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To My Family

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ABSTRACT

Foreign Institutional Investors, Monetary Policy, and Reaching for Yield with (with Boris Hofmann and Martin Schmitz)

This paper uses a unique security-level data set to demonstrate that foreign institutional investors shift their U.S. corporate bond portfolios toward bonds with higher credit spreads when U.S. monetary policy tightens, which reflects institutional factors related to nominal return targets and foreign exchange hedging. Foreign institutional investors in low-yielding jurisdictions are unable to meet their return target by only investing in their home bond market. To close this return gap, they increase their exposure to the higher yielding USD-denominated bonds. However, due to regulatory requirements and internal risk management, they hedge against the foreign exchange risk. To take advantage of the yield differential, they invest in long-term USD bonds while hedging the foreign exchange risk through short-term swaps on rolling basis. This makes the shape of the USD yield curve the key factor for the hedged return on their USD-denominated bonds, especially given the persistent premium to access the USD in the swap market since 2008. When U.S. monetary policy tightens, the USD yield curve flattens, erasing all the yield differential once the cost of hedging is applied. As a result, to improve returns on USD-denominated bonds, foreign institutional investors need to take more credit risk. This behavior has meaningful effects on corporate bond prices and issuances.

Fiscal Stimulus and Pension Contributions: Evidence from the TCJA (with Anna Zabai)

We evaluate the impact of the 2017 Tax Cut & Jobs Act (TCJA) pension tax break on sponsor contributions to defined-benefit retirement plans. We exploit cross-sectional variation in ex-ante exposure to the tax break. We find that the tax break induced an extra $2.8 billion of sponsor contributions to medium- and large-scale plans in 2017. However, we find strong evidence of reversal, both in terms of sponsor contributions and plan funding ratios by 2018. Our contributions model indicates that this reversal is consistent with more binding financial constraints in 2018 relative to 2019. Our results suggest that the TCJA did not have a long-lasting impact on corporate defined-benefit pension funds.

CHAPTER 1

FOREIGN INSTITUTIONAL INVESTORS, MONETARY POLICY, AND REACHING FOR YIELD

1.1 Introduction

The U.S. corporate bond market is the largest corporate bond market in the world. It accounts for more than 26% of the global corporate bonds and has become a crucial financing channel for firms in the U.S. (see International Capital Market Association August 2020 report). Since the 2008 financial crisis and against the backdrop of the prolonged period of low interest rates in many advanced economies (mainly the euro area and East Asia), foreign investors have funneled unprecedented amounts of funds into the U.S. corporate bond market. These foreign investors' holdings accounted for around 30% of the total outstanding amounts of corporate bonds in 2020, and the U.S. dollar value of these holdings doubled to over 4.5 trillion U.S. dollars between 2009 and 2020 (Figure 1.1). At the end of June 2020, 95% of these holdings were by private foreign investors (see Treasury International Capital System 2020 SHL Annual Survey). This makes private foreign investors the largest holders of U.S. corporate bonds. Given their dominance, it is important to understand their investment behavior and the effects of their institutional trading on prices in the U.S. corporate bond market.

We examine the determinants of reaching for yield in foreign institutional investors' holdings of U.S. nonfinancial corporate (NFC) corporate bonds from 2016 to 2020 and the implications for bond prices and issuance. Foreign institutional investors have been a driving factor for the large cross-border investment in the U.S. corporate bond market. Reaching for yield in our definition means tilting portfolios toward bonds with higher spreads relative to Treasury yields with the same maturity. More specifically, we examine how U.S. monetary policy affects risk-taking in foreign institutional investors' corporate bond portfolios by analyzing the extent to which foreign institutional investors shift the composition of their U.S. corporate bond holdings in response to changes in U.S. monetary policy.

Recent papers have illustrated reaching for yield among the two largest domestic institutional investors in the U.S. corporate bond market, i.e., insurance companies (Becker and Ivashina (2015) and Ozdagli and Wang (2020)) and mutual funds (Choi and Kronlund (2017)). Although the underlying mechanisms differ, a common theme in these studies is that institutional investors are especially prone to shift toward riskier bonds to generate higher returns when interest rates are low. However, past research has not focused on the reaching for yield behaviour by foreign investors in response to U.S. monetary policy. Studying the behaviour of these investors is important to fill this gap and to evaluate the full impact of U.S. monetary policy on credit conditions. Other papers have focused on the effect of unconventional monetary policies, namely the asset purchases programs of the Federal Reserve and the European Central Bank, on investors' bond portfolio rebalancing (Carpenter et al. (2015), Domanski et al. (2017), Fidora et al. (2020) and Koijen et al. (2021)). However, there has yet to be a study of the effects of conventional monetary policy on cross-border bond investments.

As such, we take a different perspective and study the reaching for yield of foreign institutional investors in response to U.S. conventional monetary policy. We find that as U.S. monetary policy tightens, foreign institutional investors tilt their U.S. corporate bond portfolios toward bonds with higher credit spreads and that such behaviour has price implications. Why would foreign institutional investors tilt their portfolios toward bonds with higher credit spreads when the U.S. monetary policy tightens? It is in line with efforts by foreign institutional investors in low-yielding jurisdictions to close their return gap by investing in higher yielding USD-denominated (USD) bonds. However, due to regulatory requirements and internal risk management, they hedge most of their U.S. dollar exposure.

Although using a longer-term cross-currency basis swap (or outright forward) contract broadly matches the maturity of the USD bonds, and hence fixes the basis for the term of the swap, foreign institutional investors hedge against the foreign exchange risk using short term FX swap (or outright forward) contracts that are renewed or "rolled over" at each FX contract maturity date until reaching the maturity of the respective USD bonds. They do so for two reasons. First, to take advantage of the yield differential between the higher yielding U.S. dollar and their low yielding currency. Second, short term swaps are the most liquid market for FX hedging, and so trading costs tend to be lower using these instruments compared to more tailored longer-term swaps. As a result, foreign institutional investors follow a "hedge short and invest long" strategy. In other words, they invest in long-term bonds, but hedge the currency through short-term swaps on a rolling basis.

The textbook cost of currency hedging is set only by the difference between the monetary policy rates in the U.S. and the foreign investor's home jurisdiction. However, in the post-great financial crisis environment, deviations of the cross-currency basis from zero are not uncommon (Borio et al. (2016), Du et al. (2018) and Avdjiev et al. (2019)). This introduces a premium to acquire the U.S. dollar in the swap market. For jurisdictions with large gross foreign asset positions (such as the euro area and East Asia), the premium is usually positive, increasing the cost for foreign investors acquiring USD bonds in an FX-hedged manner. This eliminates *more* than the short-term yield advantage of USD bonds over foreign institutional investors' jurisdictions' domestic bonds, amplifying the importance of the shape of the U.S. dollar yield curve for foreign investors. This makes the term spread on the USD bonds the key factor for their hedged return on their USD bonds investments. Given that the nominal term spread tends to get compressed during monetary policy tightening (Adrian et al. (2013), Hanson and Stein (2015), Crump et al. (2016), Nakamura and Steinsson (2018), and Kliem and Meyer-Gohde (2021)), foreign institutional investors need to increase their portfolio credit risk by investing in bonds with higher credit risk to improve the returns on the USD bonds, while leaving the FX hedge ratio unchanged.

To illustrate the currency hedging and term spread implications, we develop a mean-variance optimization framework where foreign institutional investors hedge their currency exposure and have a minimum required nominal return on their bond portfolio. The other key friction is that the premium (cross-currency basis) that foreign investors need to pay on top of the monetary policy interest rate differential between the U.S. and their home jurisdiction is persistently positive. The model predicts that foreign institutional investors who face a high FX hedging ratio and cannot achieve their required nominal return through investing in their home sovereign bond market reach for more yield when the U.S. monetary policy tightens. In addition, foreign institutional investors increase their allocation to the U.S. corporate bond market when the U.S. term spread widens or their home jurisdiction term and corporate spreads compress.

To test the model's predictions, we estimate a demand system for U.S. NFC bonds using euro area institutional investors as a representative of foreign institutional investors. Euro area institutional investors are particularly useful in studying the investment behaviour of foreign investors in response to U.S. monetary policy for two reasons. The first reason is the large size of euro area institutional investors' bond portfolio: at the end of 2020, euro area institutional investors held €688 billion in NFC bonds, accounting for 13% of the outstanding market reported in the U.S. Flow of Funds account. Second, the availability of detailed security-level data capturing all their bond holdings across the world at the sectoral level on a quarterly basis allows us to accurately capture their portfolio rebalancing, which is the main interest of this paper.

We use euro area investors' detailed quarterly bond holdings data to analyze how portfolio allocations and demand of U.S. NFC bonds relate to changes in U.S. monetary policy. The data comes from the ECB Sectoral Securities Holdings Statistics (SSHS). SSHS offers a comprehensive, fully integrated, granular dataset of the security holdings of euro area residents worldwide at the sectoral level. We also use the security-level holdings data of U.S. domestic investors from eMAXX. eMAXX provides comprehensive coverage of bonds predominantly held by domestic insurance companies, mutual funds, and pension funds at the security level. We combine the holdings data with bond yields and characteristics from the ESCB's Centralised Securities Database (CSDB). Our paper is the first – to the best of our knowledge – to construct a dataset that includes U.S. corporate bond holdings of *both* domestic and foreign investors at the sectoral level. The sectors in our sample collectively hold roughly 50% of the total outstanding corporate bond amount.

To estimate the demand system, we follow Koijen and Yogo (2019) and Koijen et al. (2021). This approach models weights of mean-variance portfolio as a logit function of bond yields, bond characteristics, and latent demand that represents heterogeneous expectations or constraints that are not captured by the observed characteristics. In equilibrium, it recognizes that investors' portfolio weights across securities have to add up to their outstanding values. Following this approach, we estimate a demand system for NFC bonds, modeling portfolio weights as a logit function of credit spreads, U.S and euro area monetary policy rates, bond characteristics, macro-financial variables (including term spreads), and latent demand.

Given the euro area investors' large size, an endogeneity problem may arise because a shock to their demand can have an impact on bond yields. With our dataset at hand and the persistence of investors' corporate bond investment mandates (Bretscher et al. (2020)), we use other sectors' contemporaneous bond holdings as an instrument to isolate exogenous variation in credit spreads for a given sector. To identify the effect of monetary policy rates and term spreads on the portfolio rebalancing, we use monetary policy shocks as an external instrument constructed from high-frequency price adjustments in the 3-month Libor rates and 10-year bonds for the U.S. and the euro area around the FOMC and ECB announcements, following the lead of Kuttner (2001), Bernanke and Kuttner (2005), Gürkaynak et al. (2005) and Gertler and Karadi (2015).

Consistent with the model and in contrast with U.S. life insurers, euro area institutional investors have a demand function that is decreasing in credit spreads. However, they tilt their portfolios toward corporate bonds with higher credit spreads when the U.S. monetary policy rate increases. Euro area investors also rebalance their portfolios toward U.S. corporate bonds when the U.S. term spread increases, and when the euro area term and credit spreads decrease. This emphasizes the importance of investor heterogeneity and its role in how monetary policy is transmitted.

Having documented the reaching for yield behaviour of euro area investors, we investigate whether their reaching for yield has an impact on corporate bond prices. To address this question, we examine the returns of bonds associated with their reaching for yield from the WRDS Bond Returns database. We find that during quarters that witness monetary policy tightening, bonds purchased by euro area investors in a given quarter have monthly raw returns that are 12 basis points higher than bonds not purchased by any euro area investor, and abnormal return 22 points higher. Such a pattern in raw and abnormal returns quickly reverses and is absent in other quarters. Using euro area investors' flows to the corporate bond market, our estimate of 12 basis points mean price effect implies a price elasticity of -1.67, which is similar to the demand elasticity reported in Chang et al. (2014) which is -1.46 in the cross section of U.S. stocks. Furthermore, we show that euro area investors reaching for yield have an impact on the volume of BBB-rated corporate bonds issuance. Overall, the results provide evidence corroborating that the behaviour of euro area investors reaching for yield can impact the credit conditions in the U.S.

Overall, our results highlight an important channel for the transmission of monetary policy that is relevant for practitioners and policy makers but has been overlooked by the academic literature. These findings match the February 2018 Schroders' report on global corporate bond market, which states that "investors should not be fooled into thinking that markets with higher yields in local terms offer higher return prospects. Currency hedging will neutralize much of this advantage, rendering comparisons of yields between domestic and overseas markets less meaningful."[1]

The rest of the paper proceeds as follows. In Section 1.2, we discuss the institutional settings of foreign investors and explain how the institutional features of currency hedging could lead them to engage in reaching for yield as the U.S. monetary policy rate increases. Section 1.3 presents the conceptual framework. Section 2.4 describes the data. Section 1.5 develops the empirical predictions that follow from the conceptual framework. It connects changes in holdings directly to changes in credit spreads and U.S. monetary policy rate. Section 1.6 investigates the effects of reaching for yield on corporate bond prices and issuance. Section 2.6, the conclusion, discusses broader implications.

1. "Breaking down borders in corporate bond markets" is available on Schroders' website.

1.2 Institutional Background

1.2.1 Global developments

Over the last two decades, the euro area and East Asian countries have been running large current account surpluses. By definition, a current account surplus can be mirrored by foreign investors drawing down previously purchased domestic assets, central banks intervening in FX markets and accumulating foreign currency reserves, or private institutions increasing their ownership of foreign assets. In the late 1990s and in the case of East Asian countries, the current account surplus was mainly mirrored by central bank intervention in FX markets and foreign currency reserves accumulation, primarily in U.S. dollar assets. However, over the last two decades, as central bank FX activity became limited, the role of institutional investors, namely insurance companies and pension funds, became the main vehicle mediating current account surpluses into overseas security markets, primarily the U.S. market. Institutional investors can invest in overseas security markets either through direct holdings or through domestic investment funds. Figure 1.2 shows the accumulation of current account surpluses, FX reserves, and foreign assets ownership by several types of institutional investors, over the last two decades. These figures piece together evidence that a driving factor for large cross-border investment and the associated impact on credit conditions in the United States may have been the “Global Institutional Investors Glut” rather than the “Global Savings Glut” or the “Global Banking Glut” coined by Bernanke (2005) and Shin (2012), respectively.

1.2.2 Institutional Investors in the Euro Area

European insurance companies and pension funds (ICPFs) constitute a large segment of the investor base both in their jurisdictions and globally. According to statistics from the European Central Bank, by the end of 2020, their assets exceeded € 12.2 trillion, or almost 12% of global bond markets outstanding amount[2]. Direct holdings of debt securities represent the main asset item on their balance sheets and account for more than 35% of their total financial assets, with the debt securities of U.S. issuers accounting for more than 7.7% of their direct holdings of debt securities.

ICPFs face reaching for yield incentives to generate higher returns especially when interest rates are low. The liabilities of life insurance companies and pension funds consist of promised fixed investment returns or fixed benefit promises that they have to meet, creating incentives to invest in higher-yielding bonds. In addition, their fixed-rate liabilities are long dated for several years or decades into the future. As a result, life insurers and pension funds typically have a negative duration gap, with liabilities of longer duration, and thus of more interest-rate sensitivity, than their assets. Breuer et al. (2019) estimates the duration mismatches between assets and liabilities for the nine countries (including Germany and France) with the largest life insurance sectors worldwide. It estimates the duration mismatch between assets and liabilities for German and French life insurers to be 10 and 6.5 years, respectively.

The duration mismatch depends on the level of interest rates and increases when rates fall (negative convexity). At times of low interest, this provides an incentive for

2. According to SIFMA total global bond outstanding is $123.5 trillion as of 2020.

ICPFs to "reach for duration" as well by investing in bonds with longer maturity. To better match the duration of their liabilities, insurance companies and pension funds are among the largest investors in the euro area bond markets: By the end of 2013, in notional amounts, almost 46.5% of their direct bond holdings were held in the form of euro area government domestic debt securities. By the end of 2020, this share decreased slightly to 45.4%, which accounts for 22% of the outstanding domestic government debt securities. Nevertheless, ICPFs have been facing pressure to achieve guaranteed returns because of persistently low and declining yields on fixed-income instruments, partly driven by the accommodative ECB policies that led to term spread compression. Figure 1.4 shows the gap between life insurance guaranteed returns and the domestic sovereign 10-years bond yields for EA life insurers by country of residence. The asset-weighted return gap for EA life insurers is 0.5%.

To boost returns, ICPFs can also invest in euro area corporate bonds. However, the domestic corporate bond market is relatively small. After the European sovereign debt crisis in 2012, banks dramatically reduced outstanding debt securities. This large decrease was not compensated by an equivalent increase in outstanding debt securities of other financial and non-financial firms (Koijen et al. (2021)). Furthermore, the ECB's unconventional monetary policies came with a credit spread compression, which reduced the yield attractiveness of euro area corporate bonds. It also led to fewer corporate bonds being available for sale. Figure 1.3 shows the total amount outstanding of euro area bonds by issuer type, the total assets of the EA ICPFs, and the share of the total euro area bonds held by the Eurosystem (i.e. the ECB together with the euro area national central banks). By the end of 2020, 20% of all euro area bonds were held by the Eurosystem. In particular, by the end of 2020, the Eurosystem held 27% of investment grade non-financial corporate bonds in the euro area. European investors have been already facing credit-risk limit saturations on major European corporate issuers (Dauphine' et al. (2021)). Figure 1.5 shows the gap between life insurance guaranteed returns and the overall investment returns for EA life insurers by country of residence. The asset-weighted return gap for EA life insurers is 0.7%. Many euro area countries face challenges to make investment returns in excess of guaranteed returns issued in the past as a result of the low yield environment.

The share of ICPFs' direct bond holdings allocated to debt issued by Monetary Financial Institutions (MFIs) declined from around 18.2% at the end of 2013 to 11.8% at the end of 2020. The shares of bonds issued by other financial institutions and non-financial corporations were stable around 8.5% and 9.5%, respectively. Geographically, by the end of 2013, close to 86% of the ICPFs' direct bond holdings were held in euro area debt securities, exhibiting a strong "euro area bias". However this figure declined to 79% at the end of 2020. Much of the increase in non-euro area debt allocation is reflected in an increase in exposure to the U.S. dollar denominated bonds. The share of U.S. dollar denominated bonds went up from 3.9% of direct bond holdings at the end of 2013 to 9% at the end 2020. This translates to an increase in holdings of magnitude of € 216 billion. Figure 1.6 shows the bond purchases by euro area ICPFs by issuer type from end of 2013 to end of 2020.

Indirect Holdings: Investment Funds

ICPFs can also invest in foreign bonds through euro area investment funds. At the end of 2020, ICPFs owned around €4 trillion in investment fund shares, or more than 25% of euro area investment fund assets. A significant portion of this amount is invested in foreign bonds. Precise figures on ICPFs' indirect holdings of U.S. NFC bonds are not readily available, but Carvalho and Schmitz (2021) presents an analysis that looks through the holdings of investment fund shares to estimate euro area investors' full exposures to global debt securities by sector in the aftermath of COVID-19 crisis. Although investment funds do not face the same contractual and regulatory environment as ICPFs, demand for subsistence yields during low-interest-rate times from end-investors (such as ICPFs) forces them to have nominal return targets (Breuer et al. (2019) and Bundesbank (2020)) which have an effect similar to fixed-rate liabilities (Choi and Kronlund (2017)). This creates an incentive for investment funds to reach for yield during times of low interest rates by investing in higher yielding bonds, resembling the behaviour of ICPFs.

Euro area investment funds also invest significantly in debt securities, accounting for around 37% of their total assets. In contrast to direct holdings of ICPFs, investments in debt markets are geographically more widely distributed. At the end of 2013, debt securities of non-euro area issuers accounted for more than 41% of their total debt securities holdings. However, as a result of the ECB unconventional monetary policies, the share of debt securities of non-euro area issuers went up to 55% at the end of 2020. Much of the increase in the non-euro area debt allocation is tied to the increase in their exposure to U.S. dollar denominated bonds. The share of the U.S. dollar denominated bonds went up from 17.4% of their bond holdings at the end of 2013 to 26% at the end 2020. This translates to an increase in holdings of magnitude of € 753 billion. Figure 1.7 shows the bond purchases by euro area investment funds by issuer type from the end of 2013 to the end of 2020.

1.2.3 U.S. Nonfinancial Corporate Bond Market

Corporate debt substantially increased after the 2008-09 global financial crisis (GFC), amounting to more than half of U.S. GDP in 2020. The rise in corporate debt was concentrated in nonfinancial corporate debt and was mainly created in the form of bonds exhibiting the substitution of firms from bank debt to capital market debt. Between 2008 and 2020, outstanding nonfinancial corporate bonds rose from $ 3 trillion to $ 6.6 trillion. Issuance activity has been widespread. The share of debt in the 'BBB' category rose from 28% of the outstanding nonfinancial corporate bonds to 46% over the same period, and from 36% of the investment grade bonds to 52% (Figure 1.8).

Since the 2008 global financial crisis and against the backdrop of the prolonged period of low interest rates in the euro area, European institutional investors have been funneling large funds in the U.S. nonfinancial corporate bond market. Hence they are becoming key players in this market and are influencing credit conditions in the U.S. Euro area investors' holdings of U.S. nonfinancial corporate bonds rose from $ 335 billion at the end of 2013 to more than $ 820 billion at the end of 2020, in market value. This translates to an increase from 8% of the amount outstanding to

12.7% (Figure 1.9). Although the increase seems to be driven mainly by investment funds and, to a smaller extent, by ICPFs' direct holdings, the estimates developed by Carvalho and Schmitz (2021) provide some key facts. First, ICPFs hold a significant share of euro area investment fund shares and the ICPFs' share of euro area investment fund holdings of U.S. NFC bonds (and non-euro area bonds in general) has been increasing sharply since the implementation of the ECB quantitative easing in 2015. Second, ICPFs' indirect holdings of U.S. NFC bonds (via euro area investment funds) have more than doubled since 2015. Third, ICPFs' indirect holdings of U.S. NFC bonds are larger than its direct holdings.

The U.S. corporate bond market is the largest bond market globally. It offers European ICPFs four attractive features (Dauphine' et al. (2021)). First, it offers diversification opportunities in terms of issuers. For example, at the end of 2020, the U.S. investment grade credit universe accounted for over 700 U.S.-domiciled issuers, of which about 600 have never issued in Euros. Second, it offers diversification opportunities in terms of sectors. Relative to the euro area, U.S. credit markets feature relatively higher weights of the energy and technology sectors at the expense of financial institutions. Third, it offers accessible exposure to longer maturities, either through the corporate or agency bond markets. For example, bonds with maturities of ten years and above account for 39% of the U.S. corporate bond universe, in contrast to being only 10% in the Euro universe. Fourth, the U.S. corporate bond market offers attractive spread pick-up especially on intermediate- and long-dated maturities. Figure 1.10 shows the excess returns of U.S. 10-year sovereign bonds over 10-year German Bunds and of U.S. investment-grade corporate Bonds over euro area investment-grade corporate bonds.

1.2.4 FX Hedging and Reaching for Yield

Although the U.S. sovereign debt market appears very appealing to foreign institutional investors, especially those with investment mandate constraints like life insurers and pension funds, there is the cost of hedging against the currency risk to consider. It is perceived that the European institutional investors hedge almost 100% of their exposure to foreign currency bonds (Borio et al. (2016) and Dauphine' et al. (2021)). There are two reasons that they hedge most of their FX exposure. The first is domestic regulation requirements. The EU's Solvency II directive, which came into effect in January 2016, stipulates that European insurers face 25% solvency capital charge applicable in the event of currency mismatches between insurance companies' assets and liabilities. The second is internal risk management practices. Taking FX risk would prove to be very risky without benefit. The losses incurred as a result of any depreciation in the U.S. dollar would likely outweigh any yield gains from the reduction in hedging cost. This concern is likely to materialise given the volatility of the FX market. Meese and Rogoff (1983) and Perold and Schulman (1988) suggested that short-term currency movements follow a random walk, representing a source of uncompensated risk. This would suggest that, by itself, an investment in unhedged foreign bond provides little value to investors, in light of the exchange rate volatility.

An FX hedge could be constructed either with a cross-currency basis swap or outright forward contracts. Although using a longer-term cross-currency swap (or outright forward) contract broadly matches the maturity of the USD bonds, and hence fixes the basis for the term of the swap, foreign institutional investors hedge against the foreign exchange risk using short term FX swap (or outright forward) contracts that are renewed or "rolled over" at each FX contract maturity date until reaching the maturity of the respective USD bonds. They do so for two reasons. First, to take advantage of the yield differential between the higher yielding USD and their low yielding currency. This is mainly determined by the shape of the USD yield curve. Second, short term swaps are the most liquid market for FX hedging, and so trading costs tend to be lower using these instruments compared to more tailored longer-term swaps. According to the BIS Triennial Central Bank Survey 2019, FX swaps (forward contracts) with maturity of six months or less made up 98% (95%) of their respective turnover. As a result, foreign institutional investors follow a "hedge short and invest long" strategy. In other words, they invest in long-term bonds, but hedge the currency through short-term swaps on a rolling basis.

In a textbook setting, FX cross-currency basis swap would be set only by USD-EUR interest rate differentials, giving rise to the academically revered no-arbitrage condition. However, in the post-great financial crisis environment, deviations of the cross-currency basis from zero are not uncommon (Borio et al. (2016), Du et al. (2018) and Avdjiev et al. (2019)). For jurisdictions with large cumulative current account surpluses over the past decade (such as the euro area and Far East Asia) and subsequently increasing gross foreign asset positions, the cross-currency basis is usually negative, increasing the cost for domestic investors acquiring U.S. dollar denominated assets in an FX-hedged manner. The EUR-USD swap rate in a cross-currency basis (CCB) contract with a tenor of three months is formalized as follows:

$$\text{EUR/USD Swap}_{\text{3month}} = \text{USD LIBOR}_{\text{3month}} - \text{EUR LIBOR}_{\text{3month}} - \text{CCB}_{\text{3month}}, \quad (1.1)$$

Equation 1.1 shows that the swap rate depends on the divergence of monetary policy rates in the U.S. and euro area directly through interest rate differentials. As a consequence, when the U.S. monetary policy is tightened, the cost increases for a euro area investor to hedge a USD-denominated long-term bond, erasing the entire increase in the U.S. short term rate. Thus, assuming hedging the full FX exposure, the hedged return on the U.S. Treasury bond is:

$$\text{Hedged Return} = \text{USD Term Spread} + \text{EUR LIBOR}_{\text{3month}} + \text{CCB}_{\text{3month}} \quad (1.2)$$

Equation 1.2 shows that the hedged return on the U.S. Treasury bond pins down to the sum of the term spread on the Treasury bond (the difference between long-term and short-term interest rates), the Euribor, and the cross currency basis. Given that both Euribor (Figure 1.16) and cross currency basis (Figure 1.11) have been negative since the ECB quantitative easing program in 2015, the term spread becomes the key factor for the hedged return for euro area investors. Empirical literature (Adrian et al. (2013), Hanson and Stein (2015), Crump et al. (2016), Nakamura and Steinsson (2018), and Kliem and Meyer-Gohde (2021)) documents that during recent monetary policy rate hike cycles, the nominal term spread tends to get compressed. This research shows that that the term spread tends to decline in monetary policy rate hike cycles even though it may initially rise modestly upon impact in some rate hike cycles such as 1994 and 1999. This contrasts with the 2004 and 2015 tightening cycles when the term spread dropped steadily in the face of hiking monetary policy rate. Although these rate hike cycles had boosted interest rates on the short end, they had not caused long-term rates to lift, and the term spread stayed close to historic lows and turned negative.

