contracts. We provide theoretical and empirical evidence that these regulatory frictions are an important driver of CDS premiums and notional amounts outstanding, and that their impact is particularly pronounced for safe-haven CDS premiums. We describe part of the regulation that gives banks the incentive to obtain capital relief using CDS and incorporate the basic features

into a simple model.

A variety of empirical tests support our explanation. First, derivative dealing banks are long CDS, and notional amounts of CDS are related to the amount of derivatives that banks have entered into with sovereign counterparties. Second, changes in bond yield spreads and in CDS premiums are almost unrelated for safe sovereigns. Third, proxies for incentives to use sovereign CDS for capital relief are significant in explaining CDS premiums for most safe sovereigns. Finally, evidence from corporate bonds suggests that the disconnect also carries over to safe corporate issuers. This market allows us to distinguish between financials and non-financials. In line with our hypothesis, we find a weaker connection between CDS premiums and yield spreads for non-financials, suggesting that the lack of collateralization here gives stronger incentives to use CDS contracts for capital relief purposes.

1.7 Appendix

1.7.1 CVA and capital

We outline in this appendix some background on regulation that is needed for our empirical section which relates volume of sovereign CDS outstanding to volumes of derivatives exposures of banks to sovereigns.

The Credit Value Adjustment (CVA) of a bank's derivatives position with a risky counterparty measures the difference between the value of the position with a risk-free counterparty and the same derivative with the credit-risky counterparty. It is defined by the Basel Committee (see Basel Committee on Banking Supervision, 2011) as

$$\text{CVA} = \text{LGD} \sum_{i=1}^T \mathbb{Q}(\tau \in (t_{i-1}, t_i)) \text{EE}(t_{i-1}, t_i), \quad (1.14)$$

where τ is the default time of the counterparty. LGD is the loss given default, $\mathbb{Q}$ is the risk-neutral default probability of the counterparty in the time interval $[t_{i-1}, t_i]$, and $\text{EE}(t_{i-1}, t_i)$ is the average expected exposure for the same interval. Since default of the counterparty is only costly in states derivative has positive value for the bank, the exposure is calculated as an expectation over values in these states. We give a detailed example of how to compute this exposure in Section 1.5.1.

Importantly, the probability of default is computed using CDS premia. It is defined in Basel Committee on Banking Supervision, 2011 as

$$\mathbb{Q}(\tau \in (t_{i-1}, t_i)) = \max \left[0, \left(\exp \left(-\frac{s_{i-1} t_{i-1}}{\text{LGD}} \right) - \exp \left(-\frac{s_i t_i}{\text{LGD}} \right) \right) \right],$$

where s_i is the CDS premium on the counterparty for a CDS with maturity date i . The maximum operator ensures non-negative default probabilities but it is irrelevant for our computations since we use a constant CDS premium based on the five-year rate.

Capital requirements are computed based on a VaR measure for the CVA, i.e., it depends on potential fluctuations in the CVA due to changes in counterparty credit risk. Since counterparty risk is measured through CDS premiums, CVA VaR is a function of the volatility of CDS premiums and the sensitivity of CVA to changes in the CDS premium. Two CVA VaR measures enter into the computation: One based on CDS volatility over the last year and a stressed VaR based on the largest volatility realized over the past three years. The simple (non-stressed) CVA VaR has the form:¹²

$$\text{CVA_VaR} = 3 \times \text{WorstCase} \times \text{CS01}. \quad (1.15)$$

WorstCase is given as

$$\text{annual CDS volatility} \times \sqrt{\frac{10}{252}} \times \Phi^{-1}(0.99). \quad (1.16)$$

The factor 3 is a supervisory multiplier, see Gregory, 2012. The 'credit delta' CS01 expresses the sensitivity of CVA towards a one-basis-point change in the CDS premium. To simplify calculations we assume throughout the paper that the CDS term structure is flat and that CS01 measures the risk of a parallel shift. With this assumption, and using a constant EE , CS01 is given as on page 33 of Basel Committee on Banking Supervision, 2011:

$$\begin{aligned} \text{CS01} &= EE \times 10^{-4} \\ &\times \sum_{i=1}^T \left(t_i \exp\left(-\frac{st_i}{LGD}\right) - t_{i-1} \exp\left(-\frac{st_{i-1}}{LGD}\right) \right) \frac{D_{i-1} + D_i}{2}. \end{aligned} \quad (1.17)$$

Thus, $\text{WorstCase} \times \text{CS01}$ represents a linear approximation of a move in CVA which is not surpassed with a probability of 99% over a 10-trading day period (assuming normally distributed movements of the CDS premium).

The exact same type of formula is used to compute a so-called *stressed* CVA VaR in which the maximum annual volatility observed over the last three years is plugged into the WorstCase part instead of the annual volatility computed over the last year. Having computed the CVA in both a normal version and a stressed version, the addition to risk-

¹²We follow Gregory, 2012, page 390 with this formula. Different banks might use different approaches to compute VaR. A more common way among banks with more than one counterparty would be to use historical simulation to compute the CVA VaR.

weighted asset, RWA , is conservatively set to be the *sum* of the two VaR measures:

$$RWA = 12.5 \times (CVA VaR + CVA Stressed VaR) \quad (1.18)$$

where 12.5 ensures that the added capital requirement is equal to $CVA VaR + CVA Stressed VaR$ under an 8% capital rule.

We assume in our calculations that the capital requirement is $0.1 * RWA$, but it might arguably be set even higher since the dealer banks that we are looking at have extra capital buffers related to their status as systemically important banks and their desire to stay on the safe side of binding capital requirements.

In our model, the bank has the choice between accepting a capital requirement of $x(s_t)EE$ or buying CDS protection on a notional amount equal to EE . The cost of buying protection only is worth paying if $x(s_t)$ is sufficiently high. From our calculations above, it follows that

$$x(s_t) = 0.1 \cdot 12.5 \cdot c \cdot \frac{CS01}{EE} (\sigma_1(s_t) + \sigma_3(s_t)) \quad (1.19)$$

where $\sigma_1(s_t)$ and $\sigma_3(s_t)$ are, respectively, the CDS volatility over the last year and the maximal level of the annual volatility over the last three years. This expression for $x(s_t)$ depends only on the level and volatility of CDS premiums. We are therefore able to compute values of $x(s_t)$ and see if historical data confirm that there is a potential for capital relief.

1.7.2 Trading Sovereign CDS in Practice

We discuss here some institutional features of CDS trading that motivate our assumptions regarding margin requirements for trading and the asymmetry of costs between buying and selling.

The CDS 'big bang' from April 2009 led to a standardization of the annualized CDS premium to either 100 basis points or 500 basis points, depending on the risk of the reference credit.¹³ If the 'fair' CDS premium is below the 100 basis point standard, which is common for safe-haven sovereigns, the seller of CDS protection makes an upfront payment to the protection buyer in order to compensate him for the higher payment. This upfront payment requires capital on the part of the seller and provides funding to the buyer. This leads to an asymmetry of capital cost between buyer and seller.

Even with no upfront payment, we would expect there to be a smaller margin requirement for CDS contracts on relatively safe reference credits. To see this, assume that the buyer of protection agrees to pay a CDS premium of 45 basis points over the next 5 years,

¹³See Casey (2009) for further details on the CDS big bang.

which corresponds to the average CDS premium on Germany throughout our sample period. The worst possible scenario from perspective of the protection buyer's counterparty is that the CDS premium drops by a significant amount, say, for simplicity, to zero, and the protection buyer defaults at the same time. In the extreme case where this scenario comes true immediately after the CDS contract is sold, the counterparty's foregone profit would be five times 45 basis points, which (ignoring discounting) corresponds to 2.25%. This extreme scenario highlights that the assumption of an initial margin requirement of 5% for the buyer of protection is very conservative and the initial margin for safe-haven CDS contracts is likely much smaller.

In our model, we assume that agents have to post initial margins, even if the reference entity is riskless. This is in line with real-world margin requirements, set by a regulator or Central Clearing Counterparty, which exist even for the least risky sovereigns. Selling CDS requires a margin that depends on the risk of the underlying plus a short-selling margin. The risk of the underlying is computed as a Value-at-Risk number using historical volatility. The short charge is to mitigate the risk of a joint default of the protection seller and the underlying entity.¹⁴ The initial margin to account for such jump-to-default risk can be massive and depends on the seller's CDS portfolio. The most extreme charge is imposed by CME, which requires an 80% (!) initial margin if the counterparty only sells one CDS. The initial margin declines with the number of CDS that the counterparty sells (to 20% if he has 5 transactions, to 10% with 10 transactions and 5% with 25). This massive charge may explain why arbitrageurs are not readily selling safe-haven CDS. Only those already active in the CDS market would do it, because only then would the return-to-margin be attractive. Of course, arbitrageurs can trade through major derivatives dealers like Barclays and JP Morgan, who have access to the two major clearing houses responsible for CDS clearing, but the dealer banks are responsible for the trading of their clients and will require compensation for trading on behalf of clients.

1.7.3 CVA Hedging in Practice

“CVA desks have come to account for a large proportion of trading in the sovereign CDS market and so their hedging activity has reportedly been a factor pushing prices away from levels solely reflecting underlying probability of sovereign default.”

–Bank of England, Quarterly Bulletin 2010

The new CVA capital charge has been subject to an extensive debate with respect to its usage and interpretation. The CVA capital charge was first announced in October 2010

¹⁴See Duffie, Scheicher, and Vuilleme (2015) for further details.

in the first proposal of the new Basel capital requirements (Basel III) and has given rise to many discussions since. For example, among the most frequently asked questions about Basel III is the question: 'can you confirm inclusion of sovereigns in the CVA charge and ability to use sovereign CDS as hedge', which was answered as follows by the committee in November 2011: 'The Committee confirms that sovereigns are included in the CVA charge, and sovereign CDS is recognized as an eligible hedge.'¹⁵ Hence, the new CVA capital charge applies to sovereigns too. This is an important clarification because other regulatory requirements treat sovereign bonds different from corporate bonds. It is worth noting, that while interest-rate swaps are in general moving towards central clearing, sovereigns have been exempt from this requirement. A recent article in the Financial Times¹⁶ explains that, moving forward, there can also be a tendency for central clearing of OTC derivatives with sovereign counterparties.

Another indication that our model captures a feature of the market is the debate as to whether the new CVA capital charge can cause pro cyclical effects. In particular, basing CVA VaR calculations on CDS volatility together with requiring CDS contracts as hedge has caused criticism from the financial industry. For instance, Risk magazine Carver (2011) and FT alphaville Murphy (2012) commented on this issue, arguing that this combination can create a 'doom loop'. The argument is that a higher CDS volatility causes more demand for CDS contracts which, in turn, fuels the volatility of the CDS contract. In the language of our model, a higher CDS volatility increases x , which in turn increases s^b and this can therefore increase the demand for safe-haven CDS. This higher demand further increases the CDS premium. Carver (2011) and Murphy (2012) further explain that the main problem is that there are no natural sellers of sovereign CDS to absorb this demand. Therefore, a small change in the demand for sovereign CDS can have a significant impact on prices. The problem that there are no natural sellers of sovereign CDS has also been discussed by the US treasury borrowing and advisory committee in a report from May 2010.¹⁷ Further, as discussed before, another indicator of the lack of natural sellers of sovereign CDS is the fact that derivatives dealers are in fact net *buyers* of sovereign CDS (Figure 1.3). This lack of supply combined with the demand for sovereign CDS introduced by regulation can cause distortions in the sovereign CDS market. Carver (2011) conjectures that a disconnect between CDS premiums and yield spreads for France in 2011 can be attributed to CVA VaR hedging. As a reason for this she quotes an official of the French debt management

¹⁵See document 'Basel III counterparty credit risk and exposures to central counterparties - Frequently asked questions'.

¹⁶'Germany's debt office set for derivatives clearing' – June 4, 2015. See: <http://www.ft.com/intl/cms/s/0/c84577c0-0acd-11e5-9df4-00144feabdc0.html>.

¹⁷See <http://www.treasury.gov/resource-center/data-chart-center/quarterly-refunding/Documents/dc-2010-q2.pdf>

office: 'On the demand side [for sovereign CDS] we see mostly two types of players: hedge funds and CVA desks, as they move into line with Basel III. It's possible that some of the dislocation with the cash market is due to legitimate CVA hedging'. This conjecture is exactly in line with our theory. We study the disconnect between bond yields and CDS premiums in more detail in Section 1.5.2.

A problem in studying the effect of the new regulatory requirement on CDS premiums is that the new CVA capital charge has not yet been implemented in all regional laws. While Switzerland has implemented it as of 2013, the final rules for the US are still not finished. Further, the European Banking Authority (EBA) decided to grant an exemption from the CVA capital charge for sovereigns. According to Risk magazine ('Europe goes its own way on CVA'), this exemption came as a positive surprise for European banks. For instance, Royal Bank of Scotland stopped reporting the CVA charge for sovereigns which lead to an increase in their equity capital, indicating that they were already incorporating the CVA charge in their capital requirements. However, the exemption is heavily debated (see for instance ft.com 'JP Morgan under pressure in Basel spat', or Risk magazine: 'The CVA helter skelter: European supervisors could quash exemptions') and more recently the EBA has announced to review the exemption (see Risk magazine 'CVA switchback will hit bank capital ratios, EBA says' and EBA document 'Opinion of the European Banking Authority on Credit Valuation Adjustment (CVA)').

Although European banks are exempt from the rule and US banks are not obliged to implement the rules yet, there is strong anecdotal evidence that several major dealers already hedge the new CVA capital charge. Most prominently, Deutsche Bank reported in the first half of 2013 that it 'cut the risk-weighted assets (RWAs) generated by Basel III's capital charge for derivatives counterparty risk – or credit valuation adjustment (CVA) – from 28€ billion to 14€ billion' Carver (2013). Another example is bank of America who states in its 2012 and 2013 annual reports that 'The Corporation often hedges the counterparty spread risk in CVA with CDS.' Further, Credit Suisse reports in its 2013 annual report an 'advanced CVA [that] covers the risk of mark-to-market losses on the expected counterparty risk arising from changes in a counterparty's credit spreads.' Overall, these examples show that major derivatives dealers already use sovereign CDS to obtain capital relief from the new CVA capital charge.

1.7.4 Detailed CVA Calculations for the Case of Germany

Interest-rate swaps are by far the largest market for OTC derivatives, and it is therefore likely that the bulk of banks' derivatives exposures to sovereigns are in this market. For the

case of Germany, we have data on swap-usage of the federal government. This allows us to use swaption prices, in a way we explain below, to compute an estimate of the expected exposure of banks to Germany that is related to the swap positions.

The Bundesrepublik Deutschland Finanzagentur (Bund) is a government agency in charge of organizing the borrowing and management of Germany's debt. We obtain data on the notional amount of interest-rate swaps concluded on the behalf of the German federal government from Bund. Table 1.9 contains the notional amount of the holdings of both payer-and receiver swaps, that are classified as 'capital market swaps' by the Bund.¹⁸ We now use these figures to obtain an estimate of the total expected exposure of the dealer banks due to these swaps. Our estimate is based on a relationship between the expected exposure and the value of a swaption, i.e., the right to enter into a swap at a future date. This connection is used for example in Sorensen and Bollier (1994), but it is useful to explain the basic idea in detail here. We refer to Longstaff, Santa-Clara, and Schwartz (2001) for more details on contract terms in swap and swaption contracts.

Let $S(c, r_t, t, T)$ denote the value at date t of a swap contract for the party receiving the fixed payment c per period until maturity T . r_t is a state variable which determines the term structure of interest rates at date t . In a short-rate model, it would just be the instantaneous short rate, but it could be a multidimensional state-variable as well. Let s_t denote the at-market swap rate at date t , i.e., the rate satisfying $S(s_t, r_t, t, T) = 0$. The value at date t of an at-market swap that was entered into at date 0 is then $S(s_0, r_t, t, T)$ and this value is positive precisely when $s_t < s_0$, and we write the exposure of the fixed receiver at date t as $\max(S(s_0, r_t, t, T), 0)$. This figure corresponds to the value at date t of the option to enter into a swap as a fixed receiver at the rate s_0 . We therefore approximate the expected exposure at t seen from time 0 using the value of a swaption.

We note that this is only a 'back-of-the-envelope' approximation for three reasons. First, the swaption value is a discounted value under a risk-neutral measure, and this may make it smaller than the expected undiscounted exposure under the physical measure. Second, we approximate the value of the receiver (and the payer option) using one half of the value of a swaption straddle, i.e., the combination of an option to enter as a fixed receiver and the option to enter as a fixed payer at date t struck at the forward swap rate at date 0, which is the strike rate at which these two options have the same value. One half of the straddle therefore gives us the value of a receiver swap (or a payer swap) struck at the forward swap rate, but of course the swap entered into at date 0 is struck at the at-market rate which might differ from the forward swap rate. Third, we assume the expected exposure as viewed

¹⁸These are mainly Euribor swaps. The Bund is also engaged in Eonia swaps. The amounts outstanding for these contracts are not as large as the ones for capital market swaps and we do not report them in the Table.

from date t to be constant over (future) revaluation dates and determined by the value at date t , of a 5-into-5 year swaption, i.e., an option which can be exercised in 5 years and which give the right to enter into a 5-year receiver swap.¹⁹

In sum, we approximate the expected exposure viewed from date t as:

$$EE_t = IRS\ Outstanding_t \times SwaptionValue_t. \quad (1.20)$$

The quotes in Table 1.9 refer to at-the-money swaption straddles based on Euribor rates and are obtained from the Bloomberg system. The price of the receiver swaption is half the value of the swaption straddle as explained above. We describe these quotes in more detail in the appendix. The resulting expected exposure is reported in column 6 (under EE) of Table 1.9.²⁰

Next, we use the figures for Germany to compute the amount of equity capital that is required for maintaining the swap positions if no hedging is used. This requires computing the CVA and CVA VaR, and for that we make the following simplifying assumption. We assume a constant LGD of 0.6, a flat CDS term structure based on the premium s of the 5-year contract for Germany, and a constant expected exposure computed using the swaption argument above. We compute CS01 as the first derivative of the Basel III default probability described in Section 1.3, using a flat CDS term structure based on the 5-year lag. Note that CS01 captures the sensitivity of the value of the protection leg of a CDS contract to a parallel shift in the term structure of CDS premiums. The notional amount is EE and the change is measured per basis point. We next compute historical volatilities of German 5-year CDS premiums which allows us to compute both the CVA Var and the stressed CVA Var following Equation (1.15). The results for CVA, CVA VaR, and stressed CVA VaR are reported in Table 1.9.

We first observe that the CVA VaR and stressed CVA VaR are typically more than 3 times higher than the CVA itself. The reason for this higher CVA VaR is that additionally to the CDS premium, the historical volatility is also an input parameter. That explains why, despite a lower CDS premium in 2012 relative to 2010, the CVA VaR in 2012 is higher than in 2010. Also, recall that to compute the stressed CVA VaR, we replace the year-end annualized CDS volatility with the maximum volatility over the last three years in Formula

¹⁹This is arguably an overestimation because the expected exposure on a 10-year swap contract typically peaks at 5 years. An alternative would be to use the average of swaptions with 1-9 years to maturity to enter into an IRS with 9-1 years to maturity. We did that as well and found that using this average would reduce the swaption value by 60-120 basis points.

²⁰We assume no netting between payer and receiver swaps in this calculation which might result in an overestimation of the expected exposure. However, it is likely that sovereigns do not allow for netting of their IRS positions between different banks to avoid additional exposure to the counterparty.

(1.15). As we can see in the column under stressed VaR in Table 1.9, stressed CVA VaR could be as much as three times higher than the actual VaR.

Given CVA VaR and stressed CVA VaR, the contribution to the banks' RWA is computed using Equation (1.18). Banks have to maintain a certain percentage of the RWA as equity capital. The exact percentage depends on several factors. There is a general common equity requirement of 7% of RWA, but for systemically important banks this is increased by between 1 and 2.5%. In addition, a countercyclical buffer between 0 and 2.5% may be imposed. We assume in our calculations a total required buffer of 10%, and with this assumption the banks' required equity capital is reported under Equity Capital in Table 1.9. Putting the required equity capital in relation to the expected exposure, gives us a proxy for x . As Table 1.9 shows, the lowest value for x was 0.093 at the end of 2010. In 2011 it went as high as 0.14 and converged to 0.11 in 2013 and 2014. Hence, if we again assume an initial margin requirement of 5% for both, buying and selling CDS, the equilibrium condition in Proposition 1 is fulfilled for most of the years.

1.7.5 Variable Descriptions

This appendix provides additional details about the data used for our analysis.

1. **Sovereign CDS premiums.** We obtain CDS premiums with 5 year maturity on 10 sovereigns from Markit, who provides daily mid-market quotes. We use weekly mid-market quotes in our analysis sampled every Wednesday. In line with previous research (e.g. Fontana and Scheicher (2014)), we use the CDS premium of contracts with 'CR' as restructuring clause.
2. **Sovereign Bond Yields.** Sovereign bond yields for 5-year bonds are obtained from the Bloomberg system. Bloomberg uses the latest 5-year benchmark bond to compute the yield. Yields are computed for bonds with semi-annual (Italy, Great Britain, Japan, and the United States) and annual (Spain, Austria, Finland, France, and Germany) coupon payments. The day-count convention is Actual/Actual.
3. **Risk-Free Rate Proxy.** For sovereigns, we use swap rates based on overnight lending rates with the same 5-year maturity and the same currency as the bond yield. For European sovereigns, we use Eonia swap rates, for Great Britain we use Sonia swap rates, for Japan we use Tibor swap rates, and for the United States we use OIS rates. For U.S. corporates, we use LIBOR swap rates with the same maturity as the underlying bonds. The day count convention for these swap rates is 360/Actual but

we do not correct for this difference in day-count conventions when computing yield spreads. All rates are obtained from the Bloomberg system.

4. **CDS Amounts Outstanding.** Data on amounts of CDS outstanding are obtained from the Depository Trust Clearing Corporation (DTCC) who collects information on CDS amounts outstanding. We use net notional amounts outstanding.
5. **Sovereign CDS bought by derivatives dealers.** Sovereign CDS bought is computed as the difference between gross notional of all sovereign CDS bought by derivatives dealers and gross notional of all sovereign CDS sold by derivatives dealers. The figures are obtained from DTCC who publishes weekly information on the gross amount of sovereign CDS bought by derivatives dealers and by end-users.
6. **Swaption Data.** The swaption quotes are basis point prices of swaption straddles in the respective currencies. A swaption straddle is a portfolio of a long position in a receiver swaption, which gives its owner the right but not the obligation to enter into a swap contract as fixed receiver, and a long position in a payer swaption, which gives its owner the right but not the obligation to enter a swap contract as fixed payer. Because at-the money swaptions refer to swap contracts with zero value, an application of the put-call parity shows that payer and receiver swaption have the same price. The data are obtained from the Bloomberg system.
7. **CDS Volatility.** We use the same formula as in the new Basel capital requirements to compute this variable. That is, at date t , we compute the standard deviation of the changes in the CDS premium over the past 252 trading days.
8. **G16 EDF.** We obtain 1-year expected default frequencies (EDFs) for the 16 largest derivatives dealing banks, commonly referred to as G16 banks, from Moody's Analytics. We then take the average of the 16 EDFs and orthogonalize the resulting time series on the respective yield spread of the sovereign we analyze.
9. **On-the-run spread.** The spread is computed for bonds with 10 years to maturity because estimates of the 30-year spread are noisy and suffer from the 2002-2005 period where the US Treasury reduced its debt issuance. The 10-year on-the-run yield is obtained from the FED H.15 website and the 10-year off-the-run yield is constructed as explained in Gürkaynak, Sack, and Wright (2007) and data are obtained from <http://www.federalreserve.gov/pubs/feds/2006>.
10. **Corporate bond yields.** We obtain the last traded yield on a trading day for each corporate bond that fulfills our filtering criteria from TRACE. We only use rated bonds

with 3 to 10 years to maturity and a matching CDS with XR restructuring clause.

11. **Corporate CDS.** We obtain CDS premiums with the same maturity on the same day as the corporate bonds from Markit. We only use contracts with “XR” (no restructuring) as restructuring clause.
12. **KfW spread.** We collect mid-market prices of all euro-denominated bullet bonds with an issuance volume above 1 billion issued by the KfW and the German government. We follow Schuster and Uhrig-Homburg (2015) and fit a Nelson and Siegel (1987) model to the KfW bond prices and the German government bond prices by minimizing the sum of squared, duration-weighted differences between observed and model-implied bond prices. We then use these model parameters to extract a 5-year zero-coupon yield for both time series. The KfW spread is then given as the difference between 5-year KfW zero-coupon yield and 5-year German government zero-coupon yield.
13. **Government bond turnover.** We collect data on weekly Treasury and Gilt turnover from the Federal Reserve’s and the Bank of England’s website respectively. For Gilts, due to a lack of finer measure, we use the aggregate turnover of all Gilts. For the U.S., we use the turnover of all bonds with three to six years to maturity.

1.7.6 Proof of Proposition 1

To prove Proposition 1, we proceed in three steps. First, we derive the end user’s optimal asset holdings using the Kuhn-Tucker (KT) theorem. Second, we proceed similarly to obtain the bank’s optimal asset holdings. Finally, we solve for equilibrium, distinguishing the two cases stated in the proposition. The KT theorem can be applied because the objective function is concave and the constraints are linear and therefore concave as well. Hence, a stationary point satisfying the KT conditions is a maximum.

We start by deriving the end user’s optimal asset holdings. To conform with the convention that the variables over which we optimize are non-negative, we let $\bar{e}$ denote the number of CDS contracts *sold* by the end user. The end user’s Lagrangian is then given as:

$$\mathcal{L}(e, \bar{e}, \lambda) = (e(\mu - r) - \bar{s}\bar{e} - 1/2(\sigma e)^2) - \lambda (me + n^-\bar{e} - W_0^E). \quad (1.21)$$

Therefore, the KT conditions for the end user's problem are:

$$\mu - r - \sigma^2 e - \lambda m \leq 0 \quad (= 0 \text{ if } e > 0) \quad (1.22)$$

$$-\bar{s} - \lambda n^- \leq 0 \quad (= 0 \text{ if } \bar{e} > 0) \quad (1.23)$$

$$W_0^E - me - n^- \bar{e} \geq 0 \quad (= 0 \text{ if } \lambda > 0) \quad (1.24)$$

$$e, \bar{e} \geq 0.$$

We first look for solutions for which the $e > 0$ and $\bar{e} > 0$. Equations (1.22) and (1.23) imply:

$$e = \frac{\mu - r + (\bar{s}/n^-)m}{\sigma^2}, \quad (1.25)$$

which is positive if:

$$s < \frac{1}{1-p} \left(\frac{n^-}{m} (\mu - r) + pLGD \right). \quad (1.26)$$

If $\bar{e} > 0$, Equation (1.23) implies $\lambda = -\frac{\bar{s}}{n^-}$ which is strictly positive as long as $\bar{s} < 0$ or, equivalently, $s > \frac{p}{1-p}LGD$, and in this case Equation (1.24) holds with equality. Therefore, $\bar{e}$ is given as:

$$\bar{e} = \frac{W_0^E - me}{n^-} = \frac{1}{n^-} \left(W_0^E - \frac{m}{\sigma^2} \left(\mu - r + \frac{\bar{s}m}{n^-} \right) \right). \quad (1.27)$$

and this is positive if

$$W_0^E - \frac{m}{\sigma^2} \left(\mu - r + \frac{\bar{s}m}{n^-} \right) > 0 \quad (1.28)$$

which corresponds to the requirement

$$s > s_0 := \frac{1}{1-p} \left(\frac{n^-}{m} \left(\mu - r - \frac{\sigma^2}{m} W_0^E \right) + pLGD \right). \quad (1.29)$$

Equations (1.25) and (1.27) characterize a stationary point where the end user supplies CDS contracts and is long the risky asset.

We now look for stationary points for which $e > 0$ and $\bar{e} = 0$. In this case e is given as:

$$e = \frac{W_0^E}{m}, \quad (1.30)$$

which leads to the following expression for λ :

$$\lambda = \frac{1}{m} \left(\mu - r - \frac{\sigma^2}{m} W_0^E \right). \quad (1.31)$$

Hence, λ is non-negative under regularity condition 1. To satisfy inequality (1.23), the CDS

premium needs to satisfy: $s \leq s_0$, where s_0 is defined in (1.29). Hence, if s is below s_0 , the end user does not supply CDS contracts.

Our second step is to derive the bank's optimal asset holdings. We follow the same procedure as for the end user, writing up the Lagrangian and the KT conditions for the bank's optimization problem:

$$\begin{aligned} \mathcal{L}(b, \bar{b}, \lambda_1, \lambda_2) = & (b(\mu - r) + \bar{s}\bar{b} - 1/2(\sigma b)^2) - \\ & - \lambda_1 (mb + n^+\bar{b} + x(EE - \bar{b}) - W_0^B) - \lambda_2 (\bar{b} - EE) \end{aligned}$$

From this we get the KT conditions:

$$\mu - r - \sigma^2 b - \lambda_1 m \leq 0 \quad (= 0 \text{ if } b > 0) \quad (1.32)$$

$$\bar{s} - \lambda_1(n^+ - x) - \lambda_2 \leq 0 \quad (= 0 \text{ if } \bar{b} > 0) \quad (1.33)$$

$$W_0^B - xEE - mb - \bar{b}(n^+ - x) \geq 0 \quad (= 0 \text{ if } \lambda_1 > 0) \quad (1.34)$$

$$EE - \bar{b} \geq 0 \quad (= 0 \text{ if } \lambda_2 > 0) \quad (1.35)$$

$$b, \bar{b} \geq 0.$$

We start by looking for a stationary point such that all conditions are satisfied with equality. This corresponds to a situation where the bank buys full protection ($\bar{b} = EE$) and invests $b > 0$ in the risky asset. We find

$$b = \frac{1}{m} (W_0^B - n^+ EE) \quad (1.36)$$

$$\lambda_1 = \frac{1}{m} \left(\mu - r - \frac{\sigma^2}{m} (W_0^B - n^+ EE) \right) \quad (1.37)$$

$$\lambda_2 = \lambda_1(x - n^+) + \bar{s} \quad (1.38)$$

and need to ensure that all quantities are strictly positive. By construction $\bar{b} = EE > 0$. Furthermore, under regularity condition 1, the bank's margin constraint binds and $b > 0$ as well as $\lambda_1 > 0$ are fulfilled. For $\lambda_2 > 0$ to hold, the CDS premium must satisfy the following inequality:

$$s < s_b := \frac{1}{1-p} \left(\frac{x - n^+}{m} \left(\mu - r - \frac{\sigma^2}{m} (W_0^B - n^+ EE) \right) + pLGD \right). \quad (1.39)$$

Hence, the bank demands full protection as long as the CDS premium satisfies Inequality (1.39) and the regularity condition 1 is satisfied.

Next, we consider the case where the bank is not buying full protection, so that $\lambda_2 = 0$.

Then, regularity conditions 1 and 3 imply that $\lambda_1 > 0$ and our solutions for b and $\bar{b}$ become:

$$b = \frac{1}{\sigma^2} \left(\mu - r - \frac{m\bar{s}}{x - n^+} \right) \quad (1.40)$$

$$\bar{b} = \frac{1}{x - n^+} \left(\frac{m}{\sigma^2} \left(\mu - r - \frac{m\bar{s}}{x - n^+} \right) - (W_0^B - xEE) \right). \quad (1.41)$$

b is positive for $s < \frac{1}{1-p} \left(\frac{x-n^+}{m} (\mu - r) + pLGD \right)$ and $\bar{b} > 0$ is satisfied if the following inequality holds:

$$s < s^B := \frac{1}{1-p} \left(\frac{x - n^+}{m} \left(\mu - r - \frac{\sigma^2}{m} (W_0^B - xEE) \right) + pLGD \right). \quad (1.42)$$

Hence, for $s \in (s^b, s^B)$ the bank buys CDS contracts with notional $\bar{b} \in (0, EE)$.

The final step of our proof is to compute the equilibrium CDS premium. The expression depends on whether the supply curve rises quickly enough to meet demand in the range of CDS premiums where demand is flat (i.e., the full protection case) or the supply curve crosses in the range where the demand curve has begun its descent against 0. We first find the rate at which the end user is willing to supply EE contracts. If the rate at which this occurs is below the rate at which the bank starts decreasing its demand away from full protection, the equilibrium is characterized by part (i) of Proposition 1.

The CDS premium that solves

$$EE = \frac{1}{n^-} \left(W_0^E - \frac{m}{\sigma^2} \left(\mu - r - \frac{\bar{s}m}{n^-} \right) \right)$$

is given as

$$s = s_f^e := \frac{1}{1-p} \left(\frac{n^-}{m} \left(\mu - r - \frac{\sigma^2}{m} (W_0^E - n^- EE) \right) + pLGD \right).$$

The bank demands EE CDS contracts as long as $s < s_b$ which proves part (i).

To prove part (ii), we equate the demand and supply expressions for the bank and the end user and make sure that the equilibrium point is indeed within the range in which the demand and supply functions take the assumed form. Equating supply and demand leads to the equation

$$\frac{1}{x - n^+} \left(\frac{m}{\sigma^2} \left(\mu - r - \frac{m\bar{s}}{x - n^+} \right) - W_0^B - x^+ EE \right) = \frac{1}{n^-} \left(W_0^E - \frac{m}{\sigma^2} \left(\mu - r - \frac{\bar{s}m}{n^-} \right) \right)$$

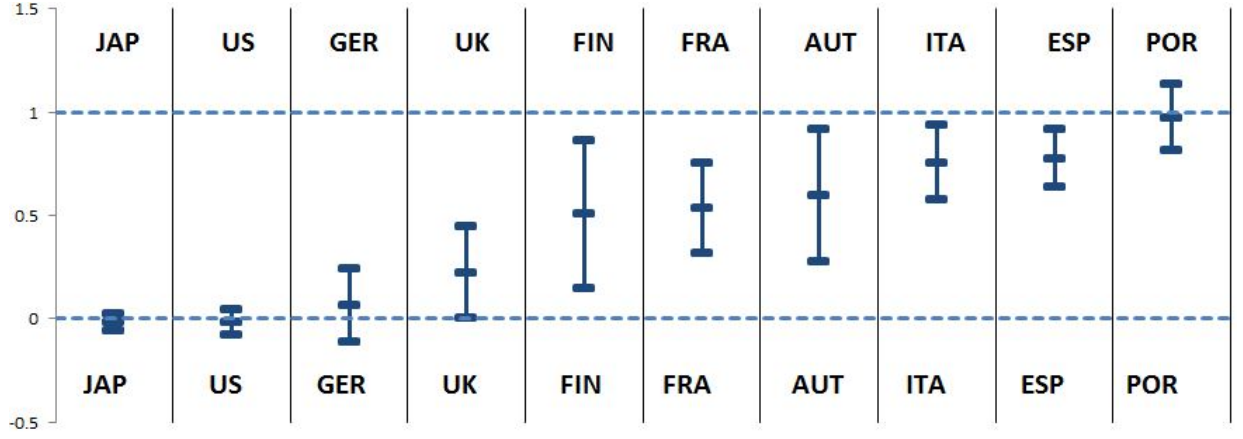
and solving for s gives:

$$s_p^e := \frac{1}{1-p} \left[\frac{1}{m \left(\frac{1}{(x-n^+)^2} + \frac{1}{(n^-)^2} \right)} \times \left((\mu - r) \left(\frac{1}{x-n^+} + \frac{1}{n^-} \right) - \frac{\sigma^2}{m} \left(\frac{W_0^B - xEE}{x-n^+} + \frac{W_0^E}{n^-} \right) \right) + pLGD \right]. \quad (1.43)$$

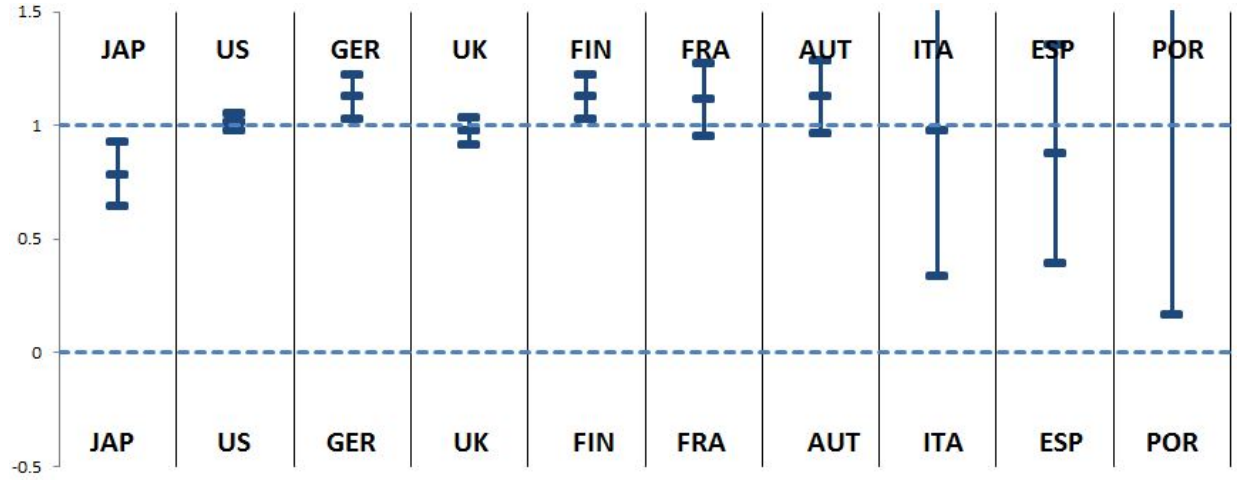
If $s^b < s_p^e \leq s^B$, we are within the domain of premiums in which the end user supplies positive protection and in which the bank does not demand full protection. Note that s^b fulfills inequality (1.29) and that

$$s^B < \frac{1}{1-p} \left((\mu - r) \min \left(\frac{x-n^+}{m}, \frac{n^-}{m} \right) + pLGD \right).$$

This completes the proof. ■



Panel A: Estimates for β^{CDS}

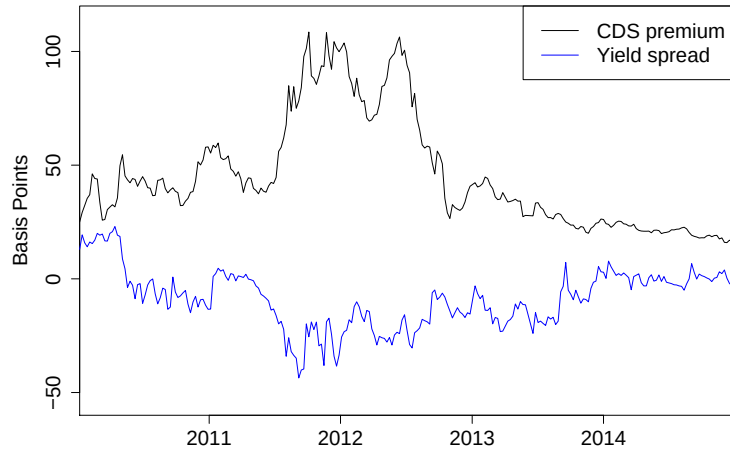


Panel B: Parameter estimates for β^{rf}

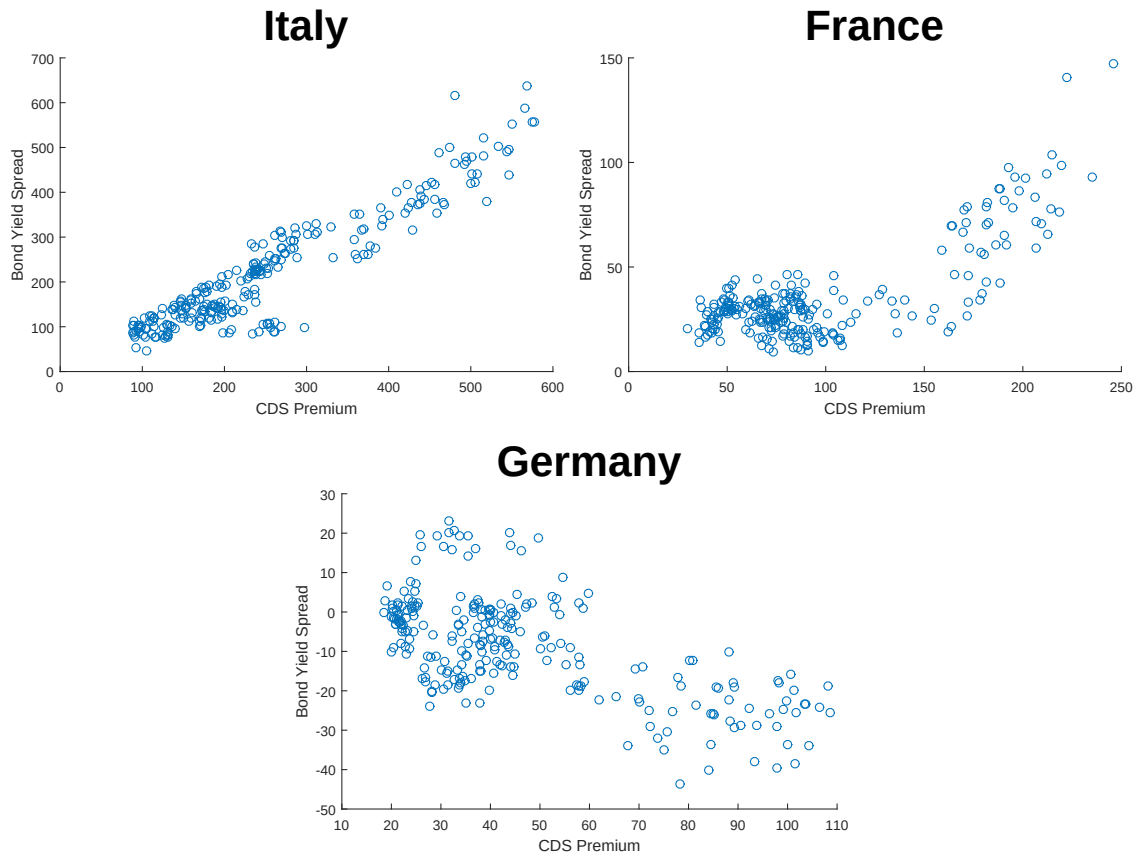
Figure 1.1: Explaining Bond Yields with Risk-Free Rates and Credit Risk. The figure shows the parameter estimates and 95% confidence interval for β^{CDS} in Panel A and for β^{rf} in Panel B for 10 different sovereigns, from the following regression:

$$\Delta Yield_t = \alpha + \beta^{CDS} \Delta CDS_t + \beta^{rf} \Delta r_{f,t} + \varepsilon_t$$

The 10 countries are sorted by β^{CDS} from lowest to highest. $Yield_t$ denotes the 5-year bond yield, $r_{f,t}$ denotes the risk-free rate proxy, measured by swap rates based on overnight lending rates in the respective currency, and CDS_t is the 5-year CDS premium. The confidence intervals are computed based on heteroskedasticity robust standard errors.