Figure 1.12 plots the coefficient of correlation between federal funds rates and term spread (based on the yield on the 10-year Treasury bond minus the yield on the 3-month Treasury bill) based on rolling regression with 20 quarters window. The coefficient has been persistently negative stabilizing around -0.43 from 2010. This means that U.S. monetary policy tightening potentially erases the entire yield differential for euro area investors as the USD yield curve flattens and term spread gets compressed. The hedged returns on the U.S. Treasury bonds turned negative for a prolonged period after the ECB quantitative easing program and the FED's monetary policy tightening in 2015. Even with yields at multi-year highs, Treasuries were returning less than their pricier German peers when euro area investors account for the steep hedging costs. Figure (1.13) shows the returns on the 10-years German Bund, and 10-years U.S. Treasury bond on unhedged and hedged (using a rolling 3-month FX swap) basis. The situation was not much different for A-rated U.S. corporate bonds. The hedged returns on AAA-rated corporate bonds for euro area investors were negative during some quarters in 2018 and 2019. The only segment in the U.S. investment grade corporate bond market that was yielding a positive hedged return for euro area investors during this period was the BBB-rated corporate bonds. Figure 1.14 shows the unhedged and hedged (using a rolling 3-month FX swap) U.S. corporate bonds yield curve for a euro area investor.

As a result, for euro area investors to improve their returns on their USD bonds, they need to reach more for yield in the USD bond market by taking extra credit risk through buying USD corporate, agencies or emerging market economies issued bonds. In other words, when U.S. monetary policy tightens, foreign institutional investors reach more for yield in clear contrast to the U.S. institutional investors who reach more for yield when U.S. monetary policy loosens. This shows significant heterogeneity in response to monetary policy between domestic and foreign institutional investors.

1.3 Conceptual Framework

This section sketches a two-period portfolio optimization model. The institutional investor pursues a total return objective of y_L which can be the rate of return on liabilities in the case of ICPFs or minimum nominal required return in the case of investment funds. At time 0, the investor chooses a global bond portfolio to achieve such return objective. To describe the portfolio re-balancing problem which arises due to the FX hedging, we solely consider bonds issued in the euro area and the U.S. At time 0, the investor's investment opportunity set consists of the following four bonds with the same maturity:

1. Euro area riskless sovereign bond with expected return $y_e + T_e$ and allocated weight w_1

2. Euro area corporate bond with expected return of $y_e + T_e + C_e$, volatility of σ_e^2 and allocated weight w_2

3. U.S Treasury riskless bond with expected return of $y_\$ + T_\$^*$ and allocated weight w_3

4. U.S corporate bond with expected return of $y_\$ + T_\$^* + C_\$$, volatility of $\sigma_\$^2$ and allocated weight w_4

Where y_e is the current short term interest rate in the euro area and $y_\$$ is the current short term interest rate in the U.S. The term spreads in the Euro-Area and the U.S. are given by T_e and $T_\$^*$, respectively. C_e is the yield spread of the euro area corporate bond over the euro area sovereign bond with the same maturity. Similarly, $C_\$$ is the yield spread of the U.S. corporate bond over the U.S. Treasury bond with the same maturity. Given the sensitivity of the term spread to the monetary policy, we model the term spread as $T_\$^* = T_\$ - \rho y_\$$, where $\rho > 0$ to capture the negative impact of monetary policy tightening on the term spread. Finally, the short-term rates and corporate spreads are assumed to be independent of one another.

FX Hedging: The euro area investor decides to currency hedge ϕ of the U.S. bonds, where $\phi \in [0,1]$. He uses a 3-month cross-currency basis swap to facilitate this hedging. The cost of hedging is $H(y_\$, y_e) = y_\$ - y_e - Z$, where $Z < 0$ to capture the persistent negative cross currency basis reflecting the premium that euro area investors need to pay in order to access the U.S. dollar in the swap market. Finally, the investor will face

FX fluctuation for the 1 - ϕ of the U.S. bonds that are not hedged. The expected return of the currency movement is F with associated risk of σ_f^2

The euro area investor has mean-variance preferences over the return on bonds, but faces the cost of FX hedging. Thus, the investor chooses its bond portfolio such that:

$$\min_{w_1, w_2, w_3, w_4} \quad w_2^2\sigma_e^2 + w_4^2\sigma_\$^2 + (1-\phi)^2(1-w_1-w_2)^2\sigma_f^2$$

$$\text{s.t.} \quad \sum_{i=1}^4 w_i r_i - \phi(1-w_1-w_2)H(y_\$, y_e) + (1-\phi)(1-w_1-w_2)F \geq y_L$$

$$\text{s.t.} \quad \sum_{i=1}^4 w_i = 1 \tag{1.3}$$

where r_i is the expected return on the respective bond. We assume short selling is allowed for simplicity. The allocation problem in Equation (2) involves constrained minimization of the portfolio variance over the bonds weights. The first-order condition yields the following solution for the optimal weight of the U.S. corporate bond, w_4^* :

$$w_4^* = \left(\frac{R_G}{C_\$}\right) \cdot \frac{(\frac{C_\$}{\sigma_\$})^2}{(\frac{D}{(1-\phi)\sigma_F})^2 + (\frac{C_e}{\sigma_e})^2 + (\frac{C_\$}{\sigma_\$})^2} \tag{1.4}$$

where R_G is the gap between the targeted nominal return and domestic sovereign bond yield and D is the spread of investing in the domestic sovereign bond over the

U.S. Treasury bond with ϕ of the FX exposure is hedged. In other words, it is the expected return on a portfolio which involves longing the euro area sovereign bond, shorting the U.S. Treasury bond and hedging ϕ of the FX exposure. For simplicity we will call D the return on the "hedged portfolio". Formally, R_G and D are defined as follows:

$$R_G = y_L - y_e - T_e \tag{1.5}$$

$$D = y_e + T_e - [y_\$ + T_\${}^* + (1-\phi)F - \phi H(y_\$, y_e)] \tag{1.6}$$

Figure 1.15 shows that D has been positive during the U.S. monetary policy tightening cycle that started at the very end of 2015, assuming an FX hedge ratio (ϕ) of 100%. Equation 1.3 implies that the optimal demand for the U.S. corporate bond is the product of two terms. The first term is the ratio of the return gap to the credit spread of the U.S. corporate bond. This captures how much investment in the U.S. corporate bond is needed to close the return gap. This is driven by the minimum return target. The second term is the risk-adjusted returns square of the U.S. corporate bond relative to the sum of the risk-adjusted returns square of other risky assets; the U.S. corporate bond, euro area corporate bond and the "hedged" portfolio. This is driven by the mean-variance optimization. We now move on to characterize the change in demand.

Implication 1: The effect of U.S. monetary policy on risk taking: Assuming a positive return gap ($R_G > 0$), a positive credit spread on the U.S. corporate bond ($C_\$ > 0$), and a positive yield on the "hedged portfolio" (D > 0), when the FX hedging ratio is low ($0 < \phi < 1-\rho$), the higher U.S. monetary policy rate leads to an increase in the attractiveness of the U.S. bonds for the euro area investor, including corporate bonds ($\frac{\partial w^*_{C_\$}}{\partial y_\$} > 0$), all else being equal.

When the FX hedging ratio is high ($\phi > 1-\rho$), the higher U.S. monetary policy rate leads the euro area investor to be discouraged from investing in the U.S. corporate bonds ($\frac{\partial w^*_{C_\$}}{\partial y_\$} < 0$). We prove this in Appendix. All else being equal, the higher the U.S. monetary policy rate, the more compressed the term spread will get, potentially erasing all the yield differential. This is a direct result of the term spread compression parameter (ρ).

Implication 2: Dynamics between U.S. monetary policy and risk taking: For investors fulfilling the conditions[3] for implication 1, if the risk-adjusted return of the U.S. corporate bond is high enough relative to the sum of the risk-adjusted return of the euro area corporate bond and the "hedge" portfolio, the higher U.S. monetary policy rate will lead to stronger demand of U.S. corporate bonds with higher credit spreads. In other words, the higher the cost of hedging, the more risk taking in the U.S. corporate bond market by the euro area investor ($\frac{\partial^2 w^*_{C_\$}}{\partial C_\$ \partial y_\$} > 0$). We prove this in Appendix. All else being equal, the higher the U.S. monetary policy rate, the more compressed the term

3. The four conditions are: (1) positive return gap ($R_G > 0$), (2) positive expected return on the "hedged" portfolio ($D > 0$), (3) a positive credit spread on the U.S. corporate bond ($C_\$ > 0$), and (4) high required FX hedge ratio ($\phi > 1-\rho$).

spread will get, potentially erasing all the yield differential. As a result, to improve their returns to close the return gap, euro area investors take more credit risk.

Implication 3: Effect of the slope of the yield curve: Steepening of the U.S. yield curve (higher U.S.term spread $T_{\$}^{*}$) leads to an increase in the attractiveness of the U.S. bonds for the euro area investor, including corporate bonds, all else being equal. On the other hand, steepening of the euro area yield curve (higher euro area term spread T_e) leads to decrease in the demand of the U.S. bonds, including corporate bonds, all else being equal.

1.3.1 Testable Predictions

This gives the following predictions that will be tested using detailed securities holding data for major euro area institutional investors in the U.S. corporate debt market. The exact empirical methodology and measurement of "reaching for yield" are described in the next Section.

Prediction 1: Euro area investors' U.S. corporate bond holdings are decreasing in the U.S. policy (cost of hedging).

Prediction 2: Euro area investors' U.S. corporate bond holdings are increasing in the U.S. bonds' credit spread in response to higher U.S. monetary policy (cost of hedging).

Prediction 3: Euro area investors' U.S. corporate bond holdings are increasing in the U.S. term spread, and decreasing in the euro area term and credit spreads.

1.4 Data

For euro area investors, we use data on security-level portfolio holdings of all 19 euro area countries from the ESCB Sectoral Securities Holding Statistics (SSHS). The data are collected by national central banks from financial investors and custodians. The dataset covers debt securities, listed shares as well as investment fund shares, all of which are in most cases identified with a unique International Securities Identification Number (ISIN). The data are collected on a quarterly basis since 2014Q1 and we use releases until 2020Q4.

The SHSS data consist of directly and indirectly reported securities. A financial institution resident in the euro area is obliged to report securities that it holds as its own investment ("direct reporting") as well as securities that it holds in custody ("indirect reporting"). Investors in the SHSS are defined by their country of domicile and sector. We follow Koijen et al. (2021) and aggregate the data to five sectors on the euro area level: monetary financial institutions (MFI) excluding monetary authorities, insurance companies and pension funds (ICPF), other financial institutions (OFI - including important intermediaries such as mutual funds which represent the largest subgroup of this sector and hedge funds), households (HH) and "others" which include non-financial corporations and general government.

For U.S. domestic institutional investors, we use detailed investors' bond holdings data from eMaxx Thomason Reuters to test the above predictions. eMAXX has been used in several papers (Becker and Ivashina (2015) and Bodnaruk and Rossi (2016)), but is still a relatively new source in international finance literature. It covers the holdings of insurance companies, mutual funds, pension funds and investment management companies.

Using the ISIN for every security, SHSS and eMAXX data are merged with individual asset characteristics obtained from the ESCB's Centralised Securities Database (CSDB) which contains information on more than six million debt and equity securities issued globally (Rousová and Caloca (2018) and Fidora et al. (2020)). Therefore, one can use information at the security-level to retrieve information related to bond yields, maturity and liquidity. In addition, we use the Wharton Research Data Services (WRDS) Bond Database to track all the corporate bonds traded over time along with their monthly returns. It provides comprehensive coverage of all traded corporate bond issues, sourced from TRACE Standard and TRACE Enhanced. We use the yield curve constructed by Gürkaynak et al. (2007) to calculate bonds' credit spread by subtracting the yield of the corresponding Treasury security with the same maturity from the yield of the corporate bond. We also use Bloomberg to retrieve euro area corporate option adjusted spread index.

Table 2.1 reports the summary statistics of the portfolio information for euro area ICPFs and investment funds, and U.S. life insurance companies. It also reports the summary statistics of bonds' average monthly return in a given quarter. Euro area investors tend to hold bonds with higher credit spreads than U.S. investors. The average credit spread for euro area investors around 3.8 %, with a standard deviation of 2.2%. On the other hand, the average credit spread of U.S. life insurers is half the average credit spread for the euro area investors with a smaller standard deviation. U.S. life insurers tend to hold bonds with slightly higher time to maturity compared to other investors. Euro area investors also tend to hold bonds with larger amount outstanding than the U.S. life insurers.

1.5 Empirical methodology for changes in holdings

In our empirical analysis we focus on the holdings of U.S. NFC bonds by euro area ICPFs and investment funds. The share of these two sectors constitutes 93% of euro area private investor holdings of U.S. NFC bonds and 78% of euro area private investor holdings of all U.S. bonds. We use the notional values in all our analyses such that market prices are not driving the results. Thus these values accurately capture new investments and portfolio shifts reflecting active choices by investors, which is the main focus of this paper.

Our analysis spans from 2016:Q1 to 2020:Q4. We choose this period for three reasons. First, the data quality of SSHS is best during this period. Second, it is a period that has witnessed a full monetary policy tightening-loosening cycle in the U.S. Figure 1.16 shows that the 3-month Libor rate gradually started to increase from the very end of 2015 to early 2019, and decreased thereafter, reaching 22 bps in 2020:Q4. On the other hand, the euro area monetary policy rate was relatively stable in the negative territory. During this period, euro area institutional investors were already facing a negative return gap as the result of the term and credit spread compression on the back of ECB unconventional monetary policies. Third, the EU's Solvency II directive came into effect in January 2016, pushing euro area insurers to hedge their currency exposures.

1.5.1 *Reaching for Yield at the Extensive Margin*

We start with studying whether the incentives of euro area investors to invest in higher-yield bonds in the U.S. is related to the U.S. monetary policy rate. We define the reaching for yield at the extensive margin to be the difference between the weight of U.S. dollar denominated non-Treasury bonds and the weight of dollar denominated Treasury bonds in the bond portfolio. We conjecture that euro area investors who are attempting to achieve a minimum nominal yield due to their liability structure, like ICPFs, increase their reaching for yield when the U.S. monetary policy increases. They do so by investing more in U.S. dollar denominated non-Treasury bonds as compared to dollar denominated Treasury bonds. To test this hypothesis, we run the following regression:

$$\mathrm{W}_t^{NT} - \mathrm{W}_t^{T} = Post_t + \beta\, y_t^{\$} + \epsilon_t \tag{1.7}$$

where W_t^{NT} is the weight of U.S. dollar denominated non-Treasury bonds in the euro area investor's global bond portfolio, W_t^{T} is the weight of U.S. dollar denominated Treasury bonds in the euro area investor's global bond portfolio, and $y_t^{\$}$ is the 3-month U.S. dollar Libor rate. To capture the impact of the Federal Reserve Secondary Market Corporate Credit Facility (SMCCF), which was created in the aftermath of the COVID-19 crisis and spurred risk taking, on euro area investors reaching for yield, we include a regression dummy $Post_t$ which is equal to one for quarters starting from 2020:Q2.

Table 1.2 reports the results of the regression characterized in Equation 1.7. Columns 1 & 2 report that a one percent increase in the 3-month U.S. dollar Libor is associated with a 0.44 and 0.75 percentage point increase in the difference between the weight of U.S. non-Treasury bonds and the weight of Treasury bonds in the bond portfolio for euro area ICPFS and investment funds, respectively. These results imply that when the U.S. monetary policy tightens, euro area investors tilt their portfolios toward non-Treasury dollar denominated bonds.

Euro area investors may be titling their portfolios toward non-Treasury dollar denominated bonds to add duration risk and not to add credit risk. To isolate lengthen the bond portfolio duration channel, we run the following regression:

$$\mathrm{W}_{m,t}^{NT} - \mathrm{W}_{m,t}^{T} = \alpha_m + Post_t + \beta\, y_t^{\$} + \epsilon_{m,t} \tag{1.8}$$

where $W_{m,t}^{NT}$ is the weight of U.S. dollar denominated non-Treasury bonds in the euro area investor's global bond portfolio with maturity of m years, and $W_{m,t}^{T}$ is the weight of U.S. dollar denominated Treasury bonds in the euro area investor's global bond portfolio with maturity of m years. We also include maturity-bin fixed effects α_m, where m is 0, 1, 2, ..., 30 years. Columns 3 & 4 of Table 1.2 report that for a given maturity, a one percent increase in the 3-month U.S. dollar Libor is associated with an 0.02 and 0.018 percentage point increase in the difference between the weight of U.S. non-Treasury bonds and the weight of Treasury bonds in the global bond portfolio of euro area ICPFS and investment funds, respectively. These results imply that when the U.S. monetary policy tightens, euro area investors tilt their portfolios toward non-

Treasury dollar denominated bonds to add credit risk.

1.5.2 *Reaching for Yield in the NFC Corporate Bond Market*

Next, we study whether the incentives of euro area investors to invest in higher-yield bonds *relative* to the market is related to the level of U.S. monetary policy rate. To measure this empirically, we compare the average credit spreads of euro investors' non-financial corporate bond holdings with the average credit spread of the aggregate non-financial corporate bond portfolio (Choi and Kronlund (2017)) and Ozdagli and Wang (2020)) in the WRDS Bond Returns database. We define the relative reaching for yield (RRFY) of the euro investors at date t as the average credit spread of the euro investors' bond portfolio relative to the average credit spread of all outstanding non-financial corporate bonds in the market:

$$RRFY_t = \frac{\sum_i H_{i,t} CS_{i,t}}{\sum_i H_{i,t}} - \frac{\sum_j V_{j,t} CS_{j,t}}{\sum_j V_{j,t}} \tag{1.9}$$

where $CS_{i,t}$ is the credit spread of bond i defined as yield spread measured as the yield-to-maturity of bond i over the Treasury yield of similar maturity, $H_{i,t}$ is the amount of bond i held by the euro area investors, and $V_{j,t}$ is the total amount of bond j outstanding in the market. Comparing the relative credit spread of euro area investors' portfolio to the market allows us to control for unobservable factors that drive variation in the market credit spreads. We focus on the average credit spread rather than the average total yield to disentangle the effect of euro area investors' reaching for yield by increasing their holdings of bonds with greater credit risk (Becker and Ivashina (2015)) from reaching for yield by lengthening the bond portfolio's duration and adding more duration risk (Ozdagli and Wang (2020)). The main advantage of our approach is that it cancels out any bias, as our RFY measure is defined as deviations of average bonds' credit spread from the average credit spread of other bonds.

Figure 1.17 plots our RRFY measure versus the 3-month U.S. dollar Libor rate for euro area investors by sector (ICPFs and investment funds). A one percentage point increase in the U.S. dollar rate is associated with a 28 and 33 point increase in the excess credit spreads on NFC bond portfolios of euro ICPFs and investment funds, respectively. These results imply that when monetary policy tightens in the U.S., euro area investors tilt their portfolios toward relatively cheaper (higher-yielding) NFC bonds.

1.5.3 A demand system for NFC bonds

We next turn to the main test of the empirical predictions derived from our theoretical framework. To bring the predictions of the model to the data, we estimate a demand system for NFC bonds to relate portfolio rebalancing to corporate credit spreads and U.S. monetary policy changes using the asset demand system developed by Koijen and Yogo (2019). This system tests how bond holdings evolve as the U.S. monetary policy rate changes over time. We conjecture that investors with a long-term investment objective due to their liability structure (ICPF) or their attempt to achieve a minimum nominal yield (investment funds) increase their reaching for yield when the monetary policy increases. We define reaching for yield as taking on more risk by acquiring higher yielding bonds, i.e., bonds with higher credit spreads.

For demand curves estimation, we extend the euro area investors' holdings data with the U.S. domestic life insurers' holdings to obtain a larger coverage of U.S. nonfinancial corporate bond holders. Domestic life insurers, who are the largest domestic holders of U.S. corporate bonds, share the same business model as euro area insurance companies and pension funds, making them a suitable placebo. To be comparable with the euro area investors, we aggregate their holdings at the sectoral level.

Formally, investor i's investment in U.S. NFC bond n is denoted by $H_{it}(n)$, and the investment in the outside asset is denoted by $H_{it}(0)$. The portfolio weight in the framework of Koijen and Yogo (2019) is then defined as:

$$w_{i,t}(n) = \frac{H_{it}(n)}{H_{it}(0) + \sum_{n=1}^{n=N} H_{it}(n)} = \frac{\delta_{it}(n)}{1 + \sum_{n=1}^{n=N} \delta_{it}(n)} \tag{1.10}$$

where $\delta_{i,t}(n) = \frac{H_{it}(n)}{H_{it}(0)}$. The portfolio weight in the outside assets is $w_{i,t}(n) = 1 - \sum_{n=1}^{n=N} w_{it}(n)$. Koijen and Yogo (2019) shows that for investors with mean-variance preferences over returns, returns which are assumed to follow a factor model, and both expected returns and factor loadings which are assumed to be affine in a set of characteristics, we can write the portfolio weight in Equation 1.10 as a logit function of the credit spread, U.S. monetary policy rate, bond characteristics, and financial macroeconomic variables $X_t(n)$. Koijen et al. (2021) show that we can characterize euro area investors' demand as follows:

$$\log(\mathrm{H}_{i,t}(n)) = \beta_{1,i}\,\mathrm{CS}_t(n) + \beta_{2,i}\,y_t^{\$} + \beta_{3,i}\,\mathrm{CS}_t(n)\cdot\,y_t^{\$} + \beta_{4,i}\,y_t^{e} + \beta_{5,i}^{'}\,\mathrm{X}_t(n) + \epsilon_{i,t}(n) \tag{1.11}$$

where the credit spread $CS_t(n)$ is bond n's risk measure defined as yield spread measured as the yield-to-maturity spread over the Treasury yield of similar maturity. $y_t^{\$}$ and y_t^{e} are the 3-month U.S. dollar and euro Libor rates, respectively. To account for the last component of the cost of currency hedging beside the U.S. dollar and euro libor rates, we include the euro-dollar cross currency basis. The component of demand that is not captured by prices, characteristics, and time-invariant characteristics, $\epsilon_{i,t}(n)$, is referred to as latent demand.

To control for time-invariant issuer characteristics, all regressions include issuer fixed effects. We also include controls for remaining time to maturity and total amount outstanding as key time-varying bond characteristics that drive demand for bonds. Investors aiming to match their liabilities or trying to achieve certain nominal yields might have a preference for certain maturities. Total amount outstanding is a measure of liquidity that potentially leads to higher or lower demand for certain bonds. To control for alternative investments to U.S. corporate bonds both in the U.S. and the EA, we include the U.S. and euro area term spreads calculated as the 10-year constant maturity minus 3-month Libor rates on the dollar and euro, respectively. As a proxy for the euro area Investment Grade corporate spreads, we use the euro area Corporate Option Adjusted Spreads index.

We capture persistent unobserved characteristics through the lagged investor i's holdings of bond n in the prior quarter t. By conditioning on initial holdings, our identification comes from time-series variation within a bond during a certain quarter. Although time fixed effects would be the most general specification, they would preclude the use of time-series variation to estimate the effect of the change in the U.S. monetary policy and the cost of hedging on the U.S. corporate bond holdings. Finally, we cluster standard errors by issuer because some companies have several traded issues over our sample.

Our main variable of interest is the interaction between the security's credit spread measure and the U.S. dollar Libor rate, CS_{it} x $y_t^{\$}$. So our key coefficient is β_3 which determines the search-for-yield effects through the credit-risk channel. A positive coefficient ($\beta_3 > 0$) would suggest that the higher U.S. monetary policy rate, the more investment is shifted toward riskier U.S. corporate bonds. In other words, this interaction captures whether the euro area investors' propensity to invest in U.S. NFC bonds with a different credit spread is affected by the U.S. monetary policy. That is, we test if euro area investors' allocations across different categories of riskiness vary positively with the U.S. monetary policy. We then interpret a positive coefficient as evidence of a search-for-yield motive since it implies that U.S. monetary policy tightening increases investment in U.S. corporate bonds with higher credit spreads. On the other side, for domestic investors, a negative coefficient ($\beta_3 < 0$) would suggest that the higher the monetary policy rate, the more investment is shifted toward safer U.S. corporate bonds.

Instrumental Variable Approach

In order to obtain consistent estimates of the parameters in Equation 1.11 using OLS, one has to assume that characteristics and prices (spreads and swap rate) are exogenous to latent demand. We can't assume this exogeniety for three reasons. First, the corporate bond market is dominated by a few key players. And so the investors included in this study cannot be assumed to be atomistic with demand shocks of non-negligible price impact. Second, correlated demand shocks could have price impact in the aggregate, which rules out any factor structure in latent demand. Third, there is a possibility that economic activity fluctuates in response to exogenous non-financial factors, and the swap rate simply reflects these changes in real activity. Consequently, bond specific credit spreads, U.S. and euro area Libor rates and term spreads, and the index of euro area corporate spreads are allowed to be jointly endogenous with latent demand. That is, a correlated positive demand shock to euro area investors can have an impact on credit and term spreads. Equation 1.11 is therefore estimated by fitting a dynamic panel data model using GMM with an instrumental variable approach.

To identify NFC bonds credit spreads, we construct an instrument based on Koijen and Yogo (2019)'s framework applied to corporate bonds in Bretscher et al. (2020). We use other investors' contemporaneous portfolio holdings as an instrument to isolate exogenous variation in credit spreads. In contrast to their paper, we define the investment holdings of an investor at the sectoral level instead of the fund level. We do so because the euro area investors' holdings are at the sectoral level. To construct this instrument, we extend the holdings of euro area investors and domestic life insurers with the holdings of domestic investment management firms, mutual funds, pension funds, and non-life insurance companies using eMaxx bond holdings.

In estimating sector i's bond demand, the instrument for credit spread of bond n is:

$$\hat{CS}_{i,t}(n) = log\left(\sum_{j\neq i} A_{j,t}\,\frac{\mathbb{1}_{j,t}(n)}{1+\sum_i \mathbb{1}_{j,t}(n)}\right) \tag{1.12}$$

where $A_{j,t}$ is the total holdings of sector j at time t and $\mathbb{1}_{j,t}(n)$ equals one if sector j at time t has positive holdings of bond n. This instrument depends only on the holdings of other sectors, which are exogenous under our identifying assumptions. The instrument can be interpreted as the counterfactual credit spread if other sectors were to hold an equal-weighted portfolio within their investment universe. For example, the life insurance sector holds an equal-weighted portfolio of NFC bonds, the mutual funds sector holds an equal-weighted portfolio of NFC bonds, and so on. In constructing the instrument for the EA ICPFs (investment funds), we drop the holdings of EA investment funds (ICPFs) from the other sector's holdings universe. We do so because the EA ICPFs own more than 25% (33%) of the bonds held by EA bonds (mixed) investment funds holdings. This makes the holdings of one sector not exogenous to the other.

Following Romer and Romer (2004), we use surprises in the three month U.S. Dollar and Euro LIBOR rates as an instruments to identify the impact of U.S. and euro area monetary policy on portfolio adjustment of euro area investors. Similar to papers that employ intraday measures of monetary policy surprises (e.g., Kuttner (2001), Bernanke and Kuttner (2005), Gürkaynak et al. (2005) and Gertler and Karadi (2015)), surprises in the LIBOR rates are measured within a tight window of 30 minutes of the FOMC and ECB announcements to ensure that the surprises in LIBOR rates solely reflect news about the FOMC and ECB decisions. Similarly, we use the cumulative surprises in the 10-year Treasury and euro area bonds around the FOMC and ECB announcements as the instruments for U.S. and euro area term spreads. To identify the euro area credit spreads index, we use the natural logarithm of ECB bond holdings of non-public sector bonds. These bonds are purchased under the ECB asset purchases programmes that started in 2009, namely asset-backed securities, corporate, and covered bond (1, 2 and 3) purchase programmes. All the variables have a first-stage t-statistic that is well above the critical value of 4.05 for rejecting the null of weak instruments at the 5 percent level (Stock and Yogo (2005)).

In order to quantify the strength of our instruments, we run a first-stage regression of the endogenous independent variable onto its instrument among other instrumental variables, swap basis, the lagged investor holdings of the bond in the prior quarter t-1, and other bond characteristics contained in the vector X_t. We estimate the first-stage regression for each sector. Formally, the first stage regression of the credit spreads is characterized as follows:

$$\mathrm{CS}_t(n) = \beta_{1,i}\hat{CS}_{i,t}(n) + \beta_{2,i}\ \hat{y}_t^{\$} + \beta_{3,i}\hat{CS}_{i,t}(n) \cdot\ \hat{y}_t^{\$} + \beta_{4,i}\ \hat{y}_t^{e} + \beta_{5,i}^{'}\hat{X}_t(n) + \epsilon_{i,t}(n) \tag{1.13}$$

where $CS_{i,t}(n)$ is the credit spread of bond n, $\hat{CS}_{i,t}(n)$ is the instrument for credit spread of bond n for investor i defined in Equation 1.12, $\hat{y}_t^{\$}$ is the instrument for the USD libor rate, $\hat{y}_t^{e}$ is the instrument for the euro libor rate, $\hat{X}_t(n)$ includes the instruments for the USD term spread, euro area term spread and euro area credit spreads index. It also includes the swap basis, the lagged investor holdings of the bond n in the prior quarter t-1, remaining time to maturity and total amount outstanding of bond n. Figure 1.18 reports t-statistic of the first stage of the credit spreads across our three sectors of interest. That is, all sectors have a first-stage t-statistic above the 4.05 (lower bound in the Figure) for rejecting the null of weak instruments at the 5 percent level (Stock and Yogo (2005)). For all other four endogenous variables, we run the first stage regression in Equation 1.13. The first-stage t-statistic is well above this critical value for all the four endogenous variables across our three sectors of interest, implying the weak instruments problem does not exist.

Results

Table 1.3 reports the estimated demand for the U.S. NFC bonds characterized in Equation 1.11. Columns 1 & 2 report the estimated demand for euro area investors and column 3 reports the estimated demand for U.S. domestic investors. Columns 1 & 2 report the impact of U.S. monetary policy on the degree to which euro area investors rebalance their bond portfolio toward riskier U.S. corporate bonds. The demand of a given bond by euro area ICPFs and investment funds is positively related to our key interaction term, the product of the credit spread on that bond and the U.S dollar Libor. This means that the higher the U.S. monetary policy rate, the larger the holdings of euro area institutional investors of higher-yielding U.S. bonds. Although we cannot reject equality, the coefficient on the bond's credit spread has a negative sign, suggesting that euro area investors are not inclined toward reaching for yield, in general. This is consistent with euro area investors taking a more cautious approach to corporate bonds. The coefficient on the U.S dollar Libor also has a negative sign which means the higher the U.S. monetary policy rate, the less attractive U.S. corporate bonds, as the monetary policy eliminates the yield advantage of the bond.

Column 3 reports the impact of U.S. monetary policy on the degree to which U.S. life insurance companies rebalance their bond portfolio toward riskier U.S. corporate bonds. In contrast with euro area investors, domestic investors have a positive coefficient on the credit spread, which means that their demand is increasing in the credit spread, and there is reaching for yield. However, they have a negative coefficient on the interaction between the 3-month U.S. LIBOR and the credit spread, which means that their reaching for yield is decreasing in the U.S. monetary policy rate. In other words, the domestic investors decrease their reaching for yield when the U.S. monetary policy is tightened. This results in heterogeneity in the effect of U.S. monetary policy on NFC bond holdings across institutional investors with different residencies.

The coefficient on the Euribor for euro area investors is positive which means that the euro area investors' demand for U.S. bonds (including corporate bonds) increases when the cost of currency hedging decreases. The coefficients on the euro area term spread and corporate credit spreads are negative which means that the euro area investors' demand for U.S. bonds (including corporate bonds) is higher when the investment opportunities deteriorate in the euro area. In contrast to domestic lifers and their inherent need to match long-term liabilities with long-term bonds, euro area institutional investors hold U.S. corporate bonds with shorter maturity. This is likely driven by "search for yield" by taking extra credit risk at shorter maturities as the result of the term premia compression during monetary tightening.

Domestic life insurers are more persistent in their NFC bond portfolio than euro area investors. Their holdings in the prior quarter $t-1$ have a higher coefficient, closer to one. This is consistent with euro area investors having a larger propensity to sell U.S. corporate bonds because of a change in the FX currency hedging cost or in times of market stress. Finally, euro area investors have a preference for bonds with a larger amount outstanding compared with domestic life insurance companies, implying a preference for more liquid bonds. This is consistent with domestic life insurers holding the least liquid bonds if they can get compensated for doing so due to their long-term liability structure (Becker and Ivashina (2015)). This makes sense especially given they do not face the same dynamic FX hedging cost as their euro area counterparts.

Robustness Check

The main theme of our identification strategy explained in Section 1.5.3 is that the instrument for the credit spread at the bond level exploits variation in the investment universe across investors and in the size of potential investors across bonds. Given that our holdings data is at the sectoral level, one concern is that bonds with small issue outstanding may have a small exogenous component of demand as the holdings of this bond may be dominated by one sector and not included in the investment universe of others. As a robustness check, we estimate the specification in Equation 1.11 for different levels of issue outstanding. Figure 1.19 reports our key interaction term at different levels of bond issue outstanding across our three sectors of interest. Across our three sectors of interest, the interaction coefficient is overall decreasing in absolute value for higher levels of issue outstanding. For the euro area investors, the coefficient is relatively stable in terms of economic and statistical significance. For domestic life insurers, the economic significant of the interaction coefficient is steadily decreasing for higher levels of bonds issue outstanding. This is likely driven by the least liquid bonds being dominantly held by domestic life insurers. These bonds have small issue outstanding and hence small exogenous component of demand.

As an alternative approach to characterize our demand system, we will combine the three components of the hedging costs $y_t^{\$}$, y_t^e and swap basis into one variable, which we call the swap rate. Formally, we characterize our demand system as follows:

$$\log(\mathrm{H}_{i,t}(n)) = \beta_{1,i}\,\mathrm{CS}_t(n) + \beta_{2,i}\,\mathrm{Swap}_t + \beta_{3,i}\,\mathrm{CS}_t(n)\cdot\mathrm{Swap}_t + \beta_{4,i}^{'}\mathrm{X}_{i,t}(n) + \epsilon_{i,t}(n) \tag{1.14}$$

where Swap_t is the 3-month euro-dollar swap rate. We identify the impact of cost of currency hedging on portfolio adjustment of euro area investors using the cumulative difference in surprises between the three month U.S. Dollar and Euro LIBOR rates as an instrument for the 3-month swap rate. The minimum first-stage t-statistic for the swap rate across our three sectors of interest is 67.92, implying a strong instrument.

Table 1.4 reports the estimated demand for the U.S. NFC bonds characterized in Equation 1.14. Columns 1 & 2 report the impact of cost of currency hedging on the degree to which euro area investors rebalance their bond portfolio toward riskier U.S. corporate bonds. Euro area ICPF and investment funds sectors' demand of a given bond is positively related to our key interaction term, the product of the credit spread on that bond and the swap rate. This means that the higher the swap rate, the larger the holdings of euro area institutional investors' of higher-yielding U.S. bonds. Similar to the results from the estimated demand system in Equation 1.11, the coefficient on the bond's credit spread have negative signs. The coefficient on the swap rate is negative which implies the higher cost of hedging, the less the attractiveness of U.S. corporate bonds, as the yield advantage gets eliminated. Column 3 reports the impact of U.S. monetary policy on the degree to which U.S. life insurance companies rebalance their bond portfolio toward riskier U.S. corporate bonds. As expected, the coefficient on the interaction term has no statistical or economic significance. This can be explained by domestic life insurers holding the majority of their bond portfolio in the domestic (corporate) bond market. Therefore, the euro-dollar swap rate is irrelevant.

The coefficient on the U.S. term spread for the euro area ICPFs is positive, which means that euro area investors' demand for U.S. bonds (including corporate bonds) is higher when investment opportunities improve in the U.S. On the other hand, and similar to the estimated demand system in Equation 1.11, the coefficients on the euro area term spread is negative, which means that euro area investors' demand for U.S. bonds (including corporate bonds) is higher when investment opportunities deteriorate in the euro area.

1.6 Implications of Reaching for Yield

In the previous sections, we have documented that euro area investors tend to reach for yield by re-balancing their portfolios toward corporate bonds with higher yield spread following U.S. monetary policy tightening. In this section, we investigate whether reaching for yield on the part of euro area investors can have implications for bond prices. We hypothesize that U.S. monetary policy tightening compresses the term spread, erasing the yield on Treasury bonds that euro area investors can earn on a hedged basis, which spurs an increase in their demand for corporate bonds with high yield spreads, thereby creating an upward price pressure in the valuations of corporate bonds.

To test the hypothesis regarding the price pressure exerted by euro area investors' reaching for yield behaviour, we examine the path of bond abnormal returns around euro area investors' purchases. We run the following regression using a framework similar to Chodorow-Reich et al. (2020):

$$Abret_{i,t+h} = \alpha\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \gamma\, \mathrm{X}_{i,t} + \epsilon_{i,t} \tag{1.15}$$

where $EAbuy_{i,t}$ is an indicator variable which equals one if euro area institutional investors in aggregate buy bond i in quarter t and $y_t^{\$}$ is the three month U.S. dollar Libor instrumented with the monetary policy shocks to LIBOR rates. To adjust for risk, the average of monthly abnormal return of bond i in quarter t, denoted as $Abret_{i,t}$, is calculated as its raw return minus the return on the benchmark portfolio to which it belongs. Using the characteristics-based procedure in Bessembinder et al. (2008) at a more granular level, the portfolio benchmarks are created based on remaining time-to-maturity and bond credit rating. At the end of every quarter, bonds are first segmented into 31 time-to-maturity groups (0, 1, 2 and up to 30 years) and then 11 credit rating groups (AAA, AA, A, BBB, ..., D and not rated). This gives a total of 341 groups of bonds. For each group, the equal-weighted return is computed and used as the benchmark portfolio return. Following Chodorow-Reich et al. (2020), we include in the control set $X_{i,t}$ the time fixed effects, coupon rate and squared coupon rate of the bond, and the change of the yield on a maturity-matched treasury .

The coefficients of interest are α on the variable EAbuy and β on the interaction term $EAbuy$ x $y_t^{\$}$. While the coefficient α measures the difference in abnormal returns between bonds bought by euro area investors and those that are not during a certain quarter, the coefficient β captures the additional difference in abnormal returns as a result of the U.S. monetary policy tightening. When the U.S. monetary policy is tightened, corporate bonds purchased by euro area investors are likely to be under price pressure. Therefore, if reaching for yield creates upward price pressure, then the coefficient on the interaction term $EAbuy\ x\ \ y_t^{\$}$ should be positive.

The results are displayed in Raw Qtr = 0 of Table 1.5. Notice that bonds purchased by euro area investors during quarters with monetary policy tightening exhibit striking positive abnormal returns, as evidence by the positive and significant coefficient on the interaction term $EAsbuy \text{x}\ y_t^{\$}$. In particular, as a result of a one percent increase in the 3-month U.S. dollar Libor rate, the difference in the monthly abnormal returns between bonds bought by euro area investors and those that are not increases by 22 basis points on a monthly basis. Then, we divide the universe of bonds into various groups: AAA/AA/A-rated bonds, BBB-rated bonds, non-investment grade bonds and non-rated (NR) bonds. Table 1.6 shows that the estimated effect for the USD Libor interaction is only statistically significant within the set of BBB-rated bonds (column 3).

To differentiate euro area investors bringing information into prices from exerting price pressure, following Coval and Stafford (2007), we look for evidence of price reversals through extending the time horizon of the dependent variable Abret in Equation 1.15 from t - 4 to t + 4. The results are presented in Table 1.5. The coefficients on the interaction term $EAsbuy \text{x}\ y_t^{\$}$ are plotted in Figure 1.20 which shows an upside-down V shape price pattern centering around Qrt 0 of euro area investors' buying, indicating that the returns reverse over the quarter that follows. To put all results above into a difference-in-difference perspective: During quarters with high U.S. dollar Libor rates, there is upward price pressure from euro area investors buying of corporate bonds. Hence, the results demonstrate that the price pressure is generated by EA purchases associated with reaching for yield as the U.S. monetary policy rate increases and term spread get compressed.

As an alternative approach to using abnormal returns to evaluate the price pressure exerted by euro area investors, we use the average monthly raw return in a given quarter. This allows us to estimate the price elasticity. We run the following regression using a framework similar to Bretscher et al. (2020):

$$ret_{i,t+h} = \alpha_i + \alpha_t + \gamma\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \epsilon_{i,t} \tag{1.16}$$

where $ret_{i,t}$ is the average of monthly abnormal return of bond i in quarter t. We include bond (α_i) and time fixed (α_t) effects. The results of the regression in Equation 1.16 are presented in Table 1.7. Similar to the results using the abnormal returns in Equation 1.15, with a one percent increase in the U.S. dollar Libor, the difference in average monthly raw returns between bonds bought by euro area investors and those that are not increases by 12 basis points. In addition, the coefficients on the interaction term $EAsbuy$x $y_t^{\$}$ which are plotted in Figure 1.21 show an upside-down V shape price pattern centering around Qrt 0 of euro area investors' buying indicates that the returns reverse over the quarter that follows.

To quantify the price pressure exerted by euro area investors, we follow Coval and Stafford (2007) and Lou (2012) and define it as the net purchases (sales) by euro area investors of a certain bond in a particular quarter scaled by the lagged bond amount outstanding from the prior quarter. Formally, $\%\Delta Demand_{i,t}$ is defined as follows:

$$\%\Delta Demand_{i,t} = -100 * \frac{\Delta Demand_{i,t}}{Outstanding_{i,t-1}} \tag{1.17}$$

The average price pressure exerted by euro area investors when they buy a U.S. corporate bond is 0.20 %. As defined in Equation (1.18) and to relate our results to the literature and to estimate price elasticity, we divide our measure of price pressure by the positive coefficient of the interaction term between EAbuy and the U.S. Libor rate of 0.12% which yields a price elasticity of -1.67. This is very similar to the demand elasticity reported in Chang et al. (2014) which is -1.46 in the cross section of U.S. stocks.

$$Elasticity = -\frac{1}{\%\Delta Demand/\Delta Return} \tag{1.18}$$

Finally, we investigate whether reaching for yield on part by euro area investors can have implications for corporate bond issuance. Similar to prices, we hypothesize that an increase in monetary policy rate spurs an increase in euro area investors' demand for corporate bonds, thereby facilitating larger issuance or even giving incentives for larger bond issuance by corporations. To test this hypothesis, we run the following regression similar to Todorov (2020) and Siani (2019):

$$log(\text{Issuance}_{i,t}) = \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \gamma \text{X}_{i,t} + \epsilon_{i,t} \tag{1.19}$$

where $log(\text{Issuance}_{i,t})$ is the natural logarithm of the issuance amount of bond i issued in quarter t. We include in the control set $X_{i,t}$ the time and issuer fixed effects, coupon rate and squared coupon rate of the bond, and the maturity of the bond. SHSS does not distinguish between bonds acquired in the primary market or the secondary market, we consider euro area investors to have participated in the issuance of a bond if they purchased such bond in the same quarter when it was issued. The first column in Table 1.8 shows that the coefficient on the interaction term EAbuy x USD Libor is positive and statistically significant.

We next explore whether there are differences in effects by types of bonds. It could be that risk-taking by euro area investors is related to monetary policy tightening, but only within certain categories of corporate bonds. We divide the universe of bonds into various groups: AAA/AA/A-rated bonds, BBB-rated bonds, non-investment grade bonds and non-rated (NR) bonds. Among the rated groups, Table 1.8 shows that the estimated effect for the USD Libor interaction is only statistically significant within the set of BBB-rated bonds (column 3). We then test the same hypothesis using the total bond issuance at the firm level by running the following regression:

$$log(\text{Issuance}_{f,t}) = \beta \, Hold_{f,t} \cdot y_t^{\$} + \epsilon_{f,t} \tag{1.20}$$

where $log(\text{Issuance}_{f,t})$ is the natural logarithm of the total bond issuance amount of by firm f in quarter t and $Hold_{f,t}$ is an indicator variable which equals one if euro area institutional investors in aggregate hold at least one bond issued by firm f in quarter t. Similar to the results at the bond level, among the rated groups, Table 1.9 shows that the estimated effect for the USD Libor interaction is only statistically significant within the set of BBB-rated issuers (column 3). This shows that reaching for yield on the part of euro area institutional investors contributed to the rapid growth of the BBB bond market. Overall, this table suggests that euro area investors' risk taking increases in the U.S. corporate bond when the U.S. monetary policy tightens.

1.7 Conclusion

Using unique security-level holdings data for euro-area investors, we study the impact of currency hedging on bond portfolio rebalancing, reaching for yield, and bond prices. We find that euro area insurance companies, pension funds, and investment funds, which are among the largest players in the U.S. non-financial corporate (NFC) corporate bond market, shift their bond portfolios toward corporate bonds with higher credit spread as U.S. monetary policy tightens.

We estimate a demand system for U.S. NFC bonds using instrumental variables to relate portfolio rebalancing to credit spreads and the stance of U.S. monetary policy. Our results highlight significant heterogeneity between euro area institutional investors and domestic life insurers. In particular, while the former increase their holdings of corporate bonds with higher credit spreads when U.S. monetary policy is tightened, the latter reduce such holdings. This result highlights fundamentally different incarnations of reaching for yield behaviour between the two types of investors reflecting the effect of currency hedging implemented by foreign investors. We further find corroborating evidence that reaching for yield on the part of euro area investors can drive overpricing of corporate bonds when U.S. monetary policy tightens. It can also facilitate larger bond issuance or even give incentives for larger issuance by U.S. corporations, especially BBB-rated corporations.

The results point to a new amplifying mechanism that contributed to the crash of the U.S. corporate bond market in the wake of the COVID crisis in 2020. When a shock hits the economy, the Federal Reserve loosens monetary policy and the yield curve steepens. This induced euro area investors to re-balance their bond portfolio back to Treasury and safer corporate bonds as their hedged return increased. This in turn led to high selling pressure in the corporate bond market, especially in the market segment with the lowest credit quality.

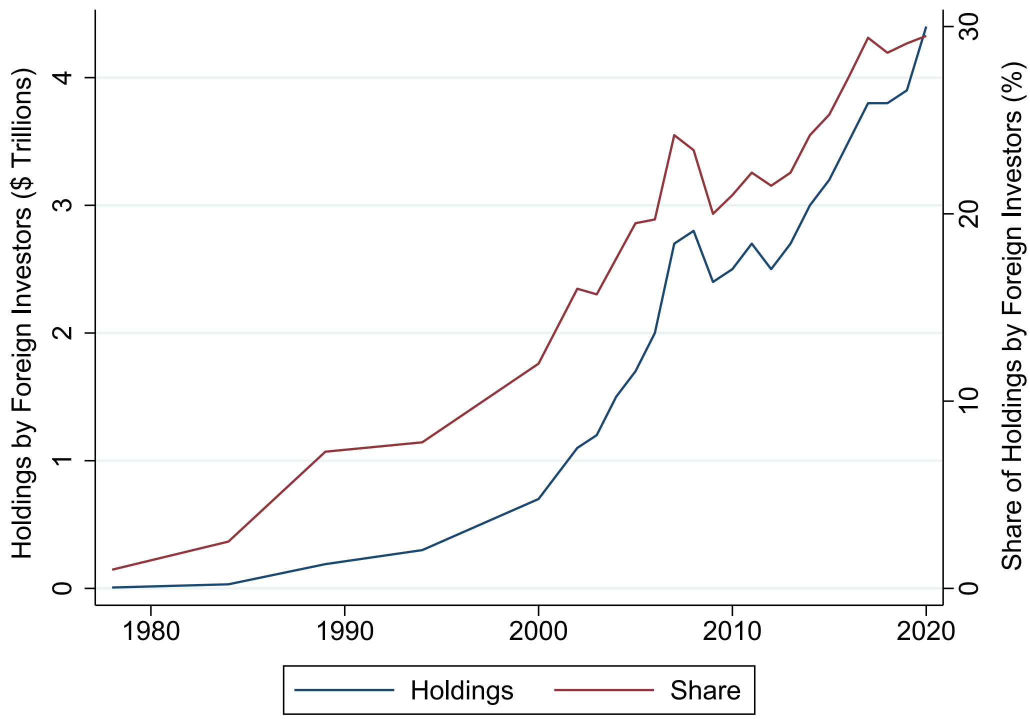
More generally, our analysis suggests that inflows by foreign institutional investors from low-yielding jurisdictions have contributed to the easing of financial conditions in the corporate bond market and have probably been a major driving force of the build up of corporate debt in the U.S. over the past decade, especially by BBB-rated corporations. At the same time, they may have had a profound impact on the transmission of the Federal Reserve's monetary policy to domestic financial conditions. Previous papers have analyzed the risk-taking channel of monetary policy working through domestic banks and institutional investors. Our paper shows how U.S. monetary policy affects the portfolio allocation of foreign investors who reach for yield. The results suggest that the FX hedging channel of foreign investors works in the opposite direction of the classical risk-taking channel for domestic investors, thus potentially weakening the effectiveness of U.S. monetary policy.

The analysis of our paper suggests interesting avenues for future research. Our framework could be applied to also analyze the extent to which other foreign investors, especially large East Asian (Japan, Korea, Taiwan) institutional investors, shift the composition of their U.S. corporate bond holdings in response to changes in U.S. monetary policy.

Breuer et al. (2019) shows that large East Asian life insurers added more than $0.4 trillion in new investments in U.S. dollar–denominated credit during 2013–18. Their combined share of the market rose from 8 percent in 2013 to 11 percent in 2018. East Asian investors are facing the same return gap challenge as euro area investors (Figure 1.22) but at a larger scale, given the tiny size of their domestic corporate bond market (Figure 1.23). In addition, Japanese investors have been facing a return gap and hedged return challenges for a longer period of time period. Figure 1.24 shows that Japanese investors have been facing negative hedged returns on their Treasury bond investments already since before the 2008 financial crisis. However, the main challenge in performing this analysis remains the availability of security-level data at the holder sectoral-level, similar to the ECB SHSS.

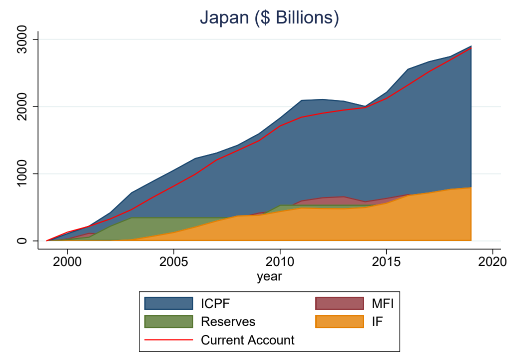
Our framework could also be used to analyze the extent to which euro area and East Asian investors adjust their holdings of U.S. dollar denominated of emerging market economies (EMEs). Exploring this question could shed light on the impact of institutional investors' risk-taking on the financial stability challenges of EMEs. Such an analysis could be helpful to understand better the continued challenges of EMEs from capital flows highlighted in the recent debate about the design of EME macro-financial stability frameworks.

Figure 1.1: Foreign holdings of U.S. corporate bonds

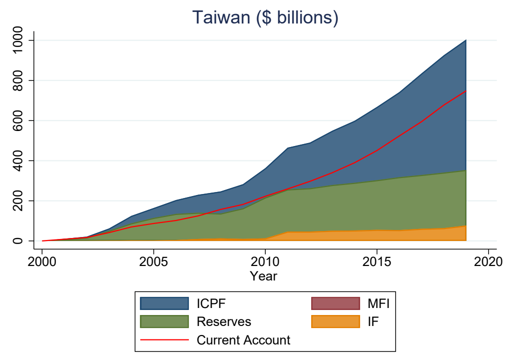


This figure reports the market value of U.S. corporate debt holdings by foreign investors and their share of the outstanding U.S. corporate bond market. The figure includes 1974, 1984, 1989, 1994, 2000, and then on an annual basis from 2002 to 2020. The data source is the Treasury International Capital System (TIC) SHL Annual Survey.

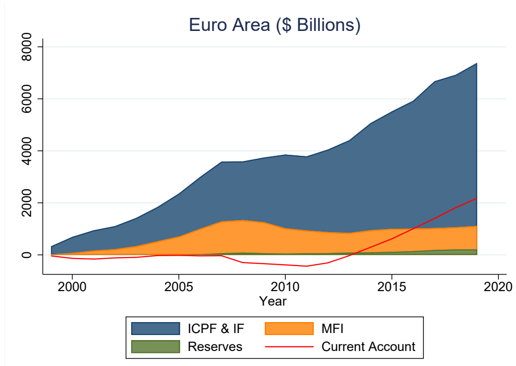
Figure 1.2: Accumulation of foreign assets



(a) Japan



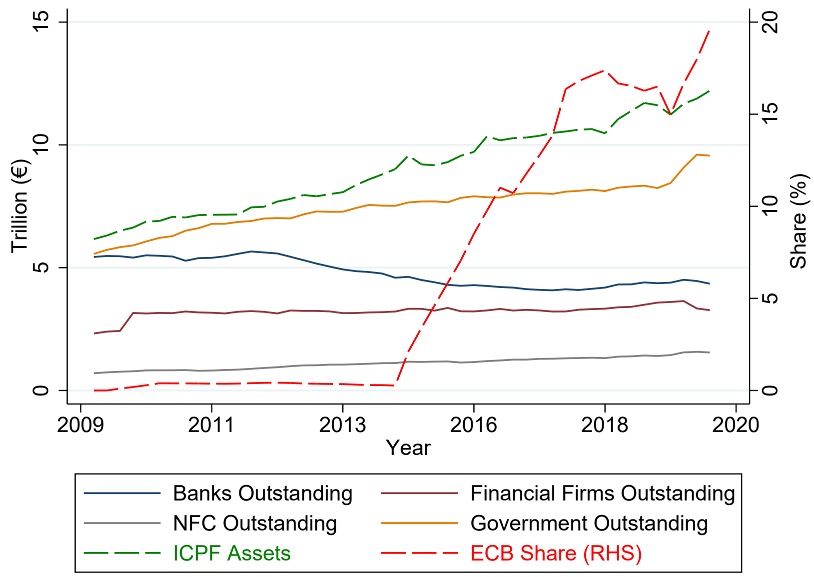
(b) Taiwan



(c) Euro Area

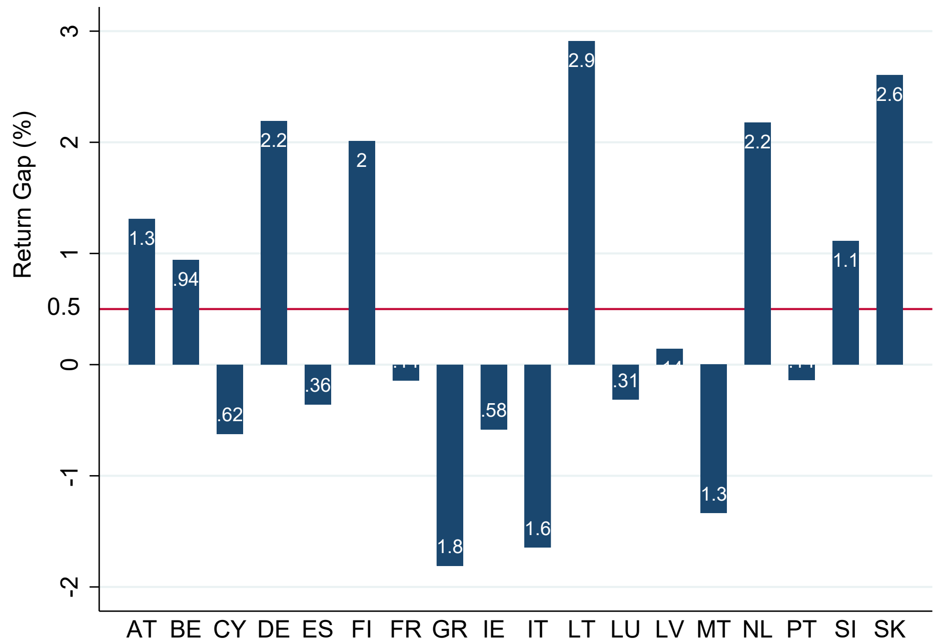
This figure reports the cumulative acquisitions of foreign assets by insurance companies and pension funds (ICPF), investment funds (IF), central banks (Reserves) and monetary financial institutions (MFI) in Japan, Taiwan, and the euro area from 2000 to 2019. For the euro area, the acquisitions of the insurance companies and pension funds are consolidated with the acquisitions of the investment funds. The data sources are Bank of Japan, the Central Bank of the Republic of China, and the European Central Bank.