Panel A



Panel B

Figure 1.2: The disconnect between CDS premiums and bond yield spreads. Panel A shows the time series of the German five-year CDS premium and bond yield spread. Panel B shows scatter plot of CDS premium and bond yield spread Italy, France, and Germany. Yield spreads are computed as the difference between 5-year bond yields and the 5-year European Overnight swap rate (Eonia). All spreads are in basis points.

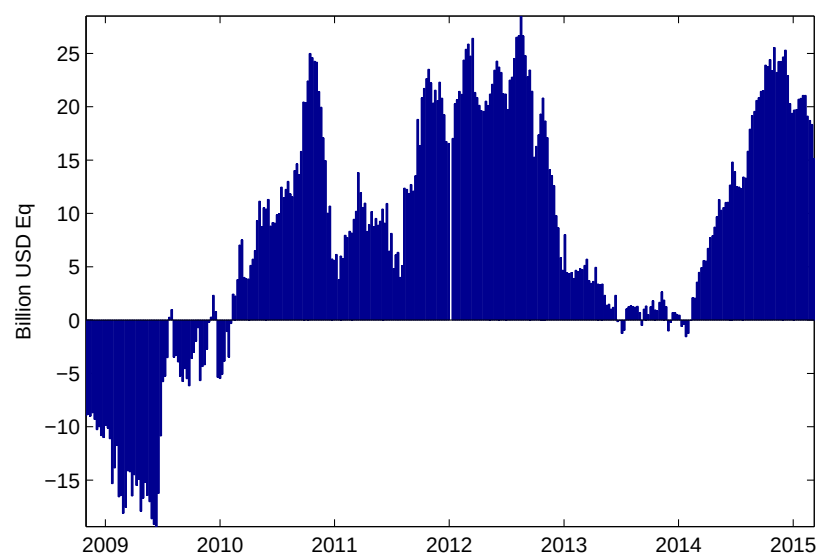


Figure 1.3: Derivatives dealers are net buyers of sovereign CDS. The Figure shows the difference between the gross amount of sovereign-CDS contracts where derivatives dealers are buying protection and the gross amount of sovereign CDS where derivatives dealers sell protection. The series is in billion US dollar and obtained from the Depository Trust & Clearing Corporation (DTCC).

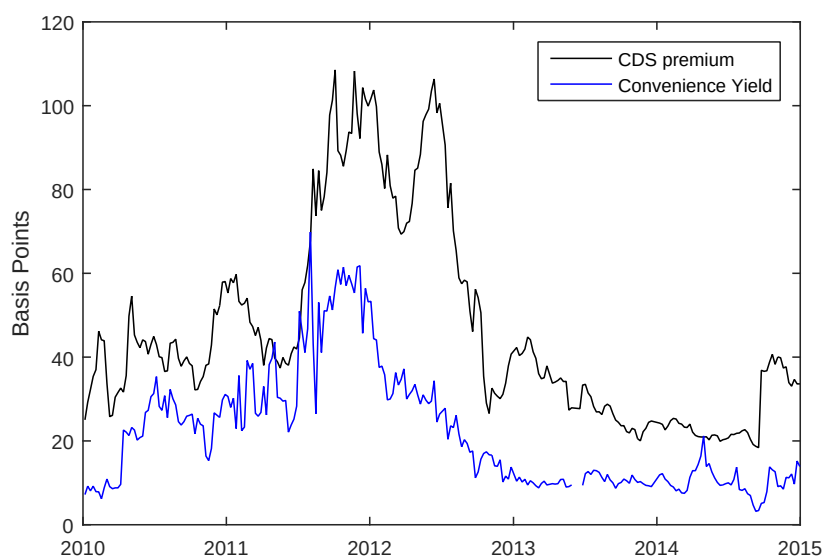


Figure 1.4: CDS premium and bond convenience yield for Germany. The Figure shows the time series of the 5-year CDS premium and a proxy for the convenience yield in the German government bonds. The convenience yield is approximated as the difference between the 3-month overnight swap rate (Eonia) and the 3-month German government bond yield, assuming that the 3-month government bond close to credit-risk free. All spreads are in basis points.

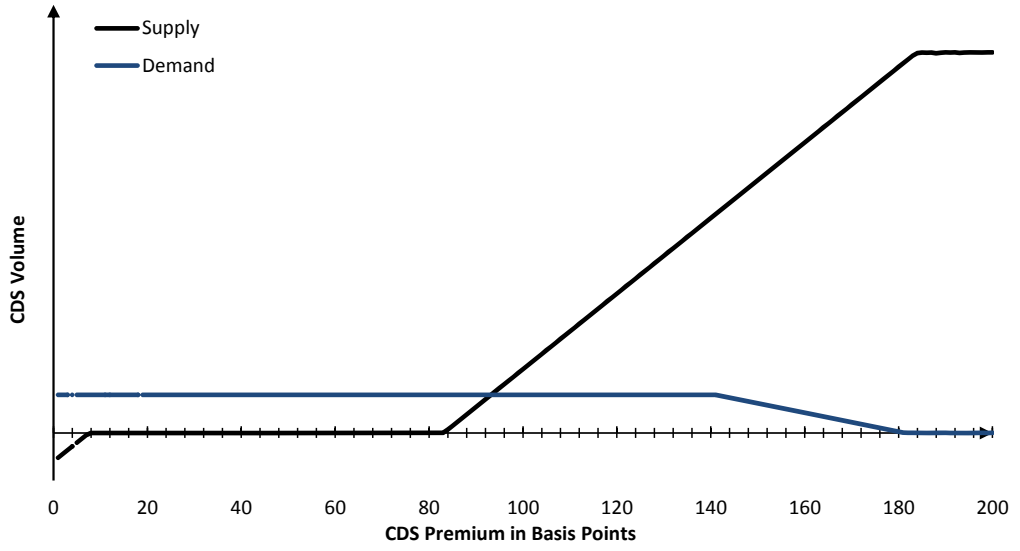


Figure 1.5: CDS supply and demand. The Figure illustrates equilibrium in the market for CDS. The black line indicates supply of CDS by the end user ($-\bar{e}$) and the blue line indicates the demand for CDS by the bank ($\bar{b}$). The market clears for a CDS premium of 93 basis points. The model parameters are: $\mu - r = 0.055$, $\sigma = 0.2$, $m = 0.2$, $n^+ = n^- = 0.05$ $W_0^E = W_0^B = 0.2$ $p = 0.75\%$ $LGD = 0.6$.

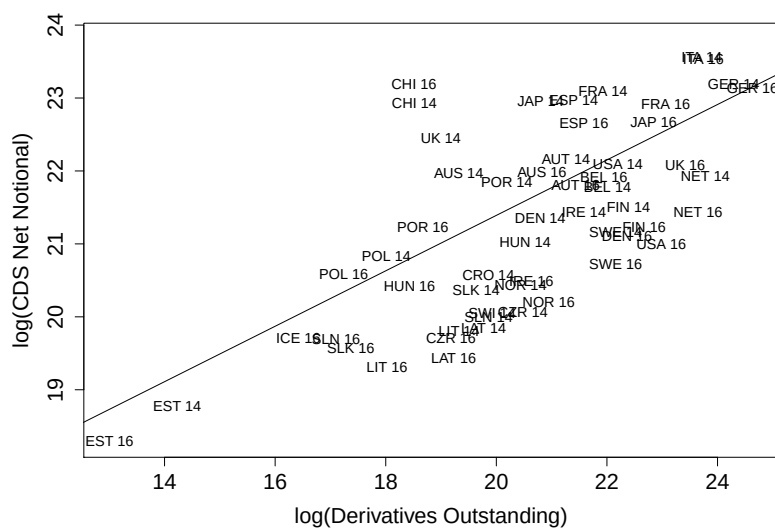


Figure 1.6: Banks' derivatives exposures and CDS volumes outstanding. This figure illustrates the relationship between the net notional amount of sovereign CDS outstanding and the fair value of all derivatives positions that European banks and banks in the UK have toward sovereigns. The fair value is the value of all derivatives positions with positive fair value, that banks have toward the respective sovereign. Data on the fair value of the derivatives positions are obtained from the EBA stress tests in December 2013 and December 2015. The net notional CDS amounts outstanding are year-end obtained from the DTCC database.

Table 1.1: CVA calculations based on EBA stress tests. OTC derivatives positions are provided by the European Banking Authority (EBA) in their stress tests from 2013 and converted to US dollar using the 2012 year-end exchange rate. Notional value (fair value) is the total value (fair value) of OTC derivatives with positive fair value, that European banks have outstanding with the respective sovereign. CDS Outst is the net notional amount of sovereign CDS outstanding. CDS refers to the 5-year CDS premium, year-end 2012. $\sigma_1(s_t)$ is the CDS volatility over the preceding year, and $\sigma_3(s_t)$ is the maximal annual volatility recorded over the preceding three years. $\frac{CS01}{EE}$ is computed using equation (1.17, and $x(s_t)$ is calculated as in equation (1.19). EE is approximated using the fair value of all derivatives with positive fair value.

Country	Mio USD			Basis Points				
	Notional Value	Fair Value	CDS Outst	CDS Premium	$\sigma_1(s_t)$	$\sigma_3(s_t)$	$\frac{CS01}{EE}$	$x(s_t)$
Germany	402,855	34,072	13,118	42	19	24	0.26%	0.150
Austria	28,403	1,644	4,224	44	40	49	0.16%	0.271
Finland	95,414	5,073	2,189	30	14	18	0.13%	0.087
France	47,938	3,210	11,742	92	46	55	0.14%	0.234
Italy	106,959	19,136	16,916	284	118	133	0.32%	0.495
Portugal	9,423	564	3,684	430	214	290	0.09%	0.821
Spain	27,691	1,883	9,259	291	118	123	0.10%	0.401
UK	7,920	19,255	5,842	42	12	30	3.08%	0.072
Japan	17,471	5,269	9,189	81	20	31	0.63%	0.091
US	77,995	54,710	3,389	36	10	19	1.56%	0.052

Table 1.2: Banks' derivatives exposures and CDS volumes outstanding. This table shows the results of regressing the logarithm of the sovereign CDS net notional outstanding on the indicated variables. $\log(FV)$ is the fair value of all derivatives positions with positive fair value, that European banks and banks in the UK have toward a sovereign. $\log(Debt)$ is the total sovereign debt outstanding for the respective country. Data on the fair value of the derivatives positions are obtained from the EBA stress tests in December 2013 and December 2015. The net notional CDS amounts outstanding are year-end obtained from the DTCC database. Amounts of debt outstanding are obtained from Countryeconomics.com. Heteroskedasticity robust standard errors are reported in parenthesis. *** indicates significance at a 1% level.

	(1)	(2)	(3)	(4)
Intercept	13.70*** (0.89)	13.73*** (1.65)	6.96*** (1.41)	6.58*** (1.59)
Intercept $\times \mathbb{1}_{\{2015\}}$		-0.05 (1.95)		0.81 (2.96)
$\log(FV)$	0.37*** (0.04)	0.37*** (0.08)	0.10*** (0.04)	0.11** (0.05)
$\log(FV) \times \mathbb{1}_{\{2015\}}$		-0.01 (0.09)		0.00 (0.09)
$\log(Debt)$			0.46*** (0.07)	0.48*** (0.07)
$\log(Debt) \times \mathbb{1}_{\{2015\}}$				-0.04 (0.16)
Observations	55	55	55	55
Adjusted R^2	0.45	0.44	0.77	0.76

Table 1.3: Explaining bond yields with risk-free rates and credit risk This table shows the results of a regression of the following form:

$$\Delta Yield_t = \alpha + \beta^{rf} \Delta r f_t + \beta^{CDS} \Delta CDS_t + \varepsilon_t.$$

$Yield_t$ is the 5-year bond yield of the most recently issued government bond, $r f_t$ denotes the risk-free rate proxy measured by 5-year overnight swap rates, and CDS_t is the 5-year CDS premium. Heteroscedasticity-robust standard errors are reported in parenthesis. Estimates of the intercept are not reported for brevity. ** Significant at 5% level, *** Significant at 1% level.

	β^{rf}	[std E]	β^{CDS}	[std E]	R^2
Japan	0.79***	[0.07]	-0.01	[0.02]	0.68
US	1.02***	[0.02]	-0.01	[0.03]	0.95
Germany	1.13***	[0.05]	0.07	[0.09]	0.8
UK	0.98***	[0.03]	0.23**	[0.11]	0.79
Finland	1.13***	[0.05]	0.51***	[0.18]	0.69
France	1.12***	[0.08]	0.54***	[0.11]	0.55
Austria	1.13***	[0.08]	0.6***	[0.16]	0.54
Italy	0.98***	[0.32]	0.76***	[0.09]	0.42
Spain	0.88***	[0.24]	0.78***	[0.07]	0.58
Portugal	1.61**	[0.72]	0.98***	[0.08]	0.56

Table 1.4: Explaining bond yields with credit risk, risk-free rates, and convenience yield. This table shows the results of a regression of the following form:

$$\Delta Yield_t = \alpha + \beta^{CDS} \Delta CDS_t + \beta^{rf} \Delta rf_t + \beta^{CY} \Delta CY_t + Controls_t + \varepsilon_t.$$

$Yield_t$ is the 5-year bond yield of the most recently issued government bond, CDS_t is the 5-year CDS premium, rf_t denotes the risk-free rate proxy measured by 5-year overnight swap rates, CY_t is a proxy for the convenience yield, measured as the difference between the 3-month overnight swap rate and the 3-month bond yield for the respective sovereign. $Controls_t$ include changes in the turnover of Treasury bonds with 3 to 6 years to maturity or Gilts with any maturity, changes in the 10-year on-the-run off-the-run spread, and changes in the spread between KfW bond yields and German government bond yields. Heteroscedasticity-robust standard errors are reported in parenthesis. * Significant at 10% level, ** Significant at 5% level, *** Significant at 1% level.

	Japan	U.S.	Germany	UK
Intercept	−0.04 (0.13)	0.07 (0.14)	−0.06 (0.27)	0.08 (0.34)
ΔCDS_t	−0.01 (0.02)	0.02 (0.08)	0.10 (0.09)	0.23* (0.13)
Δrf_t	0.48*** (0.05)	1.03*** (0.02)	1.05*** (0.04)	1.01*** (0.04)
ΔCY_t	−0.11 (0.14)	−0.01 (0.11)	−0.14** (0.07)	−0.17 (0.11)
$\Delta Turnover_t$		1.06 (0.73)		1.13* (0.64)
$\Delta OnOff_t$		−0.29* (0.15)		
ΔKfW_t			−0.25*** (0.06)	
Observations	240	253	252	189
Adjusted R ²	0.62	0.95	0.83	0.83

Table 1.5: Sovereign CDS premiums, credit risk, and regulatory proxies. The table reports parameter estimates and heteroskedasticity-robust t -statistics for regressions of the following form:

$$\Delta CDS_t = \alpha + \beta^{YS} \Delta YS_t + \beta^{Swptn} \Delta Swptn_t + \beta^{EDF} \Delta EDF_t + \varepsilon_t.$$

YS_t is the difference between 5-year bond yield and 5-year overnight swap rate in the respective currency. $Swptn_t$ is the (basis point) premium on an option to enter a 5-year swap position, as fixed payer or fixed receiver, in the respective currency, over the next 5 years. ΔEDF_t is the residual of changes in the average of the Moody's Expected Default Frequency (EDF) for the 16 largest derivatives dealing banks, regressed on changes in the yield spreads of the respective sovereign. Credit ratio denotes the ratio of the adjusted R^2 from a regression of ΔCDS_t on ΔYS_t to the adjusted R^2 from the full regression specified above. The sample period is January 2010 to December 2014, using weekly observations sampled each Wednesday. *Significant at 10% level. **Significant at 5% level. ***Significant at 1% level.

	Intercept	β^{YS}	β^{Swptn}	β^{EDF}	Adj. R^2	Credit Ratio	# Obs.
Japan	0.01 [0.03]	0.15 [0.45]	0.04* [1.88]	0.13* [1.72]	0.09	0.00	256
US	-0.08 [-0.3]	-0.05 [-0.37]	0.00 [-0.25]	0.04** [2.18]	0.01	0.00	256
Germany	-0.04 [-0.17]	0.01 [0.13]	0.04*** [2.98]	0.17*** [4.78]	0.33	0.00	256
UK	-0.22 [-1.04]	0.16*** [2.84]	0.01 [0.97]	0.11*** [4.93]	0.21	0.19	256
Finland	0.04 [0.26]	0.13*** [3.66]	0.02** [2.12]	0.14*** [6.14]	0.43	0.12	241
France	0.07 [0.17]	0.56*** [7.41]	0.05* [1.95]	0.39*** [6.89]	0.56	0.52	256
Austria	-0.10 [-0.26]	0.43*** [4.49]	0.03 [1.00]	0.26*** [4.10]	0.41	0.59	256
Italy	-0.04 [-0.04]	0.62*** [10.76]	0.13** [2.23]	0.73*** [6.14]	0.63	0.76	256
Spain	-0.22 [-0.23]	0.76*** [14.73]	0.08 [1.21]	0.38*** [3.80]	0.65	0.94	256
Portugal	0.35 [0.17]	0.58*** [12.05]	0.16 [1.00]	1.1*** [4.06]	0.65	0.86	255

Table 1.6: Summary statistics for corporate bonds and CDS. This table provides summary statistics for a sample of corporate bonds obtained from TRACE. The sample consists of all corporate bonds with a credit rating, maturities between 3 years and 10 years, and a matching credit default swap with no restructuring (docclause XR). For each trading day and each bond, the last traded yield is used. The matching CDS premium is obtained by interpolating the CDS premiums with the two nearest maturities. Swap rates based on the U.S. LIBOR curve are used as a proxy for the risk free rate. #obs per bond gives summary statistics for the time series of each bond (note that one corporation can issue several bonds). Avg Basis gives summary statistics for the average CDS bond basis, measured as the difference between CDS premium and bond yield minus risk-free rate. Avg TTM is the average time to maturity for each bond in the sample. The sample period is July 2002 to December 2014.

	Mean	SD	Min	Median	Max	N
Panel A: Aaa-Aa corporate bonds						
#obs per bond	226	295	2	88	1382	87
Avg Basis	0.10	0.18	-0.43	0.12	0.56	87
Avg TTM	5.08	1.88	3.00	4.33	9.74	87
Panel B: A corporate bonds						
#obs per bond	74	127	1.00	15	725	273
Avg Basis	-0.17	0.32	-1.90	-0.12	0.94	273
Avg TTM	5.01	1.80	3.00	4.43	9.87	273
Panel C: Baa corporate bonds						
#obs per bond	82	135.00	1	20	832	251
Avg Basis	-0.61	0.81	-8.09	-0.44	1.51	251
Avg TTM	5.18	2.03	3.00	4.39	9.98	251
Panel D: Ba-C corporate bonds						
#obs per bond	188	222	1	108	1000	158
Avg Basis	0.14	1.58	-8.73	-0.01	5.15	158
Avg TTM	5.52	1.93	3.02	5.07	9.98	158
Panel E: Aaa-Aa financials' bonds						
#obs per bond	162	218	1	56	1281	304
Avg Basis	-0.12	0.75	-3.11	-0.05	8.69	304
Avg TTM	4.88	1.75	3.00	4.27	10.00	304

Table 1.7: Link between corporate bond yields and CDS premiums. The table shows the results of a regression of the following form:

$$\Delta Yield_{i,t} = \alpha + \beta^{CDS} \Delta CDS_{i,t} + \beta^{rf} \Delta rf_t + \varepsilon_{i,t}.$$

$Yield_{i,t}$ is the bond yield of corporate bond i , $CDS_{i,t}$ is the maturity-matched CDS premium for bond i , rf_t is the maturity-matched proxy for the risk-free rate (measured as LIBOR rate). The sample period is July 2002 to December 2014. Heteroskedasticity robust standard errors, clustered on bond level are reported in paranthesis. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Aaa - Aa	A	Baa	Ba-C
Intercept	−0.00 (0.00)	−0.00 (0.00)	−0.00 (0.00)	−0.00 (0.00)
ΔCDS_t	0.42*** (0.04)	1.02*** (0.11)	0.93*** (0.05)	0.52*** (0.05)
Δrf_t	0.92*** (0.02)	0.89*** (0.02)	0.68*** (0.04)	0.02 (0.10)
Observations	19,629	20,249	20,414	29,562
Adjusted R ²	0.42	0.34	0.35	0.30

Table 1.8: Link between bond yields and CDS premiums in different episodes. The table shows the results of a regression of the following form:

$$\Delta Yield_{i,t} = \alpha + \beta^{CDS} \Delta CDS_{i,t} + \beta^{CDS} \Delta CDS_{i,t} \times \mathbb{1}_{\{Corporate\}} + \beta^{rf} \Delta rf_t + \varepsilon_{i,t}.$$

$Yield_{i,t}$ is the bond yield of bond i , $CDS_{i,t}$ is the maturity-matched CDS premium for bond i , $\mathbb{1}_{\{Corporate\}}$ is a dummy variable that equals one if the underlying is a corporate bond issuer and zero if the underlying is a financial, rf_t is the maturity-matched proxy for the risk-free rate (measured as LIBOR rate). Non-financials include bonds of non-financial corporations with Aaa or Aa rating. Financials include bonds of financial corporations with Aaa or Aa rating. Under Pre, the results for the July 2002 to June 2007 sub-period are reported. Under Crisis, the results for the July 2007 to December 2009 sub-period are reported. Under Post, the results for the January 2010 – December 2014 sub-period are reported. Heteroskedasticity robust standard errors, clustered on bond level are reported in paranthesis. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Pre	Crisis	Post
Intercept	−0.00 (0.00)	0.00 (0.00)	0.00* (0.00)
ΔCDS_t	0.98*** (0.09)	0.48*** (0.11)	0.50*** (0.07)
$\Delta CDS_t \times \mathbb{1}_{\{Corporate\}}$	−0.20 (0.27)	−0.10 (0.12)	−0.25** (0.11)
Δrf_t	0.90*** (0.02)	0.73*** (0.05)	0.96*** (0.02)
Observations	36,153	12,842	19,823
Adjusted R ²	0.43	0.22	0.57

Table 1.9: CVA calculations for Germany. All data are year end. CDS refers to the 5-year CDS premium on Germany. Worst Case is calculated as $\sqrt{\frac{10}{252}}\Phi^{-1}(0.99) \times \text{annual CDS volatility}$. CS01 is computed as first derivative of the Basel III default probability described in Section 1.3, using a flat CDS term structure based on the 5-year lag. EE is computed using Equation (1.20). Swaption Straddle is the (basis point) premium on an option to enter a 5-year swap position as fixed payer or fixed receiver over the next 5 years. Payer IRS and Receiver IRS are the amounts of Euribor-swaps outstanding by the German Federal Government. The data originally provided in Euro by the Bundesrepublik Deutschland Finanzagentur (Bund). CVA and CVA VaR are computed using Equation (1.14) and (1.15) respectively. Stressed VaR is computed based on the same formula as CVA VaR but using the maximum annual volatility over the last three years. Equity capital is computed using equation (1.18) assuming a 10% capital buffer. Margin CDS is the amount of capital that would be required to hedge the expected exposure (EE), assuming a margin requirement of $n^+ = 0.05$. $x(s_t)$ is the capital required, if the expected exposure is un-hedged and measured as a fraction of EE.

	Basis Points			Mio USD				Mio USD			Mio USD		
	CDS	Worst Case	Swaption Straddle	CS01	EE	Payer IRS	Receiver IRS	CVA	VaR	Strsd VaR	Equity Capital	Margin CDS	$x(s_t)$
2010	59	14	573	7.14	8,650	122,814	179,216	260	294	352	808	433	0.093
2011	101	23	691	9.66	12,028	152,230	195,741	311	654	655	1,636	601	0.136
2012	42	16	588	10.09	11,061	198,630	177,918	241	499	752	1,563	553	0.141
2013	26	8	674	10.19	11,116	169,329	160,504	198	257	759	1,270	556	0.114
2014	34	8	522	7.95	8,426	174,361	148,467	118	200	592	991	421	0.118

Essay 2

An Explanation of Negative Swap Spreads¹

¹We are grateful to Darrell Duffie, Wei Jiang, David Lando, Scott McDermott, Pedro Serrano, Morten Sørensen, Hyun Shin, and Savitar Sundaresan for helpful comments. Klingler acknowledges support from the Center for Financial Frictions (FRIC), grant no. DNRF102.

Abstract

The 30-year US swap spreads have been negative since September 2008. We offer an explanation for this persistent anomaly. Through a model, we show that the demand for swaps arising from duration hedging needs of underfunded pension plans, coupled with balance sheet constraints of swap dealers, can drive swap spreads to become negative. We construct an empirical measure of the aggregate funding status of Defined Benefits (DB) pension plans from the Federal Reserve's financial accounts of the United States and show that this measure is a significant explanatory variable of 30-year swap spreads, but not for swaps with shorter maturities.

2.1 Introduction

In September 2008, shortly after the default of Lehman Brothers, the difference between the swap rate (which is the fixed-rate in the swap) of a 30-year interest rate swap (IRS) and the yield of a Treasury bond with the same maturity, commonly referred to as swap spread, dropped sharply and became negative. As we explain in more detail later, this is a theoretical arbitrage opportunity and a pricing anomaly. In contrast to other crises phenomena, the 30-year negative swap spread is very persistent and still at around -40 basis points as of December 2015. In this paper, we examine the persistent negative 30-year swap spread and offer a new perspective on the possible reasons behind this anomaly. Our hypothesis is that demand for duration hedging by underfunded pension plans coupled with balance sheet constraints faced by swap dealers puts pressure on long-term swap fixed rates and ultimately turned the 30-year swap spread negative.

Negative swap spreads are a pricing anomaly and present a challenge to views that have been held prior to the financial crisis that suggested that swap spreads are indicators of market uncertainty, which *increase* in times of financial distress. This is because the fixed payment in an IRS is exchanged against a floating payment, which is typically based on Libor, and entails credit risk. Hence, even though IRS are collateralized and viewed as free of counterparty credit risk, the swap rate should be above the (theoretical) risk-free rate because of the credit risk that is implicit in Libor. Therefore, swap spreads should increase in times of elevated bank credit risk (see Collin-Dufresne and Solnik, 2001, for a treatment of this and related issues). Additionally to that, treasuries (which are the benchmarks against which swap spreads are computed) have a status as "safe haven", i.e., assets that investors value for their safety and liquidity. In times of financial distress, investors value the convenience of holding safe and liquid assets even more, which decreases the treasury yield and makes them trade at a liquidity premium or convenience yield (see,

for instance, Longstaff, 2004, Krishnamurthy and Vissing-Jorgensen, 2012a, or Feldhütter and Lando, 2008). In summary, these arguments show that the 30-year swap spread should have increased around the default of Lehman Brothers.

Contributions of the paper

We offer a demand-driven explanation for negative swap spreads. In particular, we develop a model where underfunded pension plans' demand for duration hedging leads them to optimally receive the fixed rate in IRS with long maturities. Pension funds have long-term liabilities in the form of unfunded pension claims and invest in a portfolio of assets, such as stocks, as well as in other long-term assets, like government bonds. They can balance their asset-liability duration by investing in long-term bonds or by receiving fixed in an IRS with long maturity. Our theory predicts that, if pension funds are underfunded, they prefer to hedge their duration risk with IRS rather than buying Treasuries, which may be not feasible given their funding status. The preference for IRS to hedge duration risk arises because the swap requires only modest investment to cover margins, whereas buying a government bond to match duration requires outright investment. Thus, the use of IRS allows the underfunded pension funds to invest their scarce funds in assets (such as stocks) with higher expected return.

Greenwood and Vayanos (2010) show that pension funds' demand for duration hedging in the UK can affect the term structure of British gilts by lowering long-term rates. In this sense, our paper bears a close relationship to their work. However, our approach differs from theirs since we focus on *underfunded* pension funds' optimal preference for the use of IRS for duration hedging. The model that we develop shows that the demand for IRS increases as the fund becomes more underfunded, and the sponsor combines the IRS positions with positions in the (risky) stock portfolio in the hope of potentially overcoming the underfunded status.

We provide non-parametric evidence suggesting that the swap spreads tend to be negative in periods when DB plans are underfunded. We thus illustrate a new channel that may be at work in driving long-term swap spreads down. Using data from the financial accounts of the United States (former flow of funds table) from the Federal Reserve, we construct a measure of the aggregate under-funded status of DB plans (both private and public) in the United States. We then use this measure to test the relationship between the underfunded ratio (UFR) of DB plans and long-term swap spreads in a regression setting. Even after controlling for other common drivers of swap spreads, recognized in the literature, such as the spread between LIBOR and repo rates, Debt-to-GDP ratio, dealer-banks' financial

constraints, market volatility, and level as well as the slope of the yield curve, we show that the UFR is a significant variable in explaining 30-year swap spreads. In line with our narrative, we also show that swap spreads of shorter maturities are *not* affected by changes in UFR.

One concern about using UFR as an explanatory variable for swap spreads is that the same factors that have been shown to affect swap spreads can also affect pension funds. For example, a decrease in the level of the yield curve can affect swap spreads and also increases the level of pension funds' underfunding because the present value of the funds' liabilities increases. To address this concern, we use stock returns as an instrumental variable in a two-stage least square setting. The idea here is that stock returns affect pension funds' funding status but are unrelated to swap spreads. Our results are robust to this additional test. We conclude our paper by showing the effect of pension funds' underfunding on swap spreads for two other pension systems: Japan and the Netherlands.

Related Literature

As mentioned above, Greenwood and Vayanos (2010) show that the demand pressure by pension funds lowers long-term yields of British gilts. Additionally to that, Greenwood and Vayanos (2010) mention that pension funds also fulfill their demand for long-dated assets by using derivatives to swap fixed for floating payments. They note that pension funds have “swapped as much as £50 billion of interest rate exposure in 2005 and 2006 to increase the duration of their assets” but do not investigate the impact of such demand on swap spreads any further. Their focus was on U.K. Gilt markets. Hence, our paper complements their analysis by showing that underfunded pension funds' demand for long-dated assets can have a strong impact on swap rates.

More generally, swap rates and treasury yields have been studied extensively in the previous literature. A stream of literature calibrates dynamic term-structure models to understand the dynamics of swap spreads (see Duffie and Singleton, 1997, Lang, Litzenberger, and Liu, 1998, Collin-Dufresne and Solnik, 2001, Liu, Longstaff, and Mandell, 2006, Johannes and Sundareshan, 2007, and Feldhütter and Lando, 2008, among others). Amongst these papers, the paper close in spirit to our paper is Feldhütter and Lando (2008). They decompose swap spreads into three components, credit risk in Libor, the convenience yield of government bonds, and a demand-based component. In contrast to our paper, their study focuses on maturities between one and ten years and they link the demand-based component to duration hedging in the mortgage market.

The usage of swaps by non-financial companies has been studied by, among others,

Faulkender (2005), Chernenko and Faulkender (2012), Jermann and Yue (2013). We focus on pension funds' underfunding issues, which have been studied by, among others, Sundaresan and Zapatero (1997) and Ang, Chen, and Sundaresan (2013). We add to this literature by linking changes in swap spreads to changes in pension fund underfunding.

We note that any demand-based explanation would be incomplete if there were no financial frictions for the supply of IRS. Hence, we also build on the literature of limits of arbitrage (Shleifer and Vishny, 1997, Gromb and Vayanos, 2002, Liu and Longstaff, 2004a, Gromb and Vayanos, 2010, Gârleanu and Pedersen, 2011, among many others) and especially the literature on dealer constraints and demand pressure in the derivatives market (Gârleanu, Pedersen, and Poteshman, 2009).

To the best of our knowledge, we are the first to offer a demand-based explanation for negative swap spreads. In contrast to our demand-based explanation for negative swap spreads, Jermann (2016) studies the negative swap spreads, offering frictions for holding long-term bonds as an explanation. In contrast to our paper, Jermann (2016) takes the demand for long-dated swaps as exogenously given and focuses explicitly on the risks of holding long-dated bonds to hedge the cash flows of long-dated swaps. In his model, a risk-averse derivatives dealer chooses his optimal investment in short-term government bonds, long-term government bonds, and long-dated swaps. Jermann (2016) assumes that holding bonds is costly and shows that as costs increase, the swap rate converges to the Libor rate. Since long-term Treasury yields are typically above Libor, his model predicts that there is a negative relationship between swap spreads and term spreads, where term spreads are proxied as the difference between long-dated treasuries and short-dated treasuries. He provides some empirical evidence showing the link between term spreads and swap spreads. Our explanation is distinct from his work, as the *UFR* measure of underfunded status of DB pension plans is a significant variable in explaining 30-year swap spreads but not for swap spreads with other maturities. Furthermore, controlling for term spreads leaves our main results unchanged. Holding outright long positions in bonds for under-funded pension plans to match duration has an opportunity cost in practice and this is what we stress in our work. Lou (2009) also offers derivatives dealers' funding costs as an explanation of negative swap spreads.

Finally, there is a wide range of industry research offering a variety of different reasons for the persistent negative 30-year swap spread. One frequently used explanation is the potential credit risk of US Treasuries.² The problem with this argument is that while Treasuries are linked to the credit risk of the US, swap rates are linked to the average credit risk of the banking system and a default of the US government would most likely cause

²See, for instance, <https://self-evident.org/?p=780>.

defaults in the banking system as well. A second, commonly-offered explanation, is the different funding requirements of swaps and treasuries.³ Long-term Treasury holdings are outright cash position while engaging in IRS requires only modest capital for initial collateral, typically a small fraction of the Treasury bond principal. Sophisticated investors can use repo agreements to purchase/finance Treasuries, although financing Treasury securities for 30 years would require open repo positions, which need to be rolled over for a long duration. The risk with such a strategy is that the cash lenders may refuse to renew the repo agreement. These considerations may cause pension funds to prefer swaps as opposed to a repo-financed positions in government bonds.

The roadmap of the paper is as follows. Section 2.2 of the paper provides some motivating evidence. In section 2.3, we present the swap spreads and the underlying drivers for the demand for receiving fixed rates in long-term swaps from pension funds. In section 2.4, we develop a dynamic model with stochastic interest rates, which shows that the need to match the duration of assets and liabilities can lead to a demand for receiving fixed in long-term swaps, when the pension plan is underfunded. Section 2.5 contains our empirical results. Section 2.6 concludes.

2.2 Motivating Evidence

We motivate our model, by documenting a few stylized facts: we first show in Figure 2.1 that the 30-year swap spread became negative following the bankruptcy of Lehman Brothers, and has been in the negative territory since then.

We can see from Figure 2.1 that the term structure of swap spreads track each other closely until the end of 2007 when long-term swap spreads start decreasing relative to short-term spreads. Since then, the dynamics of the 30-year swap spreads have decoupled from the dynamics of the other tenures. In the month after the default of Lehman Brothers, highlighted by the first vertical line, the 30-year swap spread drops sharply and turns negative. During that period, there is also a decline in the 10-year swap spread, while swap spreads of shorter maturities increase. Between 2008 and 2014 the 30-year swap spread slowly converges close to 0 and starts decreasing again in 2015. In August 2015, highlighted by the second vertical line, the Libor-Repo spread turns negative, which causes a decrease in swap spreads of all maturities⁴.

We perform a principal components analysis (PCA) of swap spreads before and after

³See, for instance, Van Deventer (2012).

⁴The prolonged drop in interest rates, following the crisis of 2008, increased the duration of pension liabilities and the monetary policy of the Fed also might have contributed to the overall drop in other interest rates and spreads.