Figure 1.3: Euro area bond market



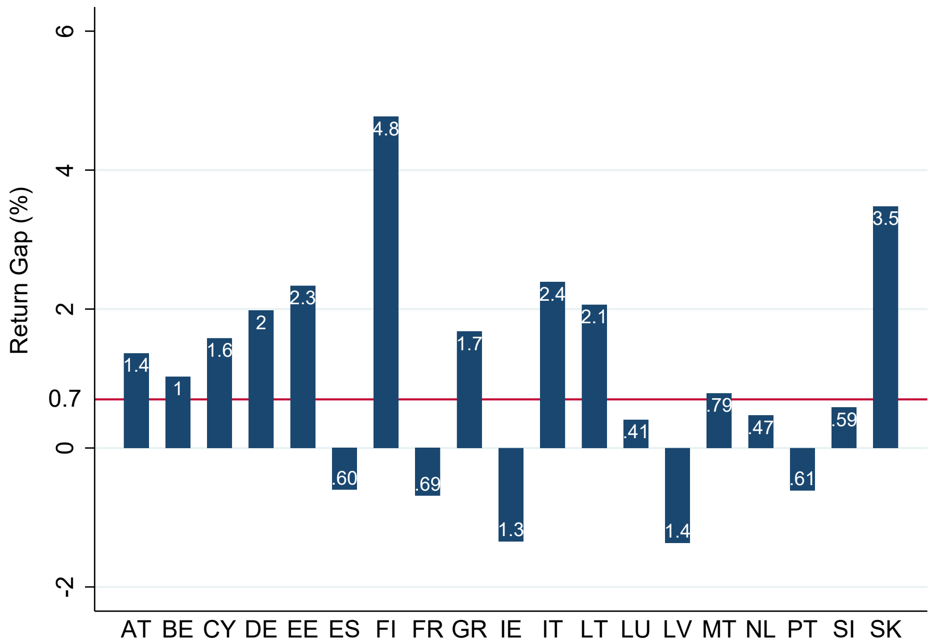
This figure reports the face value of debt outstanding for governments, banks, nonbank financial firms, and nonfinancial firms in the euro area from January 2009 to December 2020. It also reports the total assets of the euro area insurance companies and pension funds (ICPFs). Finally, it reports the share of the total euro area bonds held by the European Central Bank (ECB). It is calculated as the total euro area bonds held by the European Central Bank multiplied by the 100 divided by the total outstanding bonds in the euro area which is the sum of the debt outstanding for governments (GOVT), banks (MFI), nonbank financial firms (OFI), and nonfinancial (NFC) firms. The data source is the European Central Bank's Statistical Data Warehouse.

Figure 1.4: Return gap of euro area life insurers - domestic sovereign yields



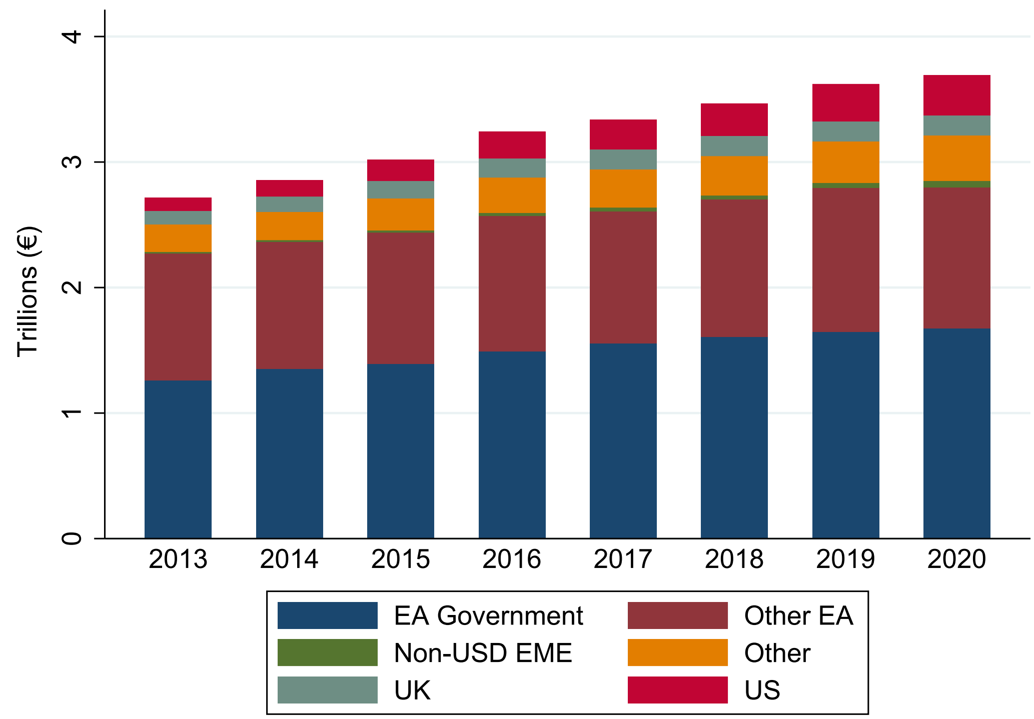
The figure reports the return gap for Q4 2018 for life insurers and composites in the euro area. The return gap is defined as the difference between the weighted average guaranteed interest rates and domestic sovereign 10-years bond yields. The horizontal red line is the average of the return gap across all the euro area countries weighted by domestic life insurance assets. Data labels use International Organization for Standardization (ISO) country codes. The data source is the European Insurance and Occupational Pensions Authority (EIOPA).

Figure 1.5: Return gap of euro area life insurers - total investment return



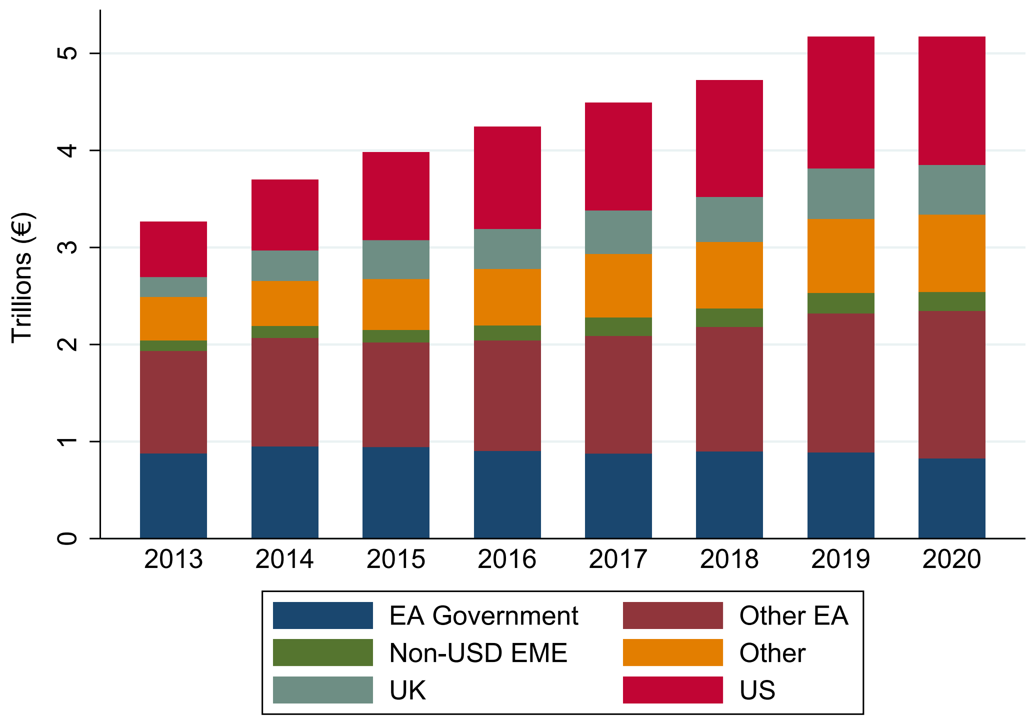
The figure reports the return gap for Q4 2018 for life insurers and composites in the euro area. The return gap is defined as the difference between the weighted average guaranteed interest rates and the return gap return on investment including unrealised gains/losses. The horizontal red line is the average of the return gap across all the euro area countries weighted by domestic life insurance assets. Data labels use International Organization for Standardization (ISO) country codes. The data source is the European Insurance and Occupational Pensions Authority (EIOPA).

Figure 1.6: Euro area insurance companies and pension funds portfolio rebalancing



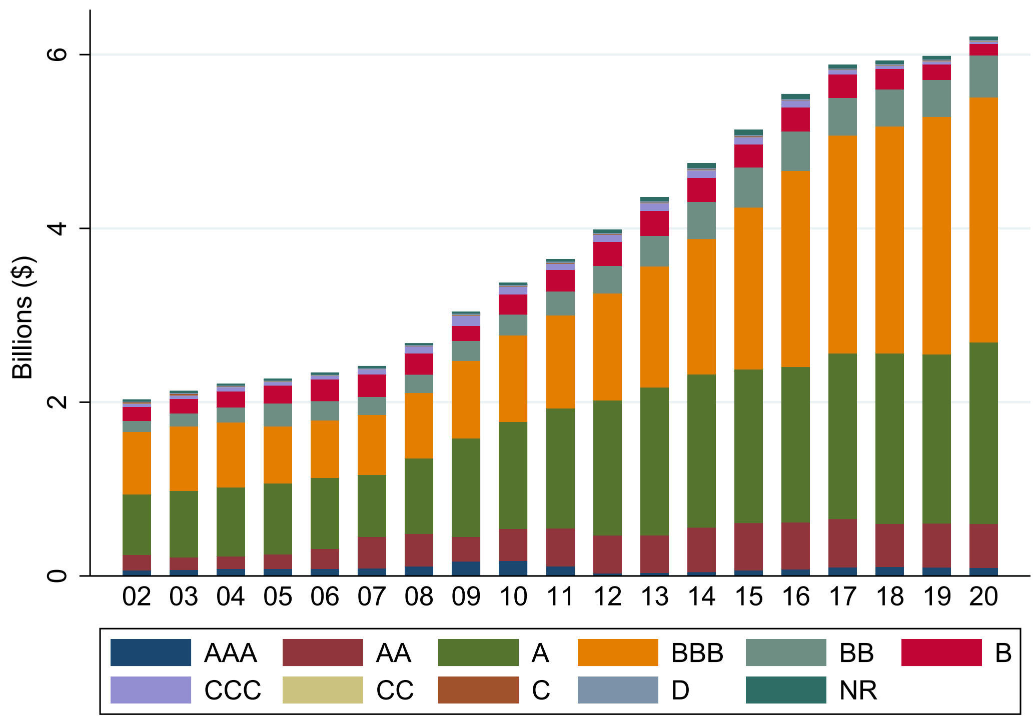
This figure plots the notional amounts of bond holdings by euro area insurance Companies and pension funds by issuer type from 2013 to 2020. "EA Government" refers to bonds issued by euro area governments, "Other EA" refers to bonds issued by monetary financial institutions, nonfinancial corporations and other financial institutions in the euro area. "UK" refers to the bonds issued by the UK government, monetary financial institutions, nonfinancial corporations and other financial institutions. "US" refers to the bonds issued by the U.S. government, monetary financial institutions, nonfinancial corporations, other financial institutions and U.S. dollar denominated bonds issued by emerging market economies (EMEs). "Non-USD EME" refers to euros or local currencies denominated bonds issued by EMEs. "Other" refers to bonds issued in other European Union countries, other advanced economies or offshore centers.

Figure 1.7: Euro area investment funds portfolio rebalancing



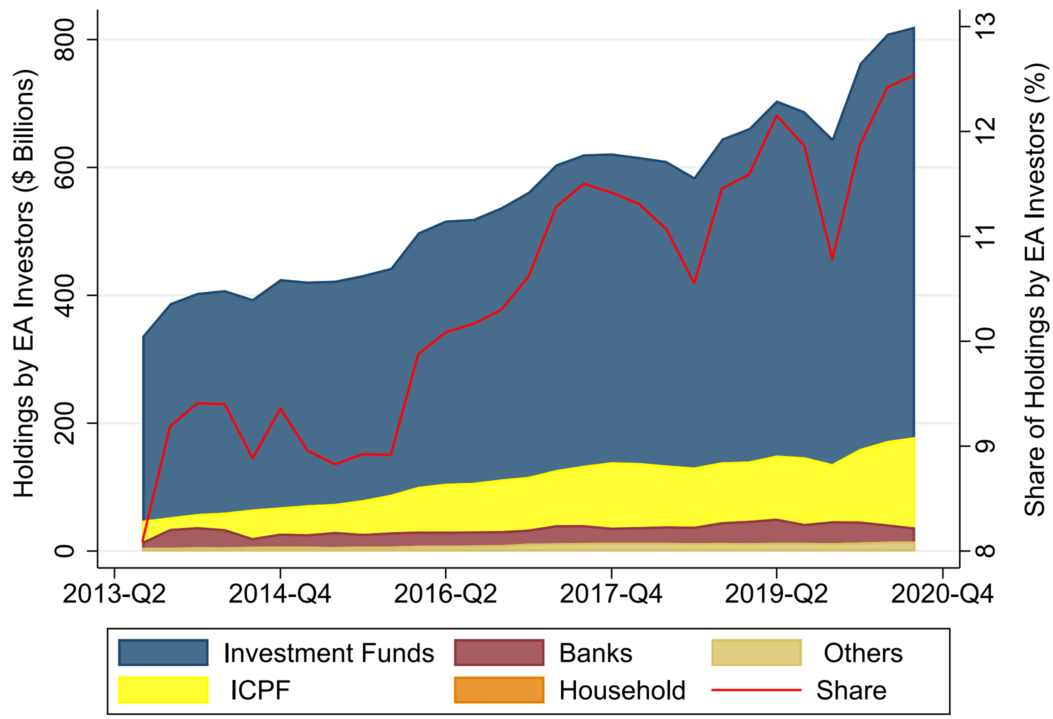
This figure plots the notional amounts of bond holdings by euro area investment funds by issuer type from 2013 to 2020. "EA Government" refers to bonds issued by euro area governments, "Other EA" refers to bonds issued by monetary financial institutions, nonfinancial corporations and other financial institutions in the euro area. "UK" refers to the bonds issued by the UK government, monetary financial institutions, nonfinancial corporations and other financial institutions. "US" refers to the bonds issued by the U.S. government, monetary financial institutions, nonfinancial corporations, other financial institutions and U.S. dollar denominated bonds issued by emerging market economies (EMEs). "Non-USD EME" refers to euros or local currencies denominated bonds issued by EMEs. "Other" refers to bonds issued in other European Union countries, other advanced economies or offshore centers.

Figure 1.8: Outstanding corporate bond market breakdown



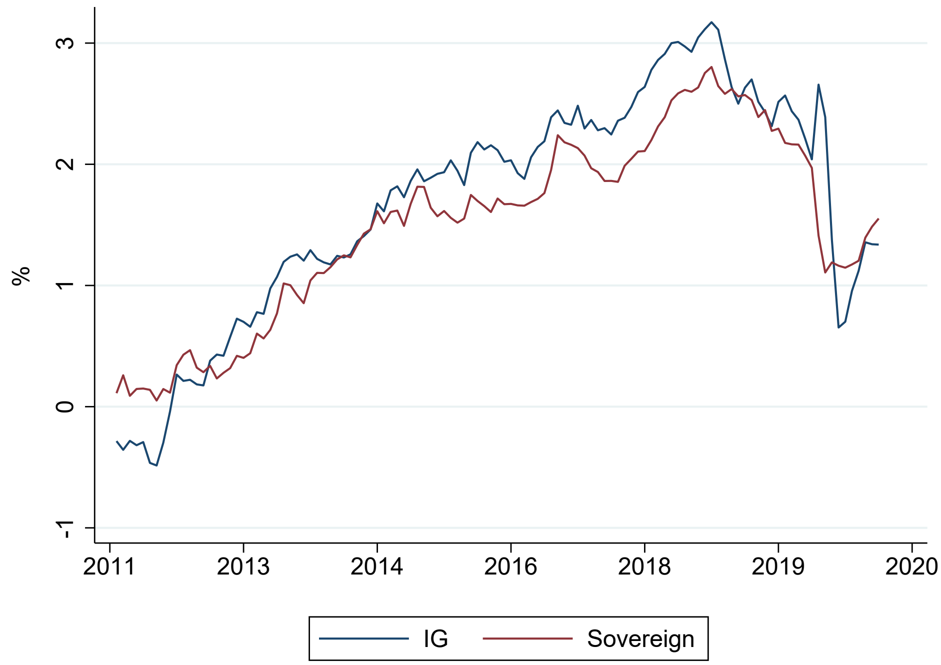
This figure plots the outstanding corporate bonds by credit rating from 2002 to 2020. The data source is Wharton Research Data Services (WRDS) Bond Returns.

Figure 1.9: The breakdown of euro area investors' holdings of U.S. NFC bonds



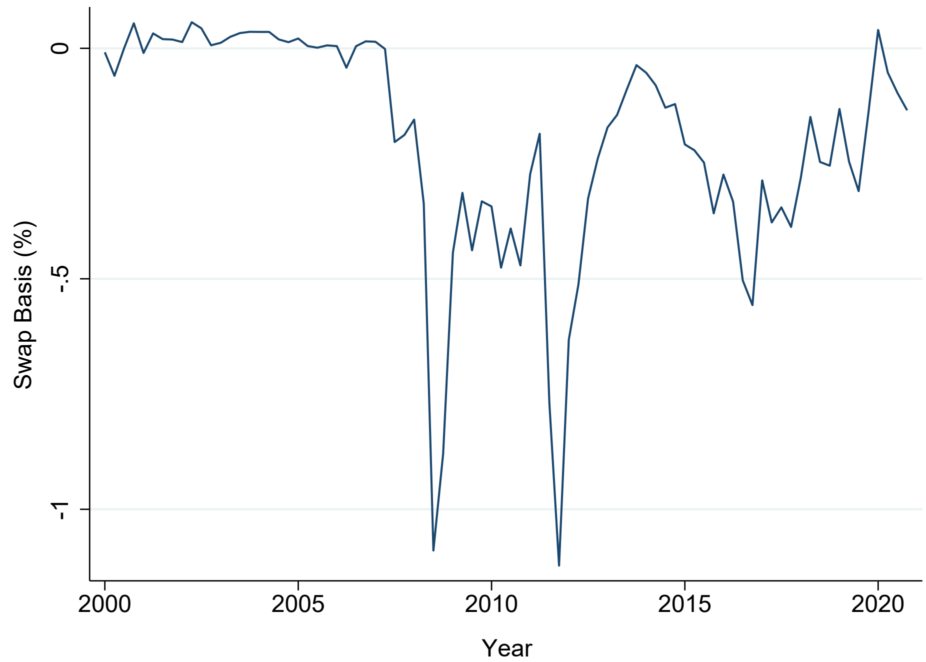
This figure plots the euro area investors' holdings U.S. nonfinancial corporate bonds by investor type from 2013:Q4 to 2020:Q4. The share outstanding is calculated as the euro area investors' holdings in Euros multiplied by the Euro-Dollar exchange rate multiplied by 100 divided by the outstanding nonfinancial corporate bond. "ICPF" stands for insurance companies and pension funds. "Banks" refers to monetary financial institutions. "Others" refers to governments and non-financial corporations. The holdings data comes from the European Central Bank's Securities Holdings Statistics. The outstanding nonfinancial corporate bonds data comes from Federal Reserve Statistical Release Z.1, Financial Accounts of the United States, Table 213 line 2.

Figure 1.10: Excess returns on U.S. bonds over euro area bonds



This figure plots excess returns on 10-year U.S. Treasury bonds over the 10-year German Bunds and on U.S. investment-grade corporate Bonds over euro area investment-grade corporate bonds from June 2011 to December 2020. The U.S. investment corporate bonds is the SP 500 Investment Grade Bond Index. The euro area investment-grade corporate bonds is the SP Eurozone Investment Grade Corporate Bond Index. The date source for the sovereign returns is the Federal Reserve Bank of St. Louis. The data source for the corporate bonds returns is SP Global Market Intelligence.

Figure 1.11: Short-term Euro/Dollar basis



This figure plots the three-month Libor cross-currency basis, measured in percentage points for euro/dollar. The data is on quarterly basis from 2000:Q1 to 2020:Q4. The date source is Bloomberg.

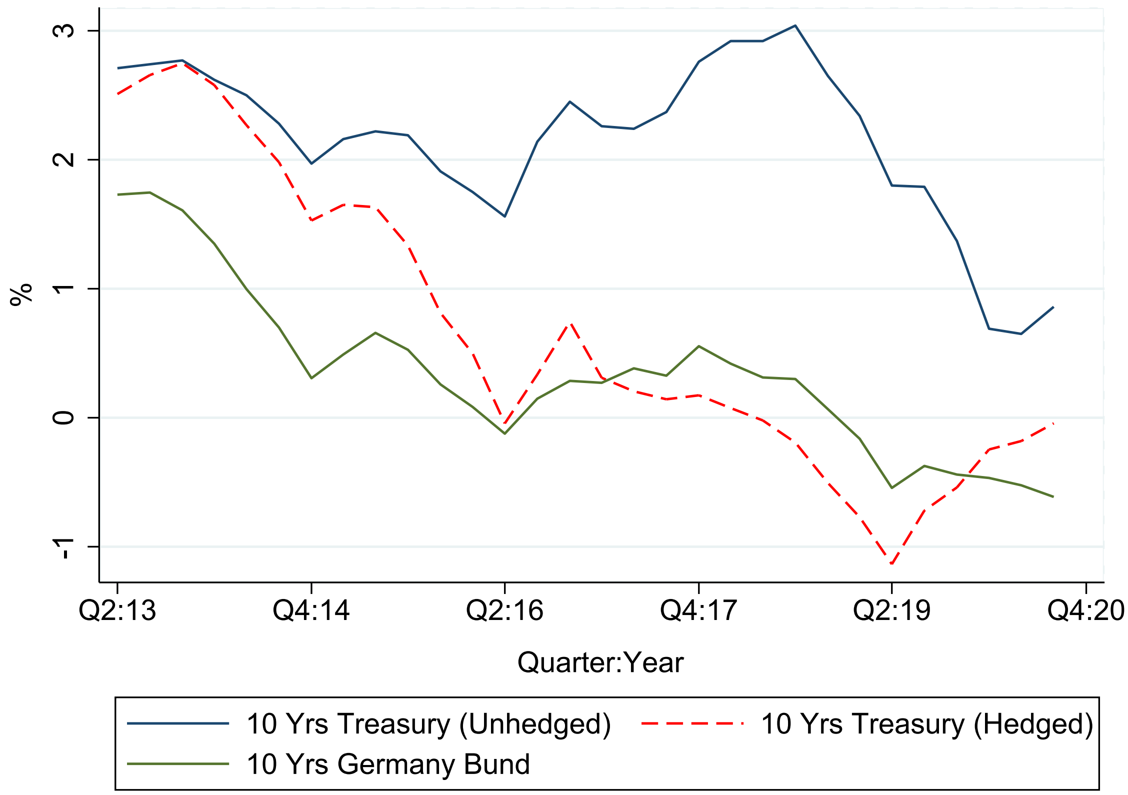
Figure 1.12: The relationship between U.S. term spread and U.S. monetary policy

This figure plots the coefficient on the federal funds rate from the following rolling window regression:

$$TS_t = \alpha + \beta\, FFR_t + \epsilon_t$$

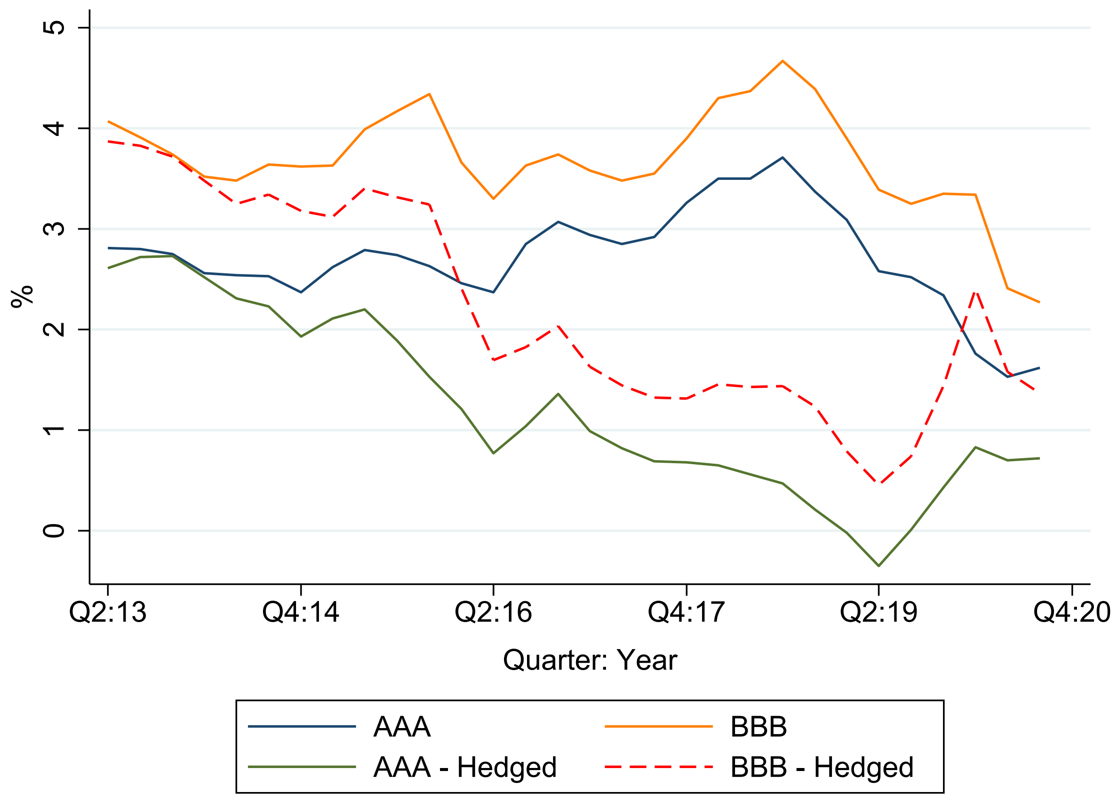
The window of the rolling regressions is 20 quarters. TS_t is the term spread in quarter t calculated as the return on the 10-years Treasury bonds minus the return on the 3-month Treasury bill and FFR_t is the federal fund rate in quarter t. The coefficients are presented in percentage points. The quarterly sample period is from 1982:Q1 to 2020:Q4. Dashed lines represent 95% confidence intervals on the point estimates for each horizon based on standard errors clustered by time.

Figure 1.13: Treasury bond returns for euro area investors



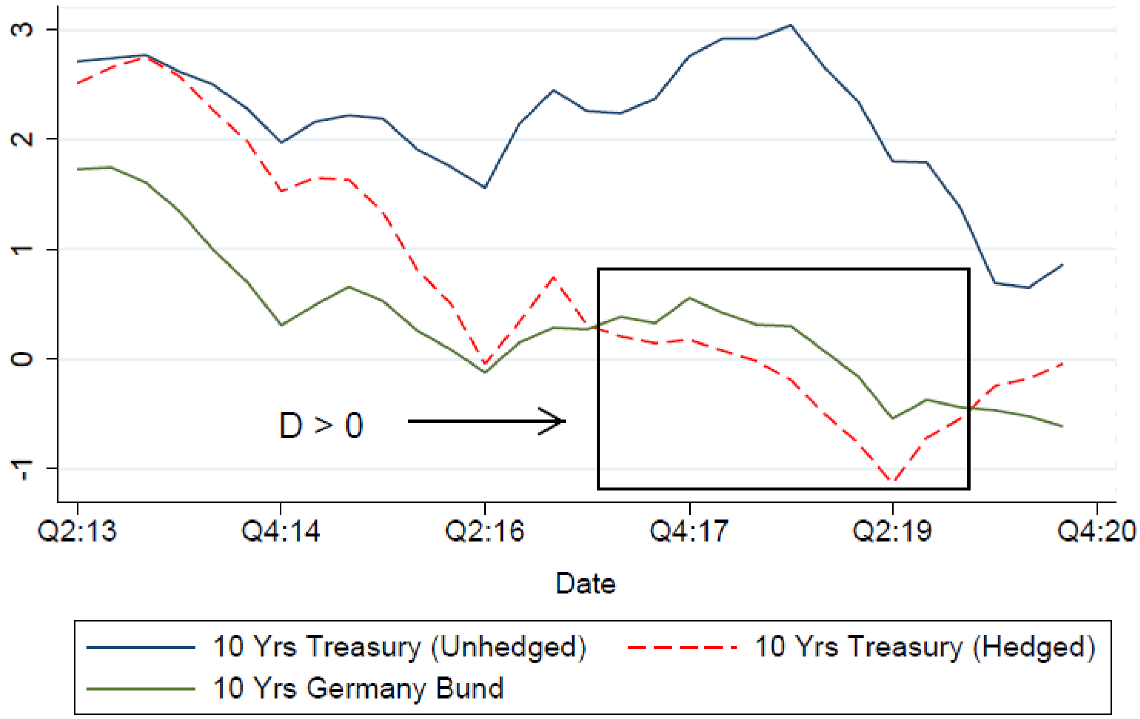
This figure plots the unhedged yields on the 10-years U.S. Treasury bonds, hedged yields on the 10-years U.S. Treasury bonds, and yields on the 10-years German Bunds on quarterly basis from 2013:Q2 to 2020:Q4. Hedged yields assume a rolling three-month Euro-Dollar cross currency swap hedge. In particular, the hedged yield is the yield on the 10-years U.S. Treasury bonds minus the 3-month Euro-Dollar swap rate. The date source for the returns on the 10-years Treasury bond and German Bund is the Federal Reserve Bank of St. Louis. The data source for the swap rate is Bloomberg.

Figure 1.14: U.S. corporate bond returns for euro area



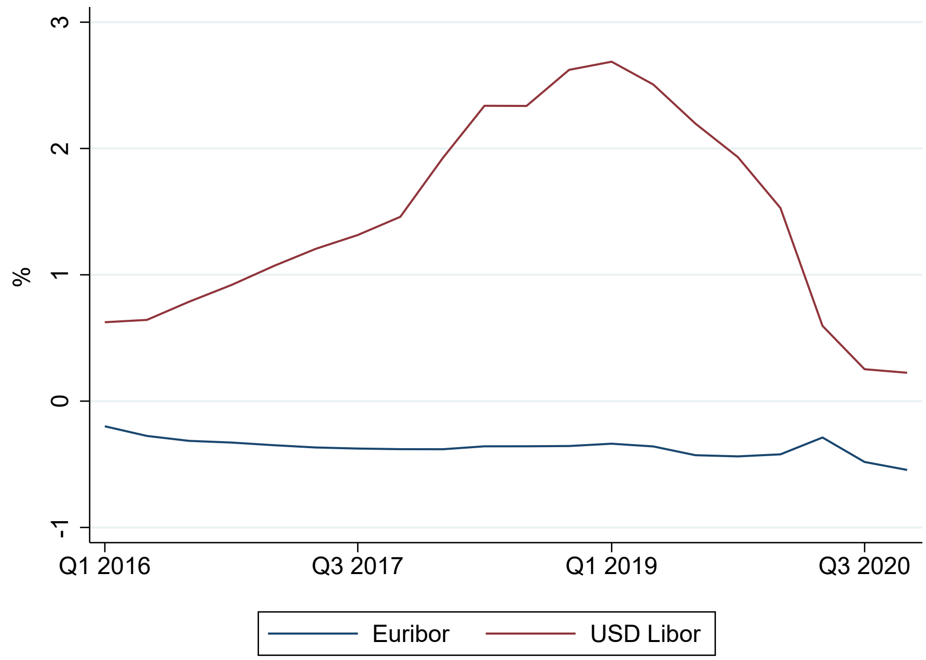
This figure plots the unhedged and hedged yields on the U.S. corporate bonds for euro area investors by credit rating. The yields are ICE BofA AAA and BBB effective yields. Hedged yields assume a rolling three-month Euro-Dollar cross currency swap hedge. In particular, the hedged yield is the effective yield minus the 3-month Euro-Dollar swap rate. The data is on quarterly basis from 20013:Q2 to 2020:Q4. The data source for the indices effective yields is the Federal Reserve Bank of St. Louis. The data source for the swap rate is Bloomberg.

Figure 1.15: Treasury bond returns for euro area investors



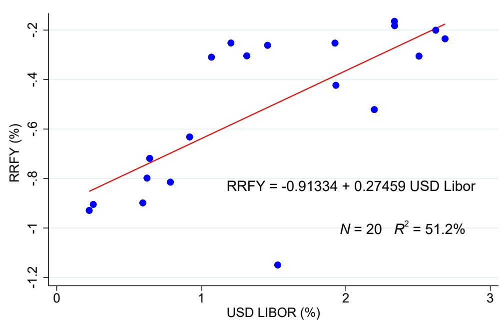
This figure plots the unhedged yields on the 10-years U.S. Treasury bonds, hedged yields on the 10-years U.S. Treasury bonds, and yields on the 10-years German Bunds on quarterly basis from 2013:Q2 to 2020:Q4. Hedged yields assume a rolling three-month Euro-Dollar cross currency swap hedge. In particular, the hedged yield is the yield on the 10-years U.S. Treasury bonds minus the 3-month Euro-Dollar swap rate. D is the return on the "hedge portfolio" which is the 10 Yrs German Bund return minus the 10 Yrs Treasury return on hedged basis. In this figure we assume the FX hedge ratio (ϕ) is 100%. The date source for the returns on the 10-years Treasury bond and German Bund is the Federal Reserve Bank of St. Louis. The data source for the swap rate is Bloomberg.

Figure 1.16: U.S. and euro area monetary policy

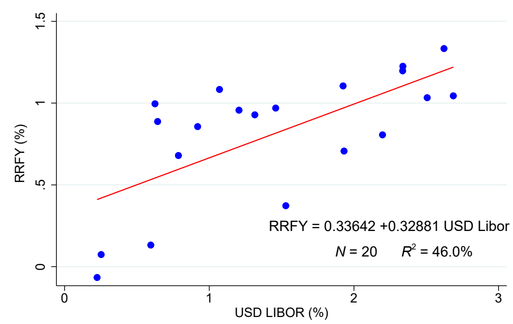


This figure plots the three months LIBOR rate on the U.S. dollar and euro as a proxy for the monetary policy rates in the two jurisdictions. The data is on quarterly basis from 2016:Q1 to 2020:Q4. The date source is the Federal Reserve Bank of St. Louis.

Figure 1.17: U.S. monetary policy and excess credit spread on EA investors' NFC bond portfolio



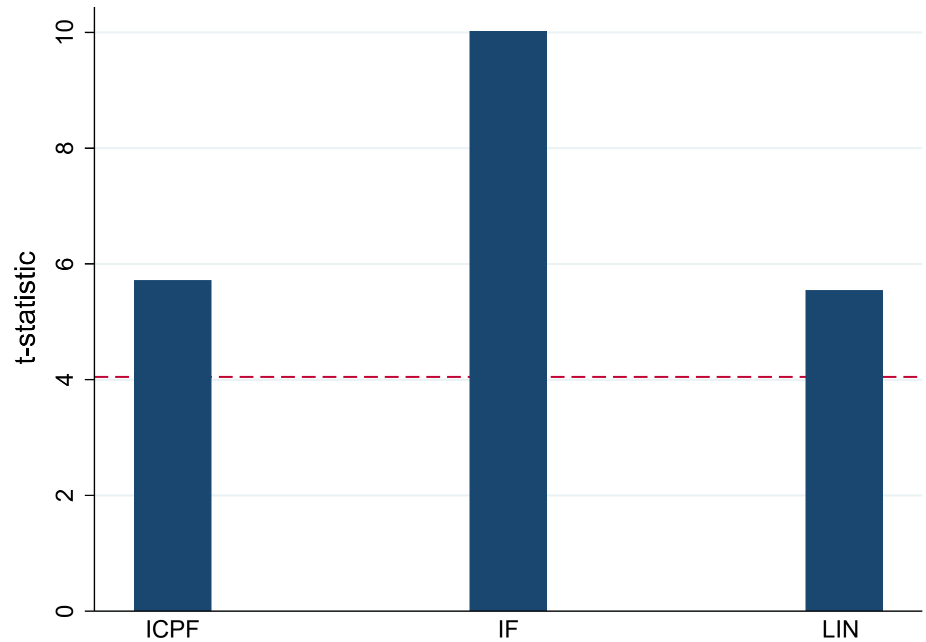
(a) EA Insurance Companies and Pension Funds



(b) EA Investment Funds

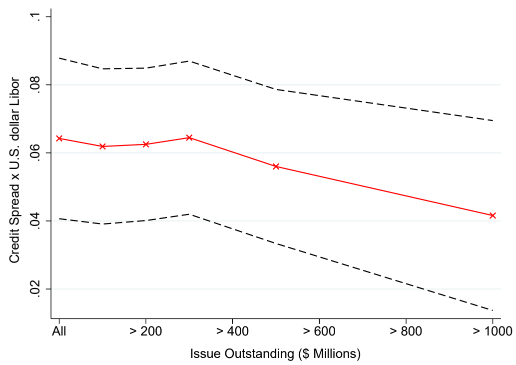
The figure plots our relative reaching for yield (RRFY) measure defined in Equation 1.9 on the three month U.S dollar Libor rate, both in percentage points. Panel A reports results for euro area insurance companies and pension funds, and Panel B reports results for euro area investment funds. The sample period is from 2016:Q1 to 2020:Q4. The data source for the holdings data is from the ECB Sectoral Securities Holding Statistics. The data source for the U.S. dollar Libor is the Federal Reserve Bank of St. Louis. The data source for all corporate bonds outstanding in the market is WRDS Bond Database.

Figure 1.18: Instrument strength: First-stage t-statistic

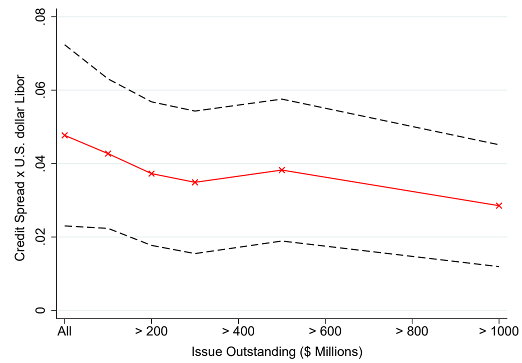


This Figure plots the minimum first-stage t-statistic across sectors on the instrument for the credit spread. The critical value for rejecting the null of weak instruments is 4.05 (Stock and Yogo (2005)). The quarterly sample period is from 2016:1 to 2020:4.

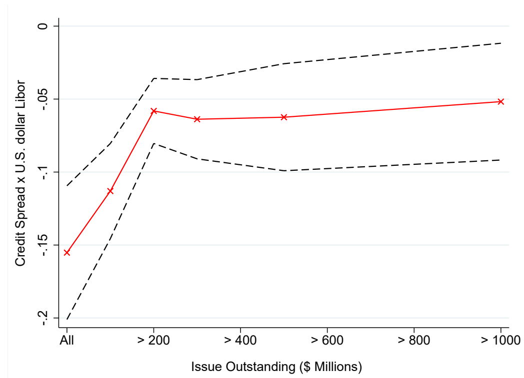
Figure 1.19: Interaction term between credit spread and U.S.dollar LIBOR



(a) EA ICPF



(b) EA Investment Funds



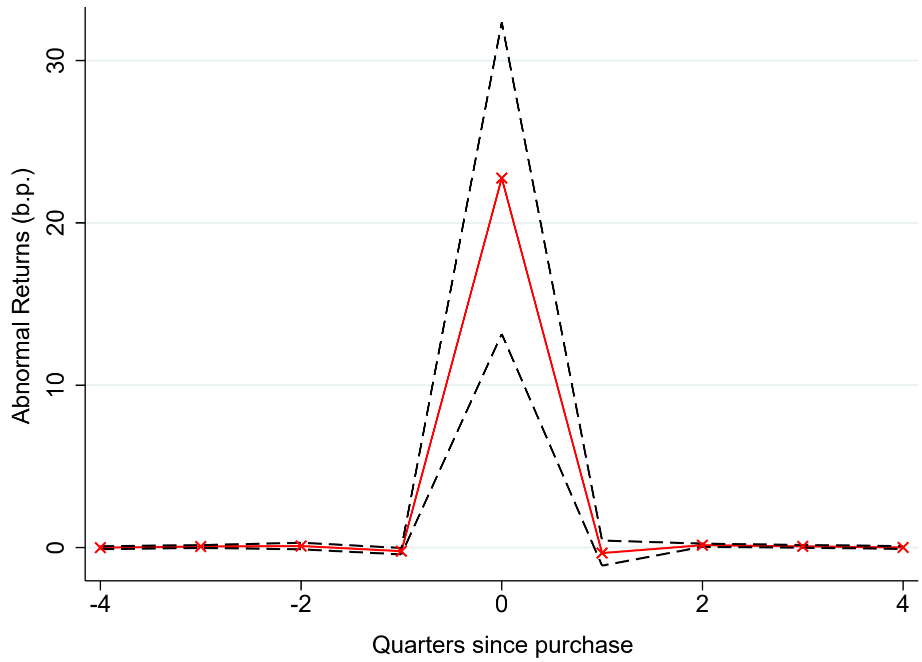
(c) U.S. Life Insurers

This figure reports the coefficient on the interaction term between bond's credit Spread and U.S.dollar Libor from the regression:

$$\log(\mathrm{H}_{i,t}(n)) = \beta_{1,i}\mathrm{CS}_{i,t}(n) + \beta_{2,i}\; y_t^{\$} + \beta_{3,i}\mathrm{CS}_{i,t}(n)\cdot\; y_t^{\$} + \beta_{4,i}\; y_t^{e} + \beta_{5,i}^{'}\mathrm{X}_t(n) + \epsilon_{i,t}(n)$$

We run the regression for different levels of issue outstanding. For euro area investors, these levels are all issue outstanding amounts, larger than € 100 million, larger than € 200 million, larger than € 300 million, larger than € 500 million and larger than € 1 billion. For the U.S. life insurers, these levels are all issue outstanding amounts, larger than $ 100 million, larger than $ 200 million, larger $ 300 million, larger $ 500 million and larger than $ 1 billion. Dashed lines represent 95% confidence intervals on the point estimates for each issue amount outstanding. Dashed lines represent 95% confidence intervals on the point estimates for each issue amount outstanding.

Figure 1.20: Quarterly abnormal return around euro area investors' purchases

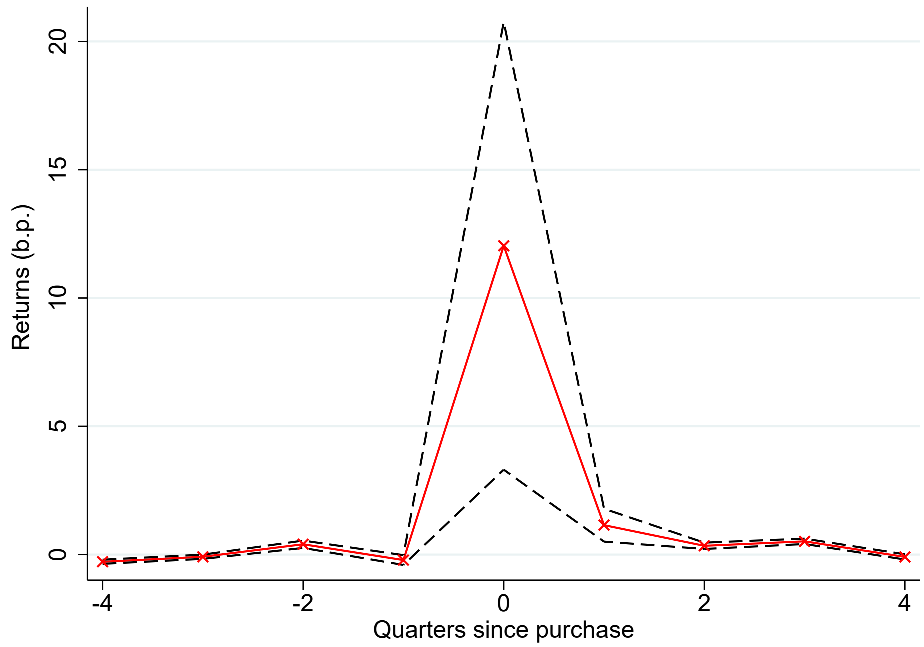


This figure plots the coefficients on the interaction term EAbuy x U.S. dollar Libor from the regression in Equation (1.15):

$$Abret_{i,t+h} = \alpha\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \gamma \mathrm{X}_{i,t} + \epsilon_{i,t}$$

The coefficients are presented in basis points. The quarterly sample period is from 2016:Q1 to 2020:Q4. Dashed lines represent 95% confidence intervals on the point estimates for each horizon based on standard errors clustered by issuer.

Figure 1.21: Quarterly raw return around euro area investors' purchases

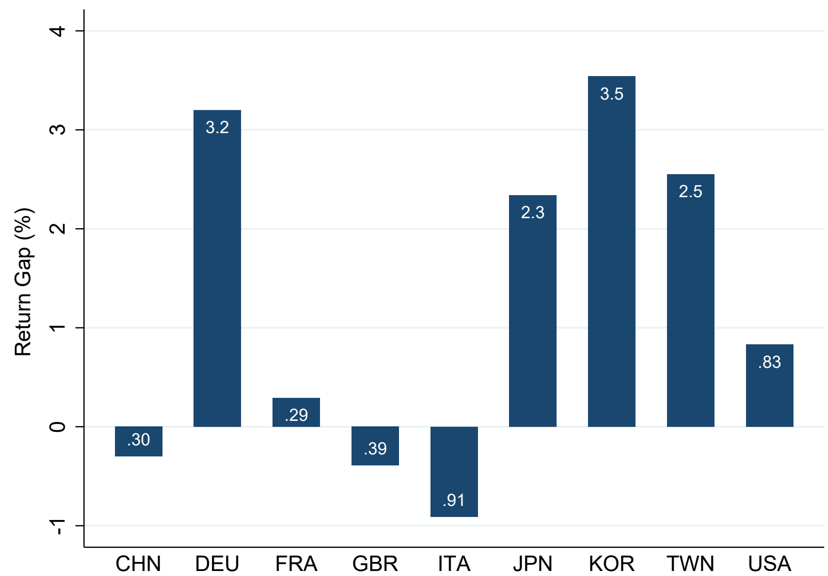


This figure plots the coefficients on the interaction term EAbuy x U.S. dollar Libor from the regression in Equation (1.16):

$$ret_{i,t+h} = \alpha_i + \alpha_t + \gamma\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \epsilon_{i,t}$$

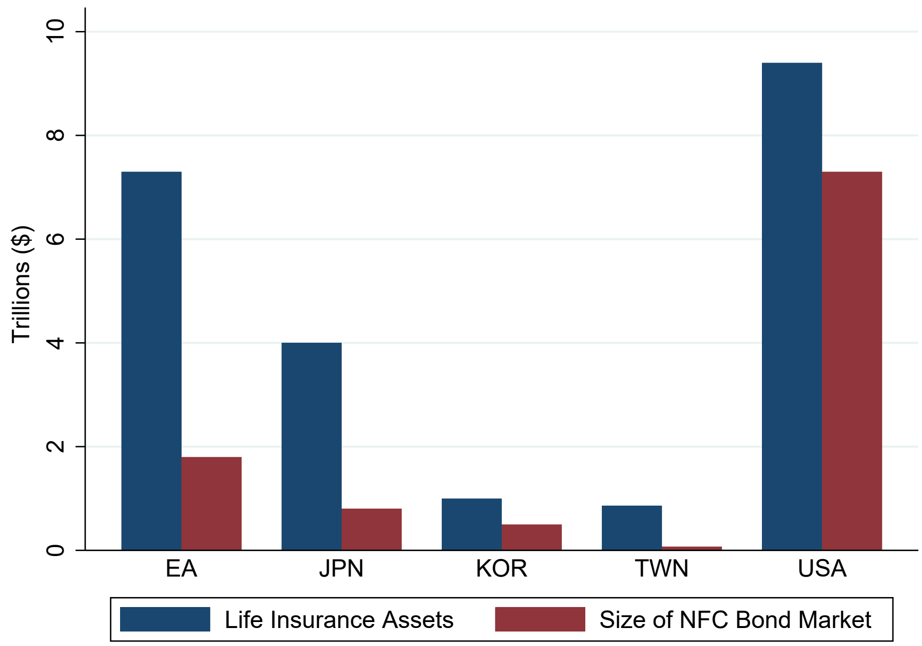
The coefficients are presented in basis points. The quarterly sample period is from 2016:Q1 to 2020:Q4. Dashed lines represent 95% confidence intervals on the point estimates for each horizon based on standard errors clustered by issuer.

Figure 1.22: Return gap for global life insurers using domestic sovereign yields



The figure plots the gap between the guaranteed returns and domestic sovereign 10-years bond yields for the nine jurisdictions with the largest life insurance sectors. The nine jurisdictions account for 73 percent of the world's life insurance premiums. Data labels use International Organization for Standardization (ISO) country codes. The data source is the October 2019 Global Financial Stability Report report by the International Monetary Fund.

Figure 1.23: Domestic nonfinancial corporate bond market and life insurance industry size



This figure plots the life insurance assets and the size of the domestic nonfinancial corporate bond market for the euro area, Japan, Koea , Taiwan and the United States at the end of 2020. The data source for the life insurance companies assets is the central bank of the jurisdiction. The data source for the outstanding nonfinancial corporate bond market is the debt securities statistics of the Bank For International Settlements.

Figure 1.24: Treasury bond returns for Japanese Investors

This figure plots the unhedged yields on the 10-years U.S. Treasury bonds and on monthly basis from 1997 to 2020. Hedged yields assume a rolling three-month Dollar-Yen cross currency swap hedge. In particular, the hedged yield is the yield on the 10-years U.S. Treasury bonds minus the 3-month Dollar-Yen swap rate. The date source for the returns on the 10-years Treasury bond is the Federal Reserve Bank of St. Louis. The data source for the swap rate is Bloomberg.

Table 1.1: Summary statistics by investor type

	Mean	Std. Dev.	10th	50th	90th	N
Euro Area ICPF:						
Credit Spread (%)	3.71	2.14	1.35	3.45	6.12	111709
Time to Maturity (30/365 convention)	9.23	8.36	1.71	6.17	25.02	111709
The par value of debt (€ Millions)	17.6	73.3	0.11	2.66	17.4	111709
Amount Outstanding (€ Billions)	1.79	49.2	0.40	1.24	3.35	111641
Euro Area Investment Funds:						
Credit Spread (%)	3.87	2.25	1.43	3.60	6.38	151718
Time to Maturity (30/365 convention)	9.66	8.56	1.63	6.42	25.02	151718
The par value of debt (€ Millions)	40.6	70.7	0.38	14.4	109	151718
Amount Outstanding (€ Billions)	1.09	31.0	0.17	0.72	2.35	150947
U.S. Life Insurance Companies:						
Credit Spread (%)	1.78	1.59	0.51	1.30	3.60	100179
Time to Maturity (30/365 convention)	10.59	9.02	2	7	26	100179
The par value of debt ($ Millions)	145.74	123.36	57.54	119.81	298.72	100179
Amount Outstanding ($ Billions)	0.66	0.61	0.25	0.50	1.25	87279
Bond Returns:						
Average Monthly Return (%)	0.63	1.72	-0.91	0.42	2.56	192933

Notes: This table summarizes bond-quarter level statistics by euro area insurance companies and pension funds (ICPFs), euro area investment funds and U.S. life insurance companies. Credit spread and bond returns are winsorized at the top and bottom 1% level. The data source for the euro area investors' holdings is the ECB Sectoral Securities Holding Statistics. The data source for the U.S. life insurerance companies' holdings is the eMAXX . Bond characteristics and yields are from the ESCB's Centralised Securities Database (CSDB). Bond returns are from WRDS Bond Returns. Sample period is from 2016Q1 to 2020Q2.