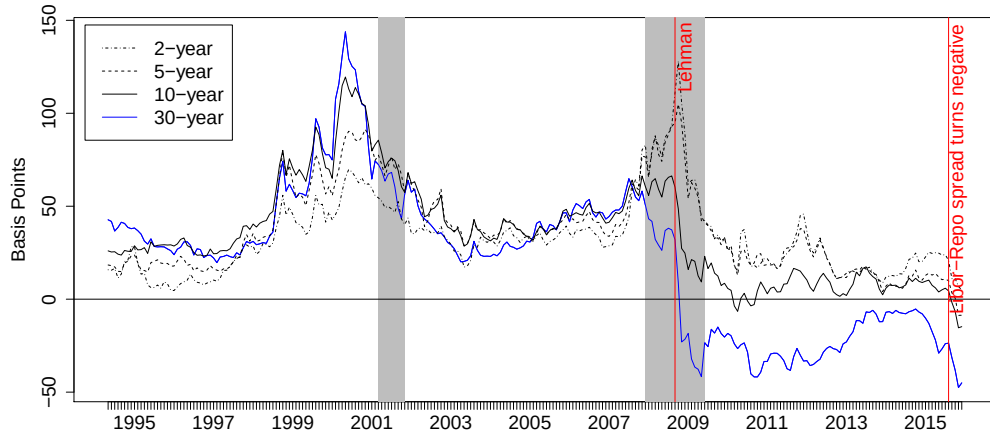


Figure 2.1: Term structure of interest rate swap spreads: The graph shows the history of swap spreads from May 1994 until December 2015. The grey shaded areas represent US recession periods. The source for our data is Bloomberg. The differences in market conventions have been taken into account in computing the spreads

September 2008 to see if there is a significant change in the PCs driving the swap spreads after the crisis, relative to the drivers prior to the crisis. The results of our PCA are shown in Table 2.1 next. We present the loadings of each PC before and after September 2008 as well as the proportion of the spreads explained by each PC.⁵ Note that prior to the crisis,

Table 2.1: OLS Regressions of Swap Spreads on Principal Components. For each of the two sub-periods, *PC1* and *PC2* are the first and second principal components of the term structure of swap spreads for this sub period. We use monthly treasury yields and swap rates from the Bloomberg system to compute swap spreads. 3-year and 7-year treasury data are supplemented using data from the FED H15 reports. The sample period is May 1994 – July 2015. $\hat{\beta}$ is the estimated loading on the respective PC.

	Pre September 2008				Post September 2008			
	PC1		PC2		PC1		PC2	
	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2
2 yr	0.33	78.00%	0.41	18.20%	0.53	97.00%	-0.03	0.10%
3 yr	0.33	74.40%	0.46	23.10%	0.56	98.10%	-0.01	0.00%
5 yr	0.4	91.10%	0.26	6.00%	0.45	97.90%	0.04	0.20%
7 yr	0.37	94.10%	0.16	2.80%	0.36	86.50%	-0.14	2.80%
10 yr	0.43	94.80%	-0.14	1.70%	0.19	64.40%	-0.26	24.80%
20 yr	0.35	78.10%	-0.42	17.30%	-0.20	24.80%	-0.71	70.20%
30 yr	0.42	77.00%	-0.57	21.50%	0.06	3.10%	-0.64	85.60%

the first PC explained more than 75% of the variations in swap spreads for all maturities.

⁵In this analysis, we include 7-year and 20-year swap spreads as well. Data for these spreads are obtained from the FED H.15 website.

The explanatory power of the second PC varied from 23.1% for 3-year swap spreads to 1.7% for 10-year swap spreads. After the crisis, the first PC became even more important in explaining the swap spreads of maturities up to five years, and less so for maturities from seven to thirty years. But the drop in its explanatory power for the 30-year swap spreads is dramatic: it fell from 77.0% to just 3.1%. In fact, the second PC became the dominant component in explaining the swap spreads for 30-year maturity, in sharp contrast with swap spreads associated with shorter maturities of 10 years or less. Similarly, but to a smaller extent, the explanatory power of the first PC decreased from 78.10% for the 20-year swap spread to 24.8%, while the explanatory power of the second PC increased from 17.3% to 70.2%. Our results in Table 2.1 demonstrate that the determinants of 30-year swap spreads underwent a big change after September 2008. This change appears to be unique for swap spreads with maturities above 10-years. To a lesser extent, we see a similar effect for the ten year swap spreads as well.

Taken together, Figure 2.1 and Table 2.1 suggest that the 30-year swap spreads behaved qualitatively different from the rest of the swap spreads after September 2008. This provides the motivation for both our theory and empirical work. We provide next a possible link between the above evidence and the funding status of defined benefit (DB) pension plans. DB Pension funds have long-dated liabilities and they use long-term interest rate swaps to hedge their duration risk in swap overlay strategies. Adams and Smith (2009) show how interest rate swaps are used by pension funds to manage their duration risk. Furthermore, CGFS (2011) documents that insurance companies and pension funds need to balance asset-liability durations and can do so using swaps.

In theory, a sophisticated investor with full access to repo financing, can buy Treasury bonds and use the repo market to obtain an almost unfunded position. This repo transaction requires an initial funding of approximately 6%.⁶ At the same time, engaging in an IRS could also require an initial margin and regular collateral posting. With the implementation of mandatory central clearing this is becoming more of an issue recently. Nevertheless, as noted earlier, financing a long-term bond for thirty years remains a less practical proposition than merely entering into an interest rate swap. As noted in a recent Bloomberg article (see Leising, 2013), US pension funds use IRS markets. Overall, pension funds may find long-term IRS as a simpler vehicle to take leverage than utilizing the repo market for duration hedging purposes.⁷

⁶This number is a first approximation that we obtained from <http://www.cmegroup.com/clearing/financial-and-collateral-management/>. They analyze haircuts for securities posted as collateral in cleared derivatives transactions. However, market participants confirm that 6% is a reasonable proxy for haircuts of Treasuries with 30 years to maturity.

⁷There may be other frictions such as taxes that may also favor IRS relative to repo. In the US, Internal Revenue Service views repo as financing that would subject the pension plan to tax filings as Unrelated

Further anecdotal evidence of pension funds' demand for IRS and resulting demand pressure is best summarized by the following quote from a recent Bloomberg article: "Pension funds need to hedge long-term liabilities by receiving fixed on long-maturity swap rates. When Lehman dissolved, pension funds found themselves with unmatched hedging needs and then needed to cover these positions in the market with other counterparties. This demand for receiving fixed in the long end drove swap spreads tighter."⁸

We provide next some motivating evidence that suggests a strong association between the funded status of pension plans and thirty year swap spreads. The size of pension funds in the United States is significant relative to the GDP of the US economy.⁹ To make the case that the demand by pension funds to receive fixed in the long-term swap contracts can potentially influence the 30-year swap spread, Figure 2.2 offers a comparison between the size of the interest rate swap market to the value of pension funds' liabilities. The solid line indicates the mark-to-market value of USD interest rate swap contracts with a maturity of more than 5 years. The dashed line illustrates the time series of the total pension liabilities in the US defined benefits plans, which are the focus of our paper.

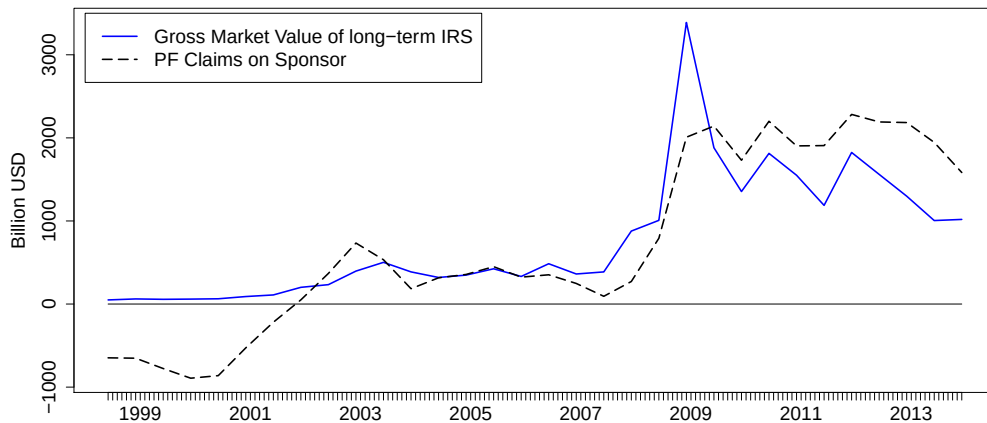


Figure 2.2: Size of pension liabilities and long-term Interest Rate Swaps: This plot illustrates that the total unfunded liabilities of private as well as state and local government employee defined-benefit (DB) pension plans are qualitatively similar to the gross market value of interest rate swaps denominated in US dollars with maturity greater than five years. The amounts are in billions of dollars, not seasonally adjusted. (Source: BIS and financial accounts of the U.S.)

Business Income (UBI). Most US pension plans will therefore avoid UBI taxes by avoiding repo and relying on IRS, which does not invoke UBIT. We thank Scott McDermott for alerting us to this point.

⁸See <http://www.bloomberg.com/apps/news?pid=newsarchive&sid=aUq.d1dYuhEA>

⁹The size of pension plan assets in the US is about \$13.60 trillion dollars as of the first quarter of 2015.

2.3 Demand for and Supply of Duration

In this section we discuss pension funds, their duration matching needs and how underfunding affects their demand for long-dated IRS. We briefly review the implications of regulations such as the pension protection act of 2006 and the diminished incentives to overfund pension plans, due to some tax policy developments. We conclude with an overview of the demand for receiving fixed in long-dated IRS as well as the supply of long-dated IRS.

2.3.1 Pension Funds' Duration Matching Needs

The most important customers in the long end of the swap curve are pension funds and insurance companies, who have a natural demand for receiving fixed for longer tenors. Pension funds have long-term liabilities towards their clients and the Pension Protection Act of 2006 requires them to minimize underfunding by stipulating funding standards and remedial measures to reduce under-funded status. This promotes the incentive to match the duration of their asset portfolios with the duration of these liabilities: any duration mismatch can produce future shortfalls. Increasing the duration of their asset portfolios could be achieved by receiving fixed in an IRS or by buying bonds with long maturities. Greenwood and Vayanos (2010) provide evidence from the 2004 pension reform in the United Kingdom where pension funds started buying long-dated gilts and more recently Domanski, Shin, and Sushko (2015) show that German insurance companies increased their holdings of German long-term bonds significantly over the past years. In line with previous research (see, for instance, Ang et al., 2013 or Ring, 2014, among many others) we document that many US pension funds are underfunded and therefore tend towards more risky investments. Using IRS instead of long-dated Treasuries for duration hedging allows pension funds to use their limited funding to invest in more risky assets such as stocks.

To illustrate that pension funds are indeed using swaps, we collect survey data from the Chief Investment Officer magazine, who conducts regular surveys on US pension funds and their investment strategies.¹⁰ In 2013, 2014, and 2015 they surveyed more than 100 US pension fund managers on their investment strategies. The question most relevant to this paper was whether the plans are using derivatives. A majority of 64.6%, 63%, and 70% of the respondents in 2013, 2014, and 2015, respectively, stated that they were currently using derivatives. In 2013 and 2014 the respondents provided additional details on their derivatives usage. In 2013 and 2014 80.9% and 79% stated that they were using interest rate swaps, among other derivatives. Furthermore, 25.4% (29%) of the respondents in 2013 (2014) stated that they were using derivatives to obtain leverage and 49.2% (39%) stated

¹⁰These surveys are available under <http://www.ai-cio.com/surveys/>.

that they are using derivatives for capital/cash efficiency.

Pension Funds' Aversion to Over-funding after 1990

During the period 1986-1990, laws were enacted in the US to discourage “pension reversions” whereby, a pension plan with excess assets can be tapped into by the sponsoring corporation to draw the assets back into the corporation. In 1986, the reversion tax rate was 10% but by 1990, this tax rate had increased to 50%. In addition, the sponsoring firm was also required to pay corporate income tax on reversions. These changes in tax policies meant that the US corporations have dramatically lower incentives to overfund their pension plans since 1990 than was the case before. This is important to note because pension plans were generally not significantly overfunded before the onset of the credit crisis of 2008 which made them vulnerable to becoming underfunded should there be a big correction in equity markets or a protracted fall in discount rates, which can cause the pension liabilities to increase (both these developments occurred after the credit crisis of 2008)

Once the plans become underfunded and the rates fall (as was the case after 2008), the plan sponsors are faced with two objectives: first to match the duration of assets with liabilities to avoid future underfunding due to market movements, and second to find assets which can provide sufficiently high returns to get out of their underfunded status. This is the context in which the long-term swaps play a role: they enable the sponsors to match duration without setting aside any explicit funding and the sponsor can then use the limited funding to invest in riskier assets in the hope of earning higher returns.

2.3.2 The Supply of Long-Dated Swaps

Investors

In general, investors could either have a demand for receiving fixed in an IRS or for paying fixed in an IRS and may use IRS for speculative and hedging purposes. In any case, the demand for IRS can depend on the level and the slope of the yield curve. The level of the yield curve matters, for instance, for agencies issuing Mortgage-Backed Securities (MBS). Agencies aim to balance the duration of their assets and liabilities. When interest rates fall, mortgage borrowers tend to execute their prepayment right, thereby lowering the duration of the agencies' mortgage portfolio. Hence, agencies want to receive fixed in an IRS to hedge this mortgage prepayment risk (see Hanson, 2014).¹¹ The slope of the yield curve may also matter for non-financial firms. According to Faulkender (2005), these firms tend

¹¹Feldhütter and Lando (2008) argue that using IRS is the predominant way for doing this as opposed to using Treasuries.

to use IRS mostly for speculation, preferring to pay floating when the yield curve is steep. Faulkender (2005) also finds that firms tend to prefer paying fixed when macro-economic conditions worsen. Overall, these papers show that there could be demand and supply effects from other investors. However, as these examples suggest, it is hard to conclude that non-financial firms have a large demand for long-dated IRS with a maturity of 30 years.¹² We therefore conclude that the demand for receiving fixed in long-dated IRS by pension funds has to be met largely by derivatives dealer-brokers.

Dealer-Brokers

A dealer-broker paying fixed in a long-dated IRS, thereby taking the opposite position than a pension fund would generally aim to hedge the interest-rate risk of his position. He can either do so by finding another counterparty willing to pay fixed or by following the hedging strategy described in Table 2.8 in the appendix where he purchases a 30-year treasury bond financed with a short-term repo transaction in order to hedge the duration risk. We discussed above that finding a counterparty willing to pay fixed in long-dated swaps is difficult and now highlight several constraints with this hedging strategy that limit the supply of long-dated IRS.

The first issue has to do with margin requirements. Financing the purchase of the long-dated government bond with short-term borrowing is subject to the risk of increasing margin requirements. For instance, Krishnamurthy (2010) documents that haircuts for longer-dated government bonds increased from 5% to 6% during the crisis. The haircut for 30-year bonds conceivably increased even more. Furthermore, Musto, Nini, and Schwarz (2014) document that the amount of repo transactions decreased sharply during the financial crisis. One possible reason for this observation is that the supply of repo financing deteriorated and hence borrowing at repo was not always possible, especially for long-term swaps. Hence, the arbitrage strategy is subject to a severe funding risk. Furthermore, engaging in an IRS requires an initial margin as well. This margin requirement increased after the financial crisis. Hence the dealer may be forced to offer a lower fixed rate on long-term swaps.

The second issue is a standard limits of arbitrage argument. As pointed out by Shleifer and Vishny (1997), Liu and Longstaff (2004a), and many others, arbitrage opportunities are subject to a risk: it is the possibility that the mispricing increases before it vanishes, thereby forcing the arbitrageur out of his position at a loss. With negative 30-year swap spreads arbitrage, we know that the mispricing vanishes after 30 years, but we do not know

¹²Insurance companies could be another big demanders for receiving fixed rate on long-term swaps, but we have no data available to characterize their demand. Additional to insurance companies, recent long-term corporate bond issuance might also create a demand for receiving fixed in long-dated interest rate swaps. This is because companies may hedge the duration risk of their bond issuance.

whether it will vanish within a much shorter and practical horizon. To benefit from negative swap spreads arbitrage a high amount of leverage is required and arbitraging negative swap spreads can therefore be seen as a case of “picking up Nickels in front of a steamroller” (Duarte, Longstaff, and Yu, 2007).¹³

2.4 Model

In this section, we develop a model that links pension funds’ underfunding to swap spreads, proceeding in three steps. First, we show that underfunded pension plans optimally take a long position in interest rate swaps, by receiving fixed and paying floating. Second, we model the supply of long-dated swaps in reduced form, assuming that derivatives dealers provide fixed rates in swaps elastically up to a certain threshold at the arbitrage-free rate, and then require additional compensation in the form of negative swap spreads above this threshold, due to their balance sheet constraints. Third, bringing the demand and supply side of our model together, we show how pension funds’ underfunding leads to decreasing swap spreads.

2.4.1 Demand For Long-Dated Swaps by Pension Funds

Model with Swaps and Safe Assets

We now present a simple model of an under-funded pension plan, which has assets A_t and a flow rate of liabilities L per unit time, that live forever. The pension plan is underfunded, i.e., $A_0 < PV(L)$. *We make this assumption to explicitly model under-funded pension plans. In addition, this implies that the fund cannot buy a perpetual bond to match the cash flows. Hence, there is a natural role for interest rate swaps, which are funded as a floating rate.* The fund can also contribute at a rate y_t per unit time. Formally, the sponsor’s value function is given as:

$$G(t, A) := \min_{\{y_t, m_t\}} \mathbb{E}_t \left[\int_t^\tau e^{-\rho s} u(y_s) ds \right], \quad (2.1)$$

¹³Another friction we abstract away from in this project is the possible presence of credit risk in US Treasuries. The reason for doing so is that it is not obvious how an increase in credit risk in US Treasuries might affect the swap spread. Clearly, an increase in treasury credit risk would increase the treasury yield and assuming all else equal, a decrease in the swap spread would result. However, it is not obvious that swap rates would be unaffected by the increase in treasury credit risk since interbank lending rates would presumably increase sharply when US credit risk increases. Therefore, it is just as likely that the swap rate would be elevated.

where we assume CRRA utility with risk-aversion coefficient $\beta > 1$: $u(y) = y^\beta$ and where the fund optimally chooses its funding rate and asset allocation between interest rate swaps and safe assets to reach a fully funded status at a future point.¹⁴ Once the plan is fully funded, it simply dedicates the cash flows from its portfolio to meet its liabilities.

We define a swap as one in which the pension plan receives a fixed dollar amount of 1 per unit time, and pays a floating rate of r_t per unit time. The derivatives dealer, who is the counterparty to this IRS, will be introduced in the following section. We abstract from credit risk, which implies that the fixed rate of the swap can be funded at the risk-free floating rate. The value of this swap is $P - 1$, where P is the value of a perpetuity which pays \$1 forever. The value of floating payments $\{r_s, s \geq t\}$ is simply 1.¹⁵ This is a stylized representation of an interest rate swap, which differs from a newly minted interest rate swap, which will always be valued at zero. Our stylized representation of the swap provides tractable and simple closed form solutions.¹⁶ The swap can be regarded as a seasoned swap, which the fund enters into for duration matching purposes. The pension funds can buy m swaps which cost $m[P - 1] < A_t < PV[L]$, where A_t is the value of the assets held by the pension funds at time t . The remaining funds $A_t - m[P - 1]$ are invested in r_t at each time. It should be noted the fund will pay the intermediation costs, δ , per unit time, which affects the cash flows. The cash flows from the swap position is: $m(1 - r - \delta)$ per unit time.

The term structure model is a simple one-factor model where the instantaneous interest rate r follows a diffusion process, as in Constantinides and Ingersoll (1984):¹⁷

$$dr = \alpha r^2 dt + sr^{3/2} dw_1 \quad (2.2)$$

For this process, the consol bond price P is:

$$P = \frac{1}{(1 + \alpha - s^2)r} = H/r, \quad (2.3)$$

where $H \equiv \frac{1}{1 + \alpha - s^2}$. We can derive the dynamics of the pension fund's asset value as:

$$dA = [Ar + m(1 - \delta) + y - L]dt - mPs\sqrt{r}dw_1. \quad (2.4)$$

The dynamic problem facing the pension sponsor is specified below: we choose A to be

¹⁴We have also solved the model under the alternative assumption that the pension fund can invest in a perpetual bond by borrowing short-term. The results for this assumption are available upon request.

¹⁵See, Cox, Ingersoll, and Ross, 1985 for a proof of this assertion.

¹⁶Modeling a swap that is zero valued is feasible, but may likely require a numerical approach in the context of our model, where the fund has to also choose asset allocation.

¹⁷In Appendix 2.7.1, we briefly characterize the term structure of zero coupon yields implied by this model.

our state variable, and formulate the HJB equation associated with the funds' optimization problem next:

$$0 = \inf_{y_t, m} \left[y^\beta - \rho G + G_A [m(1 - \delta) + rA + y - L] + \frac{1}{2} G_{AA} m^2 P^2 s^2 r \right]. \quad (2.5)$$

The fund will close out the swaps, payoff any loans, and stop contributing when the assets are sufficient to meet the present value of the liabilities. Note that when $m = y = 0$, the dynamics of assets are:

$$dA = [Ar - L]dt = r[A - \frac{L}{r}]dt = r[A - LP(1 + \alpha - s^2)]dt \quad (2.6)$$

When $A \uparrow A^*$ where $A^* = LP(1 + \alpha - s^2)$, note that $A^*r = L$, and the fund can meet its liabilities from its assets, and the value function goes to zero. This leads to the following boundary condition for the HJB equation above:

$$G(A \uparrow A^*) = 0. \quad (2.7)$$

The above condition follows from the fact that the cost of funding goes to zero when the assets are sufficient to meet the liabilities. Let us define $\Psi \equiv A^* - A$. We can now characterize the demand functions of the pension fund and the optimal funding policy.

Demand Functions

Proposition 2. *The sponsor's optimal contribution and optimal asset allocation are given as:*¹⁸

$$y = \frac{r\beta - \rho - \frac{1}{2}\beta \left[\frac{(1-\delta)^2}{(\beta-1)P^2s^2r} \right]}{(\beta-1)} \Psi \quad (2.8)$$

and

$$m^* = \frac{(1-\delta)(1+\alpha-s^2)\Psi}{(\beta-1)Ps^2}. \quad (2.9)$$

□

The proof of Proposition 2 can be found in the appendix. Note that $\frac{\partial m^*}{\partial \delta} < 0$ and $\frac{\partial m^*}{\partial P} > 0$, which makes intuitive sense. The fund's demand for IRS falls when the intermediation costs (negative swap spreads) are higher. As interest rates go down, P increases, and this leads to a higher demand for IRS.

¹⁸We require that $r\beta - \rho - \frac{1}{2}\beta \left[\frac{(1-\delta)^2}{(\beta-1)P^2s^2r} \right] > 0$. In addition, the intermediation costs represented by δ cannot be too high, i.e., $\delta < 1$.

When the underfunding is high, the fund uses more IRS and funds more aggressively: this is in fact the basic implication of our model. As β increases, the pension fund increases y much more and reduces its positions in IRS: this suggests that funding requirements imposed by regulators may have beneficial implications for the way in which pension assets are managed by the sponsors.

Model with Stocks, Swaps and Safe Assets

We now extend our model to allow the pension fund to additionally invest in a risky asset. To keep the model simple, we introduce a generic risky asset which can be interpreted as stock portfolio. The price of the stock portfolio follows a geometric Brownian motion:

$$dS = S\mu dt + S\sigma dw_2. \quad (2.10)$$

We allow the processes $\{w_1, w_2\}$ to be correlated with each other with correlation coefficient R and introduce the notation $\sigma_{12} := s\sigma R$. The fund invests in n shares of the stock portfolio, m swaps and places the remainder in risk-free asset.

Proposition 3. *The sponsor's optimal contribution and optimal asset allocation are given as:*

$$y = g^{\frac{1}{\beta-1}} \Psi, \quad (2.11)$$

$$nS = \frac{\left[\left(\frac{\mu-r}{\sigma^2} \right) + P \frac{\sigma_{12}}{\sigma^2} \sqrt{r} \left(\frac{1-\delta}{P^2 s^2 r} \right) \right]}{(\beta-1) \left(1 - \frac{\sigma_{12}^2}{s^2 \sigma^2} \right)} \Psi \equiv \lambda_1 \Psi, \quad (2.12)$$

and

$$m = \frac{\left[\left(\frac{1-\delta}{P^2 s^2 r} \right) + \frac{\sigma_{12}}{P s^2 \sqrt{r}} \left(\frac{\mu-r}{\sigma^2} \right) \right]}{(\beta-1) \left(1 - \frac{\sigma_{12}^2}{s^2 \sigma^2} \right)} \Psi \equiv \lambda_2 \Psi, \quad (2.13)$$

where g is given as:

$$g = \left[\frac{1}{1-\beta} \left((\rho + \beta(\lambda_1(\mu-r) + \lambda_2(1-\delta) - r)) - \frac{1}{2} \beta(\beta-1) (\lambda_1^2 \sigma^2 + \lambda_2^2 P^2 s^2 r - 2\lambda_1 \lambda_2 P \sqrt{r} \sigma_{12}) \right) \right]^{\beta-1}. \quad (2.14)$$

□

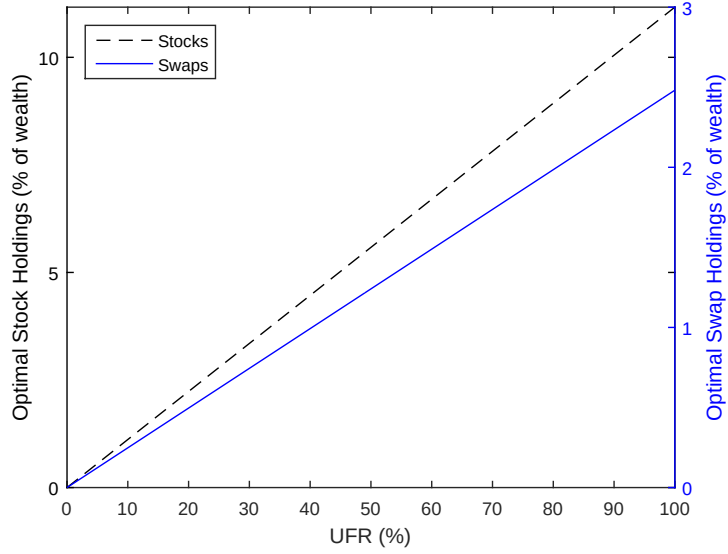


Figure 2.3: Pension funds' optimal holdings of stocks and swaps. UFR is computed as $(L/r - A)/A$. Parameter choices are: $\beta = 10$, $\rho = 0.05$, $\mu = 0.06$, $\sigma = 0.2$, $\alpha = 0.1$, $s = 0.3$, $R = 0$, $P = 50$, and $\delta = 50 \cdot 10^{-4}$.

The proof of Proposition 3 can be found in the appendix. We now illustrate the pension fund's optimal holdings of stocks and swaps in a numerical example. We define the pension fund's underfunded ratio as: $UFR = \frac{L/r - A}{A}$ and choose the following parameters for our illustrations: $\beta = 10$, $\rho = 0.05$, $\mu = 0.06$, $\sigma = 0.2$, $\alpha = 0.1$, $s = 0.3$, $R = 0$, $P = 50$, and $\delta = 50 \cdot 10^{-4}$, corresponding to a negative swap spread of 50 basis points.

As we can see from Figure 2.3, both, risky asset holdings and the amount of swaps held increase with UFR. For $UFR \downarrow 0$ the fund closes out his risky asset and swap holdings and pays off any loans. It is important to note that the pension fund increase both the exposure to swaps and risky assets as the fund gets more underfunded: the increase in risky asset is due to a desire to get out of the underfunded status in the future. On the other hand, the increase in swap position is to manage the duration risk to prevent future losses arising from interest rate changes.

2.4.2 Supply of Long-Dated Swaps by Dealer-Brokers

On the supply side of long-dated IRS, we assume that derivatives dealers decide on the amount $\mathcal{S}$ that they supply, by maximizing their discounted stream of profits. The supply is per unit time and we assume that there is a base amount $\mathcal{S}_0$ up to which the dealers supply fixed payments at the arbitrage-free rate. In this situation, the dealer pays \$1 per unit time and hedges himself by financing a perpetuity with time- t price P_t . The cost of financing

this position is $P_t r_s$ for $s \geq t$, which corresponds to the cashflows received from the pension fund. Hence, this position has a value equal to zero. In this case, the swap spread is simply equal to zero and the equilibrium swap rate is given as the yield of the perpetuity.

If the demand for receiving fixed exceeds $\mathcal{S}_0$, the dealers are not able to provide swaps at the frictionless rate anymore. The intuition behind this assumption is that dealers are facing balance sheet constraints. In our model, they dedicate a certain part of their balance sheet to swap trading. If they decide on supplying more swaps than the dedicated amount they face the risk of future costs due to binding balance sheet constraints. For instance, a dealer supplying a large amount of IRS might either not be able to take advantage of an attractive arising investment opportunity due to binding balance sheet constraints or would need to face costly unwinding of his swap positions. We model this risk using a random variable $\xi \sim \mathcal{N}(0, \sigma_\xi^2)$. Assuming that dealers have CARA utility, with risk-aversion λ , the cost of supplying an amount $\mathcal{S} > \mathcal{S}_0$ of IRS is given as $\frac{\lambda}{2} ([\mathcal{S} - \mathcal{S}_0]P)^2 \sigma_\xi^2$. To compensate for this risk, dealers charge a fee of δ on the present value of the stream of fixed payments per unit time, thereby earning a profit of $\delta(\mathcal{S} - \mathcal{S}_0)P$. Hence, the dealer is maximizing

$$\delta(\mathcal{S} - \mathcal{S}_0)P - \frac{\lambda}{2} ([\mathcal{S} - \mathcal{S}_0]P)^2 \sigma_\xi^2.$$

Taking first-order conditions leads to the following supply of IRS:

$$\mathcal{S}^* = \mathcal{S}_0 + \frac{\delta}{\lambda \sigma_\xi^2 P}. \quad (2.15)$$

We provide the detailed derivation of the supply function in the appendix.

2.4.3 Equilibrium Swap Spreads: Numerical Example

Equating Equations (2.13) and (2.15) gives the equilibrium swap spread. To illustrate the effect of pension fund's underfunding on the equilibrium swap spread, we continue the numerical example from Section 2.4.1. For the supply, we choose the following parameters: $\mathcal{S}_0 = 0.05$, $\sigma_\xi = 0.2$, and $\lambda = 0.2$. Furthermore, we choose three different values for UFR, 15%, 25%, and 35%. The equilibrium swap rate in the three different UFR regimes is illustrated in Figure 2.4.

As we can see from the figure, our model supports the intuition that higher underfunding of pension funds leads to more demand for IRS which decreases swap spreads and eventually pushes them into negative territory. Note that the pension funds' demand for IRS is almost unaffected by changes in the swap spread.

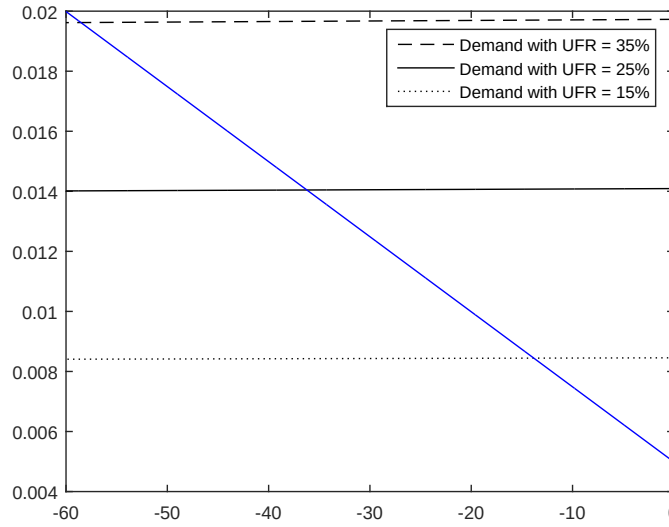


Figure 2.4: Supply and demand of IRS. This graph illustrates the demand for IRS by pension funds for different UFR and the supply of IRS by derivatives dealers. UFR is computed as $(L/r - A)/A$. Parameter choices are: $\beta = 10$, $\rho = 0.05$, $\mu = 0.06$, $\sigma = 0.2$, $\alpha = 0.1$, $s = 0.3$, $R = 0$, $P = 50$, $\sigma_\xi = 0.2$, $\lambda = 0.2$, and $S_0 = 0.005$, which corresponds to a frictionless swap spread of zero as long as $UFR < 10\%$. The amount of IRS traded is given under the assumption that pension fund's total assets are normalized to 1.

2.5 Empirical Analysis

In this section, we first describe our approach to measuring pension fund underfunding and constructing an aggregate measure for the underfunded ratio (UFR) of US pension funds. We then show that 30-year swap spreads differ in different funding regimes, using a Kolmogorov-Smirnov test. Subsequently, we run OLS regressions to test the relationship between the 30-year swap spread and UFR . We conclude this section with addressing the possible endogeneity concern that the level of the yield curve can drive both the swap spread and UFR . To account for this, we run a 2-stage least squares regression, where we use stock returns as an instrument.

2.5.1 Measuring Pension Fund Underfunding

To test our hypotheses, we first construct a measure of pension fund underfunding. We obtain quarterly data on two types of defined benefit (DB) pension plans, private as well as public local government pension plans, from the financial accounts of the US (former flow of funds) tables L.118b and L.120b. We exclude defined contribution pension plans since they cannot become underfunded and also exclude public federal DB pension plans since they are only allowed to invest in government bonds. We first note that the overall

size of the pension funds' balance sheet is 8,235 billion US dollar (as of Q3 2015), thereby capturing approximately 45% of the total assets held by all US pension funds. Furthermore, comparing the size of the pension funds' balance sheet to the size of the US dealer-brokers' balance sheet shows that it is approximately 2.5 times as large.

Table 2.2: Aggregate pension fund balance sheet as of Q3 2015. This table presents the assets and liabilities of private as well as state and local government employee defined-benefit (DB) pension plans. The amounts are in billions of dollars, not seasonally adjusted. (Source: Financial accounts of the United States)

DB Pension Fund Assets (billions)		DB Pension Fund Liabilities	
checkable deposits and currency	\$14.42	pension entitlements	\$8,430.73
total time and savings deposits	\$75.67		
money market mutual fund shares	\$85.45		
security repurchase agreements	\$7.11		
debt securities	\$1,744.41		
<i>commercial paper</i>	\$82.26		
<i>Treasury securities</i>	\$365.19		
<i>agency- and GSE-backed securities</i>	\$191.15		
<i>corporate and foreign bonds</i>	\$1,101.75		
<i>municipal securities</i>	\$4.05		
total mortgages	\$20.00		
corporate equities	\$3,141.21		
mutual fund shares	\$613.29		
miscellaneous assets	\$2,533.58		
<i>unallocated insurance contracts</i>	\$58.02		
<i>pension fund contributions receivable</i>	\$47.28		
<i>claims of pension fund on sponsor</i>	\$2,044.53		
<i>unidentified miscellaneous assets</i>	\$383.76		
TOTAL	\$8,235.12	TOTAL	\$8,430.73

Table 2.2 shows the aggregate pension fund balance sheet for the third quarter of 2015. As we can see from the table, the liabilities of these pension funds consist only of pension entitlements. On the asset side, there are three major positions. First, corporate equities, which make up more than one third of the balance sheet. Second, claims of pension fund on sponsor, which account for almost one quarter of the pension funds assets. As we describe below, these claims on sponsor are our main proxy for underfunding.¹⁹ Third,

¹⁹The Financial Accounts report assets and liabilities (and corresponding financial flows) for both private and public DB pension funds. Prior to September 2013, the assets and liabilities of DB pension plans were reported using cash accounting principles, which record the revenues of pension funds when cash is received and expenses when cash is paid out. Under this treatment, there was no measure of a plan's accrued actuarial liabilities. Rather, the liabilities in the Financial Accounts were set equal to the plans' assets. As a result, the Financial Accounts did not report any measure of underfunding or overfunding of the pension sector's actuarial liabilities, as would occur if the assets held by the pension sector fell short of or exceeded the liabilities. Starting with the September 2013 release, the Financial Accounts treat DB pensions using accrual accounting principles, whereby the liabilities of DB pension plans are set equal to the present value of future DB benefits that participants have accumulated to date, which are calculated using standard actuarial methods. This new measure is retroactively made available. Throughout, we use the accrual measures of the claims of pension funds on sponsors.

debt securities, which consist mainly of fixed-rate securities, like corporate bonds.

We use claims of pension funds on sponsors as our measure of pension funds' underfunding ratio (UFR). UFR in quarter t is computed as:

$$UFR_t = \frac{\text{Private DB claims on Sponsor}_t + \text{Public DB claims on Sponsor}_t}{\text{Private DB total financial assets}_t + \text{Public DB total financial assets}_t}. \quad (2.16)$$

The claims of pension fund on sponsors represents the difference between actuarial liabilities and pension fund assets. It reflects the amount of underfunding or overfunding of the plans. These claims (which can be positive or negative) are treated as an asset of the pension funds sector and a liability of the sponsors of the plans.²⁰ If claims of pension fund on sponsor is positive, pension funds are underfunded. Since our hypothesis is that UFR has a more significant impact on swap spreads if it is negative, we introduce the notation $UFR_t^+ := \max(UFR_t, 0)$ and $UFR_t^- := \min(UFR_t, 0)$ for the positive and negative part of the underfunding ratio respectively. Since we are using changes in UFR in our regression analysis, we also introduce the notation $\Delta UFR_t^+ := (UFR_t - UFR_{t-1})\mathbb{1}_{\{UFR_t > 0\}}$ and $\Delta UFR_t^- := (UFR_t - UFR_{t-1})\mathbb{1}_{\{UFR_t \leq 0\}}$. Note that the way we define ΔUFR_t^+ means that the measure includes a change from fully funded to underfunded periods but not from underfunded to fully-funded periods (this change is included in ΔUFR_t^-).

2.5.2 Swap Spreads in Different Underfunding Regimes

It should be noted for the end of quarter t , the Fed's flow of funds report the pension sponsors' funding status resulting from events during the end-quarter $t - 1$ to end-quarter t . This is reported roughly 2 weeks after end of quarter t . The swap spreads that we use in the paper are calculated precisely at the end of quarter t . In this sense our measure of funding status for quarter end t , UFR_t , which is based on the information from end-quarter $t - 1$ to end-quarter t is effectively a lagged measure relative to the time at which the swap spreads are collected.

Using the measure constructed above, we provide some preliminary evidence on the proposition that the demand by a significant subset of pension sponsors to receive fixed in long-term swaps has an effect on long-term swap spreads.

The top panel of figure 2.5 shows a scatter plot of the 30-year swap spreads in basis points against our measure of aggregate funding status, UFR_t , and gives a first overview of the results. The time period is between Q2 1994 and Q3 2015. The swap spreads are quarter-end observations and we are distinguishing between the negative part (solid dots) and positive part (circles) of the UFR . [It is worth recalling that when $UFR > 0$, there is

²⁰See Stefanescu and Vidangos (2014) for further details.

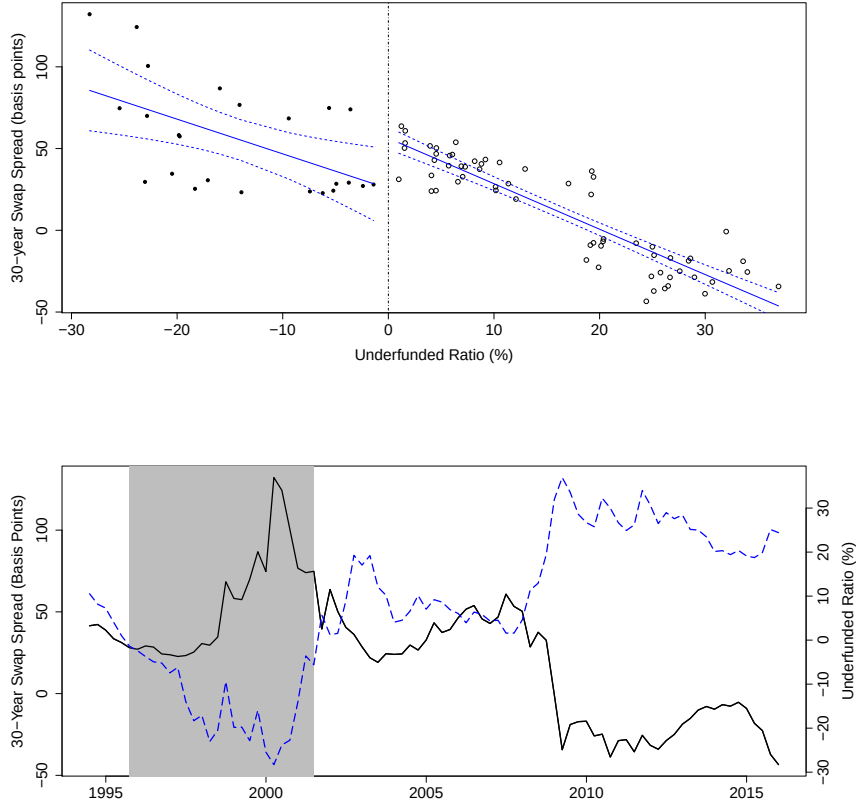


Figure 2.5: Relationship between 30-year swap spreads and the aggregate funding status of DB pension plans. The lower panel shows the time series of the two variables, wherein the black solid line is the 30-year swap spread (left-hand axis) and the blue line with dots is pension funds' underfunded ratio (right-hand side). The grey shaded areas indicate periods where pension funds are fully funded or over-funded. Data on pension fund underfunding ratios are obtained from the financial accounts of the U.S. and the underfunding ratio is computed as indicated in Equation (2.16).

underfunding]. The dashed lines indicate 95% confidence intervals. As we can see from the figure, the level of the swap spread is negatively related to the UFR for both funded and underfunded regions. However, in line with our theory, the dots are less scattered around the solid line if the UFR is positive, indicating a stronger correlation when pension funds are underfunded. As we can also see from the top panel, the intercepts of the two lines differ. Unfortunately, there are not enough data points around this cutoff to provide a powerful empirical analysis of the relationship at the cusp where the funds are just fully funded. Instead, we test below whether the distribution of swap spreads is different when pension funds are underfunded when compared to regimes in which they are fully funded. The lower panel of Figure 2.5 shows the time series plot of the same variables, illustrating that both

variables are relatively volatile without an obvious trend component. The grey shaded areas indicate periods where pension funds are fully funded or over-funded. The U.S. Economy was generating a surplus during the end this (shaded) period, with a drop in the supply of long-term government bonds, which might have partially accounted for the increase in swap spreads. The stock market boom during this period could have partially accounted for the over-funded status of the pension plans.

The evidence presented in this section helps to motivate why the funding status of pension plans, as suggested by our theory, may be a channel that could be at work in driving the swap spreads down to negative levels. We next use regression analysis to further explore this channel.

2.5.3 Regression Analysis

To shed additional light on the relationship between UFR and swap spreads we next run a regression analysis of changes in 30-year swap spreads on changes in UFR .²¹ Motivated by the no-arbitrage argument in Table 2.8, we control for the change in the difference between the 3-months Libor rate and 3-month general collateral repo rate ($\Delta LR\ spread_t$) in all regression specifications. Panels (1) and (2) of Table 2.3 show that, without additional control variables, pension fund underfunding is a significant explanatory variable for 30-year swap spreads.

Panel (1) shows that UFR for the entire sample period is statistically significant at a 1% level with a coefficient of -1.09 (t -statistic of -2.99). More importantly and in line with our theory, panel (2) shows that UFR is even more significant when only considering underfunded regimes and insignificant when pension funds are fully funded. For underfunded periods, UFR is statistically significant at a 1% level with a coefficient of -1.32 (t -statistic of -3.49) and for funded periods UFR is insignificant with a coefficient of -0.57 (t -statistic of -0.83). Note that a coefficient of -1 indicates that swap spreads fall by one basis point when pension fund underfunding increases by 1%.

We next check whether our results are robust to controlling for other factors that are likely to affect swap spreads. We start by adding four control variables, the US debt-to-GDP ratio as a proxy for the “convenience yield” of US Treasuries (Krishnamurthy and Vissing-Jorgensen, 2012a), the average Moody’s expected default frequency (EDF) of the 14 largest derivatives-dealing banks, the implied volatility in US Treasuries as proxied by the Move index, and a term factor, measuring the slope of the yield curve. These variables (as well

²¹We use changes in these variables since both are highly serially correlated. A regression of the level of the 30-year swap spread (level of UFR) on the lagged level of the 30-year swap spread (UFR) gives a highly significant coefficient of 0.95 (0.97).

Table 2.3: 30-year swap spreads and pension fund underfunding. This table reports results from regressions of quarterly changes in the 30-year swap spread on the indicated variables. ΔUFR_t is the change in the underfunding ratio of private and local government defined benefit pension funds, as defined in Equation (2.16), ΔUFR_t^+ (ΔUFR_t^-) is the change in UFR_t if $UFR_t > 0$ ($UFR_t \leq 0$) and zero otherwise. $\Delta LR Sprd_t$ is the change in the quarter-end difference between the 3-month Libor rate and 3-month General Collateral repo rate, $\Delta Debt/GDP_t$ is the change in the ratio of US public debt to GDP, ΔEDF_t is the change in the Moody's expected default frequency of the 14 largest derivatives-dealing banks (G14 banks), $\Delta Move_t$ is the change in the 1-month implied volatility of US Treasuries with 2, 5, 10, and 30 years to maturity, $\Delta TERM_t$ measures changes in the slope of the yield curve, approximated as the difference between 30-year and 3-month treasury yields. In panels (5) and (6) five additional controls are added: changes in the level of the 30-year treasury yield, changes in the mortgage refinancing rate, obtained from the mortgage bankers association, the broker-dealer leverage factor provided by Adrian, Etula, and Muir (2014), changes in VIX, and changes in the 10-year on-the-run-off-the-run spread. All variables are quarter-end. The numbers in parenthesis are heteroskedasticity-robust t -statistics. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively. The observation period is Q3 1994 – Q4 2015 with 5 missing observations between Q4 1997 and Q4 1998 due to missing repo rates.

	(1)	(2)	(3)	(4)	(5)	(6)
Intercept	-1.13 (-0.90)	-1.05 (-0.85)	-0.59 (-0.42)	-0.53 (-0.37)	-0.48 (-0.36)	-0.38 (-0.29)
ΔUFR_t	-1.09*** (-2.99)		-0.96** (-2.20)		-0.76 (-1.56)	
ΔUFR_t^+		-1.32*** (-3.49)		-1.27*** (-3.17)		-1.25*** (-2.68)
ΔUFR_t^-		-0.57 (-0.83)		-0.48 (-0.60)		-0.35 (-0.43)
$\Delta LR Spread_t$	0.06 (1.16)	0.06 (1.30)	0.01 (0.22)	0.01 (0.26)	0.01 (0.19)	0.02 (0.21)
$\Delta Debt/GDP_t$			-1.02 (-1.31)	-0.95 (-1.22)	-0.78 (-1.07)	-0.73 (-1.01)
ΔEDF_t			-0.05** (-2.12)	-0.04 (-1.65)	-0.06*** (-2.87)	-0.05** (-2.37)
$\Delta Move_t$			0.17** (2.45)	0.17** (2.49)	0.19** (2.23)	0.19** (2.09)
$\Delta TERM_t$			-0.05* (-1.76)	-0.06** (-2.01)	-0.08* (-1.96)	-0.08* (-1.99)
Add. Controls?	No	No	No	No	Yes	Yes
Observations	81	81	81	81	81	81
Adjusted R ²	0.15	0.15	0.27	0.28	0.26	0.27

as all other data used in our analysis) are described in more detail in Appendix 2.7.4. The results of these regressions are reported in Panels (3) and (4) of Table 2.3.

As we can see from the table, Debt-to-GDP is insignificant but with the expected sign: An increase in Debt-to-GDP lowers the convenience yield of treasuries, thereby lowering the swap spread. ΔEDF_t is statistically significant and an increase in derivatives dealers' expected default frequency lowers the swap spread, indicating that, as dealers become more constrained, swap spreads decrease. $\Delta Move_t$ and $\Delta TERM_t$ are both significant and an increase in uncertainty, as captured by $Move$, increases the swap spread. Most importantly, as we can see from Panels (3) and (4) of Table 2.3, controlling for these variables leads to a small drop in the statistical and economic significance of UFR , but leaves our main result unchanged. Panel (3) shows that UFR for the full sample period is still significant at a 5% level with a coefficient of -0.96 (t -statistic of -2.20). More importantly, Panel (4) shows that UFR during times of underfunding is still statistically significant at a 1% level with a coefficient of -1.27 (t -statistic of -3.17).

In panels (5) and (6) we add five additional controls to check whether our results remain robust to including more potential drivers of swap spreads. These five controls are the level of the 30-year treasury yield, the mortgage refinancing rate, the broker-dealer leverage factor by Adrian et al. (2014), the VIX index, and the 10-year on-the-run off-the-run spread. We do not report the coefficient estimates for these variables for brevity and note that none of these additional variables is statistically significant. Moreover, as we can see by comparing the R^2 values from panels (3) and (4) with the R^2 values from panels (5) and (6), adding these additional controls does not improve the explanatory power of our analysis. Furthermore, adding these controls leads to a minor drop in statistical and economical significance of $UFR +$.

As a next step, we check whether UFR is a significant explanatory variable for swap spreads with shorter maturities. To that end, we regress changes in the 2-year, 5-year, 10-year, and 30-year swap spread on the positive and negative part of changes in UFR , controlling for changes in the Libor-repo spread.²² The results of this regression are exhibited in Table 2.4.

In line with our theory, UFR^+ is only significant for the 30-year swap spread and insignificant for swap spreads with shorter maturities. We note that, in line with the no-arbitrage argument from Table 2.8, the Libor-repo spread is a significant explanatory variable for swap spreads with shorter maturities (2-year and 5-year).

²²We focus on 2, 5, 10, and 30-year swap spreads since there are no missing observations for these data and we do not need to supplement them with data from the FED H.15 reports.

Table 2.4: Pension fund underfunding and swap spreads with different maturities. This table reports results from regressions of quarterly changes in swap spread with 2, 5, 10, and 30 years to maturity on the indicated variables. ΔUFR_t^+ (ΔUFR_t^-) is the change in the underfunding ratio of private and local government defined benefit pension funds as defined in Equation (2.16), conditional on pension funds being underfunded (funded) at time t . $\Delta LR Sprd_t$ is the change in the quarter-end difference between the 3-month Libor rate and 3-month General Collateral repo rate. The numbers in parenthesis are heteroskedasticity-robust t -statistics.***, **, and * indicate significance at a 1%, 5%, and 10% level respectively. The observation period is Q3 1994 – Q4 2015 with 5 missing observations between Q4 1997 and Q4 1998 due to missing repo rates.

	2 Year	5 Year	10 Year	30 Year
Intercept	0.01 (0.01)	−0.46 (−0.47)	−0.73 (−0.66)	−1.05 (−0.85)
ΔUFR_t^+	−0.05 (−0.12)	0.04 (0.10)	−0.31 (−0.81)	−1.32*** (−3.49)
ΔUFR_t^-	0.07 (0.17)	0.04 (0.11)	−0.15 (−0.23)	−0.57 (−0.83)
$\Delta LR Spread_t$	0.35*** (5.42)	0.19*** (3.15)	0.06 (1.26)	0.06 (1.30)
Observations	81	81	81	81
Adjusted R ²	0.56	0.25	0.01	0.15

2.5.4 Two-Stage Least Squares Regression Results

One concern about using UFR as an explanatory variable for swap spreads is that the same factors that similar factors can affect both variables. For example, a decrease in the level of the yield curve can affect swap spreads and also increases the level of pension funds' underfunding, which increase because the present value of the funds' liabilities is computed using long-term interest rates. To mitigate these endogeneity concerns, we next run a 2-stage least squares regression. In a first stage, we regress ΔUFR_t on US stock returns proxied by the excess return on the CRSP value-weighted portfolio. In panels (2), (4), and (6) we drop fully funded periods and only regress ΔUFR_t^+ on stock returns. Stock returns affect UFR since pension funds are heavily invested in corporate equity (almost half their assets are invested in corporate equity according to Table 2.2) and therefore decreasing stock returns increase UFR . At the same time, there is no obvious link between stock returns and swap spreads.

At the same time, there is no obvious connection between the 30-year swap spread and stock returns. We therefore argue that the exclusion restriction is fulfilled. Furthermore, the results from a weak instrument test give a p-value far below 0.1% for all six regression specifications. Additionally to that, the results from a Hausman test give a p-value above

Table 2.5: 30-year swap spreads and pension fund underfunding (2-stage least squares).

This table reports results from a second stage regressions of quarterly changes in the 30-year swap spread on the indicated variables. In the first stage, the change in the underfunding ratio of private and local government defined benefit pension funds, ΔUFR_t (ΔUFR_t^+) is regressed on the excess returns of the CRSP value-weighted portfolio, controlling for the other indicated variables. $\Delta LR\ Sprd_t$ is the change in the quarter-end difference between the 3-month Libor rate and 3-month General Collateral repo rate, $\Delta Debt/GDP_t$ is the change in the ratio of US public debt to GDP, ΔEDF_t is the change in the Moody's expected default frequency of the 14 largest derivatives-dealing banks (G14 banks), $\Delta Move_t$ is the change in the 1-month implied volatility of US Treasuries with 2,5, 10, and 30 years to maturity, $\Delta TERM_t$ measures changes in the slope of the yield curve, approximated as the difference between 30-year and 3-months treasury yields. In panels (5) and (6) five additional controls are added: changes in the level of the 30-year treasury yield, changes in the mortgage refinancing rate, obtained from the mortgage bankers association, the broker-dealer leverage factor provided by Adrian et al. (2014), changes in VIX, and changes in the 10-year on-the-run-off-the-run spread. All variables are quarter-end. The numbers in parenthesis are small-sample and heteroskedasticity-robust t -statistics. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively. The observation period is Q3 1994 – Q4 2015 with 5 missing observations between Q4 1997 and Q4 1998 due to missing repo rates.

	(1)	(2)	(3)	(4)	(5)	(6)
Intercept	−1.14 (−0.91)	−1.61 (−1.45)	−0.60 (−0.42)	−1.18 (−1.00)	−0.48 (−0.36)	−1.00 (−0.85)
ΔUFR_t	−1.05*** (−3.02)		−0.93** (−2.29)		−0.75 (−1.62)	
ΔUFR_t^+		−1.28*** (−3.31)		−1.13*** (−2.75)		−1.32*** (−2.70)
$\Delta LR\ Spread_t$	0.06 (1.16)	0.08* (1.87)	0.01 (0.23)	0.03 (0.79)	0.01 (0.20)	0.02 (0.40)
$\Delta Debt/GDP_t$			−1.03 (−1.30)	−0.72 (−1.11)	−0.78 (−1.07)	0.15 (0.22)
ΔEDF_t			−0.05** (−2.37)	−0.05** (−2.10)	−0.06*** (−3.08)	−0.06** (−2.56)
$\Delta Move_t$			0.16** (2.49)	0.14** (2.39)	0.19** (2.21)	0.16*** (2.80)
$\Delta TERM_t$			−0.05* (−1.76)	−0.05** (−2.12)	−0.08* (−1.88)	−0.09** (−2.55)
Add. Controls?	No	No	No	No	Yes	Yes
Observations	81	62	81	62	81	62

0.6 (ranging from 0.619 for specification (1) to 0.934 for specification (6)) for all six specifications. Hence, we can reject the hypothesis that stock returns are a weak instrument and we cannot reject that the 2 SLS is as consistent as the OLS regression.

Table 2.5 shows the results of the second stage, where we use the projected UFR as explanatory variable. Overall the results from the second stage are similar to those from the OLS regression discussed before. The projected UFR is significant at a 1% level and decreases in significance as we add controls. More importantly, the projected underfunded ratio in regimes when pension funds are underfunded is even more significant (t -statistic of -3.31 without controls) and remains significant even after adding several controls (t -statistic of -2.71).

2.5.5 International Evidence

Because neither negative swap spreads nor large-scale duration hedging by pension funds are purely a U.S. phenomenon, we next investigate the effects of pension funds' underfunding in different regions. According to a recent study by the OECD, the world's five largest autonomous pension fund systems are located in the U.S., the UK, Australia, Japan, and the Netherlands, with all five systems managing more than one trillion U.S. dollars of investments (see (OECD, 2016)). Therefore, we next study the effects of pension fund underfunding in two of these regions: Japan and the Netherlands.

The Data

We obtain Japanese government bond yields and swap rates from the Bloomberg system. In contrast to the U.S., where the floating rate is paid quarterly, in Japan, a fixed rate is exchanged against a semi-annual floating payment. We then obtain private DB pension funds' claims on their sponsor as well as private DB pension funds' total financial assets from Japan's flow of funds accounts. We exclude public pension funds because, for this subcategory, the flow of funds accounts do not provide a split between DB and DC funds. Hence our measure of underfunding for Japan is constructed as:

$$UFR_t^{Jap} = \frac{\text{Private DB claims on sponsor}_t}{\text{Private DB total financial assets}_t}. \quad (2.17)$$

Quarterly data on the funding status of DB pension funds are available from Q1 2005. Panel A of Table 2.6 provides summary statistics for UFR_t^{Jap} as well as 30-year swap spreads. As we can see from the table, Japanese pension funds have been underfunded during the entire sample period. Moreover, the maximum level of UFR_t^{Jap} exceeds the maximum level of UFR_t in the U.S. by almost 10%. The higher underfunded ratio of Japanese pension funds

relative to the U.S. is not surprising, given that Japanese pension funds have been dealing with decreasing interest rates and falling stock prices for much longer than U.S. pension funds. Similarly to the U.S., Japanese pension funds try to avoid forcing their sponsors to cover losses and the usage of swaps is explicitly permitted for these funds.

Table 2.6: Summary Statistics for International Data. This table shows summary statistics of pension fund underfunding and 30-year swap spreads for Japan and the Netherlands. UFR_t^{Jap} is constructed based on Equation (2.17). 30-yr SS (Jap) is the difference between the fixed rate in a 30-year IRS where the fixed rate is exchanged against 6-month Japanese LIBOR rates and the bond yield of the most recently issued Japanese government bond with 30-years to maturity. UFR_t^{Neth} is constructed based on Equation (2.18). 30-yr SS (Ger) and 30-yr SS (Neth) are the difference between the fixed rate in a 30-year IRS where the fixed rate is exchanged against annual EURIBOR rates and the bond yield of the most recently issued German or Dutch government bond with 30-years to maturity, respectively. The sample period in Panel A is Q1 2005 – Q4 2015. The sample period in Panel B is Q1 2007 – Q4 2014. # Under counts the number of quarters where pension funds are underfunded.

	Mean	SD	Min	Median	Max	# Obs	# Under
Panel A: Summary statistics for Japan							
UFR_t^{Jap}	28.88	7.25	17.38	28.04	41.17	44	44
30-yr SS (Jap)	-0.11	11.73	-24.94	-0.65	26.34	44	–
Panel B: Summary statistics for the Netherlands							
UFR_t^{Neth}	-8.83	11.81	-34.14	-7.38	8.99	32	12
30-yr SS (Ger)	0.16	12.31	-25.85	2.04	23.10	32	–
30-yr SS (Neth)	-11.83	15.32	-47.55	-14.42	16.00	32	–

When investigating the impact of pension funds’ underfunding on swap spreads for the Netherlands, we use swap spreads as the difference between the EURIBOR swap rate and the yield of German government bonds in our main analysis and use swap spreads relative to the yield of Dutch government bonds, as a robustness check. We obtain swap rates and government bond yields from the Bloomberg system. Data for the funding status of Dutch DB pension funds are available on the DNB website, which provides data for “Liquid assets at funds’ risk” and “Estimated technical provision at funds’ risk” from Q1 2007 on. According to the Dutch pension fund regulation, a pension fund is underfunded if the ratio between the two variables drops below 105%. In that case, a plan needs to provide a proposal of how to become fully funded in the future to the Dutch supervisory authority and needs to lower the overall risk of its portfolio, which is mainly done by reducing interest rate risk.²³ Based on these arguments, we first estimate the funding gap of Dutch pension funds as the difference between 1.05 times the estimated technical provision at funds’ risk and liquid

²³ “One of the major risks that Dutch pension funds run is interest rate risk and hence their reduced ability to take risk could on the margin increase receiving pressure [...] from the Dutch pension fund community” (Deutsche Bank markets research, Dutch UFR curve adjustment)

assets at funds' risk. We then construct UFR_t^{Neth} as follows:

$$UFR_t^{Neth} = \frac{\text{Funding gap}_t}{\text{Liquid assets at funds' risk}_t}. \quad (2.18)$$

Finally, we split the measure into a positive part, which corresponds to times when pension funds are not underfunded and negative part that captures pension funds' underfunding. Panel B of Table 2.6 provides summary statistics for the Dutch UFR measure as well as 30-year swap spreads relative to German government bonds and relative to Dutch government bonds. As we can see from the table, Dutch pension funds are only rarely underfunded with a total of 12 underfunding observations.²⁴

Results

We next test the relationship between swap spreads and UFR_t for Japan and the Netherlands. To that end, we regress changes of 2-year, 5-year, 10-year, and 30-year swap spreads on the ΔUFR_t^+ and ΔUFR_t^- . In Japan, pension funds have been underfunded for the entire sample period and we therefore drop ΔUFR_t^- from the regression. Furthermore, we add the 6-month LIBOR-Repo spread as a control variable for Japan and do not control for changes in the LIBOR-Repo spread in Europe due to limited data availability.

As we can see from Panel A of Table 2.7, ΔUFR_t^+ is a significant explanatory variable for 10-year and 30-year Japanese swap spreads but not for swap spreads with shorter maturities. Both, the statistical and economic significance of UFR_t^+ are higher for 30-year swap spreads than for 10-year swap spreads. Similarly to the results for Japan, Panel B of Table 2.7 shows that, for the Netherlands, ΔUFR_t^+ is a significant explanatory variable for 30-year swap spreads, both relative to German and Dutch government bond yields and insignificant for swap spreads with shorter maturities.

²⁴We only include data up until Q4 2014 because from Q1 2015, the policy funding ratio is not based on the current ratio between assets and liabilities anymore but on the average funding ratio over the past year.

Table 2.7: Pension fund underfunding and swap spreads in other regions. This table reports results from regressions of quarterly changes in swap spread with 2, 5, 10, and 30 years to maturity on the indicated variables. In Panel A, the swap spreads are computed as the difference between the fixed rate in an IRS based on Japanese LIBOR rates and Japanese government bond yields. ΔUFR_t^+ is the change in the underfunding ratio of Japanese pension funds as defined in Equation (2.17), conditional on pension funds being underfunded at time t . There are no time periods where Japanese pension funds are fully funded. $\Delta LR Spread_t$ is the change in the quarter-end difference between the 6-month Japanese LIBOR rate and 6-month General Collateral repo rate. In Panel B, the swap spreads are computed as the difference between the fixed rate in an IRS based on EURIBOR and German government bond yields. Under 30 Year (Neth), the swap spread is computed relative to the Dutch government bond yield. ΔUFR_t^+ (ΔUFR_t^+) is the change in the underfunding ratio of Dutch pension funds as defined in Equation (2.18), conditional on pension funds being underfunded (funded) at time t . Pension funds are underfunded if the policy funding ratio drops below 105%. The numbers in parenthesis are heteroskedasticity-robust t-statistics. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively. The observation period is Q1 2005 – Q4 2015.

Panel A: Regression analysis for Japan					
	2 Year	5 Year	10 Year	30 Year	
Intercept	−0.05 (−0.08)	0.00 (0.00)	−0.20 (−0.28)	−1.06 (−0.84)	
ΔUFR_t^+	−0.28 (−1.18)	−0.28 (−1.02)	−1.00*** (−3.72)	−2.02*** (−4.45)	
$\Delta LR Spread_t$	0.23*** (3.04)	0.12 (1.01)	−0.05 (−0.39)	0.06 (0.38)	
Observations	43	43	43	43	
Adjusted R ²	0.11	0.00	0.32	0.34	
Panel B: Regression analysis for the Netherlands					
	2 Year	5 Year	10 Year	30 Year	30 Year (Neth)
Intercept	0.98 (0.30)	0.72 (0.23)	−0.29 (−0.14)	−0.90 (−0.59)	−0.52 (−0.25)
ΔUFR_t^+	1.92* (1.85)	1.65 (1.49)	0.89 (1.28)	−1.27*** (−4.29)	−1.11** (−2.31)
ΔUFR_t^-	0.25 (0.27)	0.57 (1.15)	0.76** (2.42)	−0.09 (−0.32)	−0.43 (−1.17)
Observations	31	31	31	31	31
Adjusted R ²	0.05	0.06	0.13	0.16	0.09

2.6 Conclusion

We provide a novel explanation of persistent negative 30-year swap spreads. Our explanation is based on the funding status of DB pension plans. Specifically, we argue that underfunded pension plans prefer to meet the duration needs arising from their unfunded pension liabilities through receiving fixed payments in 30-year interest rate swaps. This allows them to use their scarce funding to invest in risky assets with the hope of improving their future funding status. We present empirical evidence, which supports the view that the under-funded status of DB pension plans has a significant explanatory power for 30-year swap spreads, even after controlling for several other drivers of swap spreads, commonly used in the swap literature. Moreover, we show that the funding status does not have any explanatory power for swap spreads associated with shorter maturities.

2.7 Appendix

2.7.1 Characterizing the Term Structure

In this section, we illustrate several properties of the short rate, which we assume to follow a process suggested by Constantinides and Ingersoll (1984) with dynamics given by Equation (2.2). To illustrate the main properties of the process and get an understanding of the term structure of interest rates, we simulate the process using the following set of parameters:²⁵

$$\alpha \in \{0.1, 0.5, -0.5\}$$

$$s \in \{0.2, 0.6\}$$

$$r_0 = 0.02.$$

and characterize the term structure of interest rates under the process specified in Equation (2.2). Recall that the price of a zero coupon bond is given as:

$$p(0, T) = \mathbb{E} \left[e^{-\int_0^T r_s ds} \right]$$

and the zero coupon yield can be computed as:

$$y(0, T) = -\frac{\ln(p(0, T))}{T}.$$

We compute the zero-coupon yield using the simulated processes considering the following five parameter choices using $r_0 = 0.02$ in all specifications:²⁶ (i) the base case with $\alpha = 0.1$ and $s = 0.2$, (ii) the case where $s = 0.6$ is increased while $\alpha = 0.1$ is kept fixed, (iii) the case where $\alpha = 0.5$ is increased while the $s = 0.2$ is kept fixed, (iv) the case where both $\alpha = 0.5$ and $s = 0.6$ are increased, (v) and the case where $\alpha = -0.5$ is set to a negative negative number with $s = 0.6$.

Figure 2.6 shows the term structure of interest rates under the five parameter specifications. As we can see from the figure, the parameter α is key in characterizing the slope of the yield curve. For cases (i) and (ii) where we keep $\alpha = 0.1$ we see that the term structure of interest rates is upward sloping but almost flat, while for cases (iii) and (iv), where we increase $\alpha = 0.5$, the term structure becomes steeper. In case (v), where $\alpha = -0.5$ is set to a negative number, the term structure is downward-sloping. In contrast to α , the volatility

²⁵When simulating the process, we use 10 years of observations, 10,000 simulations and 1,000 time steps per year to discretize the sample period.

²⁶For each simulated sample path we approximate $\int_0^T r_s ds$ using the simulations and then compute the expected value of the exponential as average of all sample paths.

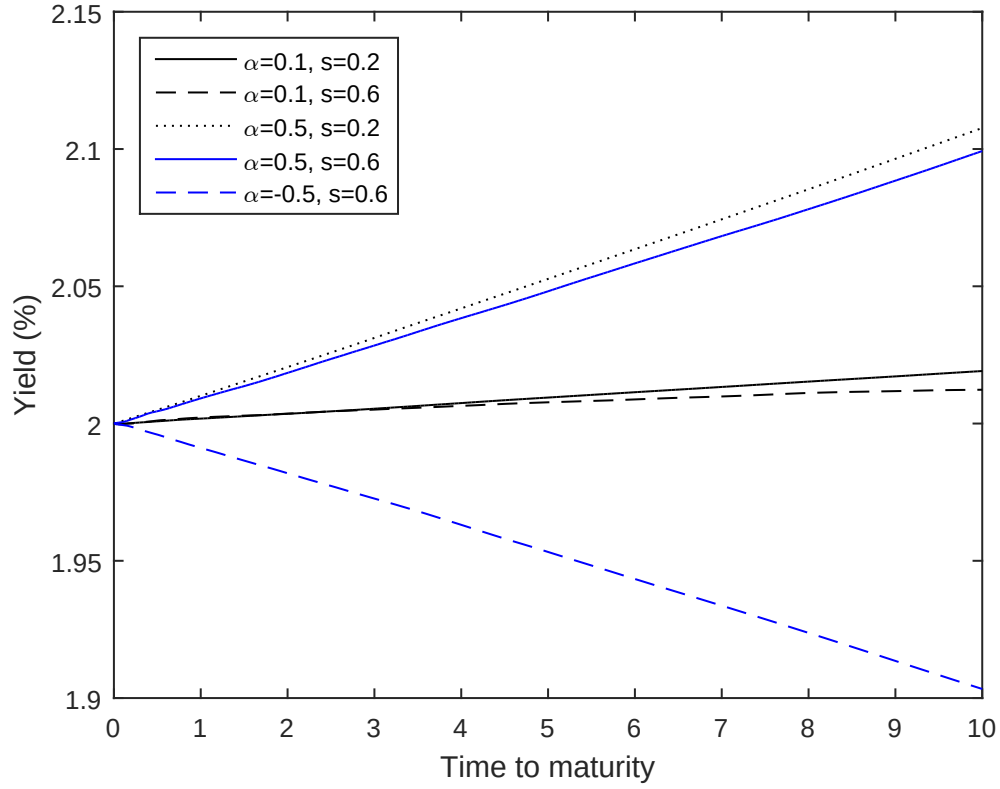


Figure 2.6: Term structure of interest rates for the short rate. The short rate follows a process suggested by Constantinides and Ingersoll as characterised by Equation (2.2) with the indicated parameters. The initial short rate is set to $r_0 = 2\%$. We use 10.000 sample paths and 1.000 time steps per year to discretize the time interval.

parameter has a minor effect on the term structure. Comparing cases (i) and (ii) as well as (iii) and (iv), we can see that an increase in s results in a flatter the term structure of interest rates.

2.7.2 Proofs and Additional Theoretical Results

Proof of Propositions 2 and 3

We first note that Proposition 2 is a special case of Proposition 3 where $\sigma_{12} = 0$ and $n = 0$. Hence, it remains to show that Proposition 3 holds. To show this result, we first state the pension fund's HJB in the general setting with stocks and swaps:

$$0 = \inf_{y,m,n} \left[y^\beta - \rho G + G_A (nS(\mu - r) + m(1 - \delta) + Ar + y - L) + \frac{1}{2} G_{AA} (n^2 S^2 \sigma^2 + m^2 P^2 s^2 r - 2mnSP\sqrt{r}\sigma_{12}) \right].$$

Taking derivatives with respect to y and m leads to the following first order conditions (FOC):

$$\begin{aligned} y &= \left(-\frac{1}{\beta} G_A \right)^{1/(\beta-1)} \\ nS &= -\frac{G_A}{G_{AA}} \frac{\mu - r}{\sigma^2} + m \frac{P\sqrt{r}\sigma_{12}}{\sigma^2} \\ m &= -\frac{G_A}{G_{AA}} \frac{1 - \delta}{P^2 s^2 r} + nS \end{aligned}$$

Plugging m into the expression for n and vice versa gives the optimal controls:

$$\begin{aligned} nS &= \frac{-\frac{G_A}{G_{AA}} \left(\frac{\mu - r}{\sigma^2} + \frac{(1 - \delta)\sigma_{12}}{s^2 P \sqrt{r} \sigma^2} \right)}{\left(1 - \frac{\sigma_{12}^2}{s^2 \sigma^2} \right)} \\ m &= \frac{-\frac{G_A}{G_{AA}} \left(\frac{1 - \delta}{P^2 s^2 r} + \frac{\mu - r}{\sigma^2} \frac{\sigma_{12}}{P s^2 \sqrt{r}} \right)}{\left(1 - \frac{\sigma_{12}^2}{s^2 \sigma^2} \right)}. \end{aligned}$$

We now guess and verify that the value function is of the following form:

$$G = g(P)(LP(1 + \alpha - s^2) - A)^\beta = g\Psi^\beta,$$

where we introduce the notation $\Psi := (LP(1 + \alpha - s^2) - A)$ and short hand $g := g(P)$ to simplify notations. Taking partial derivatives of the value function gives:

$$\begin{aligned}
G_A &= (-\beta g \Psi^{\beta-1}) \\
G_{AA} &= (\beta(\beta-1)g \Psi^{\beta-2}) \\
\frac{G_A}{G_{AA}} &= \frac{(-\beta g \Psi^{\beta-1})}{(\beta(\beta-1)g \Psi^{\beta-2})} = -\frac{\Psi}{\beta-1}.
\end{aligned}$$

Hence, under the guess, the optimal controls are given as:

$$\begin{aligned}
y &= \left(-\frac{1}{\beta} (-\beta g \Psi^{\beta-1}) \right)^{1/(\beta-1)} = g^{1/(\beta-1)} \Psi \\
nS &= \frac{\left(\frac{\mu-r}{\sigma^2} + \frac{(1-\delta)\sigma_{12}}{s^2 P \sqrt{r} \sigma^2} \right)}{(\beta-1) \left(1 - \frac{\sigma_{12}^2}{s^2 \sigma^2} \right)} \Psi =: \lambda_1 \Psi \\
m &= \frac{\left(\frac{1-\delta}{P^2 s^2 r} + \frac{\mu-r}{\sigma^2} \frac{\sigma_{12}}{P s^2 \sqrt{r}} \right)}{(\beta-1) \left(1 - \frac{\sigma_{12}^2}{s^2 \sigma^2} \right)} \Psi =: \lambda_2 \Psi,
\end{aligned}$$

where the last two Equations correspond to Equations (2.12) and (2.13). It remains to show that the function $g(P)$ stated in Equation (2.14) solves the HJB. To see this, we plug both, the guess and the optimal controls into the HJB:

$$\begin{aligned}
0 &= g^{\beta/(\beta-1)} \Psi^\beta - \rho g \Psi^\beta - \beta g \Psi^{\beta-1} (\lambda_1 \Psi (\mu-r) + \lambda_2 \Psi (1-\delta) - r(L/r - A) + g^{1/(\beta-1)} \Psi) \\
&= \left[(1-\beta) g^{1/(\beta-1)} - \rho - \beta \lambda_1 (\mu-r) - \beta \lambda_2 (1-\delta) + \beta r \right. \\
&\quad \left. + \frac{1}{2} \beta (\beta-1) (\lambda_1^2 \sigma^2 + \lambda_2^2 P^2 s^2 r - 2 \lambda_1 \lambda_2 P \sqrt{r} \sigma_{12}) \right] g \Psi^\beta
\end{aligned}$$

Dividing by $g \Psi^\beta$ and solving for g leads to Equation (2.14), which completes the proof. ■

The Supply Function

Proposition 4. *Assume that derivatives dealers can provide swaps at the frictionless rate up to a level S_0 . Afterwards, they face a random cost of $\xi \sim \mathcal{N}(0, \sigma_\xi^2)$ and are risk-averse with aversion coefficient λ towards that cost. Then, the Dealer's optimal supply of IRS is given by Equation (2.15).*

Proof of Proposition 4

Under risk-aversion λ , the dealer's expected profits of providing swaps $\mathcal{S} > S_0$ are given as:

$$\delta(\mathcal{S} - S_0)P - \frac{\lambda}{2} ([\mathcal{S} - S_0]P)^2 \sigma_\xi^2.$$

More formally, the dealer is facing the following HJB:

$$0 = \max_S \left(\delta(\mathcal{S} - S_0)P - \frac{\lambda}{2} \left([\mathcal{S} - S_0]P \right)^2 - \rho J + J_r \alpha r^2 + \frac{1}{2} J_{rr} s^2 r^3 \right).$$

Taking FOC then leads to the optimal strategy in Equation (2.15) which completes the proof. ■

2.7.3 What keeps Arbitrageurs Away?

Table 2.8: The arbitrage relationship between interest rate swaps and Treasuries.

This table provides an arbitrage argument for positive swap spreads. s_0 denotes the fixed rate in an interest rate swap with maturity T , l_t denotes the variable Libor rate in month t , c_0 denotes the coupon of a treasury bond with maturity T , and r_t denotes repo rate in month t . Since the difference between Libor and Repo rate is usually positive, the difference between swap rate and treasury yield should be positive too.

	$t = 0$	$t = 1$	...	$t = T$
Pay fixed rate s_0 in IRS	0	$-s_0$	...	$-s_0$
Receive Libor l_t from IRS	0	$+l_t$	...	$+l_T$
Buy bond with coupon c_0	-1	$+c_0$	...	$+1 + c_0$
Borrow at repo rate r_t	+1	$-r_t$	...	$-1 - r_T$
Payoff	0	$-(s_0 - c_0)$ $+(l_t - r_t)$	...	$-(s_0 - c_0)$ $+(l_T - r_T)$

In this section we show that even if negative swap spreads are a textbook arbitrage opportunity, assuming no transaction costs and institutional frictions, the arbitrage strategy explained in Table 2.8 is still a risky strategy.²⁷ As pointed out by Shleifer and Vishny (1997), Liu and Longstaff (2004a) and many others, even textbook arbitrage opportunities are subject to a risk, namely the possibility that the mispricing increases before it vanishes, thereby forcing the arbitrageur out of his position at a loss. With negative swap spreads arbitrage, we know that the mispricing vanishes after 30 years, but we do not know whether it vanishes within a much shorter and practical horizon.

²⁷We ignore potential issues with leverage constraints or frictions in the repo market and illustrate the returns to swap spreads arbitrage in a "best case".