Table 1.2: Reaching for yield at the extensive margin

	ICPF	IF	ICPF	IF
USD Libor	0.439***	0.746***	0.017***	0.021**
	(0.099)	(0.220)	(0.005)	(0.009)
Post	1.656***	2.456***	0.058***	0.061***
	(0.220)	(0.487)	(0.011)	(0.020)
N	20	20	582	620
Maturity FE	No	No	Yes	Yes

Notes: Columns (1) and (2) report regression results of Equation 1.7 where the dependent variable is the difference between the weight of U.S. dollar denominated non-Treasury bonds in the euro area investor's global bond portfolio and the weight of U.S. dollar denominated Treasury bonds in the euro area investor's global bond portfolio. Columns (3) and (4) report regression results of Equation 1.8 where the dependent variable is the weight of U.S. dollar denominated non-Treasury bonds in the euro area investor's global bond portfolio with maturity of m years and the weight of U.S. dollar denominated Treasury bonds in the euro area investor's global bond portfolio with maturity of m years, where where m is 0, 1, 2, ..., 30 years. The U.S. dollar Libor is the three months U.S. dollar Libor rate. Post is a dummy variable which takes value of one for quarters after 2020:Q1. Results are based on quarterly data from 2016-Q1 to 2020-Q4. Columns 3 & 4 include maturity-bin fixed effects. Standard errors are clustered around issuers. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.3: Estimated nonfinancial corporate bond demand by investor sector

	EA ICPF	EA IF	U.S. LI
Credit Spread X USD Libor	0.064***	0.048***	-0.155***
	(0.012)	(0.013)	(0.023)
USD Libor	-0.289***	-0.160**	0.300***
	(0.062)	(0.065)	(0.043)
Credit Spread	-0.035	-0.032	0.094***
	(0.027)	(0.021)	(0.028)
Euribor	0.755***	0.738***	-0.747***
	(0.149)	(0.143)	(0.123)
USD Term Spread	-0.032	0.044	0.119***
	(0.049)	(0.061)	(0.043)
EA Term Spread	-0.051	-0.160**	-0.0474
	(0.055)	(0.067)	(0.047)
EA Credit Spread	-0.110**	-0.128***	0.361***
	(0.047)	(0.038)	(0.045)
Maturity	-0.002*	-0.003***	0.005***
	(0.001)	(0.001)	(0.001)
Outstanding	0.114***	0.155***	0.020
	(0.012)	(0.015)	(0.013)
Swap Basis	-0.039	-0.008	-0.225***
	(0.087)	(0.072)	(0.041)
Lag Log Holdings	0.928***	0.933***	0.957***
	(0.013)	(0.011)	(0.018)
Number of observations	85689	111972	84208
Number of bonds	9171	11390	6942

Notes: This table reports GMM instrumental variables estimates of nonfinancial corporate bond demand 1.11 by investor sector. The credit spread is the bond yield spread over a treasury bond with similar maturity. The USD LIBOR is the 3-month U.S. LIBOR rate. The Euribor is the 3-month Euribor. The quarterly sample from 2016Q1 to 2020Q4. Robust standard errors clustered by issuer are reported in parentheses. "EA" refers to euro area, "ICPF" refers to insurance companies and pension funds, "IF" refers to investment funds and "U.S. LI" refers to the U.S. life insurance companies. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.4: Estimated nonfinancial corporate bond demand by investor sector

	EA ICPF	EA IF	U.S. LI
Credit Spread X Swap	0.054***	0.088***	-0.013
	(0.013)	(0.017)	(0.018)
Credit Spread	-0.033	-0.159***	0.022
	(0.042)	(0.048)	(0.030)
Swap	-0.151***	-0.116***	-0.003
	(0.027)	(0.033)	(0.028)
USD Term Spread	0.184***	0.395***	-0.040***
	(0.049)	(0.066)	(0.013)
EA Term Spread	-0.236***	-0.482***	0.073***
	(0.061)	(0.087)	(0.018)
EA Credit Spread	0.075	0.251***	-0.054*
	(0.0667)	(0.088)	(0.032)
Maturity	-0.003***	-0.003***	0.007***
	(0.0009)	(0.0006)	(0.0005)
Outstanding	0.128***	0.174***	0.080***
	(0.013)	(0.018)	(0.011)
Lag Log Holdings	0.915***	0.909***	0.883***
	(0.015)	(0.016)	(0.012)
Number of observations	85690	111971	84208
Number of bonds	9170	11389	9070

Notes: This table reports GMM instrumental variables estimates of nonfinancial corporate bond demand 1.14 by investor sector. The credit spread is the bond yield spread over a treasury bond with similar maturity. "Swap" is the three month EUR-Dollar swap rate. The quarterly sample is from 2016Q1 to 2020Q4. Issuer fixed effects are included. Robust standard errors clustered by issuer are reported in parentheses. "EA" refers to euro area, "ICPF" refers to insurance companies and pension funds, "IF" refers to investment funds and "U.S. LI" refers to the U.S. life insurance companies. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.5: Monthly abnormal return around EA investors' purchases

Quarter	EA Buy	Standard Error	EA Buy x $y_t^{\$}$	Standard Error	N
-4	-0.007	0.046	-0.021	0.068	148628
-3	-0.093	0.065	0.062	0.045	155402
-2	-0.142	0.159	0.094	0.103	162433
-1	0.343**	0.1566	-0.224**	0.101	170846
0	-36.735***	7.533	22.761***	4.899	191775
1	0.470	0.619	-0.340	0.395	162248
2	–0.261***	0.0815	0.143***	0.051	145997
3	-0.166**	0.066	0.073*	0.043	131648
4	-0.0198	0.070	0.004	0.044	118323

Notes: This table reports results for the regression:

$$Abret_{i,t+h} = \alpha\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \gamma \mathrm{X}_{i,t} + \epsilon_{i,t}$$

with h $\in$ [-4; 4]. $EAbuy_{i,t}$ is an indicator variable which equals one if euro area investors buys a bond i in quarter t. $y_t^{\$}$ is the three month U.S. Libor rate instrumented with the monetary policy shocks to Libor rates. To adjust for risk, the abnormal return of bond i in quarter t, denoted as $Abreti,t$, is calculated as the quarterly average of its raw monthly return minus the monthly return on the benchmark portfolio to which it belongs. It is presented in basis points (bps). Issuer and time fixed effects are included. Standard errors are clustered around issuers and time. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.6: Monthly abnormal return around EA investors' purchases

	All	AAA/AA/A	BBB	Non-IG	NR
EA Buy x USD Libor	21.860***	5.823	18.227***	8.694	-29.330
	(4.338)	(4.749)	(5.709)	(22.187)	(120.302)
N of Observations	187860	63518	71317	22100	30722
Number of Issuers	2857	843	1218	834	581

Notes: This table reports results for the regression in Equation 1.19:

$$Abret_{i,t} = \alpha\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \gamma \mathrm{X}_{i,t} + \epsilon_{i,t}$$

$EAbuy_{i,t}$ is an indicator variable which equals one if euro area investors buys a bond i in quarter t. $y_t^{\$}$ is the three month U.S. Libor rate instrumented with the monetary policy shocks to Libor rates. To adjust for risk, the abnormal return of bond i in quarter t, denoted as $Abreti,t$, is calculated as the quarterly average of its raw monthly return minus the monthly return on the benchmark portfolio to which it belongs. It is presented in basis points (bps). $X_{i,t}$ includes coupon rate and squared coupon rate of the bond, and the remaining time to maturity. Issuer and time fixed effects are included. Standard errors are clustered around issuers. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.7: Monthly raw return around EA investors' purchases

Quarter	EA Buy	Standard Error	EA Buy x $y_t^{\$}$	Standard Error	N
-4	0.429***	0.062	-0.281***	0.041	156118
-3	0.201***	0.062	-0.086**	0.042	163677
-2	-0.558***	0.113	0.402***	0.073	171507
-1	0.447***	0.154	-0.213**	0.098	179446
0	-13.128*	7.006	12.034***	4.452	187009
1	-1.874***	0.521	1.147***	0.327	175215
2	-0.541***	0.101	0.342***	0.063	157258
3	-0.855***	0.083	0.516***	0.053	141791
4	0.178**	0.082	-0.091*	0.052	127139

Notes: This table reports results for the regression in Equation (1.16):

$$ret_{i,t+h} = \alpha_i + \alpha_t + \gamma\, EAbuy_{i,t} + \beta\, EAbuy_{i,t} \cdot y_t^{\$} + \epsilon_{i,t}$$

with h ∈ [-4; 4]. $EAbuy_{i,t}$ is an indicator variable which equals one if euro area investors buys a bond i in quarter t. $y_t^{\$}$ is the three month U.S. Libor rate instrumented with the monetary policy shocks to Libor rates. The raw return of bond i in quarter t is denoted as $reti,t$. It is calculated as the quarterly average of its raw monthly return. It is presented in basis points (bps). Bond and time fixed effects are included. Standard errors are clustered around bonds and time. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.8: Quarterly bond-level issuance around EA investors' purchases

	All	AAA/AA/A	BBB	Non-IG	NR
EA Buy x USD Libor	0.099***	0.076	0.042***	-0.023	1.168**
	(0.021)	(0.099)	(0.012)	(0.023)	(0.589)
N of Observations	56037	6715	5931	1402	41895
Number of Issuers	1486	419	557	241	485

Notes: This table reports results for the regression in Equation 1.19:

$$log(\text{Issuance}_{i,t}) = \beta \, EAbuy_{i,t} \cdot y_t^{\$} + \gamma \, \text{X}_{i,t} + \epsilon_{i,t}$$

$EAbuy_{i,t}$ is an indicator variable which equals one if euro area investors buys a bond i in quarter t. $y_t^{\$}$ is the three month U.S. Libor rate instrumented with the monetary policy shocks to Libor rates. The dependent variable is the log of bond i issuance in quarter t. $X_{i,t}$ includes coupon rate and squared coupon rate of the bond, and the remaining time to maturity. Issuer and time fixed effects are included. Standard errors are clustered around issuers. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Table 1.9: Quarterly firm-level issuance around EA investors' purchases

	All	AAA/AA/A	BBB	Non-IG	NR
Hold x USD Libor	0.084**	0.029	0.055***	0.030	0.335***
	(0.036)	(0.046)	(0.017)	(0.027)	(0.002)
N of Observations	3088	542	874	463	1111
Number of Issuers	961	177	290	162	327

Notes: This table reports results for the regression in Equation 1.20:

$$log(\text{Issuance}_{f,t}) = Hold_{f,t} \cdot y_t^{\$} + \epsilon_{i,t}$$

$EAbuy_{i,t}$ is an indicator variable which equals one if euro area investors buys a bond i in quarter t. $y_t^{\$}$ is the three month U.S. Libor rate instrumented with the monetary policy shocks to Libor rates. The dependent variable is the log of bond i issuance in quarter t. Issuer and time fixed effects are included. Standard errors are clustered around issuers. The symbols ***, **, and * indicate significance levels at 1%, 5%, and 10%, respectively.

Appendix

Implication 1

Taking the first order condition of Equation 1.3 with respect to $y_\$$, we get:

$$\frac{\partial w^*_{C_\$}}{\partial y_\$} = \frac{(\frac{2R_G D}{(1-\phi)^2 \sigma_F^2})(\frac{C_\$}{\sigma_\$^2})(1-\rho-\phi)}{S^2} \tag{1.21}$$

where $S = (\frac{D}{(1-\phi)\sigma_F})^2 + (\frac{C_e}{\sigma_e})^2 + (\frac{C_\$}{\sigma_\$})^2 > 0$. It is the sum of the squares of the risk adjusted returns. The U.S. corporate bond's allocation is decreasing in the U.S. monetary policy rate if $\frac{\partial w^*_{C_\$}}{\partial y_\$} < 0$. This is true if the following conditions are fulfilled:

1. Positive return gap: $R_G > 0$
2. Positive yield on the "hedge portfolio": D > 0
3. Positive credit spreads: $C_\$ > 0$.
4. High hedge ratio: $\phi > 1-\rho$.

Implication 2

Taking the cross derivative of Equation 1.3 with respect to $C_\$$ and $y_\$$, we get:

$$\frac{\partial^2 w^*_{C_\$}}{\partial C_\$ \partial y_\$} = \frac{(\frac{2R_G D}{(1-\phi)^2 \sigma_\$^2 \sigma_F^2})(1-\rho-\phi)[(\frac{D}{(1-\phi)\sigma_F})^2 + (\frac{C_e}{\sigma_e})^2 - (\frac{C_\$}{\sigma_\$})^2]}{S^3} \tag{1.22}$$

The higher the U.S. monetary policy rate, the stronger the demand of U.S. corporate bonds with higher credit spreads if $\frac{\partial^2 w^*_{C_\$}}{\partial C_\$ \partial y_\$} > 0$. This is true if the following conditions are fulfilled:

1. Positive return gap: $R_G > 0$
2. Positive yield on the "hedge portfolio": D > 0
3. High hedge ratio: $\phi > 1 - \rho$.
4. Higher relative credit spreads: $(\frac{C_\$}{\sigma_\$})^2 > (\frac{D}{(1-\phi)\sigma_F})^2 + (\frac{C_e}{\sigma_e})^2$.

CHAPTER 2

FISCAL STIMULUS AND PENSION CONTRIBUTIONS: EVIDENCE FROM THE TCJA

2.1 Introduction

Governments around the world are embarking on unprecedented levels of fiscal stimulus in the wake of the Covid-19 shock and the ensuing global recession. Against this backdrop, a better understanding of the effects of *temporary* measures is warranted from both an academic and a policy perspective. In particular, there is evidence that these measures can result in both short-lived and permanent effects. The literature has focused on the impact of temporary stimulus on the household sector, either as a way to test the permanent income hypothesis (Johnson, Parker, and Souleles, 2006) or to evaluate particular programmes (Mian and Sufi, 2012; Berger, Turner, and Zwick, 2020) or to combine the two (Parker, Souleles, Johnson, and McClelland, 2013). Mian and Sufi (2012) find only a temporary effect of temporary cash subsidies on household consumption of durable goods (cars). By contrast, Berger, Turner, and Zwick (2020) show evidence of a permanent effect of temporary tax incentive for new homebuyers on home sales. Comparatively little work addresses temporary policies that target the corporate sector. House and Shapiro (2008) and Zwick and Mahon (2017) study the the effects of temporarily accelerated tax depreciation on investment, but they do not investigate whether the initial positive response was subsequently reversed. This paper contributes to filling the gap by documenting an impact of temporary fiscal stimulus on corporate-sponsored pension plans.

We study the impact of the Tax Cuts & Jobs Act of 2017 (TCJA, also known as the "Trump tax cuts") on sponsor contributions to corporate defined benefit (DB) pension funds, as well as the response of the plans' funding ratios. The TCJA resulted in a temporary tax break on pension contributions. The Act permanently reduced the statutory federal corporate tax rate from 35% to 21%, beginning in 2018. Under U.S. tax law, contributions to retirement plans made in a particular year can be deducted from previous year tax returns if they are made within a "grace period" ending by the tax return due date including extensions (in practice, mid-September).[1] As a result, sponsor contributions made in both 2017 and 2018 could be deducted from 2017 income, thereby benefiting from a higher corporate tax rate. Concretely, a late-filing sponsor contributing $1bn to its DB pension plan before mid-September 2018 – rather than after the end of the grace period, for instance, in December 2018 – would have saved an extra $140m in 2017 taxes.

Because of widespread underfunding amongst corporate DB plans, whether the temporary TCJA tax break had a permanent effect on contributions – and thus narrowed plan deficits – is an especially relevant question. Despite an ongoing shift towards defined contribution pension plans (e.g. 401(k)s), corporate DB plans accounted for about 15% of US pension assets in Q4 2018 (US Financial Accounts). In the aggregate, DB plan assets fall short of liabilities, with plan funding ratios hovering around 80% for the past decade.[2] We first address the temporary/permanent effect question through the lens of a contributions model that embeds constraints on sponsor access to external

1. See Deloitte (2018). "Considerations for accelerating deductions for qualified retirement plans".

2. Funding ratio computed as in Klingler and Sundaseran (2019) using data from the US Financial Accounts (Table L.118.b).

finance. The model is built on the premise that higher contributions today lower the costs of external finance tomorrow, by improving plan funding and thus the sponsor's balance sheet. The improvement in plan funding also reduces plan (insurance) expenses and boosts tomorrow's after-tax cash flows, further reducing the future costs of external finance. At the same time, higher contributions are a drag on today's after-tax cash flows, and thus also increase the current costs of external finance.

The model suggests that TCJA should induce an increase in 2017 contributions followed by a decline in 2018 (Claim 2.3). 2017 contributions increase because, given lower expected corporate taxes tomorrow, an extra dollar put to work to reduce plan expenses has a bigger impact on tomorrow's cash flows and finance costs. 2018 contributions are different because, in addition to lower expected taxes tomorrow, sponsors also face lower taxes today. As the value of the contributions tax shield falls, the marginal cost of 2018 contributions rises. At the same time, higher 2017 contributions have already boosted plan funding. As a result, TCJA now makes tomorrow's external finance costs less sensitive to further plan improvements brought about by extra contributions today.

The model also indicates that whether the increase in 2017 contributions is large enough to offset the 2018 decline (permanent impact) or not (reversal) depends on the time profile of sponsor financial constraints (Claim 2.3). The more financially constrained a sponsor, the bigger the impact of changes in after-tax cash flows on external finance costs, and thus the larger the effect of changes in the corporate tax rate on the marginal cost and benefit of contributions. For a sponsor that expects to be more financially constrained in 2018 than in 2019, the increase in the marginal cost of 2018 contributions induced by TCJA could (in theory) cause a large enough drop in contributions to more than offset the rise in 2017 sponsor payments.

Our empirical results suggest that the contributions induced by the temporary tax break replaced contributions that would have been made in the near future anyway. A cross-sectional regression points to an above-average impact of our proxy for tax-based incentives on 2017 sponsor contributions (by 1/3 of a standard deviation).Regressing 2018 contributions on our measure of tax-based incentives returns a coefficient that is about 1/3 of a standard deviation below pre-TCJA average. Plan sponsors do respond to tax-based incentives for contributions. At the same time, they do not appear to be constrained – in setting pension plan strategies – by the amount of cash that have at hand.

In line with the result that the TCJA affected the time profile but not the overall level of sponsor contributions, we find no evidence of a long-lasting impact on plan funding ratios. Regressions of changes in funding ratios on tax-based incentives point to a relative increase of 2.5 percentage points for sponsors subject to such incentives in 2017, and a fully offsetting decrease in 2018.

Our identification strategy exploits cross-sectional differences in tax-based incentives for plan sponsoring firms, as in Gaertner et al. (2020) and Zwick and Mahon (2017). Sponsors have other, non-tax-based, time-varying incentives to shore up underfunded pension funds through higher contributions. For instance, industry newsletters often mention a sustained rise in the costs of insuring pension benefits through the Pension Benefit Guaranty Corporation (driven by deteriorating funding ratios in a prolonged low interest rate environment) as a possible driver of higher sponsor contributions.[3]

3. See Pielichata, Paulina. 2017. "Corporate pension plans push demand for Treasury STRIPS." *Pensions & Investments*, March 30. See also Kozlowski, Rob. 2018. "2018 corporate pension

By using sponsor-level data, we exploit the fact that not all sponsoring firms would have been equally affected by the increase in tax-based incentives induced by the TCJA. For a sponsor's contribution decision to respond to tax-based incentives, two conditions need to be satisfied. First, the sponsor has to have a positive corporate income tax bill before deducting contributions (tax-paying sponsor). Second, plan funding has to be below the upper bound above which contributions stop being deductible (funding ratio below 150%). We say that a sponsor is exposed to tax-based incentives if it meets both these conditions, and split our sample into tax-exposed firms and non-tax-exposed firms. Non-exposed sponsors provide a counterfactual for outcome variables in the absence of the tax break.

One possible concern about our tax exposure measure is endogeneity to subsequent firm contribution decisions. The timing of tax-based incentives for retirement plan contributions, however, suggests that a sponsor is likely to take the pre-contribution tax bill as given when choosing how much to transfer to its pension plans. The "grace period" for tax deductibility of contributions gives a sponsor the option to wait until the end of its fiscal year before deciding on its contributions, by which point there is no residual uncertainty about ex-contribution tax expenses. Empirical evidence suggests that sponsors are likely to prefer to exercise this option and to hold off on decisions regarding deductible expenses until income uncertainty is largely resolved.[4]

Our pension plan data come from yearly IRS 5500 filings of listed Compustat firms

contribution tally to top $32 billion." *Pensions & Investments,* March 19.

4. Xu and Zwick (2018) show that most of CAPEX expenses are made in the last quarter, for tax-minimising purposes.

that sponsor medium- and large-scale DB retirement plans. [5] As an alternative source of pension data, we could have used yearly SEC 10-K filings. Unlike the IRS data, however, the SEC filings data are not well suited to assessing the impact of TCJA and its reversal. First, SEC filings report contributions made in a calendar rather than a fiscal year. Therefore, contributions made in the 2017 contributions grace period (deductible from 2017 returns, and thus subject to the TCJA tax break) would be counted as part of 2018 contributions, as the 2017 grace period falls in 2018. Because of this confounding effect, 2018 contributions measured with SEC filings would be too large. Second, SEC filings do not distinguish between mandatory and voluntary contributions, making it harder to assess whether changes in contributions were driven by changes in plan service cost or changes in tax-based incentives. Third, SEC filings do not contain information on plan funding, which is necessary to control for non-tax-based contribution incentives like the PBGC insurance premium. And fourth, SEC filings do not distinguish between domestic plans and plans pertaining to foreign subsidiaries. By contrast, the TCJA tax break applies only to contributions made to domestic plans.

Other researchers have also studied firms' response to the TCJA. In a paper closely related to ours, Gaertner et al. (2020) also consider the effect of TCJA on sponsor contributions. Our analysis differs from theirs in several ways. First, we investigate whether the initial positive response was subsequently reversed. Second, we explicitly model pension contribution incentives to discipline our empirical approach. Third, we use data from IRS 5500 filings rather than from SEC filings, so we can study both contributions made in fiscal years 2017 and 2018. By contrast, 2018 contributions in

5. See Rauh (2006) and Rauh (2008) for additional information on IRS 5500 filings.

Gaertner et al. may be deductible from either 2017 or 2018 tax returns. As a result, we can document both the effects of expectations about the upcoming change in tax-based contribution incentives and its actual impact. Fourth, we broaden the analysis to funding ratios of the sponsored pension plans.

Our results have implications for work on the incidence of corporate income taxes. In particular, ignoring "uncertainty" effects on deferred compensation may lead to underestimating the incidence of corporate tax cuts on workers. To the best of our knowledge, the literature has concentrated on the current component of workers' compensation. It estimates that, on average, around 50% of the corporate tax burden is passed on to workers through changes in wages (Arulampalam, Devereux, and Maffini, 2012; Serrato and Zidar, 2016; Fuest, Peichl, and Siegloch, 2018). Current wages, however, are only one part of workers' compensation, with pensions (i.e. deferred wages) being another. Our model indicates that a temporary increase in tax-based incentives for contributions could in principle result in a permanent improvement in funding, depending on the time profile of financial constraints. The ensuing decrease in retirement income uncertainty would thus improve workers' welfare. That said, we find no evidence for this effect in the case of TCJA.

The rest of this paper is organised as follows. Section 2.2 provides an overview of the different incentives underpinning sponsor contribution choices, including the tax-based incentives directly affected by the TCJA tax break. Section 2.3 outlines a simple model that illustrates these incentives and guides the empirical analysis. Section 2.4 describes our data and explains how we constructed key variables. Results are in Section 2.5, and Section 2.6 contains concluding remarks.

2.2 Contribution incentives and the TCJA

A DB pension plan is a promise of predictable retirement benefits from a plan sponsor (typically an employer) to participants (employees). Plans are funded by employer and employee contributions. In this section we review the main factors underpinning these transfers, and we discuss how the TCJA created tax-based incentives for sponsors to increase contributions.

Since corporate DB plans are subject to funding rules under U.S. law, the size of employer contributions depends on the funding status of the plan.[6] If a plan is overfunded, its sponsor has to contribute the present value of the expected yearly change in accrued benefits (normal or service cost), net of excess assets. Sponsors of overfunded plans have little incentive to contribute more than required, as the fiscal regime penalises them for drawing down plan assets net of liabilities.[7] Sponsors of underfunded plans, by contrast, are required by law to contribute more than the service cost. The Pension Protection Act of 2006 stipulates that plan funding should equal 100% of the plan's liabilities. As a result, minimum required contributions (MRCs) are typically set according to rules which prescribe that sponsors contribute the service cost plus a fraction of the funding shortfall (shortfall amortisation charge). MRC schedules are intended to close funding deficits over a medium-term horizon. Sponsors of underfunded plans might also choose to improve funding status by making *voluntary* contributions in

6. The rules are set out in the Employee Retirement Income Security Act of 1974 (ERISA) and the Pension Protection Act of 2006 (PPA). See Manning & Napier (2014) for a concise discussion of funding and contribution rules. Firms are fined for under-contributing.

7. Proceeds from taking excess plan assets and using them for other purposes (reversions) are subject to corporate income tax plus a 50% excise tax.

excess of MRCs. Firms subject to federal corporate income taxation (C-corporations) can deduct pension contributions from tax returns. As a result, there are tax-based incentives for sponsors to contribute more than minimum requirements. Section 404 of the Internal Revenue Code (IRC) specifies that contributions made in a particular year can be deducted from previous-year income under two conditions. First, the contribution has to be made on account of pension benefits accrued in the previous year. Second, the contribution has to be made by the employer's tax return due date, including extensions. Concretely, a firm whose fiscal year ends in December (called a calendar-year firm) has until mid-October of the current year to make contributions that are deductible from the previous-year tax return.[8] In practice, if the firm's "plan year" (the 12-month period relevant for plan reporting) also ends in December, the firm would want to make contributions before mid-September. This is because contributions made after this date would not count towards satisfying minimum funding requirements under Section 430 of the IRC.[9] There are limits to deductibility: contributions are only allowed to be tax-deductible up to the point where a plan is 150% funded.

The TCJA made plan contributions counted towards 2017 sponsor income more valuable than contributions counted towards 2018 income. The Act permanently reduced the statutory federal corporate tax rate from 35% to 21%, beginning in 2018. As a result, sponsor contributions made by calendar-year firms within the grace period between

8. Calendar-year firms can either file tax returns by April 15 (on time) or apply for a 6-month extension and file until October 15 (late).

9. All firms in our sample are calendar-year firms. Plan year and firm fiscal year match by both day and month for about 95% of the firms in our sample in each year between 2014 and 2017. The share of exact matches is 99.8% in 2018. Remaining firms have pension plan years that end a couple of months earlier than their fiscal year (e.g., if a firm's fiscal year ends in December, its plan year ends either in October or in November).

January 2017 and mid-September 2018 could be deducted from 2017 income and thus reduce the corporate tax bill at the old, higher tax rate. By contrast, contributions made after mid-September were deducted at a lower rate. As an example, a late-filing sponsor contributing $1bn to its DB pension plan before mid-September 2018 – rather than after the end of the grace period (e.g. December 2018) – would have saved an extra $140m in 2017 taxes. In this sense, the TCJA included a temporary tax break on pension contributions.

Sponsors have other incentives to shore up underfunded pension plans, with rising benefit insurance premia being an oft-mentioned driver by industry commentary (Figure 2.1).[10] The retirement benefits of private sector workers are guaranteed (up to a limit) by the Pension Guaranty Corporation (PBGC), a government agency established in the mid-1970s to protect plan beneficiaries in case of sponsor bankruptcy. In addition to a flat-rate premium which applies to all plans, there is a variable-rate premium which applies *only* to underfunded plans. Variable rate premia grow with plan deficit, so employers have incentives to make voluntary contributions in order to reduce insurance costs. Sufficiently overfunded firms are exempt from paying premia altogether.

Sponsors may also worry about the impact of unfunded pension liabilities on their cost of capital and valuations, particularly if bankruptcy risk is already a concern. Since 2006, financial accounting standards require plan sponsors to "flow through" pension fund deficits into their financial statements, meaning that employers must recognise a plan's funded status on their balance sheets (FAS 158). And credit rating agencies took pension liabilities into account even prior the change in reporting standards,

10. See footnote 3.

when the funded status of plans was disclosed in financial statement footnotes (Clifton et al., 2003; Mathur et al., 2006; Campbell et al., 2012). As a result, unfunded pension liabilities can have material effects on sponsor cost of capital and equity valuations.[11]

That said, there are opportunity costs to diverting firm resources to pension plans through contributions. In the presence of financing frictions, a reduction in internal financial resources may limit a sponsor's ability to finance investment projects. Indeed, Campbell et al. (2012) show that an increase in mandatory pension contributions – which reduces a firm's ability to rely on internal financing for investment projects – increases the cost of capital for firms facing greater constraints on external financing, a result consistent with earlier evidence of a negative relationship between contributions and firm investment (Rauh, 2006).

Sponsor contributions started rising before the TCJA (Figure 2.1), an increase which would likely have continued through 2017 even in the absence of tax-based incentives. Industry commentary tends to attribute this rebound to the sharp increase in the PBGC insurance premium.[12].

2.3 Conceptual framework

This section sketches a model of contributions. The framework is designed to formalise the tradeoffs faced by a sponsor in choosing contributions (see Section 2.2), and to

11. Ang et al. (2013) illustrate the point by referring to AT&T, whose funding status changed from $17 billion surplus in 2007 to a nearly $4 billion dollar deficit in 2008. This played a role in the decline of AT&T's equity price from 2007 to 2008.

12. Industry commentary has linked 2016 growth to both expectations of lower future corporate tax rates and to an upcoming increase in the PBCG variable premium (Pielichata, 2017; Kozlowski, 2018)

derive testable predictions about the impact of the TCJA on optimal contributions. It embeds the idea that sponsors dislike plan funding deficits because they worsen the balance sheet, thereby increasing the costs of finance. As we focus on the TCJA impact on contributions, we do not model the sponsor's investment decision endogenously.[13] However, we do model the opportunity cost of diverting internal resources away from investment by letting investment returns affect the tradeoff that underpins optimal contributions.