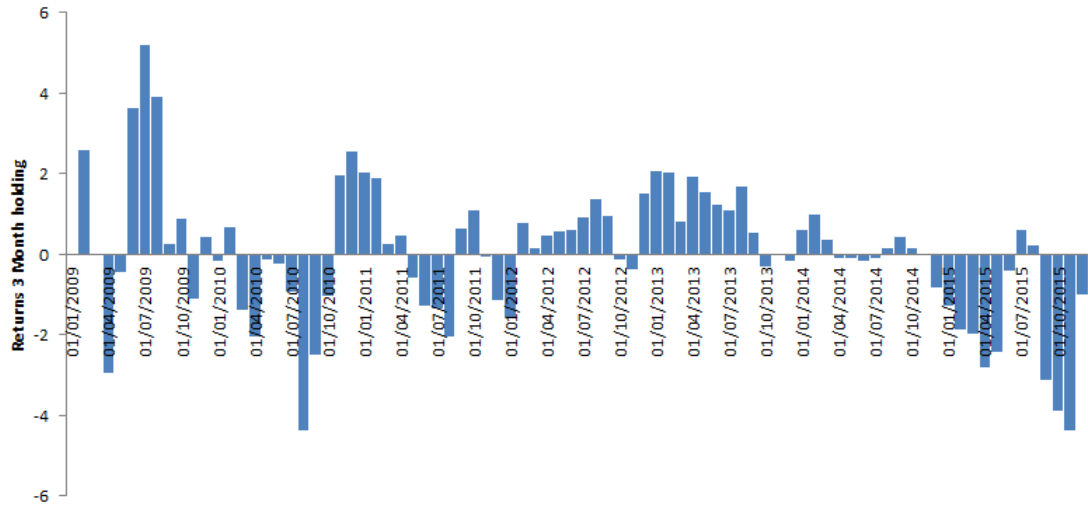
To illustrate this point we provide some stylized sample calculations to approximate the excess returns of an arbitrageur engaging in the strategy, described in Table 2.8. We assume that the arbitrageur unwinds his position before maturity and consider two cases. In the first case, we assume that the arbitrageur unwinds the position after 3 months, in the second case we assume that he unwinds after 12 months. In both cases he receives a positive carry from the strategy but is exposed to the risk that the swap spread becomes even more negative. For simplicity, we ignore the ageing of the treasury and swap and simply assume that the arbitrageur unwinds the position by engaging in an opposite transaction where he sells a treasury bond with 30-years to maturity and receives fixed in an IRS with 30-years to maturity.²⁸ Every month, the arbitrageur observes the 30-year swap spread and engages in the transaction if the swap spread is negative. We illustrate the resulting excess returns of the two strategies in Figure 2.7. The Sharpe ratio for the 3-month and 12-month strategies are 0.86% and 5.03% respectively. Note that the Sharpe ratio for investing in the US stock market for the same time period is 29.39%

2.7.4 Data Description

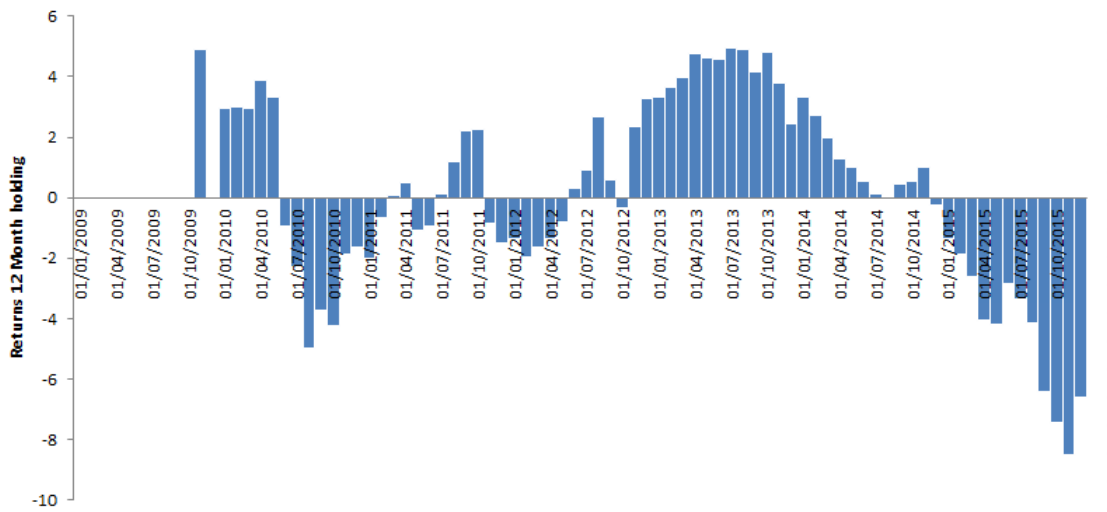
This appendix provides additional details about the data used for our analysis.

1. **Swap Spreads:** Swap rates and treasury yields for 2, 3, 5, 10, and 30 years to maturity are obtained from the Bloomberg system. The swap rates are the fixed rates an investor would receive on a semi-annual basis at the current date in exchange for quarterly Libor payments. The treasury yields are the yields of the most recently auctioned issue and adjusted to reflect constant time to maturity. For 3-year and 7-year treasury yields, we supplement the Bloomberg data with treasury yields from the FED H.15 reports due to several missing observations in the Bloomberg data. Swap spreads are computed as the difference between swap rate and treasury yield, where the swap rate is adjusted to reflect the different daycount conventions which are actual/360 for swaps and actual/actual for treasuries.
2. **Underfunded Ratio (UFR) :** Quarterly data on two types of defined benefit (DB) pension plans, private as well as public local government pension plans, are obtained from the financial accounts of the US (former flow of funds) tables L.118b and L.120b. UFR in quarter t is then computed using Equation (2.16). Next, positive and negative part are defined as $UFR_t^+ := \max(UFR_t, 0)$ and $UFR_t^- := \min(UFR_t, 0)$. Changes

²⁸This simplification leads to a duration mistake of 3 months in case one and 1 year in case two. Since swap and treasury originally have 30 years to maturity this ageing effect is neglect-able for our approximation.



(a) 3-months holding return



(b) 12-months holding return

Figure 2.7: Returns from 30-year swap spread arbitrage. The Figure shows the returns from engaging in swap spreads arbitrage. The Sharpe ratio of the two strategies are 0.86% and 5.03% respectively.

in UFR in the different regimes are computed as $\Delta UFR_t^+ := \Delta UFR_t \mathbb{1}_{\{UFR_t > 0\}}$ ($\Delta UFR_t^- := \Delta UFR_t \mathbb{1}_{\{UFR_t \leq 0\}}$).

3. **Libor-repo spread:** The 3-month Libor rate as well as the 3-month general collateral repo rate are obtained from the Bloomberg system. The Libor-repo spread is then computed as the difference between these two variables.
4. **Debt-to-GDP ratio:** Quarterly data on the US debt-to-GDP are obtained from the federal reserve bank of St. Louis which provides a seasonally-adjusted time series.
5. **Dealer-Broker EDF:** Expected default frequencies are provided by Moody's analytics and we use the equally-weighted average of the 14 largest derivatives dealing banks (G14 banks). These 14 banks are: Morgan Stanley, JP Morgan, Bank of America, Wells Fargo, Citigroup, Goldman Sachs, Deutsche Bank, Societe Generale, Barclays, HSBC, BNP Paribas, Credit Suisse, Royal Bank of Scotland, and UBS.
6. **Move Index:** The Move index is computed as the 1-month implied volatility of US treasury bonds with 2,5,10, and 30 years to maturity. Index levels are obtained from the Bloomberg system.
7. **Term Factor:** This factor captures the slope of the yield curve, measured as the difference between the 30-year treasury yield and the 3-month treasury yield. A description of these yields can be found under point 1 (swap spreads).
8. **Level:** The level of the yield curve is captured by the 30-year treasury yield. For a description of this yield see point 1 (swap spreads).
9. **VIX:** Is the implied volatility of the S&P 500 index and data on VIX are obtained from the Bloomberg System.
10. **On-the-run spread:** The spread is computed for bonds with 10-years to maturity because estimates of the 30-year spread are noisy and suffer from the 2002-2005 period where the US treasury reduced its debt issuance. The 10-year on-the-run yield is obtained from the FED H.15 website and the 10-year off-the-run yield is constructed as explained in Gürkaynak et al. (2007) and data are obtained from <http://www.federalreserve.gov/pubs/feds/2006>.
11. **Dealer Broker Leverage:** This variable captures the leverage of US broker-dealers and is described in more detail in Adrian et al. (2014). Until Q4 2009, data on this variable are obtained from Tyler Muir's website. Since the data ends in Q4 2009, we

use the financial accounts of the US data, following the procedure described in Adrian et al. (2014) to supplement the time series with more recent observations for the Q1 2010 – Q4 2015 period.

12. **Mortgage Refinancing:** Quarterly mortgage origination estimates are directly obtained from the Mortgage Bankers Association website. We use mortgage originations due to refinancing as a proxy for the mortgage refinancing rate.

Essay 3

High Funding Riks, Low Return¹

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Abstract

I develop a simple model in which hedge fund managers with access to less profitable investment strategies choose a higher exposure to funding risk in an attempt to generate competitive returns. Empirically, I find that hedge funds with a higher loading on a simple funding risk measure generate lower returns than hedge funds with a lower loading on that risk measure. In line with the model predictions, I find that (i) this underperformance is driven by a high loading on adverse funding shocks, (ii) a higher loading on funding risk predicts lower fund flows, and (iii) the results are significantly weaker for funds with less favorable redemption terms or funds with multiple prime brokers.

3.1 Introduction

Hedge funds are managed portfolios in which the returns depend on the fund's investment strategies and risk management. A good hedge fund follows alpha-generating strategies and simultaneously manages the funding risk that arises from the liability side of its balance sheet, that is, the risk of investor withdrawals and unexpected margin calls or increasing haircuts. If not managed properly, these funding risks can transform into severe losses because they can force a manager to unwind otherwise profitable positions at an unfavorable early point in time. Therefore, it is possible that hedge funds with a higher exposure to funding risk do not earn a risk premium for this additional risk, but generate lower expected returns than hedge funds with a lower exposure to that risk.

I show in a simple model that more funding risk taking is optimal for managers with access to less profitable strategies, even though it lowers expected fund returns. Utilizing a large cross-section of hedge fund returns, I find that hedge funds with a high loading on a simple funding risk measure (funds that are more exposed to common funding shocks) severely underperform hedge funds with a low loading on that measure (funds that are less exposed to common funding shocks). The empirical proxy for market-wide funding conditions is based on deviations from the covered interest rate parity (CIP), and I show that the proxy spikes when major institutional investors face tightening funding constraints. In line with the model's predictions, I document that hedge funds with a high loading on the funding risk measure experience more equity withdrawals than funds with a low loading on that measure and have a lower cash buffer against deteriorating funding conditions. Furthermore, the link between a high loading on the funding risk measure and low expected returns is less significant for funds that impose stricter redemption terms on their investors and for funds that have multiple prime brokers.

In my model, two hedge funds differ with respect to the return that they can generate from investing in an alpha-generating strategy. Funding risk arises because both funds face an exogenous risk of outflows which can force them to unwind their strategies early at a cost. Investors are initially unaware of the difference in the funds' alpha-generating strategies and withdraw from the bad fund, which is the fund with the lower alpha-generating strategy, once they can identify it. The bad fund, therefore, invests more aggressively in its funding-risky strategy to avoid being revealed as bad. Hence, if the funding shock is small, investors are unable to identify the bad fund. It is only if the funding shock is large enough that the bad fund generates losses. These losses due to the funding shock predict lower returns in the next period and enable the investors to identify the fund as the bad fund.

This mechanism gives the model's first two predictions. First, hedge funds that are

exposed to more funding risk generate lower returns. More precisely, hedge funds that generate lower returns when funding conditions worsen also generate lower future returns. Second, a higher exposure to funding risk predicts fund outflows. Hence, hedge funds with a higher exposure to funding risk have lower fund flows than hedge funds with a lower exposure to that risk. The third model prediction is that the difference in returns between funds with a high exposure to funding risk and funds with a low exposure to funding risk is lower if the size of the expected funding shock is smaller. This lower expected funding shock comes from the fund’s liabilities and could occur, for instance, if the fund imposes stricter redemption terms on its equity investors.

To proxy market-wide funding conditions faced by hedge funds, I construct an index of deviations from the CIP across several different currencies and maturities. The index (henceforth CIP^{Index}) is similar to one in Pasquariello (2014), capturing “dislocations in international money markets” and is strongly related to other proxies of funding liquidity, such as the Treasury-Eurodollar (TED) spread and the dealer-broker leverage factor constructed by Adrian et al. (2014).² Furthermore, deviations from the CIP are an ideal measure of the funding conditions faced by hedge funds for two reasons. First, they point toward a deviation from the law of one price which would not occur if major dealer banks had ample funding to take advantage of the mispricing. Second, they indicate the shortage of one currency relative to another, which suggests that major international investors face tightening funding constraints. These tightening constraints are likely passed on to hedge funds either through their prime brokers or via equity withdrawals from major institutional investors and can force funds to unwind otherwise profitable positions at a loss.

I use $CIPD_t$, defined as $CIP_{t-1}^{Index} - CIP_t^{Index}$, in my analysis to keep consistent with the notion that a high loading on unexpected funding shocks corresponds to high risk. To test my hypothesis that a higher loading on funding risk predicts lower returns, I obtain hedge fund returns and other fund characteristics for the January 1994 – May 2015 sample period from the TASS hedge fund database. Using the returns of these funds I then form decile portfolios based on their loading on CIPD over the past three years and rebalance the portfolios on a monthly basis. I find that hedge funds with a low loading on CIPD outperform hedge funds with a high loading on CIPD by a large margin. The risk-adjusted return of the difference portfolio that is long the hedge fund portfolio with the lowest loading on CIPD and short the hedge fund portfolio with the highest loading on CIPD has a risk-adjusted monthly return of 0.54% (t -statistic of 2.46). This result demonstrates that a high loading on funding risk indeed predicts poor fund performance. Instead of being a “priced

²This strong link to other funding risk proxies distinguishes CIP^{Index} from other previously used liquidity measures such as the noise measure Hu, Pan, and Wang (2013) or the Pastor and Stambaugh (2003) stock market liquidity measure.

risk factor,” funding risk, as measured by CIPD, has the opposite effect: a *higher* loading on CIPD predicts *lower* risk-adjusted returns.

To rule out the possibility that fund-specific characteristics drive this result, I perform two additional tests. First, I repeat the analysis forming style-neutral portfolios by fixing the percentage of hedge funds within a certain style in each of the decile portfolios; doing so leaves the main result unchanged. The difference portfolio – which is long hedge funds with a low loading on CIPD and short hedge funds with a high loading on CIPD – generates a monthly risk-adjusted return of 0.42% (t -statistic of 2.58). Second, I run Fama and MacBeth (1973) regressions of risk-adjusted hedge fund returns on β^{CIPD} , controlling for fund age, fund size, redemption notice period, lockup provision, investment style, minimum investment, management fee, and incentive fee. Even after controlling for all these fund characteristics, β^{CIPD} is a statistically significant explanatory variable (t -statistic of 2.78) for risk-adjusted hedge fund returns.

Because my model implies that lower returns due to an adverse funding shock predict lower subsequent returns I next investigate to which extent the results are driven by the negative part of CIPD. To that end, I split CIPD into a negative part, $CIPD^-$, which captures worsening funding conditions, and a positive part, $CIPD^+$, which captures improving funding conditions. I then repeat this sorting procedure twice, once only using $CIPD^-$ and once only using $CIPD^+$. In line with my theory, I find that hedge funds with a high loading on $CIPD^-$ severely underperform hedge funds with a low loading on $CIPD^-$. The difference portfolio – which is long funds with a low loading on $CIPD^-$ and short funds with a high loading on $CIPD^-$ – generates a monthly risk-adjusted return of 0.58% (t -statistic of 2.64), which is higher and more significant than the return of the CIPD-sorted difference portfolio described above. In contrast, there is no significant difference between hedge fund returns that are sorted based on their loading on $CIPD^+$.

The second testable prediction of my model is that the high loading on market-wide funding shocks enables hedge fund investors to distinguish bad funds from good funds and therefore triggers subsequent withdrawals. I investigate this prediction by checking whether hedge funds with a high loading on CIPD experience lower fund flows than hedge funds with a low loading on CIPD. Indeed, the difference between fund flows for hedge funds with a low loading on the funding risk proxy and fund flows for hedge funds with a high loading on that proxy is positive and statistically significant at a 1% level (t -statistic of 2.64). To disentangle the effect of a higher loading on CIPD from the effect of lower past returns, I repeat the sorting procedure conditional on past returns. Doing so lowers the significance of the result to a 5% level (t -statistic of 2.40), but leaves the main inference intact.

The third testable prediction of my theory is that the effect of a higher loading on

funding shocks is less pronounced for funds with a lower risk of investor redemptions or forced deleveraging due to their prime brokers. To investigate this hypothesis, I perform the following three tests. First, I split hedge funds into two different subsamples, one with redemption notice period of one month or less and one with redemption notice period above one month. Second, I split hedge funds into one subsample of funds with a lockup provision and one subsample of funds without a lockup provision. Finally, I split the sample into funds that have more than one prime broker and funds that only have one prime broker. The difference portfolio earns a higher risk-adjusted return for funds with a shorter redemption notice period, funds without a lockup provision and funds with only one prime broker compared to funds with a longer redemption notice period, funds with a lockup provision, and funds with more than one prime broker respectively.

In addition to my main findings, I address the concern that the higher return of funds with a low loading on funding shocks is driven solely by a few severe crisis episodes. To that end, I split the full sample period into crisis periods and normal periods based on two criteria. First, I use anecdotal evidence to classify 19 months as crisis episodes and find that the difference portfolio that is long hedge funds with a low loading on CIPD and short hedge funds with a high loading on CIPD generates a monthly risk-adjusted return of 0.45 (t -statistic of 2.09) during normal periods and a monthly risk-adjusted return of 1.10% (t -statistic of 1.74) during crisis episodes. Second, I classify NBER recession periods as crisis periods and find that the difference portfolio generates a monthly risk-adjusted return of 0.52% (t -statistic of 2.29) during normal times and a monthly risk-adjusted return of 0.46% (t -statistic of 0.90) during crisis periods.

Finally, I address the following four common biases in reported hedge fund returns: double counting, return smoothing, backfilling bias, and dropout bias. First, to address concerns about double-counting, I remove 14% of the funds in the database, which are likely to be subsidiaries of the same fund. Second, to address return smoothing, I use the return un-smoothing technique proposed by Getmansky, Lo, and Makarov (2004) and replace the original returns with the un-smoothed returns. Third, to address backfilling bias, I drop all returns reported before a fund was added to the database. Fourth, to address dropout bias, I add a delisting return of 0.00% after the last reported return for each hedge fund. This dropout return is motivated by the finding in Aiken, Clifford, and Ellis (2013) that fund returns for delisted funds are, on average, 0.5% lower than for funds that continue reporting. I then repeat the main analysis, sorting hedge funds into deciles based on their loading on CIPD and CIPD⁻, using this bias-cleaned subsample. While the alphas of each decile portfolio drop sharply, the main result remains virtually unchanged: The difference portfolio that is long hedge funds with a low loading on CIPD (CIPD⁻) and short hedge

funds with a high loading on CIDP ($CIPD^-$) earns a risk-adjusted return of 0.64% (0.67%) with a t -statistic of 2.29 (3.01).

3.2 Related Literature

The main finding of this paper is that hedge funds that face more risk, as proxied by a higher factor loading, do generate lower expected returns. This finding fits into a large literature on the cross-section of hedge fund returns and risk factors. Sadka (2010) and Teo (2011) document that stock market liquidity, as approximated by the Sadka (2006) liquidity measure and the Pastor and Stambaugh (2003) liquidity measure is a significant risk factor in the cross-section of hedge fund returns. Hu et al. (2013), Buraschi, Kosowski, and Trojani (2013), Bali, Brown, and Caglayan (2014), Agarwal, Arisoy, and Naik (2015), Agarwal, Ruenzi, and Weigert (2016), Gao, Gao, and Song (2016), construct other risk measures and show that they are significant risk factors in the cross section of hedge fund returns. In contrast to these studies, I find that a higher loading on CIPD predicts *lower* fund returns instead of higher future returns.

My finding that a higher risk exposure leads to lower risk-adjusted returns is similar to the finding in Titman and Tiu (2011) that hedge funds with returns that are less-well explained by common risk factors deliver higher risk-adjusted returns. In contrast to Titman and Tiu (2011), I find that a higher loading on a single risk measure predicts lower subsequent returns. This finding is in line with Chen and Lu (2015) who construct a funding-liquidity measure based on stock returns and find that hedge funds with a lower loading on their funding risk measure outperform funds with a higher loading on that measure. Chen and Lu (2015) focus on establishing a new funding risk measure based on stock market returns, while I use a simple measure which is based on a deviation from the law of one price. I also document that my high funding risk, low return finding is linked to manager's access to alpha-generating strategies. Jurek and Stafford (2015) show that the average alpha of the hedge fund industry can be explained by taking downside risk, as approximated by a put writing strategy. While Jurek and Stafford (2015) find that hedge funds are profiting from taking downside market risk, I find that exposure to deteriorating funding conditions lowers expected returns.

My finding is also related to the literature on two hedge-fund-specific funding risks. First, the risk of equity withdrawals which has been empirically studied by, among others, Aragon (2007), Klebanov (2008), and Hombert and Thesmar (2014). These authors find that hedge funds that offer less-favorable redemption terms to their equity investors outperform funds offering more favorable redemption terms. In line with these papers, I find that hedge funds

with less favorable redemption terms and with a high loading on $CIPD^-$ are delivering higher risk-adjusted returns than funds without more favorable redemption terms and a high loading on $CIPD^-$.³ Second, the risk of adverse funding conditions being passed on from the prime broker. This risk has been studied by, among others, Aragon and Strahan (2012), Mitchell and Pulvino (2012), and Ang, Gorovyy, and Van Inwegen (2011). Aragon and Strahan (2012) find that hedge funds for which Lehman Brothers was the prime broker suffered a large funding shock in 2008. Mitchell and Pulvino (2012) show that short-term financing through prime brokers was a general issue for hedge funds during the financial crisis. Ang et al. (2011) show that hedge fund leverage is counter-cyclical to the leverage of major dealer-brokers and that it decreased significantly during the financial crisis. In line with these studies, I find that a higher loading on $CIPD^-$ is less relevant for funds with multiple prime brokers, that is, funds that face a lower risk of tightening funding conditions.

The theoretical part of my paper is related to the literature on how the institutional environment surrounding hedge funds restricts their optimal risk taking. My model is closest to that of Liu and Mello (2011), who show that the “fragile nature of hedge fund equity” limits a manager’s ability to profit from funding-risky positions. In a similar spirit, Shleifer and Vishny (1997) find that withdrawals can occur precisely when the manager needs cash the most and Chen, Goldstein, and Jiang (2010) show that funds with illiquid asset holdings face a higher withdrawal risk. In addition to outflows, a second source of funding risk is borrowing constraints, which can arise if funds have to collateralize their positions and can force a manager to unwind otherwise profitable strategies early. These constraints have been studied by, among others, Gromb and Vayanos (2002), Liu and Longstaff (2004b), Brunnermeier and Pedersen (2009), Gârleanu and Pedersen (2011), and Gromb and Vayanos (2015). Finally, my model is also close to the literature that incorporates funding frictions into a hedge fund manager’s investment decisions (see Dai and Sundaresan, 2011, Buraschi, Kosowski, and Sritrakul, 2014, Pangeas and Westerfield, 2009, Lan, Wang, and Yang, 2013, and Drechsler, 2014, among many others).

The remainder of this paper is organized as follows. In Section 3.3, I develop a simple model to derive the main hypotheses, and in Section 3.4, I describe the hedge fund data and the CIP deviation measure. Section 3.5 presents my main results, which are complemented with additional results that provide a closer link to the theory in Section 3.6. Section 3.7 shows the results of various robustness checks. Section 3.8 investigates whether it is possible to make money on a CIPD-based hedge fund strategy. Section 3.9 concludes.

³More recently, Aiken, Clifford, and Ellis (2015) point out that additionally to the redemption terms reported in commercial hedge fund databases, hedge funds are often using discretionary liquidity restrictions, such as gates and side pockets, which effectively alter redemption terms. Hence, using factor loadings as an indirect measure of funding risk can provide additional insights, compared to reported fund characteristics.

3.3 The Model

The purpose of this section is to show through a model that hedge fund managers with access to less profitable strategies choose a smaller cash buffer against worsening funding conditions, thereby exposing them more to funding risk, even though this higher exposure lowers expected fund returns.

3.3.1 Model Setup

There exist two hedge funds which differ only with respect to their access to alpha-generating strategies. A “good” fund g , with access to a strategy that yields a return of α_g and a “bad” fund b , with access to a strategy with return $\alpha_b < \alpha_g$. To keep the model simple, the returns α_g and α_b are known constants and both strategies are subject to the same unwinding cost c if not held until maturity. More precisely, one dollar invested in strategy $i \in \{g, b\}$ is worth $1 + \alpha_i$ dollars if the strategy is held until maturity and $1 - c$ dollars if the strategy is unwound early. Additionally to their strategies, both funds can invest in a risk-free asset (cash) that yields a return of zero and is not subject to unwinding costs. Both funds are risk neutral and have a total of one unit of capital that they can allocate between their alpha-generating strategies and their risk-free asset.

Investing in the alpha-generating strategy is risky because of an uncertain funding shock λ which occurs before the maturity of the risky strategy and can be interpreted as a fraction of investors withdrawing their money because of a sudden liquidity need. The funding shock λ is uniformly distributed on the interval $[0, \bar{\lambda}]$, where $0 < \bar{\lambda} \leq 1 - c$ to avoid situations in which the funding shock can lead to bankruptcy of the fund, and both funds face the same risk of a funding shock. The expected terminal fund wealth for fund $i \in \{g, b\}$ is therefore given as follows:

$$\mathbb{E}[W_i] = \frac{1}{\bar{\lambda}} \left[\int_0^{1-\theta_i} (1 - \lambda + \theta_i \alpha_i) d\lambda + \int_{1-\theta_i}^{\bar{\lambda}} \left(\theta_i - \frac{\lambda - (1 - \theta_i)}{1 - c} \right) (1 + \alpha_i) d\lambda \right]. \quad (3.1)$$

The first integral is up to the cash holding $1 - \theta_i$. If the realization of the funding shock does not exceed the fund’s cash holding, the funding shock does not lead to any unwinding costs. The second integral starts at $1 - \theta_i$ and goes up to $\bar{\lambda}$. In this region, the funding shock forces the manager to unwind part of his strategy early.

Hedge fund investors cannot directly observe a fund’s type $i \in \{g, b\}$ or its risk taking θ_i and initially invest the same amount in both funds. However, investors know the parameter values α_g , α_b , and c , the distribution of the funding shock, and can observe the realization of the funding shock as well as the terminal fund wealth. From that, investors can infer the

good fund's wealth-maximizing investment θ_g^* and the good fund's terminal wealth. If the terminal wealth of the bad fund is below that of the good fund, the bad fund is revealed. I assume that being revealed as bad fund is associated with a cost γ for the manager, which can be interpreted as investors withdrawing from the bad fund and the fund losing the continuation value. Hence, γ can be interpreted as “performance-based withdrawals”. Fund i therefore faces the following optimization problem:

$$\max_{\theta_i} \{ \mathbb{E}[W_{2,i}] - \gamma p^i(\theta_i) \}, \quad (3.2)$$

where $p^i(\theta_i)$ is the probability of being revealed as bad fund, which is equal to zero for the good fund and strictly positive for the bad fund.

Next, I introduce two parametric conditions that I assume to hold throughout this section.

Condition 1. (a) *The maximal funding shock $\bar{\lambda}$ satisfies the following inequality:*

$$\bar{\lambda} \geq (\alpha_g - \alpha_b) \frac{(\alpha_g + c)}{(1 + \alpha_g)\alpha_g c}. \quad (3.3)$$

(b) *The continuation value γ satisfies the following inequality:*

$$\gamma > \frac{\bar{\lambda}(\alpha_g - \alpha_b)(c\bar{\lambda}[c(1 + \alpha_g) + \alpha_b(\alpha_g + c)] - (\alpha_g + c)(\alpha_b + c))}{2(1 - c)(\alpha_g + c)\alpha_b(c + \alpha_b)(\alpha_g(1 + \alpha_g)c\bar{\lambda} + (\alpha_g + c)\alpha_b) - \alpha_g(\alpha_g + c)}. \quad (3.4)$$

Although these two expressions are complex, they have a simple interpretation and I show in a numerical illustration in the appendix that they are only imposing a mild restriction on the model parameters. The intuition behind Condition 1 is as follows. Because the maximal amount that the bad fund can invest in its strategy is one, Inequality (3.3) is necessary to ensure that it is possible for the bad fund to mimic the returns of the good fund by taking on additional risk. Inequality (3.4) ensures that the cost of being revealed is large enough to give the bad fund an incentive to deviate from its wealth-maximizing strategy in order to mimic the returns of the good fund. Inequality (3.4) can be obtained by computing $p^b(\theta^M) = \frac{1 - \theta_b^M}{\bar{\lambda}}$ and solving the following inequality for γ :

$$\mathbb{E}[W_2 | \theta = \theta_b^M] - \gamma p^b(\theta^M) > \mathbb{E}[W_2 | \theta = \theta_b^*] - \gamma.$$

3.3.2 Results and Testable Predictions

Computing the two integrals in Equation (3.1) and taking the first order condition (FOC) leads to the wealth-maximizing investment θ_i^* in the risky asset:⁴

$$\theta_i^* = 1 - \frac{c(1 + \alpha_i)\bar{\lambda}}{\alpha_i + c}. \quad (3.5)$$

Because $p^i(\theta^i) = 0$ for the good fund, Equation (3.5) gives the solution to the good fund's optimization problem. The solution to the bad fund's optimization problem is stated in the following proposition.

Proposition 5. *Assume that Condition 1 holds. Then, fund b invests $\theta_b^M = \frac{\alpha_g}{\alpha_b}\theta_g^* < 1$ in its strategy. It increases its investment in the strategy relative to the wealth-maximizing investment by:*

$$\theta_b^M - \theta_b^* = \left(\frac{\alpha_g}{\alpha_b} - 1 \right) + \bar{\lambda} \frac{(1 - c)(\alpha_g - \alpha_b)c}{(\alpha_g + c)(\alpha_b + c)} > 0. \quad (3.6)$$

The proof of this proposition is straightforward. If the bad fund invests θ^M in its risky asset and the funding shock $\lambda < 1 - \theta^M$, then both funds generate the same return. Furthermore, because $\theta^M > \theta_b^*$, the bad fund does not invest more than θ^M in its risky asset. An immediate consequence of Equation (3.6) is that the risk taking of Fund b in excess of its optimal investment in the strategy increases for lower α_b and for higher $\bar{\lambda}$. This additional risk taking corresponds to a higher exposure to funding risk because the fund keeps a lower cash buffer against the funding shock.

The return of fund i can be computed as $R_i = W_i + \lambda - 1$, which is the percentage change in fund wealth, adjusted for early withdrawals and is given as:

$$R_i = \begin{cases} \theta_i \alpha_i, & \text{if } \lambda \leq 1 - \theta_i \\ \frac{(\alpha_i + c)(1 - \lambda) - c\theta(1 + \alpha_i)}{1 - c}, & \text{if } \lambda > 1 - \theta_i. \end{cases} \quad (3.7)$$

Defining $\beta^i := \frac{\text{cov}(R_i, \lambda)}{\text{Var}(\lambda)}$, the following proposition shows that funds with a higher sensitivity to the funding shock (measured by a more negative β^i) have lower expected returns and face more investor withdrawals.

Proposition 6. *Assume that Condition 1 holds and that $\beta^i < \beta^j$. Then the following statements hold:*

⁴The second derivative of Equation (3.1) with respect to θ is a negative constant, which ensures that taking the FOC indeed leads to the optimal investment.

(a) *Fund i has a lower expected return than fund j .*

(b) *In expectation, fund i faces more performance-based withdrawals than fund j .*

To prove this proposition, note that the good fund generates higher expected returns than the bad fund and that, in expectation, the bad fund faces more performance-based withdrawals than the good fund. Computing the betas for the good fund and the bad fund shows that $\beta^g > \beta^b$, which proves the proposition. The proposition provides two testable predictions: Funds with a higher exposure to funding risk (i) generate lower expected returns and (ii) face withdrawals. Note that more funding risk in the model is taken on indirectly by investing more aggressively in the alpha-generating strategy and leaving a lower cash buffer. It is this indirect exposure to funding risk that I later capture by computing the loadings of fund returns on changes in a market-wide funding risk measure.

Finally, the following proposition shows that past return sensitivity to the funding shock becomes less informative as the funds' exposure to the funding shock decreases.

Proposition 7 (Prediction 3). *Assume that Condition 1 is satisfied and that $\beta^i < \beta^j$. Then $\mathbb{E}[R_j] - \mathbb{E}[R_i]$ is decreasing in $\bar{\lambda}$.*

What remains to show in order to prove this proposition is that the difference between $\mathbb{E}[R_g]$ and $\mathbb{E}[R_b]$ is decreasing in $\bar{\lambda}$. Taking the first derivative of this difference with respect to $\bar{\lambda}$ shows that the expression is falling in $\bar{\lambda}$ and proves the proposition. This proposition delivers the third testable prediction: A higher sensitivity to the common funding shock is less informative when the maximal size $\bar{\lambda}$ of the funding shock is smaller. The intuition behind this prediction is that, if $\bar{\lambda}$ is small, then the difference between α_g and α_b needs to be small as well, otherwise the bad fund would not be able to mimic the returns of the good fund. Empirically, a lower $\bar{\lambda}$ could come from either the equity side of the funds' balance sheet, due to lockups and less favorable redemption terms, or from the liability side of the balance sheet, where a hedge fund with multiple prime brokers is less susceptible to an adverse funding shock than a manager with only one prime broker.

3.4 The Data

3.4.1 Hedge Fund Data

The hedge fund data for my analysis are obtained from the May 2016 version of the Lipper TASS hedge fund database. Hedge funds report voluntarily to this database, and one concern with these self-reported returns is survivorship bias because poorly performing funds might

just decide to drop out of the database. To mitigate this concern, I follow the common practice and use both live hedge funds (which are still reporting to TASS as of the latest download) and graveyard funds (which stopped reporting). Because the graveyard database was not established until 1994, I focus my analysis on the January 1994 – May 2015 period. Following the literature on hedge funds (see, for instance, Cao, Chen, Liang, and Lo, 2013 and Hu et al., 2013, among others), I apply three filters to the database. First, I require funds to report returns net of fees on a monthly basis. Second, I drop hedge funds with average assets under management (AUM) below 10 million USD.⁵ For funds that do not report in USD, I use the appropriate exchange rate to convert AUM into USD equivalents. Third, I require that each fund in my sample reports at least 36 monthly returns during my sample period.

Panel A of Table 3.1 provides summary statistics for all hedge funds in the filtered sample. For variables that change over time, I first compute the time-series average and then report cross-sectional summary statistics in the table. The first two rows of Panel A show that the average fund in the database reports a positive return of 0.58% per month with a standard deviation of 3.07. On average, funds have 146 million U.S. dollar in AUM, ranging from the minimum of 10 million up to 7,158 million. AUM is defined as the value of all claims that equity shareholders have on the fund, that is, the difference between the value of all long positions (including cash) and the value of all short positions (including borrowing). Furthermore, the average fund in the database reports 90 monthly returns and is 47 months old.

TASS also provides information on when each hedge fund began reporting to the database, which I use to compute the percentage of backfilled returns – 46.51% on average, with a high standard deviation of 32.7% across funds. In my main analysis, I include backfilled return observations and show later that the results are robust to dropping backfilled observations. The next two variables provide an overview of the funds’ risk of withdrawals. The first variable is a dummy variable that equals one if the fund has a lockup provision and zero otherwise. Nineteen percent of the funds in the database have a lockup provision. The second variable is the funds’ redemption notice period which indicates how long it takes for equity investors to withdraw their money. The variable varies across funds from 0 to 12 months, with an average of approximately one month. The last two variables in Panel A show the manager’s compensation. In line with the often-mentioned 2/20 rule, the median management and the median incentive fee of funds in my sample are 1.5% and 20% respectively.

⁵I also experimented with different requirements for AUM, such as 5 Mio USD and 20 Mio USD, which left the results unchanged.

Table 3.1: Hedge fund summary statistics. This table provides summary statistics of average hedge fund returns in the TASS database, as well as key fund characteristics. AUM is the fund’s assets under management and converted in USD for funds that report in a different currency (using the appropriate exchange rate). “Reporting” and “Age” are the number of monthly return observations and the average number of past return observations, respectively. “Backfilled” is a dummy variable that equals one if the fund return in a given month is backfilled. “Lockup” is a dummy variable that equals one if the fund has a lockup provision. “Notice” is the number of months that investors have to notify the manager before withdrawing capital from the fund. Panel B reports summary statistics of hedge fund returns per style. The sample period is January 1994 to May 2015.

	N	Mean	SD	Min	Median	Max
Panel A: Summary statistics for all hedge funds						
Return (mean)	8,541	0.58	0.64	-6.68	0.54	5.80
Return (SD)	8,541	3.07	2.62	0.00	2.30	45.74
AUM (mio USD)	8,541	146.26	320.79	10.00	53.92	7158.02
Reporting (Months)	8,541	97.63	49.72	36.00	85.00	257.00
Age (Months)	8,541	50.53	30.45	17.50	42.50	365.00
Backfilled	8,541	0.46	0.33	0.00	0.40	1.00
Lockup	8,541	0.19	-	-	-	-
Notice (Months)	8,541	1.07	1.12	0.00	1.00	12.17
Management Fee	8,480	1.41	0.74	0.00	1.50	22.00
Incentive Fee	8,046	13.43	8.67	0.00	20.00	50.00
Panel B: Hedge fund returns for different styles						
Convertible Arbitrage	170	0.49	0.49	-1.24	0.53	1.81
Emerging Markets	445	0.78	0.84	-3.14	0.72	5.58
Equity Market Neutral	315	0.47	0.47	-1.08	0.40	2.64
Event Driven	474	0.76	0.67	-3.92	0.72	5.35
Fixed Income Arbitrage	251	0.56	0.60	-2.88	0.61	2.11
Fund of Funds	2,987	0.32	0.47	-5.20	0.31	3.03
Global Macro	337	0.72	0.77	-6.68	0.74	5.64
Long Short Equity	1,812	0.82	0.67	-2.11	0.76	4.89
Managed Futures	402	0.68	0.65	-3.99	0.60	3.80
Multi-Strategy	1,019	0.73	0.58	-2.61	0.78	5.73
Other	329	0.65	0.79	-1.75	0.58	5.80

Panel B of Table 3.1 summarizes average hedge fund returns for the different styles. As we can see, average monthly returns range from 0.82% for long-short equity to 0.32% for funds of funds. There are a total of 2,987 funds of funds in my sample. I run my main analysis using all 8,541 funds and later show that my results are robust to splitting the sample into hedge funds and funds of funds. Summary statistics for hedge fund returns in different years can be found in Appendix 3.10.1 (Table 3.10). These yearly summary statistics show that the number of funds varies from a minimum number of 711 in 1994 up to 5,720 in 2009. Hence, splitting the overall sample of hedge funds into different subcategories can result in a

relatively small sample during some years. Later, in my analysis, I account for this problem by sorting hedge funds into quintiles instead of deciles to ensure a sufficient number of funds per portfolio.

3.4.2 Deviations from the Covered Interest Rate Parity

In this section, I construct a simple measure of market-wide funding conditions faced by hedge funds, which is based on deviations from the covered interest rate parity (CIP). The idea behind the CIP is that the theoretical forward exchange rate between currency A and currency B can be computed using the following no-arbitrage argument. One can either invest one unit of currency A at time t in a money-market account with interest rate $r^A(t, T)$ or exchange this one unit of currency A into currency B , putting it into a money-market account with interest rate $r^B(t, T)$. To avoid arbitrage opportunities from borrowing in one currency and investing into the other, the theoretical forward rate should be given as:

$$Fwd_{A/B}^*(t, T) := FX_{A/B}(t) \left(\frac{1 + r^A(t, T)}{1 + r^B(t, T)} \right), \quad (3.8)$$

where $FX_{A/B}(t)$ denotes the spot exchange rate from currency A to currency B .

The measure of deviations from the CIP, which is closely related to the measure of “dislocations in international money markets” constructed in Pasquariello (2014), is based on the following nine currency pairs: CHFUSD, EURUSD, GBPUSD, JPYUSD, CHFEUR, GBPEUR, JPYEUR, CHFGBP, JPYGBP, as well as spot rates and forward rates with 7, 30, 60, 90, 180, 270, and 360 days to maturity. In each of the currencies, LIBOR rates with the same maturity as the forward rates are used as a proxy for the risk-free rate. All data for constructing CIP deviations are obtained from the Bloomberg system.⁶ Deviations from the CIP for currency pair A/B with maturity $T - t$ are then computed as the absolute difference between the observed forward exchange rate $Fwd_{A/B}$ and the theoretical forward exchange $Fwd_{A/B}^*$ implied by Equation (3.8):

$$CIP_{i,t} = |\ln(Fwd_{A/B}(t, T)) - \ln(Fwd_{A/B}^*(t, T))| \times 10^4. \quad (3.9)$$

The expression is multiplied by 10^4 to obtain a mispricing in basis points. In total, this leads

⁶Spot and forward exchange rates are the London closing rates (4:00 pm). LIBOR rates are the ICE LIBOR rates which are released at 11:45 am London time.

to 63 different currency-maturity pairs, which are aggregated into one index as follows:

$$CIP_t^{Index} = \frac{1}{n_t} \sum_{i=1}^{n_t} CIP_{i,t}, \quad (3.10)$$

where n_t denotes the number of available mispricings at time t .

Deviations from the CIP occur if demand pressure for currency forwards is not met by a sufficient amount of arbitrage capital. The demand for currency forwards is driven by an imbalance between international funding supply and investment demand. Such an imbalance points towards a shortage of one currency relative to another (Bottazzi, Luque, Pascoa, and Sundaresan, 2012 and Ivashina, Scharfstein, and Stein, 2015, among others). The most prominent example of such a shortage is the USD shortage in 2011, in which foreign banks experienced tightening funding conditions in the U.S. money markets. The amount of available arbitrage capital decreases when major dealer banks face tightening funding constraints and can therefore no longer supply currency derivatives at the arbitrage-free rate. These tightening funding conditions can be passed on to hedge funds either through equity withdrawals by major institutional investors needing liquidity or through prime brokers passing their own tightening funding conditions to their hedge fund clients. Hence, deviations from the CIP point to a situation in which hedge funds face tightening funding conditions.

There are two main criticisms of using CIP_t^{Index} as measure for funding conditions. First, using LIBOR as a proxy for the risk-free rate can be problematic because LIBOR is an unfunded lending rate that can contain a credit-risk component and because LIBOR rates are potentially biased due to misreporting.⁷ I address these concerns in Appendix 3.10.1, where I construct CIP_t^{Index} using overnight lending (OIS) rates instead of LIBOR rates and find that the main results remain intact when using this alternative index, even though OIS rates for most currencies are only available from 2002 on. The second concern is that deviations from the CIP are not driven by dislocations in international money markets but by trading costs in currency markets. I address the concern that CIP_t^{Index} might be driven by currency market liquidity in Appendix 3.10.1, where I repeat my main analysis controlling for the currency liquidity measure constructed in Karnaukh, Ranaldo, and Söderlind (2015).⁸ The main results remain unchanged after controlling for FX liquidity.

Figure 3.1 shows the time series of month-end CIP_t^{Index} , where the blue lines highlight major market events and the grey-shaded areas are U.S. recession periods. The first larger

⁷Tuckman and Porfirio (2003) argue that the credit-risk component in LIBOR is one of the primary drivers of CIP deviations. Eisl, Jankowitsch, and Subrahmanyam, 2013 investigate LIBOR misreportings.

⁸The correlation between CIPD and changes in the FX liquidity measure is 0.15. I am grateful to Valeri Sokolovski for his help with updating this measure.

spike in CIP^{Index} occurs in September 1998, when Long-Term Capital Management (LTCM) was bailed out. Afterwards, the measure starts spiking again at the onset of the financial crisis, showing a small increase during the Quant crisis in August 2007, a larger spike during the bailout of Bear Stearns in March 2008, and a major spike in September 2008, when Lehman Brothers filed for bankruptcy. The next major spike of the measure occurs during the onset of the European debt crisis in autumn 2011. The blue line labelled “Euro Crisis” marks June 2011, when the rating agency Moody’s put several European banks on watch for possible downgrades, which lead to tightening funding conditions for these banks. This event was followed by more negative news about European sovereigns, which subsequently lead to the European debt crisis. The measure remains elevated until July 2012 when Mario Draghi delivered his famous speech declaring that “the ECB is ready to do whatever it takes to preserve the euro. And believe me, it will be enough.”⁹ The most recent spike of the measure occurs in January 2015, when the Swiss National Bank decided to lift its currency peg.¹⁰

I introduce the notation $CIPD_t := CIP_{t-1}^{Index} - CIP_t^{Index}$ to be consistent with the notion that lower returns during an unexpected funding shock correspond to a high factor loading. Using this notation, I now investigate whether $CIPD_t$ is related to the following four measures of funding liquidity and market uncertainty: (i) changes in the difference between the 3-month USD LIBOR rate and the 3-month U.S. treasury yield, commonly referred to as the TED spread, ΔTED_t , (ii) changes in the implied volatility of the S&P 500 index, ΔVIX_t , (iii) stock returns of the nine largest investment banks, Ret_t^{IB} , and (iv) the dealer-broker leverage variable introduced by Adrian et al. (2014), $Leverage_t$ (more details on these variables can be found in Appendix 3.10). The results of regressing $CIPD_t$ on these variables are exhibited in Table 3.2, where I first focus on the relationship between $CIPD_t$ and the first three variables, which are available on a monthly basis. As we can see from columns (1)–(3) of Table 3.2, ΔTED_t , ΔVIX_t , and Ret_t^{IB} all have a significant effect on $CIPD_t$. Increases in TED_t and VIX_t correspond to worsening funding conditions and are negatively related to $CIPD_t$, while higher bank stock returns correspond to improving funding conditions and are positively related to $CIPD_t$. The TED spread is the strongest explanatory variable and explains 36% of the variation in $CIPD_t$ in a univariate regression. Column (4) shows that combining the three independent variables explains 41% of the vari-

⁹A verbatim of the speech is available on the ECB website [Link].

¹⁰Another observation from Figure 3.1 is that CIP^{Index} became more volatile after the default of Lehman Brothers and even comparably small events like the lifting of the Swiss Currency Peg triggered large spikes. One possible explanation for this observation can be the implementation of the Volcker rule which explicitly forbids banks to engage in proprietary trading, such as arbitrage. Hence, a major group of arbitrageurs who used to enforce the CIP is not allowed to do so anymore which causes smaller events to have a larger impact on the index.

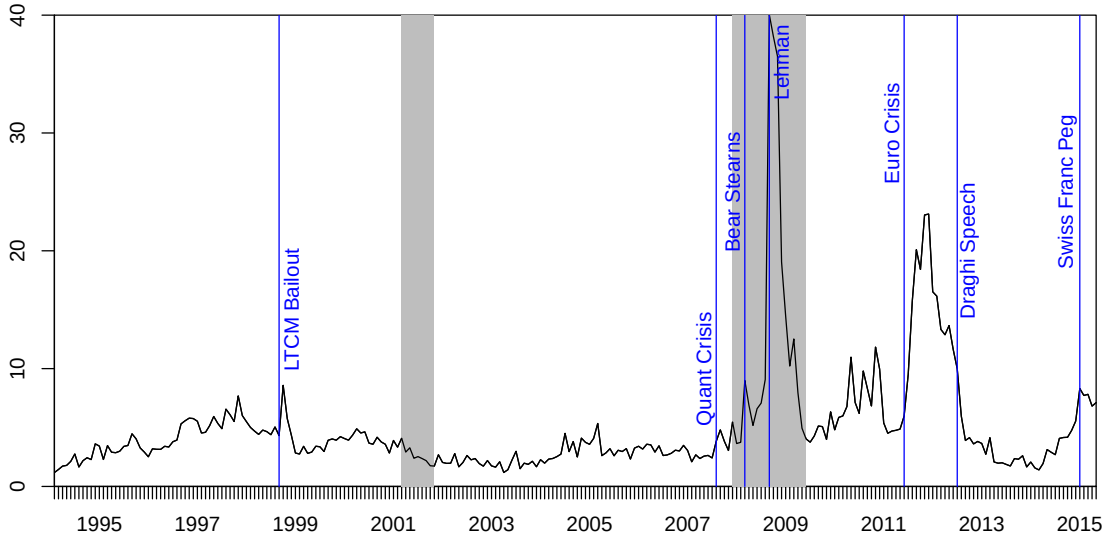


Figure 3.1: Time Series of the Covered-Interest Rate Parity (CIP) Deviation Index. This figure shows the time series of the CIP deviation index. The index is constructed as an equal-weighted average of nine of the most liquid currency pairs with seven different maturities, ranging from one week to one year, based on Equations (3.8)–(3.10). All observations are month-end. The highlighted events (blue vertical lines) are the bailout of Long-Term Capital Management (LTCM) in September 1998, the quant crisis in August 2007, the bailout of Bear Stearns in March 2008, the default of Lehman Brothers in September 2008, the onset of the European debt crisis in June 2011 (marked by rising concerns about European banks), Mario Draghi’s speech in July 2012 declaring that the ECB will do “whatever it takes” to preserve the Euro, and the Swiss National Bank lifting the currency peg to the Euro in January 2015. The two grey-shaded areas are US recession periods.

ance in $CIPD_t$ and confirms that ΔTED_t is the most significant explanatory variable of the three. Column (5) shows that $Leverage_t$, which is only available on a quarterly basis, explains 67% of the variation in $CIPD_t$. Combining all four explanatory variables, column (6), shows that $Leverage_t$ is the most significant explanatory variable, followed by ΔTED_t . Overall, the results confirm that $CIPD_t$ is capturing tightening funding conditions.¹¹

3.4.3 Hedge Fund Risk Factors

I now briefly describe the seven hedge fund risk factors, proposed by Fung and Hsieh (2004), that I use as benchmarks to compute risk-adjusted returns and show that these factors are only weakly related to $CIPD_t$. Again, more details about the data used to construct these factors are available in Appendix 3.10. The first two factors are related to stock markets,

¹¹An overview of the correlation between $CIPD_t$ and other commonly used liquidity proxies can be found in Appendix 3.10.1 (Table 3.15). One distinct feature of $CIPD_t$ compared to these other liquidity proxies is its strong correlation with ΔTED_t and $Leverage_t$.

Table 3.2: Properties of the Covered Interest Rate Parity Deviation Index. This table shows the results for regressions of $CIPD_t := CIP_{t-1}^{Index} - CIP_t^{Index}$ on other proxies of funding liquidity and market uncertainty. The four different explanatory variables are (i) the change in the differences between the 3-month USD LIBOR rate and the 3-month US treasury yield (ΔTED_t), (ii) average returns of the nine major U.S. investment banks (Ret_t^{IB}), (iii) changes in the option-implied volatility of the S&P 500 Index (ΔVIX_t), and (iv) the dealer-broker leverage factor ($Leverage_t$) introduced by Adrian et al. (2014). The sample period is January 1994 to May 2015. In columns (1)–(4), observations are month-end; in columns (5)–(6) observations are quarter-end. Newey-West t -statistics are reported in parenthesis. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	(1)	(2)	(3)	(4)	(5)	(6)
Intercept	−0.02 (−0.21)	−0.03 (−0.20)	−0.02 (−0.18)	−0.02 (−0.25)	−0.42 (−1.17)	−0.01 (−0.03)
ΔTED_t	−0.07** (−2.16)			−0.06** (−2.10)		−0.05*** (−2.81)
Ret_t^{IB}		0.10** (2.06)		0.02 (0.76)		0.06** (2.21)
ΔVIX_t			−0.22* (−1.89)	−0.12* (−1.67)		−0.07 (−1.30)
$Leverage_t$					−0.23*** (−6.38)	−0.13*** (−6.09)
Observations	257	257	257	257	85	85
Adjusted R ²	0.36	0.12	0.12	0.41	0.67	0.82

capturing U.S. stock market excess returns (MKT) and the returns from a small-minus big portfolio (SMB). These factors are proxied by the first two Fama-French factors. The next two factors are related to fixed income markets; Fung and Hsieh (2001) suggest using the monthly change in the 10-year U.S. treasury constant maturity yield (YLD) and the monthly change in the Moody’s Baa yield less 10-year Treasury constant maturity yield (BAA) as risk factors capturing interest-rate risk and credit risk. Finally, Fung and Hsieh (2001) also propose three trend-following factors, constructed from trading strategies in lookback straddles one for bonds (BD), one for currencies (FX), and one commodities (COM). The pairwise correlation between $CIPD_t$ and MKT, SMB, YLD, BAA, BD, FX, and COM is 0.17, 0.03, −0.14, −0.07, −0.12, −0.17, and −0.13 respectively. The entire correlation matrix can be found in Appendix 3.10.1 (Panel B of Table 3.15).

Sadka (2010) points out that YLD and BAA are not capturing excess returns and are therefore not suitable to compute risk-adjusted hedge fund returns. I therefore follow Sadka (2010) and replace these two factors with tradable factors in my performance analysis in the following section. In particular, I use excess returns of the Merrill Lynch Treasury bond

index with 7-10 years to maturity over the one-month risk-free rate as a tradable YLD factor. The correlation between YLD and the tradable YLD factor is -69% . Similarly, I use the difference between returns of the corporate bond index of BBB-rated bonds with 7-10 years to maturity and returns of the above Treasury bond index as a tradable BAA factor. The correlation between BAA and the tradable BAA factor is -76% . In the following, I replace YLD and BAA with the two tradable factors to compute risk-adjusted returns.

3.5 Results

In this section, I test my main hypothesis: Hedge funds with a higher exposure to funding risk generate lower returns than hedge funds with a lower exposure to funding risk. To do so, I sort hedge funds into deciles based on their loading on CIPD. Every month, for each Fund i , I run a regression of hedge fund returns over the past 36 months on CIPD, controlling for excess returns of the (stock) market portfolio:¹²

$$R_{i,t} = \alpha + \beta^{CIPD} CIPD_t + \beta^{Mkt} R_t^{Mkt} + \varepsilon_t. \quad (3.11)$$

Based on β^{CIPD} , I then put each hedge fund in one decile portfolio. The decile portfolios are rebalanced every month, repeating the sorting procedure. The first portfolio (P1) has the highest loading on CIPD while the last portfolio (P10) has the lowest loading on CIPD.

Panel (a) of Figure 3.4 shows the monthly risk-adjusted returns of the 10 portfolios, controlling for the seven Fung and Hsieh (2004) risk factors. As we can see from the Figure, funds with the lowest loading on CIPD (Portfolio 10) earn a monthly risk-adjusted return of 0.50% , which corresponds to an annual alpha of 6.00% . The risk-adjusted returns of hedge funds in the different deciles decrease almost monotonically. Funds in Portfolio 9 earn a monthly risk-adjusted return of 0.38% followed by funds in Portfolio 8, which earn a monthly risk-adjusted return of 0.33% . In contrast, funds with the highest loading on CIP deviations (Portfolio 1) do not earn any risk-adjusted returns, having a monthly alpha of 0.00% .

Although a long-short trading strategy in different hedge funds is not possible, it is still instructive to look into the performance of the difference portfolio that is long hedge funds with a low loading on CIPD and short hedge funds with a high loading on CIPD. The monthly risk-adjusted return of the difference portfolio is 0.50% , illustrated by the black

¹²Controlling only for returns of the market portfolio in the first step has been common practice in the literature (see Sadka, 2010 and Hu et al., 2013 for hedge funds, or Ang, Hodrick, Xing, and Zhang, 2006, among many others, for stocks) and ensures a sufficient amount of degrees of freedom for each independent variable.

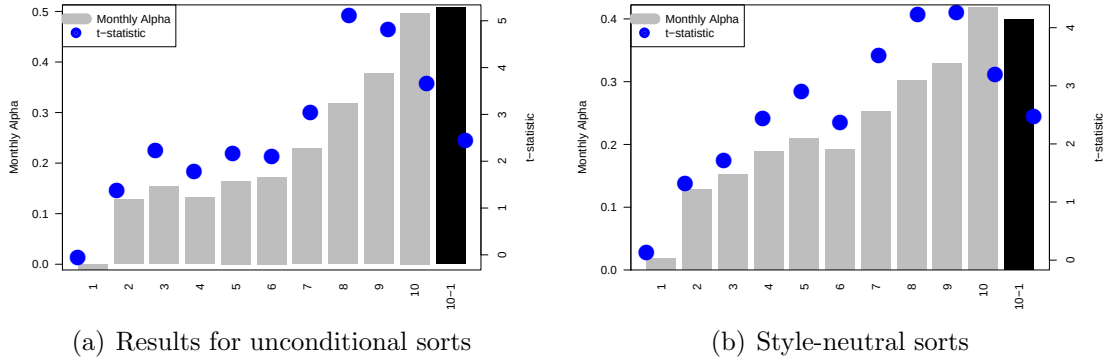


Figure 3.2: Risk-adjusted returns of CIPD-beta sorted hedge fund portfolios. Each month hedge funds are sorted into 10 equally-weighted portfolios according to their historical beta to the CIPD measure, constructed in Section 3.4.2. Funds in Portfolio 1 have the highest loading on CIPD, funds in Portfolio 10 have the lowest loading on CIPD. For each fund, the CIPD beta is calculated using a regression of monthly fund returns on CIPD controlling for the returns of the stock market portfolio, using the 36 months prior to portfolio formation. The bars represent monthly risk-adjusted portfolio returns, calculated using the Fung and Hsieh (2004) seven-factor model, where the YLD and BAA factors are replaced by factor-mimicking tradable portfolios. The blue dots are Newey-West t -statistics of the respective risk-adjusted returns. The black bar displays the risk-adjusted return of the difference portfolio, which is long hedge funds in portfolio 10 and short hedge funds in portfolio 1. Panel (a) shows the results for unconditional sorts. Panel (b) shows the results for style-neutral sorts. The sample period is January 1994 to May 2015, including all 8,541 hedge funds from the TASS database.

bar in Panel (a) of Figure 3.4. The blue dots in Figure 3.4 show Newey-West t -statistics of the respective portfolio alphas and indicate that the results are statistically significant.¹³ In particular, the risk-adjusted return of the difference portfolio is significant at a 5% level (t -statistic of 2.39). Hence, the null hypothesis that funds in Portfolios 10 and 1 generate the same risk-adjusted returns can be rejected.

One obvious concern about these results is that they might simply be driven by different hedge fund styles. For example, convertible arbitrage and fixed income arbitrage are investment styles that rely heavily on leverage and Table 3.1 shows that these styles also generate lower average returns than other styles. To address this concern, Panel (b) of Figure 3.4 reports the results for style-neutral portfolios, where I repeat the sorting procedure, conditional on each of the ten portfolios consisting of the same percentage of hedge fund styles. More precisely, I first split the overall sample of hedge funds into the 11 different styles and sort each of these subsamples into decile portfolios, based on their loading on CIPD. For each decile, I then merge the 11 different style portfolios, which ensures that each portfolio

¹³Throughout the paper the number of lags to compute Newey-West t -statistics is determined using the automatic lag selection described in Newey and West (1994).

has the same percentage of styles.

As we can see from Panel (b) of Figure 3.4, the results remain almost unchanged when portfolios are style neutral. Funds with the lowest loading on $CIPD_t$ (Portfolio 10) earn a monthly risk-adjusted return of 0.42% and the risk-adjusted returns decreasing almost monotonically to the portfolio consisting of funds with the highest loading on CIPD (Portfolio 1), which generates a monthly alpha of 0.02%. The difference portfolio, which is long Portfolio 10 and short Portfolio 1, generates a monthly risk-adjusted return of 0.38% (t -statistic of 2.33). Because my results are almost unchanged when forming style-neutral portfolios and because forming style-neutral portfolios is an uncommon approach in the hedge fund literature, I report my main results using unconditional sorts and provide additional details for conditional sorts in Appendix 3.10.1 (Panel A of Table 3.12).¹⁴

More details and additional results for the unconditional sorts are reported in Table 3.3. The first column of the table reports the Fung and Hsieh alphas discussed previously. The second column reports the risk-adjusted returns relative to the seven Fung and Hsieh factors and five additional factors. I add more risk factors because the seven factors might not be sufficient to capture all the risks that funds with different investment styles can be exposed to. The five factors that I add to the seven-factor model are the following. First, because fund returns in a subsequent month could be a consequence of an institutional momentum effect (see, for instance, Lou, 2012 and Vayanos and Woolley, 2013), I add the UMD momentum factor from Kenneth French's website. Second, because the CIPD is related to currency risks, I add the two currency risk factors proposed by Lustig, Roussanov, and Verdelhan (2011), which capture currency returns of a U.S. dollar investor and a carry trader, respectively. Finally, I add the excess returns of the S&P GSCI Commodity Index and the MSCI Emerging Markets Index to ensure that the risks of funds investing in commodities or emerging markets are captured as well. As we can see from the second column of Table 3.3, adding these risk factors does not have a significant impact on risk-adjusted returns. If anything, risk-adjusted returns of the ten decile portfolios tend to increase moderately. Most importantly, the difference portfolio generates a monthly risk-adjusted return of 0.48% (t -statistic of 2.44) which is almost identical to the risk-adjusted returns relative to the Fung and Hsieh factors.

¹⁴In my model, a higher loading on CIPD is equivalent to lower past returns. To rule out the possibility that my finding is only driven by lower past returns, I also form past-return-neutral portfolios in which I first split the overall sample of hedge funds into deciles based on their past returns over the last 36 months and then sort each of these subsamples into decile portfolios, based on their loading on CIPD. For each decile, I then merge the 10 different past return deciles. This procedure leads to qualitatively similar results as the unconditional sorts. The conditional difference portfolio which is long hedge funds with a low loading on CIPD and short hedge funds with a high loading on CIPD is generating a risk-adjusted return of 0.44 (t -statistic of 2.44). More details for this split can be found in Appendix 3.10.1 (Figure 3.7 and Table 3.12).

Table 3.3: Risk-adjusted returns and other characteristics of CIPD-sorted portfolios.

Hedge funds are sorted into deciles based on their beta to the CIPD measure described in Section 3.4.2. Beta is calculated using a regression of monthly hedge fund returns on CIPD, controlling for the stock market portfolio, and using the 36 months prior to portfolio formation. α^{FH} is the intercept of regressing the portfolio returns on the seven Fung Hsieh risk factors, α^{Add} is the intercept of regressing hedge fund returns on the seven Fung Hsieh factors plus five additional factors, β^{Mkt} and β^{CIPD} are the portfolio loadings on the stock market portfolio and on CIPD, respectively, R_{FH}^2 is the adjusted R^2 of regressing the portfolio returns on the seven Fung Hsieh factors. Under post-sorting, all quantities are computed using the returns of the formed hedge fund portfolios. Under pre-sorting, the average factor loadings of individual hedge funds, prior to portfolio formation, are reported. The seven Fung Hsieh factors are the market excess return (MKT), a size factor (SMB), tradable factors to mimic monthly changes in the 10-year Treasury constant maturity yield (YLD) and monthly changes in the Moody's Baa yield less 10-year Treasury constant maturity yield (BAA), as well as three trend-following factors: BD (bond), FX (currency), and COM (commodity). The five additional factors are a stock market momentum factor, the two currency risk factors proposed by Lustig et al. (2011), excess returns of the S&P GSCI Commodity Index, and excess returns of the MSCI Emerging Market Index. The sample period is January 1994 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Post-sorting					Pre-sorting	
	α^{FH}	α^{Add}	β^{Mkt}	β^{CIPD}	R_{FH}^2	β^{Mkt}	β^{CIPD}
P1	0.00 [-0.02]	0.06 [0.49]	0.52*** [8.65]	0.17*** [5.05]	0.64	0.41*** [9.25]	1.25*** [6.45]
P2	0.14 [1.46]	0.15** [2.22]	0.30*** [7.94]	0.15*** [4.03]	0.63	0.30*** [11.83]	0.51*** [5.28]
P3	0.16** [2.33]	0.17*** [3.11]	0.26*** [9.10]	0.12*** [5.93]	0.69	0.25*** [11.43]	0.31*** [4.42]
P4	0.14* [1.85]	0.13** [2.47]	0.22*** [8.68]	0.10*** [4.69]	0.65	0.22*** [10.81]	0.18*** [3.53]
P5	0.17** [2.25]	0.18*** [3.15]	0.21*** [8.69]	0.10*** [4.40]	0.60	0.21*** [10.63]	0.09** [2.27]
P6	0.17** [2.14]	0.19*** [3.01]	0.20*** [7.34]	0.10*** [4.37]	0.59	0.18*** [10.05]	0.02 [0.43]
P7	0.23*** [3.12]	0.24*** [3.96]	0.17*** [8.26]	0.08*** [3.97]	0.51	0.18*** [9.35]	-0.06* [-1.76]
P8	0.33*** [5.20]	0.32*** [5.82]	0.19*** [10.64]	0.06*** [2.67]	0.56	0.21*** [9.86]	-0.15*** [-3.47]
P9	0.38*** [4.86]	0.40*** [5.03]	0.22*** [10.32]	0.01 [0.29]	0.47	0.28*** [9.54]	-0.34*** [-4.88]
P10	0.50*** [3.71]	0.53*** [4.32]	0.31*** [5.67]	-0.02 [-0.57]	0.42	0.44*** [8.89]	-1.05*** [-6.20]
P10-P1	0.50** [2.39]	0.48** [2.44]	-0.21** [-2.36]	-0.19*** [-3.42]	0.30	0.03 [0.41]	-2.29*** [-7.07]

Columns 3 and 4 of Table 3.3 show the post-sorting β^{Mkt} and β^{CIPD} of the ten decile portfolios. The post-sorting β^{CIPD} is significantly different in the portfolio with the highest loading on CIPD (P1), which has a β^{CIPD} of 0.17 (t -statistic of 5.05), than in the portfolio with the lowest loading on CIPD (P10), which has an insignificantly negative beta of -0.02 (t -statistic of -0.57). Furthermore, the difference portfolio has a β^{CIPD} of -0.19 (t -statistic of 3.42) and the post-sorting betas are monotonically decreasing from Portfolio 1 to Portfolio 10. In addition to the significant difference in CIPD loadings, Portfolio 10 also has a lower loading on the market portfolio than Portfolio 1. This observation is in line with Titman and Tiu (2011) who argue that hedge funds with access to a true alpha-generating strategy choose a lower loading on well-known risk factors. Column 5 of Table 3.3 reports the R^2 from regressing the decile portfolio returns on the seven Fung-Hsieh risk factors.¹⁵ In line with the results of Titman and Tiu (2011), returns of hedge funds with a lower loading on CIPD are less-well explained by common risk factors. Finally, the last two columns of Table 3.3 show the pre-sorting β^{Mkt} and β^{CIPD} .

3.5.1 Other Explanations?

I now address the question of whether the difference in returns between funds with a low loading on CIPD and funds with a high loading on CIPD is simply driven by other fund characteristics, such as age, size, redemption terms, or managerial incentives.¹⁶ To that end, I run Fama and MacBeth (1973) regressions of risk-adjusted hedge fund returns on their CIPD beta, controlling for various fund-specific characteristics. To run the Fama-MacBeth regression, I compute the risk-adjusted excess return of each hedge fund, using the following equation:

$$R_{i,t}^{\perp} = R_{i,t}^{Exc} - (\beta_i^{Mkt} R_t^{Mkt} + \beta_i^{SMB} R_t^{SMB} + \beta_i^{YLD} R_t^{YLD} + \beta_i^{BAA} R_t^{BAA} + \beta_i^{BD} R_t^{BD} + \beta_i^{FX} R_t^{FX} + \beta_i^{COM} R_t^{COM}), \quad (3.12)$$

where fund-specific betas are computed using the entire time series of hedge fund returns. I then follow the common practice (see, e.g. Klebanov, 2008 or Hu et al., 2013) and assign the CIPD betas of the respective portfolios to each fund instead of using the rolling estimates of each individual fund. In particular, a fund that is in Portfolio i at time t and in Portfolio j at time $t + 1$ gets β^{CIPD} of Portfolio i at time t and β^{CIPD} of Portfolio j at time $t + 1$. I

¹⁵For brevity, I do not report the loadings on all seven Fung-Hsieh factors separately. These results can be found in Appendix 3.10.1 (Table 3.11).

¹⁶An overview of the average fund characteristics for the 10 CIP-beta-sorted portfolios can be found in Appendix 3.10.1 (Table 3.13).

then run regressions with the following control variables:

$$\begin{aligned}
R_{i,t}^\perp = & \gamma_0 + \gamma^{CIPD} \beta_{i,t-1}^{CIPD} \\
& + \gamma^{Age} Age_{i,t-1} + \gamma^{Size} \ln(AUM_{i,t-1}) + \gamma^{Notice} Notice_i + \gamma^{Lockup} DLockup_i \\
& + \gamma^{MinInvest} MinInvest_i + \gamma^{MgFee} MgFee_i + \gamma^{IncFee} IncFee_i + \varepsilon_{i,t}, \quad (3.13)
\end{aligned}$$

gradually adding the controls in the second and third line. $Age_{i,t-1}$ and $\ln(AUM_{i,t-1})$ are Fund i 's age and log-size at time $t - 1$. $Notice_i$ and $DLockup_i$ are Fund i 's redemption notice period (in months) and a dummy variable that equals one if Fund i has a lockup provision and zero otherwise. $MinInvest_i$, $MgFee_i$, and $IncFee_i$ are variables capturing the minimum investment, the management fee, and the incentive fee for Fund i .

The regression results are exhibited in Table 3.4. In Column (1) I run a regression without controlling for any fund-specific characteristics. In Column (2), I add fund age, size, redemption notice period, lockup provision, and backfilled dummy, as well as style dummies as controls. In Column (3) I run the full regression (3.13), controlling for minimum investment, management fee, and incentive fee. The table shows that β_{t-1}^{CIPD} is a significant explanatory variable for the cross-section of risk-adjusted hedge fund returns, even after controlling for fund-specific characteristics. In all three specifications, β_{t-1}^{CIPD} is statistically significant at a 1% level. However, the economical and statistical significance of β_{t-1}^{CIPD} decreases with the amount of control variables added to the regression. Without additional controls an increase in β_{t-1}^{CIPD} of 0.10 corresponds to a decrease of 0.48% in monthly risk-adjusted returns and the effect is statistically significant at a 1% level (t -statistic of -3.11). Adding the first four control variables lowers the economic significance of an increase of 0.10 in β_{t-1}^{CIPD} to 0.46 with a t -statistic to -2.97 . In the third specification an increase of 0.10 in β_{t-1}^{CIPD} corresponds to a decrease of 0.41 in risk-adjusted returns (t -statistic of -2.61).

In addition to the results for β^{CIPD} , it is also worth noting that all the controls have the expected signs. Aggarwal and Jorion (2010) document that younger hedge funds tend to outperform older hedge funds, which is in line with the negative coefficient on age. Aragon (2007) finds that hedge funds with lockups and longer redemption notice periods outperform hedge funds without lockups and with shorter redemption notice periods, which is also the case in my sample and reflected by the positive coefficients on Notice and Lockup. Moreover, higher managerial incentives (in the form of higher fees) tend to increase returns, which is in line with Agarwal, Daniel, and Naik (2009).

Table 3.4: Results using cross-sectional regressions. This table reports the results of Fama and MacBeth (1973) regressions of the cross section of monthly hedge fund alphas (relative to the Fung-Hsieh seven factor model). In the first specification, the beta on CIPD, estimated over the past 36 months, is used as an independent variable. In the second specification, fund age in months, the log of the fund size, the fund's redemption notice period, a dummy variable that equals one if the fund has a lockup provision and is zero otherwise, and an investment style dummy are added as controls. In the third specification, minimum investment, fund management fee, and fund incentive fees are added as control variables. The evaluation period is January 1997 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	(1)	(2)	(3)
Intercept	0.500*** [5.53]	0.540*** [4.55]	0.313*** [2.83]
CIPD beta	-4.830*** [-3.11]	-4.590*** [-2.97]	-4.093*** [-2.61]
Fund age (months)		-0.001 [-1.35]	-0.001 [-1.51]
Log(fund size)		-0.014 [-0.72]	-0.023 [-1.16]
Redemption notice (months)		0.032 [1.29]	0.039* [1.83]
Lockup dummy		0.124*** [3.45]	0.100*** [2.90]
Minimum Investment (mio USD)			0.013** [2.19]
Management Fee (%)			0.047 [1.53]
Incentive Fee (%)			0.010*** [3.61]
StyleDummies	No	Yes	Yes

3.5.2 Fund-Specific Funding Risk and β^{CIPD}

I conclude this section by investigating whether fund-specific funding risk, such as equity withdrawals and stricter redemption terms can explain factor loadings. To that end, I run cross-sectional regressions of $\beta_{t,i}^{CIPD}$ on the following variables:

$$\begin{aligned}
\beta_{i,t}^{CIPD} = & \gamma_0 + \gamma^{Flow} Flow_{i,t-1} + \gamma^{Lockup} DLockup_i + \gamma^{Notice} Notice_i + \\
& + \gamma^{MinInvest} MinInvest_i + \gamma^{Leveraged} DLeverage_i + \\
& + \gamma^{Age} Age_{i,t-1} + \ln(AUM_{i,t-1}) + \gamma^{PastRet} Ret_{i,t-1} + \varepsilon_{i,t},
\end{aligned} \tag{3.14}$$

gradually adding the controls in the second and third line. See Section 3.5.1 for a description of the independent variables.

The regression results are exhibited in Table 3.5. As we can see from Panel (I), fund-specific funding risk is significant in explaining hedge fund loadings on CIPD. Funds with higher past flows and funds with a lockup provision have lower loadings on CIPD than funds with higher past flows and funds without lockup provision, respectively. Panel (II) of Table 3.5 shows that adding a fund's minimum investment, a leverage dummy, and fund style dummies do not lower the significance of past flows and lockup provisions. Surprisingly, hedge fund leverage is not significant in explaining hedge fund factor loadings, which could be due to the fact that funds not using leverage are typically of a particular investment style. Finally, Panel (III) shows that larger funds tend to have a higher loading on CIPD and funds with higher past returns tend to have a lower loading on CIPD.

3.6 Link to the Model Predictions

So far, I have established that a higher loading on CIPD predicts lower risk-adjusted returns, which is in line with the main model prediction. I now test the other model predictions in the data and obtain the following results. First, I find that it is a higher loading on past funding shocks, as proxied by the negative part of CIPD, that drives the results; this is additional evidence for the first model prediction. Second, in line with the second model prediction, I find that hedge funds with a high loading on CIPD experience significantly lower fund flows than funds with a low loading on CIPD. Third, in line with the third model prediction, I find that the effect of a high past loading on CIPD is weaker for funds that are less exposed to funding risk. Finally, I also document that the returns of hedge funds with a high CIPD-loading are more sensitive to past flows than the returns of hedge funds with a low CIPD-loading. This last finding is indirect evidence that hedge funds with a lower loading on CIPD hold a higher cash buffer against unexpected withdrawals.

3.6.1 Negative Shocks Driving the Results

According to my model, low returns during a funding shock predict lower future returns. However, the model does not give any predictions about the relationship between fund returns and improving funding conditions. Hence, I next investigate whether it is the link between deteriorating funding conditions and hedge fund returns that causes the cross-sectional difference in performance. To do so, I split CIPD into a positive and negative part and repeat the sorting procedure described above. Recall that CIPD is defined as

Table 3.5: Cross-sectional regressions of CIPD betas. This table reports the results of Fama and MacBeth (1973) regressions of the cross section of monthly hedge fund beta on CIPD, estimated over the past 36 months on the indicated variables. In Panel (I) past fund flows, a lockup dummy, which equals one if the fund has a lockup provision and zero otherwise, and the funds' redemption notice period (in months) are used as independent variables. In Panel (II), the funds' minimum investment, a leverage dummy, and investment style dummies are added as controls. In Panel (III), fund age in months, the log of the fund size, and the past returns over the last 36 months are added. The evaluation period is January 1997 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	(I)	(II)	(III)
Intercept	0.10** [2.06]	0.106** [2.16]	0.004 [0.07]
Flow	-0.05** [-2.51]	-0.049** [-2.54]	-0.053** [-2.43]
Lockup Dummy	-0.04*** [-2.71]	-0.039*** [-2.70]	-0.030** [-2.04]
Redemption notice (months)	-0.01 [-1.15]	-0.010 [-1.07]	-0.014 [-1.25]
Minimum Investment		-0.001 [-1.34]	-0.004** [-2.15]
Leveraged		0.005 [0.42]	0.010 [0.69]
Age			0.000 [-0.17]
Size			0.027*** [3.63]
Past Return			-0.254** [-2.32]
StyleDummies	No	Yes	Yes

$CIP_{t-1}^{Index} - CIP_t^{Index}$, and therefore a lower CIPD corresponds to deteriorating funding conditions. Every month t , for each Fund i , I run the following regression using the past 36 months of return observations:

$$R_{i,t} = \alpha + \beta^{CIPD-} \min(CIPD_t, 0) + \beta^{CIPD+} \max(CIPD_t, 0) + \beta^{Mkt} R_t^{Mkt} + \varepsilon_t \quad (3.15)$$

and then perform two tests. In the first test, I sort hedge funds into decile portfolios based on β^{CIPD-} and in the second test, I sort them based on β^{CIPD+} .

The results of these two tests are exhibited in Figure 3.3, where Panel (a) shows the results for sorting on negative CIPD (deteriorating funding conditions; henceforth CIPD⁻) and Panel (b) shows the results for sorting on positive CIPD (improving funding conditions;

henceforth $CIPD^+$). Comparing Panel (a) of Figure 3.3 to Panel (a) of Figure 3.4 shows that the results even improve when only using the negative part of CIPD. The difference portfolio that is long hedge funds with a low loading on $CIPD^-$ and short hedge funds with a high loading on $CIPD^-$ delivers a monthly risk-adjusted return of 0.62 (t -statistic of 2.78). In contrast to that, we see an opposite pattern when sorting on $CIPD^+$. Here, hedge funds with a higher loading are generating higher risk-adjusted returns than hedge funds with a lower loading. However, the difference portfolio generates an insignificant risk-adjusted return of -0.27 (t -statistic of -1.63).