The data only allows us to test the impact of TCJA on sponsor contributions deducted from 2017 and 2018 tax returns. Contributions counted towards 2019 tax returns are affected by the CARES Act, which gave DB sponsors the option to wait until January 2021 to make contributions deductible from 2019 returns. We thus assume there are three periods, $t = 0,1,2$. In periods $t = 0,1$ (corresponding to 2017 and 2018), the management of a firm sponsoring a DB pension fund chooses plan contributions, c_t, to maximise the value of the firm. We capture the impact of the CARES Act of 2019 on contributions by assuming that sponsors did exercise the option to wait, and we let $c_2 = 0$.[14]

The pension plan funding status affects insurance costs, contribution requirements and the costs of external finance. Plan funding depends positively on contributions and negatively on the service cost, s_t. The service cost is determined by previous decisions about wages and by factors outside of management's control (e.g. interest

13. We work with a separable specification for the costs of external finance that is linear in cash flows. As a result, if we explicitly introduced an investment decision in the model, investment would not be affected by contributions.

14. Because of the linear separable specification for the cost of finance, contributions at time t do not depend on expected contributions at time $t+1$, so there is no loss of generality.

rates), and it is therefore exogenous to current contribution and investment decisions. Letting z_t represent plan surplus – the difference between plan assets and plan liabilities – the law of motion of the funding status is:

$$z_{t+1} = z_t + c_t - s_t + \omega_t, \tag{2.1}$$

where ω_t is a catch-all random variable capturing all uncertainty about pension assets and liabilities (e.g. uncertainty about investment returns). The sponsor does not observe ω_t before choosing contributions. The funding shock ω_t is i.i.d. over time, with bounded support $\omega_t \in \left[\underline{\omega}, \bar{\omega}\right]$.

If its pension plan is underfunded, the firm has pay the variable PBGC insurance premium. The insurance premium, $q(z_t)$, is piece-wise linear:

$$q(z_t) \equiv \max\left[0, -\bar{q} z_t\right], \ \bar{q} \in (0,1), \tag{2.2}$$

with derivative $q'(z_t) = -\bar{q}$ if $z_t < 0$ and 0 otherwise. In addition, a sponsor with an underfunded pension plans must contribute more. Regulatory requirements mandate that contributions have to be at least as high as the service cost,

$$c_t \geq \max[s_t, s_t - z_t] \equiv \Psi(z_t), \tag{2.3}$$

so an underfunded sponsor with $z_t < 0$ has to contribute more that the service cost. The lower bound on contributions, $\Psi(z_t)$, is piece-wise linear with derivative equal to $\Psi'(z_t) = -1$ if $z_t < 0$ and 0 otherwise.

If a plan is underfunded, the sponsor suffers a loss. We think of the loss as a reduced form representation of the costs of obtaining external finance. Rather than modelling external finance costs endogenously, we follow Gomes (2001) and Whited (2006) and assume that when contributions are large relative to the sponsor's internal resources, the firm can only go ahead if it obtains external funds at a premium. External finance costs depend on plan funding and cash flows, x_t. In addition, the plan sponsor faces some uncertainty about the costs of finance. Concretely, we define external finance costs as

$$R(x_t, z_t) = r_0 - r_{x,t} x_t + r_z \frac{z_t^2}{2} 1_{z_t < 0} \quad \text{if } x_t < 0 \tag{2.4}$$

and 0 otherwise, with $r_0, r_z > 0$. The linear term in (2.4) implies that a larger external finance need (a more negative x_t) makes external finance more expensive. The quadratic term denotes the underfunding penalty. Conditional on there being an external finance need, a larger plan funding shortfall (a more negative z_t) increases the cost of external finance. We thus let $R_x(x_t, z_t) = -r_{x,t}$ if $x_t < 0$ and 0 otherwise; and $R_z(x_t, z_t) = z_t < 0$ if $x_t, z_t < 0$, and 0 otherwise. As r_z plays no role in our results, we set it to 1 (see the Appendix).

The external finance cost function parameter $r_{x,t}$ represents the sensitivity of external finance costs to changes in financing needs. We can thus interpret it as the shadow value of relaxing an external finance constraint, with a larger $r_{x,t}$ implying that the sponsor is more constrained. In Section 2.2 above we argued that in the presence of financing frictions, a reduction in internal financial resources may limit a sponsor's ability to finance investment projects. In order to allow for this possibility we assume that $r_{x,t} = r(a_t)$, and we interpret the shock a_t as the productivity of investment/investment returns. A higher a_t realisation is associated with a higher $r_{x,t}$ realisation. The higher investment returns, the more valuable relaxing the external finance constraint (equivalently, the tighter the constraint). Formally, we let $r_{x,t} \in [\underline{r}(a_t), \bar{r}(a_t)]$, with $\underline{r}(a_t) > 0$ for all a_t. We further assume that $r_{x,t}$ is i.i.d. and independent of ω_t, for all t, and that a_t is a mean-preserving shock, so the mean of $r_{x,t}$ is constant.[15]

The sponsor chooses contributions to maximise the value of the firm, which is equal to the present discounted value of expected cash flows. Cash flows are given by:

$$x_t = y_t - (1-\tau_t)\big(c_t + q(z_t)\big), \tag{2.5}$$

where y_t denotes the component of cash flows that does not depend on contributions. We take it as exogenous but dependent on the shock $y_t = f(a_t)$. The higher a_t, the larger y_t. τ_t is the corporate tax rate. Letting Λ^j denote the (constant) discount factor applied to cash flows received in period $t+j$, for some $\Lambda \in (0,1)$, the firm's problem is thus given by:

$$\max_{\{c_0,c_1\}} \quad \mathbb{E}_t\left[\sum_{j=0}^{2} \Lambda^j \big(x_{t+j} - R\big(x_{t+j}, z_{t+j}\big)\big)\right], \tag{2.6}$$

subject to the law of motion of plan funding (2.1), the expression for the variable insurance premium (2.2), the regulatory requirement on contributions (2.3), the definition of external finance costs (2.4), the expression for cash flows (2.5) and $c_2 = 0$. A typical specification for the discount factor would take $\beta^j u'\big(C_{t+j}\big)/u'(C_t)$ for $\beta \in (0,1)$, where $u'(C_t)$ is the marginal utility of consumption of a representative household at date

15. For example, $r_{x,t}$ could be a uniform random variable with support $\big[\underline{r} + a_t, \bar{r} + a_t\big]$.

$t+1$. By assuming a constant discount factor we are implicitly assuming linear utility.

The first order condition for contributions illustrates the intertemporal tradeoff faced by the sponsor:

$$(1-R_x(x_t,z_t))(1-\tau_t)$$
$$=\lambda_t+\mathbb{E}_t\left\{\Lambda\left[-(1-\tau_{t+1})q'(z_{t+1})\right](1-R_x(x_{t+1},z_{t+1}))-\lambda_{t+1}\Psi'(z_{t+1})-\Lambda R_z(x_{t+1},z_{t+1})\right\}. \tag{2.7}$$

Here, λ_t denotes the Lagrange multiplier on the period-t regulatory requirement on contributions (2.3). The left-hand side of this equation is the marginal cost of current contributions. Higher contributions today lower current cash flows (but less than one for one, thanks to the contributions tax shield) and possibly raise the cost of external finance, $-R_x(x_t,z_t)\geq 0$. The right-hand side represents the marginal benefit of contributing today. Higher current contributions relax the current regulatory constraint ($\lambda_t\geq 0$). They also increase next-period plan surplus. In turn, this raises next-period cash flows by lowering the PBGC premium ($-q'(z_{t+1})\geq 0$), and reduces the need to rely on external finance and the corresponding costs. At the same time, a higher future surplus slackens the regulatory constraint by lowering the minimum required contribution ($-\lambda_{t+1}\Psi'\geq 0$). In addition, a higher future surplus reduces the costs of future external finance, $-\Lambda_{t+1}R_z(x_{t+1},z_{t+1})\geq 0$.

Modelling the impact of TCJA. We now derive a prediction for the impact of the TCJA on the time profile of contributions of underfunded sponsors. To that end, we introduce a distinction between the tax rate at which contributions can be deducted, τ_t^c, and the corporate tax rate, τ_t. As a result, the tax rate entering the left-hand side of the first order condition (2.7) need not be the same as the tax rate entering the right-hand side. We assume that until period $t-1$, both tax rates are equal and constant at the level τ. In period t, it is announced that the corporate tax rate will decline from then on, $\tau_{t+j} = \tau(1-\Delta)$ for all $j \geq 0$, with $\Delta \in (0,1)$. By contrast, the tax rate relevant for contributions stays at the old level in period t, before dropping down in all subsequent period, $\tau_t^c = \tau$ and $\tau_{t+j}^c = \tau(1-\Delta)$ for all $j \geq 1$.

TCJA causes a steepening in the time profile of underfunded sponsor contributions, with c_t increasing and c_{t+1} decreasing. Consider an underfunded sponsor that expects to continue to be underfunded in the immediate aftermath of the TCJA. The firm's pension plan is underfunded in t, $t+1$ and $t+2$. As a result, the firm contributes more than the minimum requirement, $\lambda_{t+j} = 0$ for $j = 0,1,2$. The firm relies on external finance, $x_{t+j} < 0$ for $j = 0,1,2$.[16] A higher Δ increases the marginal benefit of contributing today (see (2.7)). Intuitively, given lower expected corporate taxes tomorrow $(t+1)$, an extra dollar put to work to reduce plan expenses (by lowering PBGC payments) has a bigger impact on tomorrow's cash flows and finance costs. As a result, current contributions, c_t, increase in Δ.

Future contributions, c_{t+1}, are different because in addition to lower expected taxes tomorrow, sponsors also face lower taxes today. As Δ rises and the value of the contributions tax shield falls, the marginal cost of c_{t+1} rises. Higher Δ raises both the marginal costs and the marginal benefit of future contributions, c_{t+1}. At the same time, higher c_t contributions have already boosted plan funding z_{t+1}, in turn

16. Underfunded sponsors that contribute more than the minimum requirement account for about 76% of firms in our sample 2014 to 2018. All firms in our sample are listed firms that can be expected to rely on external finance.

decreasing the expected funding gap, $\mathbb{E}_{t+1}[z_{t+2}]$. And the smaller the funding gap, the less responsive the costs of external finance to further funding improvements brought about by additional $t+1$ contributions. As a result, the TCJA has an indirect negative effect on the marginal benefit of contributions, which dampens the direct (positive) effect. The marginal cost channel then dominates and c_{t+1} falls. We formalise in the Claim below. The TCJA causes an underfunded sponsor with a need for external finance to increase $t+1$ contributions and to decrease $t+1$ contributions.

Proof. See the Appendix. □

The overall impact of the TCJA on sponsor contributions depends on the time profile of the sponsor's financial constraint. If a sponsor is more financially constrained in period $t+1$ than it expects to be in period $t+2$, then $t+1$ contributions decrease by more than t contributions increase (reversal). To see why this is the case, consider the first order conditions for $t+1$ contributions:

$$(1-\tau(1-\Delta))\left(1+r_{x,t+1}\right)=\Lambda\left(1+\mathbb{E}_{t+1}\left[r_{x,t+2}\right]\right)(1-\tau(1-\Delta))\bar{q}-\Lambda\mathbb{E}_{t+1}[z_{t+2}]. \tag{2.8}$$

The marginal cost of contributions (the left-hand side of (2.8)) increases in Δ. As the corporate tax rate decreases, the negative impact of higher contributions on current after-tax cash flows x_t rises. And the more constrained the sponsor, the larger the negative impact on current free cash flows, $x_t-R(x_t,z_t)$. The marginal benefit of contributions (the right-hand side of (2.8)) has two terms. The first one increases in Δ, because a lower corporate tax rate raises the positive impact of lower PBGC payments on after-tax cash flows. Like the marginal cost, this term is also larger for a more constrained sponsor, because an increase in after-tax cash flows x_{t+1} has a larger impact on future free cash flows, $x_{t+1} - R(x_{t+1}, z_{t+1})$. The second term, $-\Lambda R_z = \Lambda \mathbb{E}_{t+1}[z_{t+2}]$, depends on the corporate tax rate only through plan funding, z_{t+2}. By (2.1), z_{t+2} is linear in the sum of past contributions, $c_t + c_{t+1}$. Using (2.1) and (2.8) we thus obtain:

$$\frac{d(c_t + c_{t+1})}{d\Delta} = \frac{-(1 + r_{x,t+1})\tau + \Lambda\left(1 + \mathbb{E}_{t+1}[r_{x,t+2}]\right)\tau \bar{q}}{\Lambda},$$

which is negative (reversal) if the sensitivity of the marginal cost of contributions with respect to Δ, $(1 + r_{x,t+1})\tau$, is larger than the sensitivity of the marginal benefit, $\Lambda\left(1 + \mathbb{E}_{t+1}[r_{x,t+2}]\right)\tau \bar{q}$.[17] As both the marginal cost and marginal benefit of contributions are more sensitive to the corporate tax rate when the sponsor is more constrained, but the former accrues earlier than the latter, reversal happens when $r_{x,t+1}$ is sufficiently large relative to $\mathbb{E}_{t+1}[r_{x,t+2}]$. If $1 + r_{x,t+1} > \Lambda \bar{q}\left(1 + \mathbb{E}_{t+1} r_{x,t+2}\right)$, then the positive impact of TCJA on sponsor contributions in period t is more than fully reversed in period $t+1$, so the overall impact of the TCJA on sponsor contributions is negative. Otherwise, the overall impact is positive.

Proof. See the Appendix. □

We next turn to taking the predictions formalised in Claims 2.3 and 2.3 to the data.

17. This result will continue to hold even if the underfunding penalty is not quadratic, as long as $R_z < 0$ so the denominator is positive.

2.4 Data and construction of variables

Our plan-sponsor level data comes from Schedules SB and H of the electronic IRS 5500 filings from the Department of Labor. All employers sponsoring funds with more than 100 employees must file Schedules SB and H of the IRS 5500 Form on an annual basis.[18] [19] We match the plans with Compustat employers to obtain sponsor-level information.

An alternative source of pension data for Compustat firms are annual 10-K forms filed with the SEC. Unlike IRS data, SEC filings data do not include minimum required contributions, making it harder to disentangle the voluntary component of contributions from the mandatory. In addition, pension variables obtained from SEC filings (contributions, plan assets and liabilities) do not distinguish between domestic plans and plans pertaining to foreign subsidiaries. By contrast, the TCJA tax break only applies to contributions made to domestic plans. Similarly, the PBGC premium only applies to funding shortfalls of domestic plans.

2.4.1 The sample

Our sample starts in 2014 to avoid possible confounding effects from the Transportation Bill of June 2012 (Moving Ahead for Progress in the 21st Century, or MAP-21). MAP-21 allowed single-employer plans to discount liabilities using a rolling average of yields

18. See Rauh (2006) and Rauh (2008) for additional information on IRS filings.

19. Plans with less 100 participants must file Schedule SF. This form includes very limited information on funding ratio, number of participants and investment income. Compustat firms, which are listed companies, usually don't sponsor such small plans.

over the previous 25 years instead of over the previous 2. With interest rates at historical lows, the change amounted to an increase in the discount factor, which boosted plan funding ratios and lowered contribution incentives.[20] The sample ends in 2018 to avoid confounding effects from the Coronavirus Aid, Relief and Economic Security (CARES) Act of March 2020. As part of a broader effort to mitigate the Covid-19 shock, CARES afforded DB plan sponsors the option to defer 2020 contributions (deductible from 2019 returns) until January 2021.

To ensure that all sponsors have an equal amount of time to respond to the TCJA tax break, we restrict the sample to plans sponsored by firms whose fiscal year ends in December. About 79% of sponsors (585 firms) in our matched sample are calendar year firms. We end up with a sample of 4,105 plan-year observations and 2,506 firm-year observations (some employers have multiple plans) that were matched to Compustat.[21] According to the financial accounts of the United States, the assets held by our sample plans in 2017 represent about 30% of total private DB plan assets as of 2017 Q4 (single- and multi-employer). They account for 43% of the total assets held by all single-employer pension plans that filed the IRS Form 5500. We turn to the construction of variables and the corresponding summary statistics next.

20. van Binsbergen and Brandt (2016) calculate that reported liabilities fell to half of their market value in 2012.

21. The number of firms filing IRS 5500 is decreasing over time, consistent with an ongoing shift away from DB plans in the U.S. private sector.

2.4.2 Outcome variables

We study the impact of the TCJA tax break on voluntary sponsor contributions, total sponsor contributions and plan funding.

Contributions

As discussed in Section 2.2, tax-based incentives affect only the *voluntary* component of sponsor contributions. We compute voluntary contributions by a particular sponsor to a particular plan by subtracting mandatory contributions from total contributions.[22] We define the mandatory component of pension contributions as the sum of minimum required contributions (both legacy and current) and of special contributions made to avoid restrictions on the timing of benefits payment for underfunded plans.[23] As a firm may sponsor multiple plans, we aggregate over all the plans sponsored by the same firm to obtain sponsor-level contributions (Voluntary Contributions and Total Contributions, respectively).[24]

As larger firms naturally tend to contribute more (for example, because they have higher service costs), we scale both our contribution variables – voluntary and total – by sponsor size, captured by sponsor assets at the beginning of the current year (Assets (book)). Normalising by firm assets is standard in papers studying either pension

22. Contribution figures reported on Schedule SB as year t contributions take into account transfers made by the sponsor up to the point of filing year-t tax returns, and thus include any transfers made within the "grace period" for contributions in year $t+1$.

23. The PPA imposes benefit restrictions that constrain sponsors of underfunded plans from improving or accelerating the payment of benefits. For example, plans are not allowed to pay lump sum benefits if they are less than 60% funded.

24. See Tables 2.9 and 2.10 for more detail on the construction of variables.

contributions from the sponsor's perspective (e.g. Rauh, 2006) or the impact of tax-based incentives on other firm choices, such as capital expenditures (e.g. Zwick and Mahon, 2017; Xu and Zwick, 2018).[25]

Voluntary contributions represent 0.02% of sponsor assets at the mean and 0.002% of assets at the median (Table 2.1). Total pension contributions are 0.31% of assets at the mean and 0.05% at the median. Both total and voluntary contributions grew in 2016 and in 2017. They declined sharply in 2018 (Figure 2.2, left panel).

Funding ratios

To assess whether the TCJA tax break had an impact on plan funding, we consider the change in funding ratios between 2016 and 2017. We chose this period because contributions made up until the end of the contribution "grace period" (in theory, mid-October 2018; in practice, mid-September 2018), are counted towards 2017 contributions for financial reporting purposes and thus flow into 2017 assets and funding. We compute the funding ratio for a particular plan-sponsoring firm in any given year (Funding Ratio) in a few steps. First, for each plan sponsored by a particular firm, we sum of reported plan assets (Assets) and sponsor contributions (Total Contributions) net of credit balances (Credit Balances).[26] Reported plan assets are measured at year-end market value, and they *do not* include contributions. Second, we aggregate the resulting plan-level asset measure over all the plans sponsored by the firm, and we

25. Other normalisations are appropriate when thinking about contributions from a plan's perspective (e.g. contributions as a share of plan assets or as a share of service cost).

26. Credit balances arise when an employer chooses to credit current voluntary contributions towards satisfying future minimum funding requirements and the ensuing minimum required contributions.

thus obtain the funding ratio numerator. To get the denominator, we sum plan-level liabilities (Liabilities) over all the plans sported by the firm. Liabilities are the present discounted value of future pension benefits accumulated to year-end. MAP 21 allows sponsors to discount plan liabilities using an average of market rates on corporate bonds over the past 25 years.[27] Plans in our sample are 107.6% funded at the mean and 104.4% funded at the median, with a standard deviation of 16%.

2.4.3 *Explanatory variables*

Tax-based incentives

For the TCJA tax break to affect voluntary contributions, two conditions need to be satisfied. First, the firm has to have a positive corporate income tax bill before deducting contributions (tax-paying sponsor).[28] Second, the funding ratio has to be below the 150% bound above which contributions stop being deductible (funding ratio below 150%), for at least one of the sponsored plans.

We say that a sponsor is *exposed* to tax-based incentives – including the TCJA tax break – if it meets both these conditions. We define the tax exposure of sponsor s at time t (Tax Exposure) as a dummy variable which is equal to 1 if Gross Tax > 0 and if Funding Ratio $< 150\%$ for at least one plan i of sponsor s. Here, Gross Tax denotes the

27. With interest rates at historical lows, these regulatory discount rates are higher than the discount rates used in the Financial Accounts of United States, which are based on AAA-rated corporate bond rates (Stefanescu and Vidangos, 2014). As a result, average funding ratios in our sample are higher than funding ratios derived from the flow of funds (Figure 2.2, centre panel). Financial Accounts data point to average funding of 85.5% between 2014 Q1 and 2018 Q4.

28. Gaertner et al. (2018) also employ this condition to assess the impact of TCJA. Zwick and Mahon (2017) use it to assess the impact of tax-based incentives on firm investment.

Federal corporate tax bill of sponsor s before deducting pension contributions. Since we do not observe Gross Tax, we obtain it by adding back the contribution deduction to the corporate tax bill from Compustat. Concretely, Gross Tax = Net Tax + $\tau \times$ sum of Total Contributions over sponsored plans, where Net Tax is the Federal corporate income tax expense from Compustat and τ is the statutory corporate income tax rate.

By using Tax Exposure as a proxy for sponsor exposure to the TCJA tax break, we assume that the gross corporate tax bill (Gross Tax) is exogenous to the sponsor's contribution decision. This assumption is justified by the timing of tax-based incentives for retirement plan contributions, which suggests that a sponsor is likely to take the pre-contribution tax bill as given when choosing how much to (voluntarily) transfer to its pension plans. To account for any possible endogeneity and as a robustness check, we will rerun the analysis by relaxing the definition of tax exposure of sponsor s at time t (Tax Exposure) as a dummy variable which is equal to 1 *only* if Funding Ratio < 150% for at least one plan i of sponsor s.

Tax-based incentives for sponsors to contribute could be captured by other proxies. These include estimates of corporate marginal tax rates (Graham, 1996a,b) and measures of tax exposure based on sponsor tax credits such as net operating loss carryforwards and investment tax credits.[29] According to the latter set of proxies, a firm is *not* exposed to tax-based incentives if its accumulated tax credits are large enough to cause it not to report any taxable income. Our tax-based incentives measure is positively correlated with marginal corporate tax rates, and negatively correlated a set of dummies capturing

29. Net operating losses arise when taxable corporate income falls short of applicable deductions. They can be carried forward, meaning that losses occurred in a particular year can be used to abate taxable corporate income in subsequent years. In this sense, past net operating losses result in current tax credits.

lack of exposure due to tax credits (see Table 2.2).

There are disadvantages to using corporate marginal tax rate estimates or exposure measures based on accumulated tax credits in order to capture the impact of the TCJA tax break on sponsor contributions. First, marginal tax rates may not be the relevant tax rates for sponsor contribution decisions. There is evidence that firms may prefer to use simple heuristics such as statutory and effective tax rates to evaluate incremental decisions, rather than harder-to-estimate marginal tax rates (Graham, Hanlon, Shevlin, and Shroff, 2017). This suggests that our tax-exposure measure, which is based on the statutory tax rate, is a more suitable proxy than the marginal tax rate for capturing the impact of tax-based incentives on sponsor contributions. Second, the tax credit dummies might incorrectly classify some sponsors as not exposed to the TCJA tax break. This is because the exposure measures based on accumulated tax credits reported in Compustat include tax credits accrued to foreign subsidiaries, as well as domestic subsidiaries which are unconsolidated for tax purposes (THO, 1988). By contrast, pension contributions are deducted from corporate income net of income from such subsidiaries, so sponsors may be subject to tax-based incentives even if the no-exposure dummies are equal to 1.

2.4.4 Controls for other contribution incentives

As we argued in Section 2.2, contribution incentives are affected by insurance premia. As PBGC insurance premia depend on plan funding, we include funding ratios as a control in our regressions. We also add controls for sponsor bankruptcy risk – because pension deficits flow through to sponsor balance sheets – and for the opportunity cost of diverting internal financial resources to shoring up pension benefits.

To control for sponsor bankruptcy risk, we use the Altman's Z-score, a weighted average of standard business ratios (working capital, operating earnings, sales, and retained earnings). To account for the opportunity cost of diverting internal financial resources to funding pension benefits, we use sponsor cash flows excluding contributions (Cash Flows), capital expenditures (CAPEX), earnings distribution to investors (Payout) and Tobin's Q (i.e. the market-to-book ratio of firm assets).

2.4.5 Other controls

We control for plan performance by including investment returns (Return on Investment) and liability discount rates (Discount Rate). Both these variables vary at the sponsor level. The data come from Schedules H and SB of the IRS filings, respectively. Return on Investment is a weighted average of returns over sponsored plans, with weights proportional to plan assets. Plan returns are calculated as investment income divided by beginning-of-year investable assets (measured ex-contributions). Discount rates are the interest rates used to compute the present discount value of the pensions liability of a particular plan. Discount rates for US corporate DB plans are regulated and decoupled from expected plan returns. Under MAP-21, funds discount using an average corporate bond yield over the past 25 years, with a corridor around this average.[30] The discount rate at the sponsor level is computed as the weighted average of discount rates across all sponsored plans, with weights proportional to plan liabilities.

30. The corridor was ±20% in 2014, ±25% in 2015 and ±30% since 2016. See Novick, Hunt, Ransenberg. 2012. "Corporate Pension Funding Update". *Blackrock White Papers*.

Discount rates are 6.07% and 6.08% at the mean and median, respectively. By contrast, the average yield of a 30-year Treasury bond was 2.95% over our sample period.

Finally, we include a proxy of the relative importance of DB plans for a particular sponsor, the idea being that the larger the relative importance of DB plans in a firm's pension benefits, the more likely the sponsor to shore up those plans (for example, in order to retain current employees). We proxy the relative importance with the ratio of the total number of participants in DB plans to the current number of employees of the firm (DB Plans Share). The "significance" measure is 1.15 and 0.84, at the mean and the median, respectively.

2.5 The effect of TCJA on pension plans

2.5.1 Identification strategy

Our empirical strategy exploits cross-sectional variation in sponsor exposure to tax-based incentives to assess the impact of the TCJA tax break. We use non-tax-exposed sponsors as a control group to assess the counterfactual level of voluntary and total pension contributions in the absence of the tax break for the tax-exposed firm.[31] This allows us to estimate of the marginal impact of the TCJA tax break on contributions and funding. The identification strategy depends on the assumption that tax-exposed (treatment) and non-tax-exposed control) firms do not differ across dimensions other than tax-based incentives that may affect voluntary contributions during the sample

31. Given the firm's other incentives to shore up underfunded pension plans, it would be difficult to estimate counterfactual outcomes using aggregate data.

period.

Exposure to tax-based incentives is not random in our sample. Table 2.3 reports the correlation of our measure of tax-exposure with other variables that are likely to affect pension contribution patterns: plan funding (Funding Ratio), profitability metrics (Return on Investment and Discount Rates), PBGC premia, proxies for sponsor bankruptcy risk (Altman's Z) and the opportunity cost of internal resources (Cash Flows, CAPEX, Payout, Tobin's Q). Tax-exposed firms have more underfunded pension plans, higher PBGC variable premium, higher payout payout and higher pre-contributions cash flows, which all push for higher contributions. At the same time, tax-exposed sponsors have higher pension liability discount rates, which would tend to reduce sponsor incentives to contribute. The tax-exposed also display lower CAPEX, which could be associated with relatively lower contributions if resulting from more binding constraints on external finance. On balance, it is not obvious that the significant correlates in Table 2.3 will bias our estimates in a specific direction. To account for all possible biases, we include the observable correlates as controls in our empirical specifications.

2.5.2 The TCJA tax break and contributions

The conceptual framework outline in Section 2.3 indicates that TCJA should have has a positive impact on 2017 contributions. By contrast, it should have had a negative impact on 2018 contributions (Claim 2.3). Therefore we should expect 2017 contributions from tax exposed firms to be higher than those of non-tax-exposed firms. We should expect the opposite for 2018 contributions. The model also suggests that whether the drop in 2018 contributions is large enough to offset the 2017 increase depends on the time profile of sponsor financial constraints (Claim 2.3). Empirically, we can gauge whether this was the case or not by comparing the 2017 (differential) response of tax-exposed firms to tax-based incentives to their 2018 response.