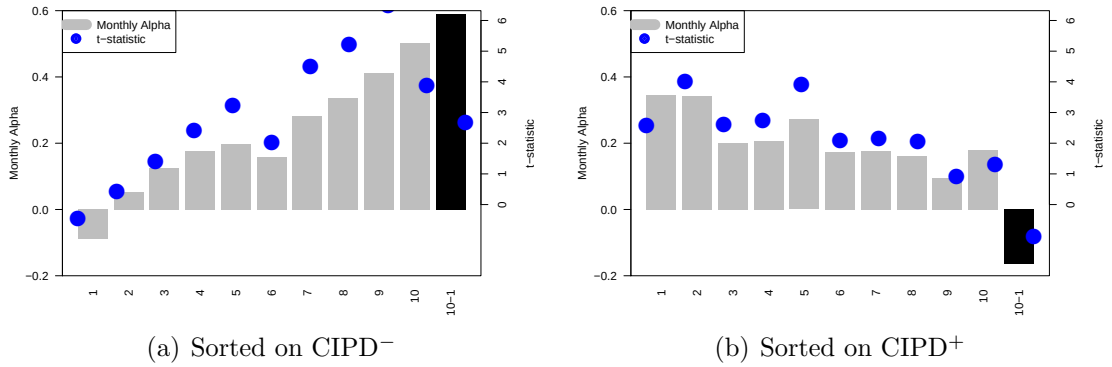


Figure 3.3: Results for hedge fund portfolios sorted on $CIPD^-$ and $CIPD^+$. Each month hedge funds are sorted into 10 equally-weighted portfolios according to their historical beta to the negative part of CIPD (Panel (a)) and their historical beta to the positive part of CIPD (Panel (b)). For a detailed description of the sorting procedure as well as the computation of risk-adjusted returns see the caption of Figure 3.4. The grey bars represent monthly risk-adjusted portfolio returns, calculated using the Fung and Hsieh (2004) seven-factor model, where the YLD and BAA factors are replaced by factor-mimicking tradable portfolios. The blue dots are Newey-West t -statistics of the respective risk-adjusted returns. The black bar displays the risk-adjusted return of the difference portfolio, which is long hedge funds in Portfolio 10 and short hedge funds in Portfolio 1. The sample period is January 1994 to May 2015, including all 8,541 hedge funds from the TASS database.

Panel A of Table 3.6 provides additional details for the two difference portfolios (the results for portfolios 1-10 are omitted for brevity). The first row of Panel A confirms that using $CIPD^-$ instead of CIPD leads to marginally stronger results. The difference portfolio earns a higher risk-adjusted return and the post-sorting β^{CIPD} (which is computed using only $CIPD^-$ is of the same magnitude as in baseline case. The second row of Panel A shows that this is not the case when sorting on $CIPD^+$, where the difference portfolio generates an insignificant return and β^{CIPD} (which is computed using only $CIPD^+$) is insignificant. Furthermore, Panel A also confirms that adding more risk factors to the Fung and Hsieh seven factor benchmark model leaves the main inference unchanged.

Table 3.6: Additional results. Hedge funds are sorted into portfolios based on their beta to the CIPD measure, described in Section 3.4.2, and based on different modifications of CIPD. For a detailed description of the sorting procedure and the different variables see the caption of Table 3.3. Each row reports the results for a difference portfolio. Panel A reports the results for hedge funds that are sorted into deciles based on their loading on the negative part of CIPD (1) and on the positive part of CIPD (2). Panel B shows the results for different subsamples of the hedge fund database, where funds are sorted into quintiles based on their loading on CIPD⁻. The sample is split into hedge funds with a redemption notice period longer than one month and hedge funds with a redemption notice period shorter than one month (rows (1) and (2)), hedge funds with a lockup provision and hedge funds without a lockup provision (rows (3) and (4)), and hedge funds which use more than one prime broker and hedge funds which only use one prime broker (rows (5) and (6)). Panel C shows the results for a bias-cleaned modification of the database; dropping all backfilled observations, adding a delisting return of 0.00% after the last reported return, dropping potential duplicates, and un-smoothing the returns using the procedure described in Getmansky et al. (2004). Panel (1) shows the results for CIPD-sorted portfolios, Panel (2) for CIPD⁻-sorted portfolios. The sample period is January 1994 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Post-sorting					Pre-sorting	
	α^{FH}	α^{Add}	β^{Mkt}	β^{CIP}	R_{FH}^2	β^{Mkt}	β^{CIP}
Panel A: Sorting on increases and decreases in CIPD^D							
(1) Sort on CIPD ⁻	0.62*** [2.78]	0.63*** [3.04]	-0.30*** [-4.07]	-0.18*** [-5.59]	0.37	-0.09 [-1.21]	-4.00*** [-5.86]
(2) Sort on CIPD ⁺	-0.27 [-1.63]	-0.29** [-2.38]	0.17* [1.86]	-0.13 [-1.33]	0.11	0.17*** [2.89]	-3.84*** [-6.19]
Panel B: Results for funds with different liquidity risk							
(1) Longer notice	0.37** [2.27]	0.35** [2.41]	-0.18*** [-3.59]	-0.13*** [-4.67]	0.28	-0.06 [-1.10]	-2.35*** [-5.92]
(2) Shorter notice	0.55*** [3.29]	0.57*** [3.64]	-0.27*** [-4.70]	-0.14*** [-5.70]	0.44	-0.11 [-1.45]	-3.01*** [-5.86]
(3) Funds with lockup	0.20 [1.32]	0.25 [1.64]	-0.11** [-2.25]	-0.16*** [-5.02]	0.19	-0.03 [-0.61]	-2.73*** [-5.90]
(4) Funds without lockup	0.50*** [3.09]	0.51*** [3.50]	-0.26*** [-4.68]	-0.15*** [-6.25]	0.44	-0.11 [-1.59]	-2.76*** [-5.88]
(5) More than one PB	0.15 [0.73]	0.14 [0.72]	-0.22*** [-2.88]	-0.27*** [-5.75]	0.27	-0.05 [-0.91]	-2.30*** [-5.52]
(6) Only one PB	0.39** [2.43]	0.42** [2.57]	-0.21*** [-3.72]	-0.16*** [-4.51]	0.39	-0.08 [-1.13]	-2.98*** [-6.08]
Panel C: Results for different robustness checks							
(1) Bias-cleaned CIPD	0.64** [2.29]	0.57** [2.45]	-0.30*** [-2.72]	-0.24*** [-4.02]	0.26	0.03 [0.37]	-2.33*** [-7.12]
(2) Bias-cleaned CIPD ⁻	0.67** [2.31]	0.68*** [2.75]	-0.38*** [-5.01]	-0.22*** [-5.96]	0.39	-0.09 [-1.21]	-4.00*** [-5.86]

3.6.2 Fund Flows

Hedge funds with a high loading on CIPD expose their investors to more funding risk and generate lower risk-adjusted returns. My theory suggests that once a hedge fund generates a lower return during a funding shock, investors realize that they invested into a fund that takes higher risks to generate its returns and withdraw their money. Hence, the second

model prediction is that funds with a high loading on CIPD experience lower flows than funds with a low loading on CIPD.¹⁷ In testing this model prediction, it is important to disentangle fund flows that occur due to a higher exposure to funding risk from fund flows that simply occur due to poor past performance.¹⁸

To investigate the second testable model prediction, I compute the flow in month t for each Fund i as:

$$Flow_{i,t} := \frac{AUM_{i,t} - AUM_{i,t-1}}{AUM_{i,t-1}} - R_{i,t}, \quad (3.16)$$

where I adjust the change in AUM for returns over the same period (as is common in the mutual funds literature, see, for instance, Chevalier and Ellison, 1997). I then compute average portfolio flows as:

$$Flow_t^{PF} := \frac{\sum_{i=1}^{n_t} Flow_{i,t} AUM_{i,t-1}}{\sum_{i=1}^{n_t} AUM_{i,t-1}}, \quad (3.17)$$

where n_t is the number of funds in the portfolio at time t . One issue with this measure of portfolio fund flows is that outflows and inflows might occur gradually because lockups and unfavorable redemption terms can keep investors from withdrawing immediately. If funds move between portfolios frequently, the flow measure is not related to the fund's sensitivity to CIPD. Since the average CIPD-sorted (CIPD⁻-sorted) fund spends 52% (53%) of its time in the same decile portfolio, I split the sample into quintiles instead, where the average fund spends 65% (65%) of its time in the same portfolio.

The resulting average flows for the quintile portfolios, as well as the difference between fund flows for the portfolio with the lowest funding risk and the portfolio with the highest funding risk, are exhibited in Panel A of Table 3.7. The first row shows the results for funds that are sorted based on CIPD and the second row shows the results for funds that are sorted on CIPD⁻. In both cases, funds in the portfolio with the highest loading on CIPD are on average subject to outflows while funds in the portfolio with the lowest loading on CIPD are on average subject to inflows. However, apart from one exception, the fund flows for the quintile portfolios are not significantly different from zero. In contrast to that, there is a significant difference between fund flows to hedge funds with a low loading on the

¹⁷The notion that investors are slow in changing their investments in different funds is in line with the idea of Gârleanu and Pedersen (2015) who argue that search costs for asset management and noise allocators make it difficult for investors to distinguish good funds from bad funds.

¹⁸A large literature details fund flows in response to past performance. See Chevalier and Ellison (1997) and Sirri and Tufano (1998) who document that investor flows are convex in past performance for mutual funds. See Baquero and Verbeek (2015), Ding, Getmansky, Liang, and Wermers (2015), and Agarwal, Green, and Ren (2015) for a discussion of hedge fund investors.

Table 3.7: Average flows for CIPD-sorted hedge fund portfolios. Hedge funds are sorted into quintiles according to their loading on CIPD (sort on β^{CIPD}) and on their loading on the negative part of CIPD (sort on β^{CIPD-}). For a detailed description of this sorting procedure see the caption of Table 3.3. Average monthly flows for these portfolios are then computed according to Equations (3.16) and (3.17). *Difference* reports the mean difference for flows of P9-10 and flows of P1-2. Panel A reports the results for unconditional sorts. Panel B reports the results for sorts that are conditional on past performance. In this sort, every month, the overall sample of hedge funds is first split into deciles based on the funds' average past return over the last 36 months. Afterwards, each of the ten portfolios is sorted into quintiles based on the individual funds' loading on the funding risk measure. Finally, for each quintile, the ten different past return deciles are merged. Newey-West *t*-statistics are reported in square brackets. ***, **, and * indicate significance at a 1% and 10% level. The sample includes all 8,541 funds in the TASS database and the sample period is January 1994 to May 2015.

	P1-2	P3-4	P5-6	P7-8	P9-10	<i>Difference</i>
Panel A: Unconditional results						
sort on β^{CIPD}	-0.20 [-1.03]	0.01 [0.04]	0.11 [0.53]	0.17 [0.84]	0.24 [1.18]	0.44*** [2.64]
sort on β^{CIPD-}	-0.22 [-1.09]	0.01 [0.05]	0.15 [0.69]	0.14 [0.68]	0.32* [1.80]	0.55*** [2.78]
Panel B: Conditional on past returns						
sort on β^{CIPD}	-0.25 [-1.27]	0.08 [0.41]	0.23 [1.13]	0.11 [0.55]	0.17 [0.84]	0.41** [2.40]
sort on β^{CIPD-}	-0.28 [-1.38]	0.08 [0.44]	0.22 [0.97]	0.15 [0.76]	0.21 [1.11]	0.49** [2.52]

funding risk measure and fund flows to hedge funds with a high loading on that measure. For portfolios sorted on CIPD the difference is 0.44% per month and statistically significant at a 1% level (*t*-statistic of 2.64). For portfolios sorted on $CIPD^-$ the difference is 0.55% per month and also statistically significant at a 1% level (*t*-statistic of 2.78).

To ensure that this difference in fund flows is not simply driven by the funds' past returns, I repeat the analysis conditional on the funds' past performance. To do so, I proceed in three steps. First, I split the overall sample of hedge funds into deciles based on their average past return over the last 36 months. Second, for each of the ten portfolios, I form quintiles based on their loading on the funding risk measure. Finally, for each quintile, I merge the ten different past return deciles. This procedure ensures that funds in each quintile have comparable past returns. Panel B of Table 3.7 shows the results for this conditional sort. As we can see from the table, forming quintiles conditional on past returns lowers the economical and statistical significance of the result marginally. For portfolios sorted on CIPD the difference in flows drops to 0.41% per month (*t*-statistic of 2.40). For portfolios

sorted on $CIPD^-$ the difference drops to 0.49% per month (t -statistic of 2.52). Overall, this test confirms that the difference in fund flows for funds with a different loading on funding risk is not simply driven by a difference in past returns.

3.6.3 Fund-Specific Funding Risk

The third model prediction is that the difference between hedge funds with a different sensitivity to funding shocks is less pronounced for funds whose liabilities are less exposed to funding shocks. To investigate this model prediction, I consider three different proxies for the riskiness a fund's liabilities, repeating my main analysis for different subsamples of the hedge fund database. The first proxy is the fund's redemption notice period, the second proxy is whether a fund has a lockup provision, and the third proxy is the number of prime brokers used by different hedge funds. To ensure a sufficient number of funds in each quantile, I follow Teo (2011) and form quintile portfolios instead of decile portfolios. I run the analysis for funds sorted based on $CIPD^-$, which is closer to my theory than sorting on $CIPD$.¹⁹

First, I divide the sample based on the funds' redemption notice period. The first subsample consists of funds with favorable redemption terms, which have a redemption notice period of one month or less (recall from Table 3.1 that the median redemption notice period is one month). The second subsample consists of funds with less-favorable redemption terms, which have a redemption notice period of more than one month. Panels (a) and (b) of Figure 3.4 show the results for the two subsamples. Hedge funds with less favorable redemption terms and a high loading on $CIPD^-$ (portfolio 1) are still able to generate positive risk-adjusted returns, while hedge funds with less-favorable redemption terms and a high loading on $CIPD^-$ are generating negative risk-adjusted returns. The lower the loading on $CIPD^-$ becomes, the smaller the difference between the two subsets of funds.²⁰ The first two rows of Panel B in Table 3.6 provide additional details and the exact numbers for the difference portfolio.

Second, I split the sample into funds with a lockup provisions and funds without a lockup provision. A lockup provision requires that all new capital invested in the fund cannot be withdrawn before a pre-specified period (typically one year). Funds with a lockup provision are therefore less susceptible to equity withdrawals and therefore the effect of a higher loading on $CIPD^-$ should be less pronounced. Panels (c) and (d) of Figure 3.4 show the results for these two subsamples.²¹ Funds with a lockup provision and with a low loading on

¹⁹Using $CIPD$ instead of $CIPD^-$ leaves the results almost unchanged.

²⁰Note that funds with less-favorable redemption terms overall generate higher risk-adjusted returns, which is in line with the findings of Aragon (2007).

²¹One drawback of this split is that the sample of funds with lockup provision is smaller than the sample

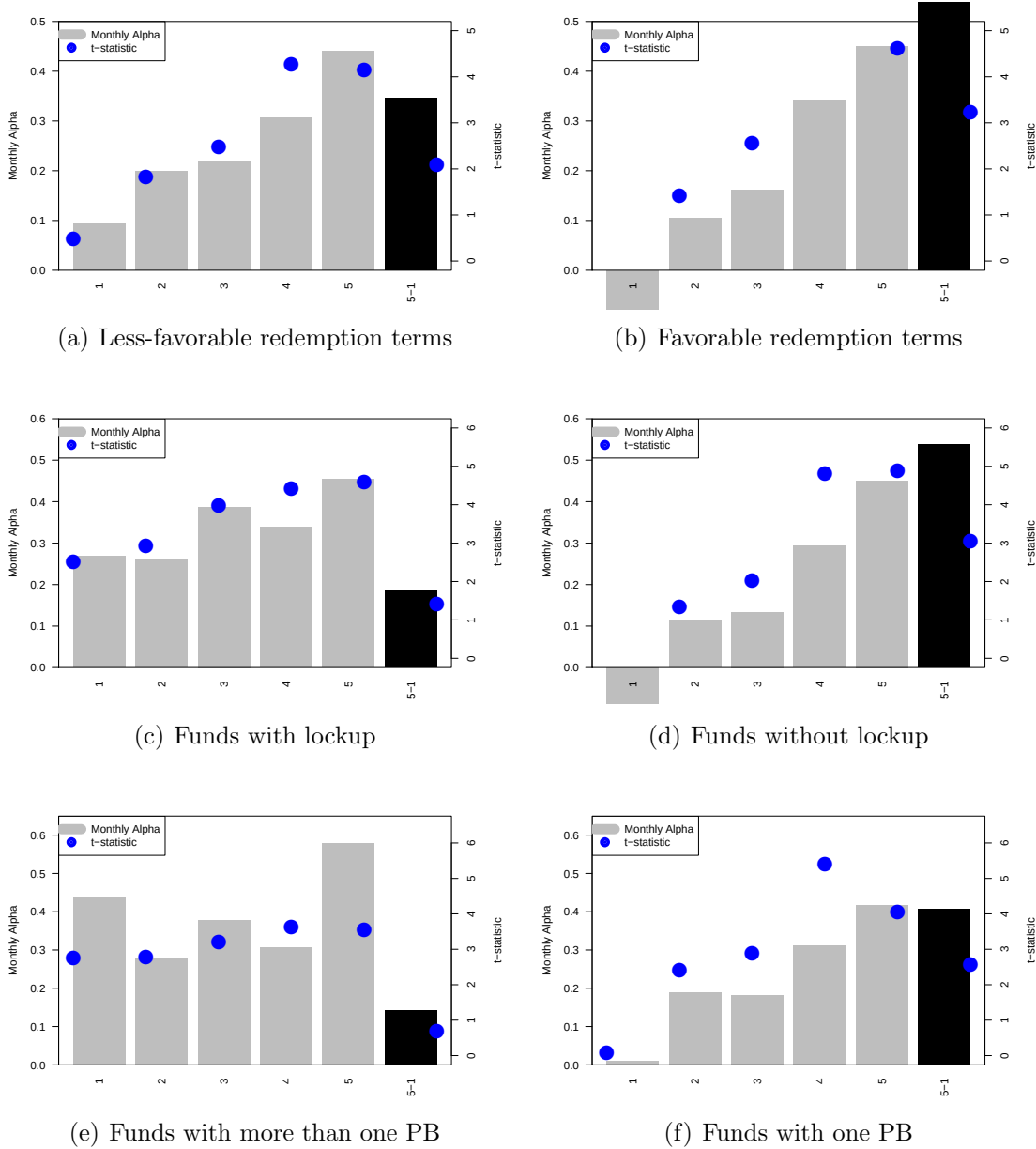


Figure 3.4: Results for different subsamples of CIPD⁻-sorted hedge fund portfolios
This figure presents the results of applying the sorting procedure described in the caption of Figure for different subsamples of the database using CIPD⁻ as sorting variable. Panels (a) and (b) compare the results for hedge funds with redemption notice periods of more than one month (less-favorable redemption terms) and funds with redemption notice period less than one month (favorable redemption terms). Panels (c) and (d) compare the results for hedge funds with lockup provision and funds without lockup provision. Panels (e) and (f) compare the results for hedge funds with more than one prime broker and hedge funds with only one prime broker. The sample period is January 1994 to May 2015, including all 8,541 hedge funds from the TASS database.

CIPD⁻ are still able to generate positive risk-adjusted returns while funds without a lockup provision and with a high loading on CIPD⁻ are generating a negative alpha. Most notably, the difference portfolio generates almost three times higher returns for funds without a lockup provision compared to funds with a lockup provision. The third and fourth row of Panel B in Table 3.6 provide additional details and exact parameter estimates.

Third, I also repeated my analysis splitting the sample into hedge funds that use only one prime broker (funds facing more funding risk) and hedge funds with more than one prime broker (funds facing less funding risk). The drawback of this split is that the TASS database only provides information on prime brokers for live hedge funds (which are still reporting to the database as of the latest version). Hence, applying this method induces survivorship bias and decreases the number of available funds. To overcome these issues, I combine the prime broker data from the most recent version of the database with data used in Aragon and Strahan (2012).²² The results of these splits are exhibited in Panels (e) and (f) of Figure 3.4. As we can see from the figures, the results are insignificant for funds with more than one prime broker and significant for funds with only one prime broker. The fifth and sixth row of Panel B in Table 3.6 provide additional details, confirming that the difference portfolio earns an almost three times higher risk-adjusted return for funds with only one prime broker, compared to funds with more than one prime broker.

3.6.4 High CIPD⁻-Loading, Low Cash Buffer

I conclude this section by testing whether hedge funds with a higher loading on CIPD⁻ also have a lower cash buffer. To that end, I perform a double-sort on CIPD⁻-loadings and fund flows over the past month, using Equation 3.16 to compute the flow. To ensure a sufficient amount of funds in each portfolio, I form 5 x 5 portfolios. The idea behind this double sort is that hedge funds with a lower cash buffer have a higher return sensitivity to inflows or outflows because they need to sell assets to accommodate outflows and can use unexpected inflows to accommodate margin calls or other unexpected liquidity needs.

Panels (a) and (b) of Figure 3.5 show the results for funds with the highest loading on CIPD⁻ and funds with the lowest loading on CIPD⁻ respectively. As we can see from the graph, low-loading funds generate positive alpha, even when they face outflows. Moreover, the difference in alphas between funds with the lowest flows and funds with the highest flows is insignificant for funds with a low loading on CIPD⁻. In contrast to that, funds with a high loading on CIPD⁻ generate a negative alpha when faced with outflows and the difference between the funds with the lowest past flows and the funds with the highest past flows is

of funds without lockup provision (recall that only 19% of the funds in the sample have a lockup provision).

²²I am grateful to George Aragon for providing me with these data.

positive and highly significant.

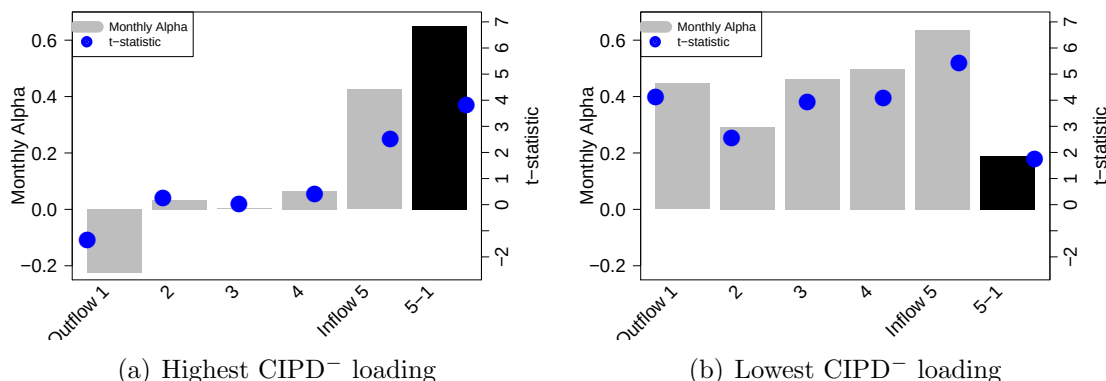


Figure 3.5: Risk-adjusted returns of past-flow sorted hedge fund portfolios. Each month hedge funds are sorted into 25 equally-weighted portfolios according to their historical beta to the $CIPD^-$ measure, constructed in Section 3.4.2, and their flows in the previous month. The figure shows the results for hedge funds with the highest loading on $CIPD^-$ (Panel (a)) and funds with the lowest loading on $CIPD^-$ (Panel (b)). For a detailed description of the sorting procedure as well as the computation of risk-adjusted returns see the caption of Figure 3.4. The grey bars represent monthly risk-adjusted portfolio returns, calculated using the Fung and Hsieh (2004) seven-factor model, where the YLD and BAA factors are replaced by factor-mimicking tradable portfolios. The blue dots are Newey-West t -statistics of the respective risk-adjusted returns. The black bar displays the risk-adjusted return of the difference portfolio, which is long hedge funds with the highest past flows and short hedge funds with the lowest past flows. The sample period is January 1994 to May 2015, including all 8,541 hedge funds from the TASS database.

3.7 Robustness Checks

In this section, I test the robustness of my main result. In Section 3.7.1, I investigate whether few crisis episodes are responsible for the difference in returns between funds with a high loading on $CIPD^-$ and funds with a low loading on $CIPD^-$. In Section 3.7.2, I address common biases in the hedge fund database and show that my main result is robust to these biases.

3.7.1 Robustness to Removing Major Crisis Episodes

To check whether my results are only driven by few major crisis episodes, I split the sample into periods of crisis and normal times. I use two different splits. First, I use anecdotal evidence about crisis periods to identify 19 months which are plausibly periods with severe deteriorations in funding conditions for hedge fund managers. The crisis periods are

August-September 1998 (the period of the Russian debt crisis and the LTCM bailout), August-October 2007 (the months of the quant crisis), August 2008 - March 2009 (the time around the default of Lehman Brothers), August - December 2011 (the first part of the European debt crisis), and April - May 2012 (the second part of the European debt crisis). Second, I classify NBER recession periods as crisis periods and the remaining periods as non-recession periods.

As we can see from Panel (I) of Table 3.8, the risk-adjusted returns of the difference portfolio that is long hedge funds with a low loading on CIPD and short hedge funds with a high loading on CIPD generates a large alpha of 1.10 during the 19 crisis months and a significant, positive alpha of 0.45 during normal times. Similarly, the difference portfolio where hedge funds are sorted on $CIPD^-$, generates an alpha of 1.28 in crisis periods and 0.53 in normal times. Note that the alpha during normal times only shows a minor drop of 0.05 for both difference portfolios when compared to the alpha for the entire sample period. Panel (II) shows that removing NBER recession periods also leaves the main result unchanged. The alpha of the difference portfolio for hedge funds sorted on CIPD as well as for hedge funds sorted on $CIPD^-$ is above 0.50 during normal times and significant at a 5% level (t -statistic of 2.29 and 2.44 respectively). Hence, the findings in Table 3.8 indicate that the results remain intact, even after removing severe crisis episodes.

Table 3.8: Crisis versus noncrisis periods. This table shows the risk-adjusted returns of the difference portfolio which is long portfolio 10 and short portfolio 1 for crises and non-crises periods. The decile portfolios are formed based on the individual funds' loading on CIPD and $CIPD^-$, respectively. See the caption of Table 3.3 for a description of the sorting procedure and for the risk-adjustment. Under (I), anecdotal evidence is used to classify crisis periods. The following 19 months form the crises periods: August-September 1998, August-October 2007, August 2008 - March 2009, August 2011 - January 2012. The remaining 200 months form the quiet period. Under (II) crisis periods are defined as NBER recession periods, which are March 2001 – November 2001 and December 2007 – June 2009. The remaining periods are quiet periods. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	(I) Anecdotal		(II) NBER Recession	
	Crises	Normal	Crises	Normal
sort on β^{CIPD}	1.10*	0.45**	0.46	0.52**
	[1.74]	[2.09]	[0.90]	[2.29]
sort on β^{CIPD^-}	1.28	0.53**	0.87	0.55**
	[1.40]	[2.46]	[1.61]	[2.44]
Number of Observations	19	201	28	192

3.7.2 Biases in Reported Hedge Fund Data

I now discuss four of the most common biases in hedge fund data and show that my results are robust to them. The four biases are the following: backfill bias, dropout bias, return smoothing, and double counting. Backfill bias arises because once a hedge fund starts reporting to the TASS database, it is allowed to enter past returns to the database as well. Clearly, only funds with high past returns would use that option which biases returns upward. Dropout bias arises because hedge funds can choose to stop reporting to the database if they perform poorly. Return smoothing arises because hedge funds investing in illiquid securities might report returns from investments in month t only in month $t+1$ since prices move infrequently (see Asness, Krail, and Liew, 2001 and Getmansky et al., 2004). Double counting could occur because the returns of different subsidiaries of the same fund could be reported as different entities in the database. Bali et al. (2014) document that approximately 16% of the funds in the TASS database are duplicates.

To clean the database for these biases, I proceed in four steps. First, to address double counting, I compute the pairwise correlation between the returns of all funds in the database that have at least 10 observations in common. I truncate the returns of all funds at 20% and -20% to avoid dropping funds that are strongly correlated due to a common jump in their returns. I then drop all funds with a return correlation above 99%. Doing so leads to a drop of 14% of the observations in the database (from 8,541 funds to 7,348), which is similar to the 16% mentioned by Bali et al. (2014).

Second, to address concerns about return smoothing, I use the un-smoothing technique proposed by Getmansky et al. (2004). Let $R_{i,t}^o$ denote the observed return of Fund i at time t and $R_{i,t}$ the true return of Fund i at time t . Then, assuming that return-smoothing does not exceed more than two periods, observed returns and true returns are linked by the following equation:

$$R_{i,t}^o = \theta_{i,0}R_{i,t} + \theta_{i,1}R_{i,t-1} + \theta_{i,2}R_{i,t-2}, \quad (3.18)$$

where $\sum_{k=0}^2 \theta_{i,k} = 1$. For each Fund i , the parameters $\theta_{i,k}$ ($k = 0, 1, 2$) are estimated using the entire time series of observed returns.²³ I then replace the observed returns with the estimated un-smoothed returns and compute the risk-adjusted un-smoothed returns of the 10 decile portfolios.

Third, to address backfill bias, I utilize the information available in the TASS database and drop returns that have been reported prior the fund's inception date to the database.

²³The estimation procedure is based on maximum likelihood, similar to estimating a moving-average model, assuming that demeaned returns are normally distributed.

As summarized in Table 3.1, on average 43% of hedge fund returns are backfilled. Therefore, dropping all backfilled observations could significantly change the results.²⁴

Finally, when addressing dropout bias it is important to distinguish survivorship bias from dropout bias (Aiken et al., 2013). While concerns about survivorship bias can be mitigated by using both hedge funds that are currently reporting to the database and funds that have stopped reporting to the database (which I do in my analysis), dropout bias arises because poorly-performing hedge funds can choose to stop reporting to the database. Using a proprietary dataset of hedge funds, not reporting to any database, Aiken et al. (2013) document that hedge funds that stop reporting to the database continue to exist but deliver returns that are, on average 0.5% lower than the returns of funds that continue reporting to the database. To address this concern, I add a delisting return 0.00% (which is 0.58% lower than the average fund return) after a fund stops reporting to the database.²⁵

Figure 3.6 shows the results for applying the sorting procedure described in Section 3.5 to this modified dataset, where Panel (a) reports the results for CIPD-sorted portfolios and Panel (b) reports the results for CIPD⁻-sorted portfolios. As we can see from the figure, the risk-adjusted returns of the ten individual portfolios drop sharply. However, the main result remains unchanged: hedge funds with a lower exposure to funding risk generate significantly higher returns than hedge funds with a higher exposure to funding risk. As reported in Panel C of Table 3.14, the difference portfolio which is long hedge funds with the lowest loading on CIPD (CIPD⁻) and short hedge funds with the highest loading on CIPD (CIPD⁻) is generating a risk-adjusted return of 0.64% (0.67%) with a *t*-statistic of 2.29 (2.31).

3.8 Making Money on the Strategy?

I now investigate whether it is possible to make money on the high funding risk, low return finding and whether the gains of using CIPD loadings as an investment criterion for hedge fund investments are better than looking into past returns or past risk-adjusted returns. To that end, because it is not possible to short hedge funds, I present the returns of investing in the long leg of three different investment strategies. For each of the three different strategies, hedge funds are sorted into decile portfolios based on their returns over the past 36 months, using (i) raw returns and investing in the portfolio with the highest past returns, (ii) using Fung-Hsieh seven factor alphas and investing in the portfolios with the highest past

²⁴Note that dropping all backfilled information is a conservative approach since hedge funds that start reporting to the TASS database might already have reported to other databases. Hence, not all backfilled observations classified as backfilled by my method are “truly” backfilled.

²⁵I also experimented with more negative dropout returns such as -25% and -50% which decreased the alphas of all ten portfolios sharply, but left the result for the difference portfolio unchanged.

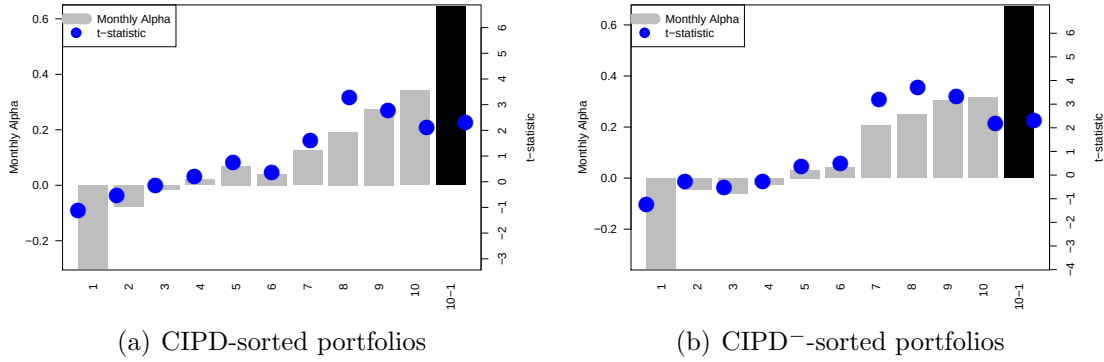


Figure 3.6: Robustness to common biases in reported hedge fund data. This figure presents the results of applying the procedure described in Figure 3.4 to a modification of the hedge fund dataset, adjusted for four common biases in reported hedge fund data. The four adjustments are: dropping all backfilled returns (returns before the fund was added to the TASS database), adding a delisting return of 0.00% after the last reported return for each hedge fund that dropped out of the database, removing hedge funds which are classified as duplicates by the algorithm described in Section 3.7, and using the un-smoothing procedure described in Getmansky et al. (2004). Panel (a) shows the results for CIPD-sorted portfolios and Panel (b) shows the results for CIPD⁺-sorted portfolios. The sample period is January 1994 to May 2015.

alpha, and (iii) using CIPD⁺ and investing in the portfolio with the lowest CIPD⁺-loading. Additionally to reporting the results for monthly rebalancing, I also report the returns of a portfolio that is rebalanced on an annual basis at the beginning of each year. This is because monthly rebalancing of a hedge fund portfolio might not be feasible in practice.

The first four columns of Table 3.9 show that the portfolios sorted based on past returns and past risk-adjusted returns generate similar returns as the portfolio sorted on CIPD⁺ if monthly rebalancing is used. However, on an annual rebalancing frequency, the CIPD⁺-sorted portfolios generate higher, more significant returns than the portfolios sorted based on past returns. Moreover, in contrast to the results for past-return-sorted portfolios, the returns of CIPD⁺-sorted portfolios remain virtually unchanged for a longer holding period. As we can see from the table, a portfolio of hedge funds with the lowest exposure to CIPD⁺ generates an annual risk-adjusted return of 6.84% (t -statistic of 4.69).

To investigate whether these high risk-adjusted returns are robust to the common biases in reported hedge fund returns described in Section 3.7.2, I repeat the analysis described above using the bias-cleaned database constructed in Section 3.7.2. The results for these regressions are reported in the last four columns of Table 3.9. As we can see from the table, both alphas and raw returns of all portfolios drop sharply after addressing these reporting biases. However, the CIPD⁺-sorted hedge fund portfolio still generates a risk-adjusted return of 3.60% using annual rebalancing. Hence, considering past CIPD⁺ loadings for

Table 3.9: Returns from different long-only strategies. This table shows the raw returns and risk-adjusted returns of three different hedge fund portfolios. In each of the three rows, hedge funds are sorted based on their return characteristics over the past 36 months and the returns of the decile portfolio with the highest expected returns are reported. Under past returns, hedge funds are sorted into deciles based on their past return. Under past alpha, hedge funds are sorted based on their Fung-Hsieh seven-factor alpha, computed over the past 36 months. Under beta CIPD⁻, hedge funds are sorted based on their past loading on CIPD⁻. Columns 1–4 report the returns and risk-adjusted returns for monthly rebalancing and annual rebalancing (every January), using the raw hedge fund database. Columns 5–8 report the returns and risk-adjusted returns for monthly rebalancing and annual rebalancing (every January), using the bias-cleaned database, following the four steps described in the caption of Figure 3.6. Newey-West *t*-statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively. The sample period is January 1994 – May 2015.

	1 month (raw)		12 months (raw)		1 month (cln)		12 months (cln)	
	R^{Exc}	α^{FH}	R^{Exc}	α^{FH}	R^{Exc}	α^{FH}	R^{Exc}	α^{FH}
Past Return	0.82*** [3.39]	0.45*** [2.61]	0.46 [1.59]	0.15 [0.83]	0.56* [1.86]	0.08 [0.38]	0.32 [0.98]	-0.09 [-0.46]
Past Alpha	0.84*** [3.38]	0.57*** [4.16]	0.46* [1.75]	0.27* [1.87]	0.54* [1.92]	0.17 [1.05]	0.26 [0.90]	-0.01 [-0.08]
beta CIPD-	0.66*** [4.42]	0.52*** [3.61]	0.72*** [4.36]	0.57*** [4.69]	0.51*** [3.11]	0.30** [2.09]	0.51*** [2.77]	0.30** [2.36]

picking hedge fund investments is more valuable than simply looking into past returns or past alphas.

3.9 Conclusion

The main finding of this paper is that hedge funds with a higher exposure to funding risk, as proxied by a higher loading on a simple funding risk measure, underperform hedge funds with a lower exposure to that risk. This finding is surprising upon initial examination because it contrasts with a basic principle: higher risk should correspond to higher (expected) returns. Although this rule may hold for traded assets, it can be violated for hedge funds, which are actively managed portfolios whose returns depend on a manager's skill and proper risk management. The results in this paper point toward a situation in which more risk taking indicates less managerial skill, which lowers expected returns rather than increasing them.

To formalize this explanation, I develop a simple model that illustrates how a higher exposure to a common funding shock can lead to lower subsequent returns. The model delivers three testable predictions that are supported by the data. First, managers with a high loading on CIPD underperform. This underperformance is even more severe when

sorting only on deteriorating funding conditions, as proxied by $CIPD^-$, and insignificant when sorting on $CIPD^+$. Second, in the model, the higher exposure to funding shocks enables investors to infer the quality of the manager and cause them to withdraw their investments. In line with this prediction, I find that hedge funds with a higher loading on CIPD experience significantly lower flows than fund with a lower loading on CIPD. Finally, the difference in returns between low-loading and high-loading funds is smaller for funds that have less favorable redemption terms and relationships with more than one prime broker.

3.10 Appendix

Data Description

This appendix provides additional details about the data used for my analysis.

1. **BAB factor:** This is the betting against beta factor described in Frazzini and Pedersen (2014). The data are obtained from Lasse Pedersen's website: <http://www.lhpedersen.com/data>.
2. **Commodity risk:** The commodity risk factor is constructed using the returns of the S&P GSCI index over the one-month risk-free rate. Data for this index comes from datastream.
3. **Currency risk factors:** These factors capture currency returns of an U.S. dollar investor and the returns of a carry trader. The data are obtained from Adrien Verdelhan's website: <http://web.mit.edu/adrienv/www/Data.html>
4. **Dealer broker leverage:** This variable captures the leverage of U.S. broker-dealers and is described in more detail in Adrian et al. (2014). Until Q4 2009, data on this variable are obtained from Tyler Muir's website. Since the data ends in Q4 2009, I use the financial accounts of the U.S. data, following the procedure described in Adrian et al. (2014) to supplement the time series with more recent observations for the Q1 2010 – Q4 2015 period.
5. **Emerging markets risk:** The emerging markets risk factor is constructed using the returns of the MSCI emerging market index over the one-month risk-free rate. Data for this index comes from datastream.
6. **Fixed income risk factors:** To construct the first tradable factor (YLD), I take the difference between the Merrill Lynch treasury bond index with 7-10 years to maturity

over the 1-month risk-free rate. For the second factor (BAA), I use the difference between the Merrill Lynch corporate bond index with BBB-rated bonds and 7-10 years to maturity over the treasury bond index. The data on the two bond indices are obtained from the Bloomberg system, the one-month risk-free rate is obtained from Kenneth French's website.

7. **FX liquidity measure:** The measure is the one developed in Karnaukh et al. (2015) and represents an equally-weighted index, measuring the liquidity of U.S. dollar exchange rate for developed countries. The measure combines information from relative bid-ask spreads and high-low currency prices. The data are available at <http://rfs.oxfordjournals.org/content/early/2015/05/12/rfs.hhv029/suppl/DC1>. I use the procedure described by the authors to update the time series to May 2015.
8. **Investment bank stock returns:** I follow Ang et al. (2011) and use the stock returns of the 9 largest investment banks, which are: Bear Stearns, Citibank, Credit Suisse, Goldman Sachs, HSBC, JP Morgan, Lehman Brothers, Merrill Lynch, and Morgan Stanley. These returns are obtained from the Bloomberg system.
9. **Noise measure:** This is the noise measure developed by Hu et al. (2013). The data are obtained from Jun Pan's website: <http://www.mit.edu/~junpan/>.
10. **PS liquidity factor:** This is the Pastor and Stambaugh (2003) stock market liquidity factor, obtained from Lubos Pastor's website: <http://faculty.chicagobooth.edu/lubos.pastor/research/>.
11. **TED spread:** The treasury eurodollar spread is the difference between the 3-month U.S. Libor rate and the 3-month U.S. treasury rate. Both rates are obtained from the Bloomberg system.
12. **Trend following factors:** The three Fung and Hsieh trend-following are capturing returns from trend followers in the bonds, currency, and commodities markets. The factors are obtained from David Hsieh's website: <https://faculty.fuqua.duke.edu/~dah7/HFData.htm>.
13. **U.S. stock market returns:** The first stock market risk factor (MKT) is the monthly return of the CRSP market portfolio in excess of the one-month treasury yield. The second stock market risk factor (SMB) is the difference of returns between small and big stocks (SMB). A third, additional stock market risk factor (UMD) is the momentum factor that is long stocks with high past returns and short stocks with low past returns (UMD). Data on all three factors are obtained from Kenneth French's website.

14. **VIX index:** Is the implied volatility of the S&P 500 index and data on VIX are obtained from the Bloomberg System.

3.10.1 Additional Results

This appendix presents additional details and new results that have been omitted in the main part of the paper. Section 3.10.2 provides additional details that complement the analysis in the main part of the paper. Section 3.10.3 provides an additional test that sheds more light on the relationship between CIPD and the noise measure developed by Hu et al. (2013). The section shows that combining both measures leads to even stronger results than using any of them separately. Finally, Section 3.10.4 shows that the main result, that a higher loading on a funding risk measure leads to lower returns, is robust to using different funding risk measures.

3.10.2 Additional Details

Table 3.10 shows yearly summary statistics for the returns of all hedge funds in the sample. As we can see from the table, the years 2008 and 2011 have been especially bad years for fund managers, with negative average returns. Table 3.11 shows the loadings on all seven Fung-Hsieh factors for the CIPD-sorted decile portfolios and the difference portfolio. Row (1) in Panel A of Table 3.12 provides more details and the exact parameter estimates for style-neutral sorts that supplement the results exhibited in Panel (b) of Figure 3.4. As we can see from the table, the results are almost unchanged by fixing the allocation to styles among the decile portfolios. Figure 3.7 shows the results for two additional tests. Panel (a) shows the results for the past-return-neutral sorts described in the main part of the paper. Row (2) in Panel A of Table 3.12 provides additional details and the exact parameter estimates for this test. Panel (b) of Figure 3.7 shows the results for a different subsample of hedge funds, removing funds that reportedly invest in FX markets. The third row (3) in Panel A of Table 3.12 shows additional details and the exact parameter estimates. Overall, the main result is robust to these modifications.

Table 3.13 presents an overview of the characteristics of the funds in the different CIP-beta-sorted decile portfolios. As we can see from the table, funds in the top and bottom decile have very similar characteristics in terms of their size, age, redemption terms, management fees, as well as in their style allocation. However, the table also shows that funds in the middle portfolios tend to have slightly different characteristics. Most notably, portfolios 3-7 tend to consist of larger funds and consist of more than 30% funds of funds while top and bottom portfolio only consist of approximately 10% funds of funds.

Table 3.14 provides additional details and factor loadings for the difference portfolio that is long hedge funds with a low loading on $CIPD^-$ and short hedge funds with a high loading on $CIPD^-$. As we can see from the table, the raw excess returns of the difference portfolio are positive but insignificant. Controlling for the two stock market risk factors or the two bond market risk factors leads to statistically significant risk-adjusted returns. Furthermore, the significance of the risk-adjusted returns increases, the more factors are added.

Relationship Between $CIPD_t$ and Other Liquidity Measures

In this section, I study the correlation between $CIPD_t$ and other common liquidity measures. The goal of this section is to illustrate that $CIPD_t$ is strongly correlated with other measures of funding risk faced by hedge funds (ΔTED and *Leverage*), while other liquidity measures only show a weak correlation. Panel A of Table 3.15 shows the correlations between seven different liquidity measures. These seven measures are the betting against beta factor proposed by Frazzini and Pedersen (2014), the Pastor and Stambaugh (2003) liquidity factor, changes in the TED spread, the dealer-broker leverage factor by Adrian et al. (2014), changes in the 10-year on-the-run off-the-run spread, changes in the Hu et al. (2013) noise measure, changes in the 10-year on-the-run off-the-run spread, as well as changes in CIP^{Index} and changes in $CIP^{Index,OIS}$, which is an alternative measure of CIP deviations, constructed using OIS rates instead of Libor rates.

In line with the regression results from Section 3.4.2, the table confirms that $CIPD_t$ is strongly correlated with ΔTED_t and *Leverage*_{*t*}, which are the two other proxies for market-wide funding conditions faced by hedge funds. On the other hand, $CIPD_t$ is only weakly correlated with *BAB*_{*t*} and *PS*_{*t*} (correlation weaker than 10%). Note that ΔTED and *Leverage* also show a weak correlation with these two stock-market factors. Furthermore, the correlation between $CIPD_t$ and $\Delta Noise$ is -0.22 , indicating that the two variables are only weakly related. $\Delta Noise_t$ only has a weak correlation of 0.19 with ΔTED and is most strongly correlated to $\Delta On10Yr_t$ (correlation of 0.55).

The table also shows that $CIPD_t^{OIS}$ has similar properties compared to $CIPD_t$. The correlation with ΔTED_t and *Leverage*_{*t*} is -0.74 and -0.81 respectively. As $CIPD_t$, the measure based on OIS rates also has a correlation weaker than 10% with *BAB*_{*t*} and *PS*_{*t*} and is almost uncorrelated to $\Delta Noise_t$. Note that the sample period for the measure based on OIS rates is only from January 2002 to May 2015 while the other sample periods are from January 1994 to May 2015. I show later that using $CIPD_t^{OIS}$ as a proxy for hedge fund funding risk leads to similar results as using $CIPD_t$.

Panel (b) of Table 3.15 shows the correlation between the seven Fung and Hsieh factors as well as their correlation with $CIPD_t$. As we can see from the table, the correlation among

factors is generally low. Most importantly, the correlation between $CIPD_t$ and the seven Fung and Hsieh factors is low, ranging from 0.17 between stock market returns and $CIPD_t$ to -0.17 between the FX trend-following factor and $CIPD_t$.

Cumulative Excess Returns of $CIPD_t$ -sorted Decile Portfolios

To get a better understanding of the decile excess returns, Figure 3.8 plots the time series of cumulative excess returns of the top and bottom decile portfolios for $CIPD$ -sorted portfolios together with CIP^{Index} . As we can see from the figure, the portfolio with the highest loading on $CIPD$ (high funding risk) generates major losses during crises episodes and these losses are partly recovered afterwards. Furthermore, the returns of the high funding risk portfolio are more volatile and generally lower than those of the portfolio with a low loading on $CIPD$. More specifically, the high funding risk portfolio suffers large losses around the LTCM crisis in 1998, around the default of Lehman Brothers in 2008, and during the European debt crisis in 2011/2012. In contrast to that, the low funding risk portfolio provides stable returns during crisis periods, with moderate losses during the 2008 crisis. However, as is clear from the figure, the difference in returns between these two portfolios is not purely driven by these few crises episodes.

3.10.3 Relationship to the Noise Measure

I now show that $CIPD$ is capturing a different risk than the noise measure constructed by Hu et al. (2013), who show that their measure is a priced risk factor in the cross section of hedge fund returns and that a higher loading on that measure implies higher returns. I apply a double-sorting procedure to incorporate the information content of the two measures. In a first step, I repeat the procedure described in Section 3.5 and compute β^{Noise} for each fund in the database using a rolling regression window of 36 months. I then sort hedge funds into quintile portfolios based on their sensitivity to changes in the noise measure. I put funds with the lowest loading on $\Delta Noise_t$ (funds that I expect to perform poorly) in the first portfolio and funds with the highest loading on $\Delta Noise_t$ in the fifth portfolio. Afterwards, I split each of the five noise-sorted portfolios into five $CIPD$ -sorted portfolios, based on β^{CIPD} computed in Section 3.5. Here, I put funds with the highest loading on $CIPD$ (funds that I expect to perform poorly) in the first portfolio and funds with the lowest loading on $CIPD$ in the fifth portfolio.

This conditional double sort results in 25 hedge fund portfolios. The risk-adjusted returns of these 25 portfolios (relative to the Fung Hsieh seven-factor model), as well as the returns of the difference portfolios, are exhibited in Table 3.16. As we can see from the table, the

double sort confirms that the noise measure is a risk factor in the cross-section of hedge fund returns, even conditional on CIPD and when returns are risk-adjusted for the seven Fung and Hsieh factors. Two out of the five difference portfolios generate a significant risk-adjusted return. The table also confirms that CIPD is capturing a different aspect of market conditions than $\Delta Noise$. All five difference portfolios generate a positive and statistically significant risk-adjusted return.

The number in the bottom-right corner of Table 3.16 is the risk-adjusted return of the difference portfolio that is long hedge funds with the highest loading on $\Delta Noise_t$ and the lowest loading on CIPD and short the portfolio with the lowest loading on $\Delta Noise_t$ and the highest loading on CIPD. This portfolio generates a striking risk-adjusted return of 0.99 per month (t -statistic of 5.03). Hence, combining the information content of the noise measure with the information content in CIPD leads to even stronger results than just using any of the two measures separately.

3.10.4 Different Funding Measures

I now repeat my analysis for several alternative funding risk measures. First, I use a different variation of CIPD, where CIP^{Index} is constructed using OIS rates instead of LIBOR. The advantage of using this measure is that OIS rates is that they do not contain a credit-risk component and are not susceptible to manipulations like the LIBOR rates. The drawback is that OIS rates for most currencies are only available from January 2002 on. Hence, using this alternative index leads to a six year shorter sample period. Second, I use the original CIPD but add the FX liquidity proxy, constructed by Karnaukh et al. (2015), as an additional control variable to ensure that my results are not driven by currency market illiquidity. Third, I use changes in the difference between the 3-months U.S. LIBOR and 3-month OIS rate (henceforth LIBOR-OIS spread) instead of CIPD as sorting variable. The advantage of this measure is that it is easy to construct and clearly capturing funding conditions faced by major banks. The drawback is that the time series starts only in 2002 and shows virtually no variation before 2007. Finally, I compute average flows, defined as the average flow of all hedge funds in my sample, and use changes in this measure instead of CIPD to form decile portfolios.

Figure 3.9 shows the results for these four additional tests. As we can see from Panel (a) and (b), using different modifications of CIPD leaves the main result intact: hedge funds with a high loading on CIPD generate lower returns than hedge funds with a low loading on CIPD. Panels (c) and (d) show that qualitatively similar results can be obtained for different funding risk measures. In particular, hedge funds with a strong loading on changes

in the LIBOR-OIS spread underperform hedge funds with a weak loading on changes in the LIBOR-OIS spread. For sorts based on changes in average flows the results are insignificant but qualitatively similar: hedge funds that generate low returns when the average hedge fund experiences outflows generate lower returns than hedge funds that perform well during times of average outflows. In future work, I plan to further investigate the impact of average fund flows on hedge fund performance and, more broadly, on asset prices.

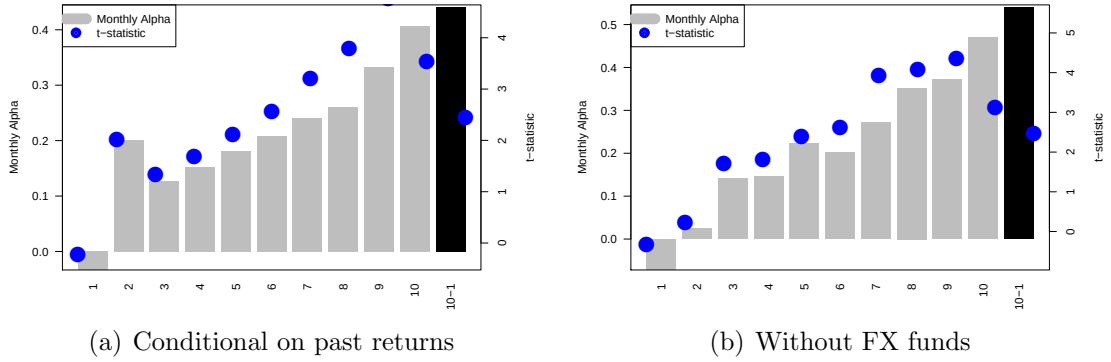


Figure 3.7: Results for different modifications of the CIPD-sort. Each month hedge funds are sorted into 10 equally-weighted portfolios according to their historical beta to CIPD. In panel (a) the sort is performed conditional on past performance. In this sort, every month, the overall sample of hedge funds is first split into deciles based on the funds' average past return over the last 36 months. Afterwards, each of the ten portfolios is sorted into deciles based on the individual funds' loading on the funding risk measure. Finally, for each quintile, the ten different past return deciles are merged. Panel (b) reports the results of an unconditional sort where hedge funds that report that they are investing in FX markets are dropped. For a detailed description of the sorting procedure as well as the computation of risk-adjusted returns see the caption of Figure 3.4. The grey bars represent monthly risk-adjusted portfolio returns, calculated using the Fung and Hsieh (2004) seven-factor model, where the YLD and BAA factors are replaced by factor-mimicking tradable portfolios. The blue dots are Newey-West t -statistics of the respective risk-adjusted returns. The black bar displays the risk-adjusted return of the difference portfolio, which is long hedge funds in Portfolio 10 and short hedge funds in Portfolio 1. The sample period is January 1994 to May 2015, including all 8,541 hedge funds from the TASS database.

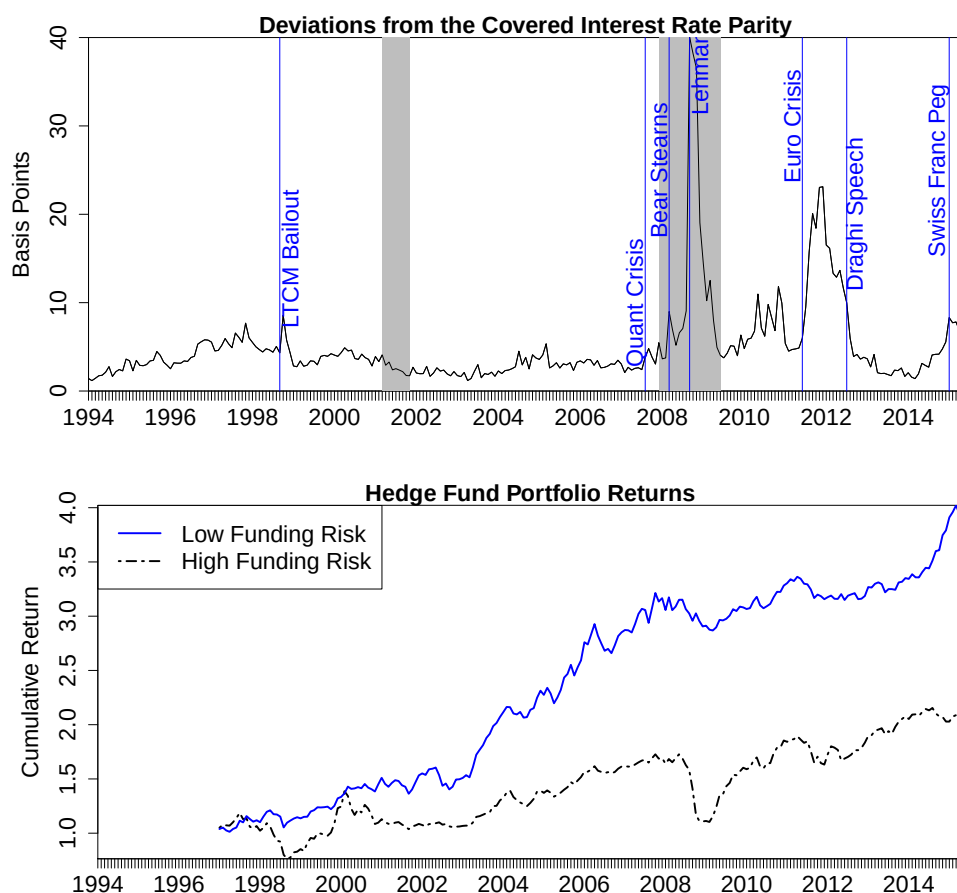


Figure 3.8: Cumulative excess returns from investing in high and low loading funds. This figure shows the cumulative excess returns of hedge funds with a strong loading (solid line) and weak loading (dashed line) on changes in the covered interest rate parity deviation index (ΔCIP_t^D), constructed in Section 3.4.2. See the caption of Figure 3.4 for a description of the sorting procedure. The high (low) loading portfolio is the first (tenth) decile portfolio.

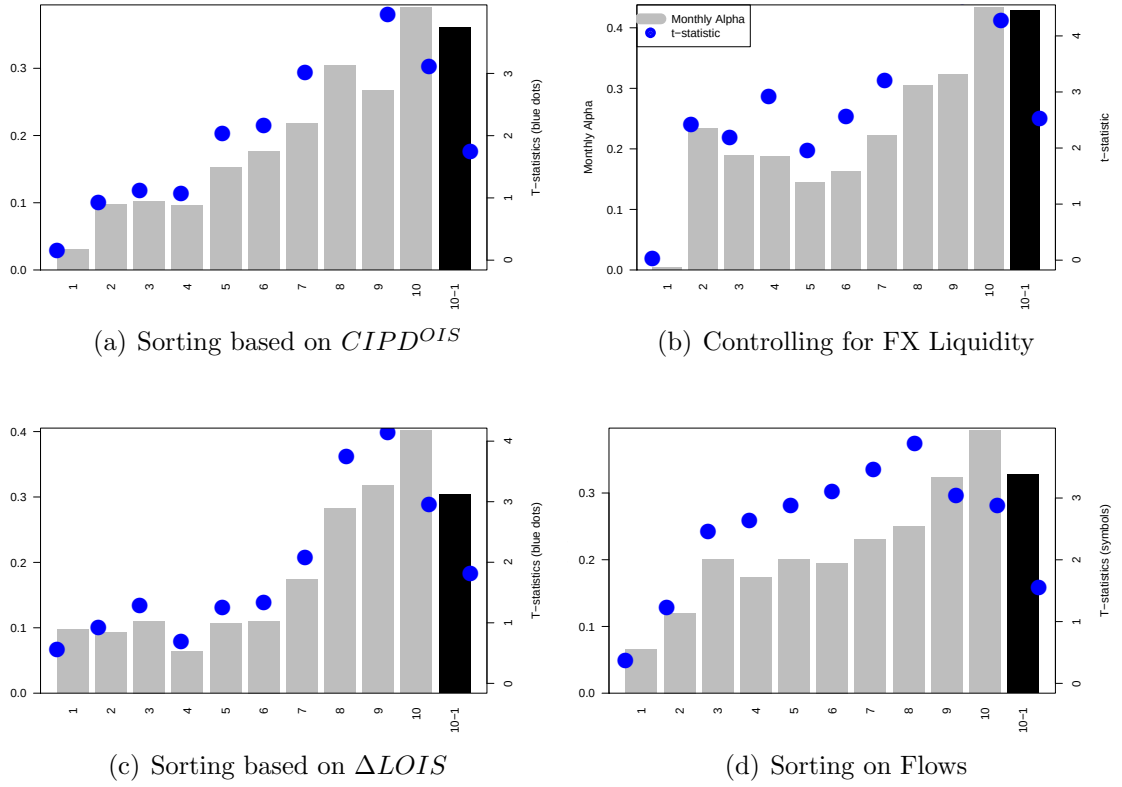


Figure 3.9: Results for different modifications of the funding risk measure. Each month hedge funds are sorted into 10 equally-weighted portfolios according to their historical beta to different modifications of the funding risk measure. Panel (a) shows the results for sorts based on $CIPD^{OIS}$. Panel (b) reports the results, when sorts are performed controlling for the Karnaukh et al. (2015) FX liquidity measure. Panel (c) shows the results for sorts on changes in the difference between the 3-month U.S LIBOR rate and the 3-month U.S. OIS rate. Panel (d) shows the results for sorts on changes in average fund flows. For a detailed description of the sorting procedure as well as the computation of risk-adjusted returns see the caption of Figure 3.4. The grey bars represent monthly risk-adjusted portfolio returns, calculated using the Fung and Hsieh (2004) seven-factor model, where the YLD and BAA factors are replaced by factor-mimicking tradable portfolios. The blue dots are Newey-West t -statistics of the respective risk-adjusted returns. The black bar displays the risk-adjusted return of the difference portfolio, which is long hedge funds in Portfolio 10 and short hedge funds in Portfolio 1. The sample period is January 1994 to May 2015, including all 8,541 hedge funds from the TASS database.

Table 3.10: Hedge fund summary statistics. This table provides summary statistics of average hedge fund returns in the TASS database separately for every year. The sample period is January 1994 to May 2015.

	N	Mean	SD	Min	Meadian	Max
1994	679	0.01	1.66	-10.62	0.05	10.94
1995	877	1.51	1.70	-6.58	1.30	16.80
1996	1,132	1.68	1.47	-3.96	1.44	11.25
1997	1,397	1.52	1.40	-12.34	1.39	11.74
1998	1,647	0.51	2.25	-12.85	0.59	15.66
1999	1,973	2.25	2.88	-9.99	1.60	32.54
2000	2,288	1.01	2.09	-23.07	1.00	23.22
2001	2,715	0.66	1.92	-21.83	0.58	48.43
2002	3,237	0.37	1.39	-17.05	0.29	15.91
2003	3,848	1.38	1.68	-14.47	0.97	40.23
2004	4,553	0.77	0.92	-5.34	0.61	10.98
2005	5,205	0.83	1.19	-9.44	0.64	27.68
2006	5,568	0.99	1.11	-6.14	0.85	23.72
2007	5,860	0.88	1.42	-15.46	0.73	43.38
2008	5,941	-1.27	2.40	-22.23	-1.23	14.95
2009	5,554	1.15	2.38	-100.00	0.88	18.93
2010	5,251	0.67	1.29	-34.55	0.62	26.85
2011	4,967	-0.19	1.25	-23.93	-0.18	10.24
2012	4,453	0.53	1.35	-48.05	0.54	24.97
2013	3,834	0.65	1.41	-20.01	0.63	29.76
2014	3,297	0.31	1.69	-62.16	0.32	34.53
2015	2,718	0.78	1.51	-21.14	0.74	12.93

Table 3.11: Factor loadings for CIPD-sorted portfolios. Hedge funds are sorted into deciles based on their beta to the CIPD measure described in Section 3.4.2. Beta is calculated using a regression of monthly hedge fund returns on CIPD, controlling for the stock market portfolio, and using the 36 months prior to portfolio formation. The seven Fung Hsieh factors are the market excess return (MKT), a size factor (SMB), tradable factors to mimic monthly changes in the 10-year Treasury constant maturity yield (YLD) and monthly changes in the Moody's Baa yield less 10-year Treasury constant maturity yield (BAA), as well as three trend-following factors: BD (bond), FX (currency), and COM (commodity). The sample period is January 1994 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Intercept	β^{CIPD}	β^{Mkt}	β^{SMB}	β^{YLD}	β^{BAA}	β^{BD}	β^{FX}	β^{COM}	R^2
P1	0.00 [0.03]	0.07 [2.02]	0.36 [8.56]	0.25 [3.43]	0.24 [2.98]	0.65 [7.01]	-1.95 [-1.55]	2.19 [2.83]	0.45 [0.42]	0.65
P2	0.15 [1.72]	0.10 [2.13]	0.21 [7.41]	0.12 [3.92]	0.16 [3.28]	0.41 [7.49]	0.17 [0.28]	1.22 [2.53]	0.48 [0.80]	0.64
P3	0.17 [2.51]	0.07 [1.69]	0.18 [9.43]	0.11 [5.41]	0.11 [3.46]	0.33 [8.97]	-0.53 [-0.77]	1.01 [2.92]	0.30 [0.65]	0.70
P4	0.14 [2.05]	0.06 [1.83]	0.15 [7.46]	0.10 [3.97]	0.08 [2.45]	0.29 [6.26]	-0.29 [-0.51]	0.98 [2.98]	0.32 [0.84]	0.66
P5	0.18 [2.70]	0.07 [1.52]	0.15 [8.34]	0.10 [4.97]	0.07 [2.22]	0.22 [4.89]	-0.15 [-0.26]	0.79 [2.39]	0.27 [0.69]	0.62
P6	0.18 [2.29]	0.06 [2.02]	0.15 [7.75]	0.08 [3.70]	0.08 [2.59]	0.24 [5.78]	-0.74 [-0.88]	0.92 [3.11]	0.56 [1.31]	0.60
P7	0.24 [3.11]	0.06 [2.16]	0.12 [6.93]	0.06 [3.03]	0.04 [1.16]	0.19 [4.57]	-0.72 [-0.91]	0.78 [2.54]	0.60 [1.35]	0.53
P8	0.33 [5.33]	0.04 [1.63]	0.17 [8.13]	0.08 [4.73]	0.06 [1.53]	0.13 [3.30]	-0.03 [-0.05]	0.83 [2.28]	0.45 [0.96]	0.56
P9	0.38 [4.88]	0.01 [0.34]	0.20 [6.74]	0.09 [3.51]	0.03 [0.38]	0.08 [1.53]	0.13 [0.20]	1.22 [3.39]	0.45 [0.64]	0.47
P10	0.50 [3.70]	0.02 [0.40]	0.32 [6.48]	0.10 [1.83]	-0.11 [-0.84]	-0.02 [-0.15]	0.53 [0.46]	2.56 [3.94]	1.33 [1.36]	0.41
P10 - P1	0.49 [2.36]	-0.06 [-1.65]	-0.04 [-0.63]	-0.16 [-1.42]	-0.34 [-2.20]	-0.66 [-3.68]	2.47 [1.81]	0.37 [0.42]	0.88 [0.74]	0.30

Table 3.12: Supplementing additional results. Hedge funds are sorted into portfolios based on their beta to the CIPD measure, described in Section 3.4.2, and based on different modifications of CIPD. For a detailed description of the sorting procedure and the different variables see the caption of Table 3.3. Each row reports the results for a difference portfolio. The results for the individual portfolios are omitted for brevity. Panel A reports the results for hedge funds that are sorted into deciles based on their loading on CIPD. In (1) the sorting is conditional on the funds' investment style, in (2) the sorting is conditional on the funds' performance over the past 36 months, in (3) funds that report that they invest in FX markets are dropped. Panel B shows the results for hedge funds that are sorted into deciles based on their loading on different funding risk proxies. In (1) hedge funds are sorted into deciles according to their loading on $CIPD^{OIS}$, a modified version of CIPD that is constructed using OIS rates instead of LIBOR rates. In (2) hedge funds are sorted into CIPD-portfolios, controlling for the Karnaukh et al. (2015) FX liquidity measure. In (3) hedge funds are sorted based on their loading on changes in the difference between the 3-month U.S. LIBOR rate and the 3-month U.S. OIS rate. In (4) hedge funds are sorted based on changes in aggregate hedge fund flows. The sample period is January 1994 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Post-sorting					Pre-sorting	
	α^{FH}	α^{Add}	β^{Mkt}	β^{CIPD}	R_{FH}^2	β^{Mkt}	β^{CIPD}
Panel A: Additional results							
Style neutral	0.38** [2.33]	0.39*** [2.60]	-0.16*** [-2.82]	-0.15*** [-2.87]	0.31	0.03 [0.42]	-2.00*** [-6.99]
Past return neutral	0.44** [2.44]	0.42** [2.42]	-0.19*** [-2.64]	-0.19*** [-3.58]	0.33	0.00 [-0.02]	-1.93*** [-6.83]
Without FX investors	0.54** [2.48]	0.53** [2.40]	-0.19** [-2.50]	-0.21*** [-3.79]	0.25	0.06 [0.74]	-2.31*** [-6.66]
Panel B: Results for different robustness checks							
(1) $CIPD^{OIS}$	0.36* [1.74]	0.23 [1.30]	-0.23*** [-3.29]	-0.04 [-1.29]	0.27	-0.09 [-1.24]	-1.49*** [-4.47]
(2) FX Liquidity	0.44** [2.44]	0.42** [2.42]	-0.19*** [-2.64]	-0.19*** [-3.58]	0.33	0.00 [-0.02]	-1.93*** [-6.83]
(3) $\Delta LOIS$	0.31* [1.82]	0.14 [0.90]	-0.34*** [-8.44]	-0.03 [-0.87]	0.48	-0.28*** [-6.67]	-30.82*** [-5.66]
(4) Flows	0.33 [1.54]	0.19 [1.23]	-0.36*** [-6.19]	-0.12 [-0.41]	0.39	-0.29*** [-4.17]	-3.84*** [-12.46]

Table 3.13: Characteristics of the CIP-deviation-sorted hedge fund portfolios. This table reports the average characteristics and average allocations within hedge fund style for the 10 CIP-beta-sorted portfolios from Table 3.3. See Table 3.1 for a description of the different variables.

	P1	P2	P3	P4	P5	P6	P7	P8	P9	P10
Panel A: Characteristics										
AUM (mio USD)	260.31	395.76	467.86	566.03	396.32	328.82	357.69	391.96	396.62	343.16
Reporting (months)	138.55	140.24	140.58	141.19	139.23	138.08	138.88	136.73	135.17	139.59
Age (months)	87.04	89.15	89.46	87.86	87.68	86.37	85.28	85.01	85.51	87.90
Backfilled	0.26	0.28	0.30	0.30	0.29	0.29	0.30	0.31	0.29	0.29
Lockup?	0.24	0.24	0.23	0.20	0.19	0.19	0.20	0.21	0.23	0.25
Notice (Months)	1.02	1.15	1.20	1.26	1.28	1.19	1.16	1.10	1.05	1.00
Management Fee	1.52	1.44	1.39	1.37	1.38	1.37	1.34	1.36	1.39	1.47
Incentive Fee	17.30	16.54	15.35	15.03	14.47	14.66	15.55	16.60	17.60	18.34
Panel B: Allocation within hedge fund style (%)										
Convertible Arbitrage	2.12	2.85	3.90	3.13	2.98	3.25	3.63	2.80	1.60	1.41
Emerging Markets	14.10	8.88	5.17	3.59	3.02	2.41	2.71	3.79	5.83	9.45
Equity Market Neutral	2.01	2.68	3.00	3.55	2.86	3.28	3.99	4.36	5.43	4.53
Event Driven	4.34	5.91	7.47	7.98	8.75	9.98	10.66	11.12	9.09	3.98
Fixed Income Arbitrage	4.10	3.95	3.28	3.49	3.06	3.13	3.72	3.17	3.38	1.71
Fund of Funds	10.77	23.34	36.58	43.53	46.51	43.87	37.12	29.52	19.18	10.00
Global Macro	5.47	4.00	2.95	2.47	2.39	2.35	2.57	3.25	3.81	5.40
Long Short Equity	35.29	30.21	23.37	18.06	16.72	15.13	16.49	22.62	31.38	40.24
Managed Futures	11.88	7.17	4.34	3.90	3.86	3.96	4.41	5.58	8.67	15.76
Multi-Strategy	5.99	7.62	6.75	7.65	7.63	10.00	12.12	11.62	8.44	4.69
Other	3.93	3.38	3.19	2.64	2.23	2.65	2.59	2.17	3.19	2.82

Table 3.14: Factor loadings and alphas for the CIPD⁻-sorted difference portfolio. Hedge funds are sorted into portfolios based on their beta to the negative part of the CIPD measure, described in Section 3.4.2. For a detailed description of the sorting procedure see the caption of Table 3.3. The table reports the results of regressing the returns of the difference portfolio – which is long hedge funds with a low loading on CIPD⁻ and short hedge funds with a high loading on CIPD⁻ – on the indicated variables. The independent variables are the excess returns of the U.S. stock market portfolio (MKT), a size factor (SMB), tradable factors to mimic monthly changes in the 10-year Treasury constant maturity yield (YLD) and monthly changes in the Moody’s Baa yield less 10-year Treasury constant maturity yield (BAA), the three Fung and Hsieh trend-following factors: BD (bond), FX (currency), and COM (commodity), excess returns of the MSCI Emerging Market Index (EM), excess returns of the S&P GSCI Commodity Index (GSCI), and the two currency risk factors proposed by Lustig et al. (2011) (Cncy US and Cncy Carry). The sample period is January 1994 to May 2015. Newey-West *t*-statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	(1)	(2)	(3)	(4)	(5)	(6)
Alpha	0.36 [1.31]	0.56** [2.37]	0.46** [2.17]	0.60*** [2.74]	0.62*** [2.78]	0.63*** [3.04]
Mkt		-0.28*** [-3.59]		-0.16*** [-3.18]	-0.16*** [-2.96]	-0.05 [-0.62]
SMB		-0.18 [-1.56]		-0.18 [-1.48]	-0.18 [-1.52]	-0.12 [-1.37]
YLD			-0.21 [-1.65]	-0.25* [-1.81]	-0.27** [-2.01]	-0.26** [-2.25]
BAA			-0.91*** [-4.95]	-0.65*** [-3.40]	-0.61*** [-3.17]	-0.57*** [-3.64]
BD					0.93 [0.75]	0.13 [0.13]
FX					0.94 [1.27]	0.97 [1.25]
COM					-0.17 [-0.13]	0.44 [0.34]
UMD						-0.12** [-2.44]
EM						-0.17*** [-2.91]
GSCI						-0.02 [-0.80]
Cncy US						0.22** [2.50]
Cncy Carry						-0.01 [-0.11]
Adj. R^2	0	0.27	0.28	0.37	0.37	0.45

Table 3.15: Correlation between $CIPD_t$ and other variables. Panel A shows the correlation between $CIPD_t$ as well as $CIPD_t^{OIS}$ and other common liquidity measures. The other measures are the betting against beta factor (BAB_t) constructed in Frazzini and Pedersen (2014), the Pastor and Stambaugh (2003) stock market liquidity factor (PS_t), changes in the treasury-eurodollar spread (ΔTED_t), the dealer-broker leverage factor suggested by Adrian et al. (2014) ($Leverage_t$), changes in the 10-year on-the-run off-the-run spread ($\Delta On10Yr_t$), and changes in the Hu et al. (2013) noise measure ($\Delta Noise_t$). Panel B shows the correlation matrix of the 7 Fung Hsieh hedge fund risk factors with $CIPD_t$. The 7 risk factors are the market excess return (MKT), a size factor (SMB), changes in the ten-year Treasury constant maturity yield (YLD), changes in the Moody's Baa yield less ten-year Treasury constant maturity yield (BAA), as well as three trend-following factors: BD (bond), FX (currency), and COM (commodity). The sample period is January 1994 to May 2015, all observations are month-end.

Panel A: Correlation between $CIPD_t$ and other liquidity measures							
	BAB_t	PS_t	ΔTED_t	$Leverage_t$	$\Delta On10Yr_t$	$\Delta Noise_t$	$CIPD_t$
PS_t	0.06						
ΔTED_t	-0.06	-0.15					
$Leverage_t$	0.00	-0.06	0.68				
$\Delta On10Yr_t$	-0.03	-0.13	0.12	0.21			
$\Delta Noise_t$	-0.08	-0.12	0.19	0.43	0.55		
$CIPD_t$	0.10	0.07	-0.60	-0.82	-0.07	-0.22	
$CIPD_t$	0.08	-0.05	-0.74	-0.81	0.05	0.02	0.78
Panel B: Correlation between $CIPD_t$ and hedge fund risk factors							
	MKT	SMB	YLD	BAA	BD	FX	COM
SMB	0.24						
YLD	0.07	0.14					
BAA	-0.32	-0.25	-0.42				
BD	-0.25	-0.07	-0.12	0.24			
FX	-0.2	-0.02	-0.06	0.22	0.29		
COM	-0.17	-0.07	0.01	0.14	0.18	0.34	
$CIPD$	0.17	0.03	-0.14	-0.07	-0.12	-0.17	-0.13

Table 3.16: Combining Noise and CIPD. This table shows the results of a conditional double sort. In a first step, all hedged funds are sorted into five different portfolios based on their sensitivity to changes in the noise measure. Funds with the highest loading on $\Delta Noise$ are in portfolio 5 and funds with the lowest loading on $\Delta Noise$ are in portfolio 1. In a second step, each of the five portfolios is split into five more portfolios based on their loading on CIPD. Funds with the highest loading on CIPD are in portfolio 1 and funds with the lowest loading on CIPD are in portfolio 5. The figure in the bottom-right corner shows the risk-adjusted returns of the difference portfolio that is long hedge funds with the highest loading on $\Delta Noise_t$ and the lowest loading on CIPD and short the portfolio with the lowest loading on $\Delta Noise_t$ and the highest loading on CIPD. All figures are risk-adjusted returns using the Fung Hsieh seven factor model. The sample period is January 1994 to May 2015. Newey-West t -statistics are reported in square brackets. ***, **, and * indicate significance at a 1%, 5%, and 10% level respectively.

	Low N1	N2	N3	N4	High N5	N5-N1
High CIPD1	-0.31* [-1.68]	-0.06 [-0.42]	0.00 [0.00]	0.08 [0.84]	0.23 [1.18]	0.54** [2.22]
CIPD2	0.22*** [2.92]	0.07 [0.65]	0.20** [2.19]	0.11 [0.98]	0.39** [2.27]	0.18 [0.93]
CIPD3	0.25*** [2.81]	0.21*** [2.95]	0.18** [2.08]	0.19*** [2.61]	0.44*** [3.31]	0.19 [1.20]
CIPD4	0.35*** [3.65]	0.36*** [5.07]	0.23*** [3.16]	0.17** [2.16]	0.29** [2.51]	-0.06 [-0.50]
Low CIPD5	0.36*** [2.81]	0.35*** [4.19]	0.30*** [4.08]	0.27** [2.33]	0.67*** [3.55]	0.31** [2.20]
CIPD5-CIPD1	0.68*** [3.71]	0.41*** [3.01]	0.30** [1.98]	0.19* [1.72]	0.45* [1.80]	0.99*** [5.03]

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