A graphical analysis suggests that the TCJA tax break had a positive impact on 2017 contributions and a negative impact on 2018 contributions, as expected. We split the sample into two groups according to tax-based incentives, proxied by our tax exposure measure. The first group includes firms that have pre-pension contribution tax-based incentives (i.e. Tax Exposure = 1) and the second group includes firms that have no pre-pension contribution tax-based incentives (i.e. Tax Exposure = 0). The left panel of Figure 2.3 plots average yearly voluntary pension contributions from 2014 through 2018, for both tax-exposed and non-tax-exposed firms. The difference between voluntary contributions from tax-exposed and non-tax-exposed sponsors was relatively stable prior to the TCJA tax break (2014 to 2016). In 2017, contributions from tax-exposed sponsors increased by 0.008% of sponsor assets. By contrast, contributions from non-tax-exposed sponsors increased by 0.0024%. In 2018, pension contributions from tax-exposed firms decreased significantly relative to those of non-tax-exposed firms. Given the more permanent nature of changes in other time-varying contribution incentives (such as increases in the PGBC variable premium), it is difficult to argue that this increase/decrease pattern can be accounted for by something other than TCJA and the ensuing temporary tax break.

Regression analysis confirms the findings of the graphical analysis on impact. We estimate the following cross-sectional specification:

$$\frac{\text{Voluntary Contributions}_{s,t}}{\text{Assets (book)}_{s,t-1}} = \alpha_t + \beta_t \text{Tax Exposure}_{s,t} + \delta_t Z_{s,t} + \varepsilon_{s,t}, \text{ for } t = 2014, ..., 2018. \tag{2.9}$$

Here, Tax Exposure is our measure of tax-based incentives and Z is a vector of controls which includes the observable correlates of tax-based incentives. The β_t coefficients are the coefficients of interest, as they capture the impact of tax deductibility of contributions on contributions in each of our sample years. We plot them on the right panel of Figure 2.3. A priori, we would expect the impact of tax-based incentives in 2017 to be above pre-TCJA average, $\beta_{2017} > \sum_{t=2014}^{2016} \hat{\beta}_t/3$, and the impact of tax-based incentives in 2018 to be below, $\beta_{2018} < \sum_{t=2014}^{2016} \hat{\beta}_t/3$.

Tax-based incentives had a larger impact on contributions in 2017 than in the three years pre-TCJA. The 2017 estimate of the tax exposure coefficient, $\hat{\beta}_{2017}$, is positive and significant (Table 2.4, columns (1) and (2)). This result is robust to including controls for the observable correlates of our tax exposure measure (column (2)), assuaging concerns about identification. According to our preferred specification (with controls, column (2)), voluntary contributions from tax-exposed sponsors were 0.037 percentage points larger than their counterpart from non-tax-exposed firms. By contrast, the average impact of tax-based incentives on voluntary contributions prior to the TCJA, $\sum_{t=2014}^{2016} \hat{\beta}_t/3$, was around 0.022 percentage points, making the 2017 impact about one third of a standard deviation higher than the pre-TCJA average.

The impact of tax-based incentives on 2018 contributions was below pre-TCJA average, with a large enough deviation to fully offset the above-average 2017 effect (reversal). The 2018 estimates of the tax exposure coefficient, $\hat{\beta}_{2018}$, are at the minimum level over our five-year sample period (Table 2.4, columns (1) and (2)), implying that tax-based incentives to contribute were at their weakest right after the end of the tax break. In our preferred specification (with controls, column (2)), the impact of tax-based incentives in 2018 amounted to 0.008 percentage points. At around one third of a standard deviation lower than pre-TCJA average, this decline fully offset the 2017 increase. We interpret this as evidence that tax-exposed firms shifted planned future contributions from 2018 to 2017. Through the lens of our model, the fact that 2018 tax-based incentives completely reversed the effects of 2017 incentives on contributions is consistent with sponsor financial constraints being relatively more binding than future expected constraints (Claim 2.3).

These results are robust to changing the definition of tax exposure to a dummy-variable equal to 1 if only the firm sponsors at least one plan with funding ratio below 150% (Tables 2.5 and 2.6). These results are also robust to including sector fixed effects. Estimates with sector fixed effects are qualitatively similar to estimates without (Table 2.7). The second column of Table 2.7 changes the dependent variable to total pension contributions. We continue to find a positive impact of the TCJA tax break in 2017, followed by a reversal in 2018.

In dollar values, our estimates imply a \$2.8bn to \$5.0bn increase in voluntary contributions to medium- and large-scale plans associated with the tax break, depending on whether or not the specification includes controls (the impact is larger with controls). Our estimates report the TCJA impact in percentage points, so we multiply by tax exposed sponsor assets to obtain a dollar figure. Accordingly, the tax break impact on voluntary contributions in dollars in sample is given by $\left(\hat{\beta}_{2017} - \sum_{t=2014}^{2016} \hat{\beta}_t/3\right) A_{2016}/100$, where $A_{2016} = \sum_{s=1}^{381} A_{s,2016}$ represents the total assets of the 381 tax-exposed sponsors in our sample at the end of 2016 (beginning of 2017). We obtain a \$1.3bn increase in voluntary contributions for the specification without controls and a \$2.3bn increase for the specification with controls. By assumption, the TCJA had no impact on contributions from the non-tax-exposed. Assuming that our sample is representative of the broader population of firms submitting Schedule SB of the IRS 5500 filings – some of which are not listed, and therefore do not appear in Compustat – we extrapolate to estimate the TCJA impact on the voluntary contributions of *all* sponsors of middle- and large-scale plans. To that end, we multiply the in-sample estimates by the ratio of total voluntary contributions by Schedule SB filers to total voluntary contributions by sponsors in our sample, which is equal to \$6.7bn/\$3.1bn. To compute the tax break impact on total contributions for firms in sample, we repeat the same steps using the estimates in Table 2.7 instead. This returns a \$15.3bn increase in voluntary contribution for the specification without controls and a \$37bn increase for the specification with controls. Total contributions by firms in our sample amount to \$50bn, while total contributions by Schedule SB filers are equal to \$107.7bn. This implies a \$33bn to \$79.7bn increase in total contributions to medium- and large-scale plans associated with the tax break.

2.5.3 *The TCJA tax break and funding ratios*

In this section, we study other effects of the TCJA tax break on firms and their DB retirement plans. We examine whether or not the tax break had an impact on funding ratios. We find that our results on contributions carry over to plan funding ratios.

Our estimates suggest that the TCJA tax break had a short-lived impact on plan funding. While the TCJA increased 2017 funding ratios, by 2018 they were already back where they would have been in the absence of the intervention. We estimate the following specification:

$$\Delta \text{Funding Ratio}_{s,t,t-1} = \alpha_t + \beta_t \text{Tax Exposure}_{s,t} + \delta_t Z_{s,t} + \varepsilon_{s,t}, \quad \text{for } t = 2017, 2018. \tag{2.10}$$

Here, Funding Ratio is defined as in section 2.4.2 and Z_t is a vector of controls which includes pre-TCJA plan funding status (Funding Ratio in 2016), the actual investment return on plan assets and the change in discount rates between t and $t-1$. Results are reported in columns (1)-(4) of Table 2.8. Tax-exposed firms experienced a increase of 2.5 to 3.4 percentage points in the funding status of their corporate pension plans between 2016 and 2017 (relative to non-tax-exposed firms), depending on whether or not the specification includes controls. Firms that were tax exposed in both 2017 and 2018 saw a relative decrease of 2 percentage points in the funding status of their corporate pension plans between 2017 and 2018. Columns (5) and (6) report the results of estimating a variant of (2.10) which considers the change in funding ratios between end-2016 and end-2018, again focusing on firms that were tax exposed in both 2017 and 2018. The coefficient of 2017 Tax Exposure is not significant, confirming that the temporary increase tax incentives for contributions associated with the TCJA had no long-lasting impact on funding ratios.

2.6 Conclusion

This paper contributes to the literature studying the effects of temporary fiscal stimulus on the corporate sector by documenting that sponsor contributions to retirement plans respond to tax-based incentives. We first develop a simple model to derive conditions under which temporary changes in tax-based incentives may result in permanent changes in contributions and plan funding. We then take these predictions to the data using the Tax Cuts & Jobs Act of 2017 (TCJA).

We use TCJA as a source of exogenous variation in tax-based incentives for contributions. The TCJA permanently lowered the federal corporate tax rate from 35% to 21% beginning in 2018. In turn, this resulted in a temporary incentive for sponsors to raise contributions reported in 2017, as they could then be deducted from federal income tax bills at the older, higher tax rate. We identify firm response to the TCJA contributions tax break by exploiting cross-sectional variation in sponsors' exposure to tax-based incentives.

Our results support the conclusion that the policy change induced an intertemporal substitution of higher contributions today for lower contributions tomorrow, and therefore it did not permanently improve the funding status of US private sector DB plans. We find that contributions and funding ratios increased – relative to what their levels would have been in the absence of the tax break – in 2017, the tax break year. That said, 2018 contributions and funding ratios fell relative to counterfactual levels. On balance, pension plan funding ended up where it would have been in the absence of the tax break by 2018.

Our results have implications for work on the incidence of corporate income taxes.

In particular, ignoring "uncertainty" effects on deferred compensation may lead to underestimating the incidence of corporate tax cuts on workers. Estimates of the share of the corporate tax burden passed on to workers focus on wages. Wages, however, are only one part of workers' compensation, with pensions being another. Our model indicates that a temporary increase in tax-based incentives for contributions could in principle result in a permanent improvement in funding, depending on the time profile of financial constraints. The ensuing decrease in retirement income uncertainty would thus improve workers' welfare. That said, we find no evidence for this effect in the case of TCJA.

A corporate tax change could also affect workers' welfare through changes in expected pension benefits (which would be reflected in plan service costs and mandatory as opposed to voluntary contributions) rather than changes in uncertainty about those benefits. There is evidence that the TCJA corporate tax cut affected the current component of workers' compensation, with firms with greater expected tax savings from the TCJA more likely to announce bonus payments to workers (Hanlon, Hoopes, and Slemrod, 2019). Whether similar findings also apply to deferred compensation is a question we leave to future research.

Appendix

Proof of Claim 2.3

Conjecture that there exists an underfunded sponsor that must rely on external finance, $z_t < 0$ and $x_t < 0$ for $t = 0,1,2$. As a result, $R_x = r_{x,t}$ and $R_z = -z_t > 0$. Since the sponsor is underfunded, it must contribute more than the minimum requirement in each of these three periods, so $\lambda_t = 0$ for $t = 0,1,2$. Moreover, the sponsor pays a positive PBGC insurance premium, $q(z_t) = -\bar{q} z_t > 0$ for $t = 0,1,2$. The first order condition for period-0 contributions (2.7) and the law of motion for plan surplus (2.1) then imply:

$$\left(1+r_{x,0}\right)(1-\tau) = \Lambda(1-\tau(1-\Delta))\bar{q}\mathbb{E}_0\left[\left(1+r_{x,1}\right)\right] - \Lambda r_z(z_0+c_0-s_0+\mathbb{E}_0[\omega_0]).$$

The left-hand side of this equation is the marginal cost of time-0 contributions, which does not depend on Δ. The right-hand side is the marginal benefit. The marginal benefit is decreasing in c_0 and increasing in Δ. As a result, contributions are increasing in Δ, $\frac{dc_0}{d\Delta} = \frac{\left(1+\mathbb{E}_0\left[r_1^x\right]\right)\tau\bar{q}}{r_z\mathbb{E}_0\left[r_1^x\right]} > 0$. Because of the linear nature of the model and the separability of the finance costs function, $\frac{dc_0}{d\Delta}$ is equal to the partial derivative of the marginal benefit of contributions with respect to Δ.

Moving one period forward, we have:

$$\left(1+r_{x,1}\right)(1-\tau(1-\Delta)) = \Lambda(1-\tau(1-\Delta))\bar{q}\mathbb{E}_1\left[1+r_{x,2}\right] - \Lambda r_z(z_1(\Delta)+c_1-s_1+\mathbb{E}_1[\omega_1]),$$

which emphasises that period 1 plan funding depends on Δ. Using the law of motion of plan surplus (2.1) to substitute out $z_1(\Delta)$, we have

$$\left(1+r_{x,1}\right)(1-\tau(1-\Delta)) = \Lambda(1-\tau(1-\Delta))\bar{q}\mathbb{E}_1\left[1+r_{x,2}\right]$$

$$-\Lambda r_z(z_0+c_0(\Delta)-s_0+\omega_t+c_1-s_1+\mathbb{E}_1[\omega_1]).$$

The left-hand side of this equation is the marginal cost of time-1 contributions, which is increasing in Δ. The right-hand side is the marginal benefit. The marginal benefit is decreasing in c_1. Δ affects the marginal benefit through two channels: directly, because of the PBGC premium term, and indirectly because plan funding is an increasing function of Δ. Letting μ_{r_x} denote the (constant) mean of the distribution of $r_{x,0}$, we can write:

$$\begin{aligned}\frac{dc_1}{d\Delta} &= \frac{\tau}{\Lambda r_z \mathbb{E}_1\left[r_{x,2}\right]}\left(\left(\mathbb{E}_1\left[r_{x,2}\right]-r_{x,1}\right)-\left(1+\mathbb{E}_1\left[r_{x,2}\right]\right)\left(1-\Lambda\bar{q}\right)\right)-\frac{\left(1+\mathbb{E}_0\left[r_{x,1}\right]\right)\tau\bar{q}}{r_z \mathbb{E}_t\left[r_{x,1}\right]} \\ &= -\frac{1}{\Lambda}\frac{\tau}{\mu_{r_x} r_z}\left(1+r_{x,1}\right)<0\end{aligned}$$

This is because $\frac{dc_0}{d\Delta}$ is equal to the partial derivative of the marginal benefit of contributions with respect to Δ, which is constant over time. As a result, the higher marginal cost dominates.

There remains to verify the conjecture above. Using the FOC for contributions (2.7) and the law of motion for plan surplus (2.1) we obtain that

$$z_1=\omega_0-\mathbb{E}_0[\omega_0]+\frac{\left(1+\mathbb{E}_0\left[r_{x,1}\right]\right)(1-\tau(1-\Delta))\bar{q}}{r_z\mathbb{E}_0\left[r_{x,1}\right]}-\frac{1-\tau}{\Lambda r_z\mathbb{E}_0\left[r_{x,1}\right]}\left(1+r_{x,1}\right),$$

and

$$z_2=\omega_1-\mathbb{E}_1[\omega_1]+\frac{\left(1+\mathbb{E}_1\left[r_{x,2}\right]\right)(1-\tau(1-\Delta))\bar{q}}{r_z\mathbb{E}_1\left[r_{x,2}\right]}-\frac{1-\tau(1-\Delta)}{\Lambda r_z\mathbb{E}_1\left[r_{x,2}\right]}\left(1+r_{x,1}\right).$$

These expressions show that the plan is underfunded when the funding shock is sufficiently below its mean and the sensitivity of external finance costs to cash flows is sufficiently above its mean. Let μ_ω denote the constant mean of the funding shock distribution.

Provided that:

$$\frac{\left(1+\mu_{r_x}\right)(1-\tau(1-\Delta))\bar{q}}{r_z\mu_{r_x}}-\frac{1-\tau(1-\Delta)}{r_z\mu_{r_x}}(1+\bar{r})>0, \tag{2.11}$$

there exists some $r_{x,1}, r_{x,2}$ that are sufficiently above their mean to ensure that $z_1, z_2 < 0$.

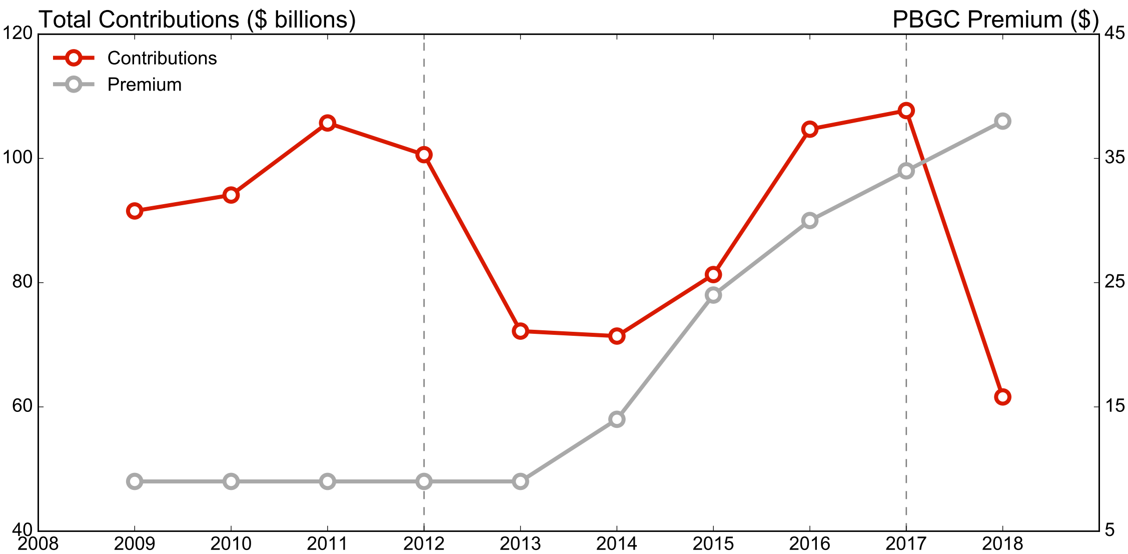
There remains to be verified that the sponsor is relying on external finance in all three periods, $x_t < 0$ for $t = 0,1,2$. By the definition of cash flows, (2.5), x_t decreases in contributions c_t for all $t = 0,1,2$. Since contributions in turn increase in the (contemporaneous) service cost for all $t = 0,1,2$, we can always find a level of s_t such that $x_t < 0$. This concludes the proof.

Proof of Claim 2.3

By the proof of Claim 2.3, the total response of contributions to the TCJA is given by:

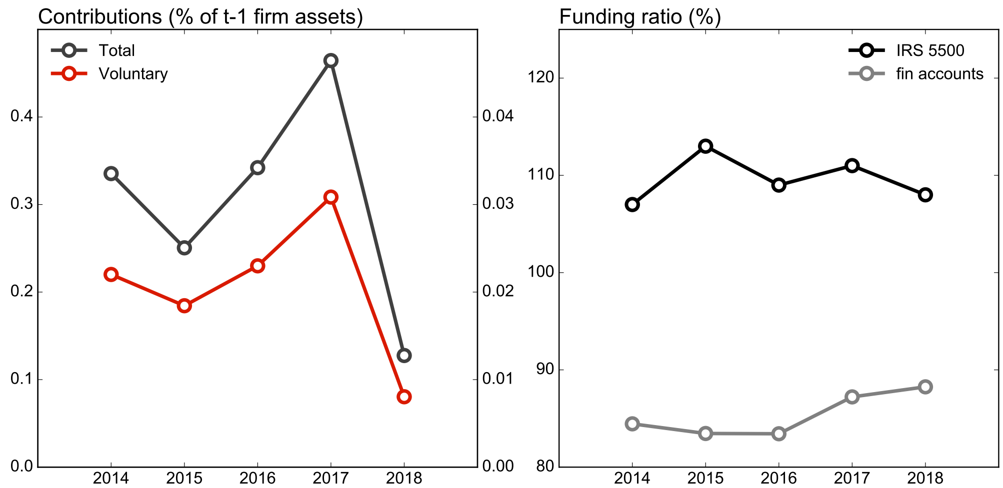
$$\begin{aligned}\frac{dc_0}{d\Delta}+\frac{dc_1}{d\Delta} &= \frac{\tau}{\Lambda r_z\mathbb{E}_1\left[r_2^x\right]}\left(\left(\mathbb{E}_1\left[r_2^x\right]-r_1^x\right)-\left(1+\mathbb{E}_1\left[r_2^x\right]\right)\left(1-\Lambda\bar{q}\right)\right)\\ &< 0 \quad \text{iff} \quad \frac{1+r_1^x}{1+\mathbb{E}_1\left[r_2^x\right]} > \Lambda\bar{q}.\end{aligned}$$

Figure 2.1: Aggregate Contributions and the PBGC Variable Premium



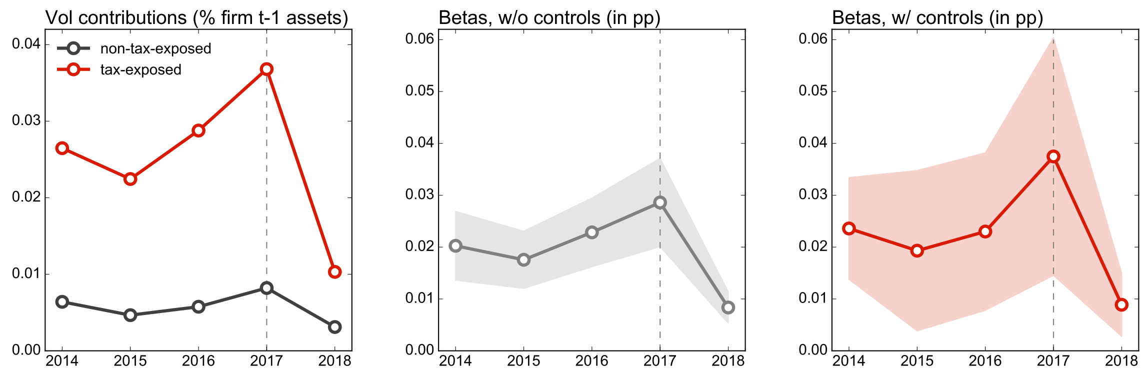
Aggregate contributions (total) of all SB filers in billions of dollars. The PBGC variable rate premium is in dollar per $1000 of funding shortfall. It is computed as PBGC Premium (Plan) = $\max\left[0, \frac{R_t}{1000}(\text{Vested Benefits}-\text{Assets})\right]$, where R is the variable-rate premium set by the PBGC according to the schedule in column (5) of Table 2.1, panel (b) and Vested Benefits represent the share of accumulated pension benefits (Liabilities) that members will receive irrespective of continued participation in the plan. Summing over plans returns the total variable premium paid by a particular sponsor in any given year (PBGC Premium). The first reference line is 2012 when Moving Ahead for Progress (MAP) was passed. The second reference line is 2017, the year of the TCJA tax break.
Sources: IRS 5500 filings, Schedule SB; PBGC website.

Figure 2.2: Aggregate Contributions and Funding



Aggregate contributions (total and voluntary) are averages across sponsors. The aggregate funding ratio is an average across sponsors of the ratio of total sponsor pension assets (sum over plans) over total sponsor pension liabilities (sum over plans).
Sources: IRS 5500 filings, Schedule SB; Financial Accounts of the United States, Table L.118.b; authors' calculations.

Figure 2.3: Voluntary Contributions and the Tax Cuts & Jobs Act of 2017



Voluntary contributions are averages across tax-exposed (treatment) and non-tax-exposed sponsors (control). Shaded areas represent 95% confidence intervals around estimates.
Sources: IRS 5500 filings, Schedule SB; authors' calculations.

Table 2.1: Summary Statistics

(a) Cross-section variation

	Mean	Std. Dev.	10th	50th	90th	N
Voluntary Contributions (%)	0.02	0.046	0.00	0.002	0.06	2417
Total Contributions (%)	0.31	0.66	0.00	0.050	0.86	2417
Funding Ratio (%)	107.58	16.02	92.33	104.42	126.31	2457
PBGC Premium (%)	0.004	0.02	0.00	0.00	0.005	2417
Return on Investment (%)	5.06	6.88	-1.16	6.18	14.43	2390
Discount Rate (%)	6.07	0.31	5.67	6.08	6.48	2459
Altman's Z-Score	2.42	4.85	0.60	1.86	3.64	1001
CAPEX (%)	4.26	3.78	0.21	3.35	8.94	2161
Payout (%)	3.97	4.83	0.05	2.32	9.77	2118
Cash Flows (%)	8.63	6.84	1.19	8.13	16.30	2063
Tobin's Q	1.21	2.66	0.50	0.72	1.72	1023
DB Plans Share	1.15	1.11	0.16	0.84	2.35	2363

(b) Time variation

	(1)	(2)	(3)	(4)	(5)
Year	Plans	Firms	TE	NTE	PBGC
2014	900	557	431	126	14
2015	848	521	406	115	24
2016	823	504	383	121	30
2017	782	482	381	101	34
2018	752	442	298	144	38

Notes: Panel (a) presents plan-level and sponsor-level summary statistics for our sample. There are 4,105 plan-year observations and 2,506 firm-year observations during the period 2014-2018 (some firms sponsor multiple plans). All plans in the sample are middle- and large-scale plans covering more than 100 employees. Plan-level data are from IRS 5500 filings. Sponsor-level data are from Compustat. Voluntary and Total Contributions, PBGC Premium, CAPEX, Payout and Cash Flows are scaled by beginning-of-year sponsor balance sheet assets. Voluntary Contributions, Total Contributions and PBGC Premium are winsorized at the top 1% level. Funding Ratio, Cash Flows, Tobin's Q, Altman's Z, CAPEX, Payout, Return on Investment, Discount Rate and DB Pension Plans Significance are winsorized at the top and bottom 1% level. Panel (b) shows time variation in DB Pension Plans Significance, Tax Exposure and PBGC Premium. Columns (1) and (2) show the number of retirement plans and sponsoring firms in each sample year. Columns (3) and (4) break the sample down by tax-exposure. Column (5) shows the Pension Benefit Guarantee Corporation variable premium rates. Rates are quoted per $1000 of unfunded vested benefits for single-employer plans.

Table 2.2: Tax Exposure and Other Proxies for Tax-Based Incentives

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Marginal Tax Rate Before Interest	1.77***						
	(6.41)						
Marginal Tax Rate After Interest		0.62**					
		(3.14)					
Net Operating Loss			-0.061				
			(-1.50)				
Net Operating Loss Dummy 1				-0.989***			
				(-9.02)			
Net Operating Loss Dummy 2					-0.62***		
					(-7.64)		
Net Operating Loss Dummy 3						-0.62***	
						(-7.62)	
Investment Tax Credit							-8.51
							(-0.67)
Observations	2256	2405	1391	2506	2506	2506	1929
Constant	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: This table presents pooled probit regression estimates of the impact of different proxies for tax-based incentives to make pension contributions on the probability that a sponsor is tax exposed. The dependent variable is Tax Exposure, a dummy variable = 1 if (i) a firm has a positive ex-contributions tax bill and (ii) the firm sponsors at least one plan with funding ratio below 150%. Marginal Tax Rate Before Interest is a simulated corporate marginal tax rate based on income before interest expense has been deducted. Marginal Tax Rate After Interest is a simulated corporate marginal tax rate based on income after interest expense has been deducted. See `https://faculty.fuqua.duke.edu/~jgraham/read.html` for more detail. Net Operating Loss (NOL) and Investment Tax Credit are scaled by beginning-of-year sponsor balance sheet assets. NOL Dummy 1 is a dummy variable which is = 1 if a sponsor has a positive carryforward balance and it pays no current U.S. income tax. NOL Dummy 2 is a dummy variable which is = 1 if if a sponsor has a positive carryforward balance and it reports no pre-tax income. NOL Dummy 3 is a dummy variable which is = 1 if a sponsor does not report any pre-tax income. z-statistics obtained using robust standard errors in parentheses.

Table 2.3: Tax Exposure and Plan- and Sponsor-Level Characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Funding Ratio	-0.63***									
	(-3.61)									
PBGC Premium		917.42**								
		(2.90)								
Investment Return			0.80							
			(1.94)							
Discount Rate				0.30***						
				(3.47)						
DB Plans Significance					-0.04					
					(-1.42)					
CAPEX						-1.47*				
						(-2.39)				
Tobin's Q							0.001			
							(0.08)			
Non-Pension Cash-Flows								2.82***		
								(6.09)		
Altman's Z-Score									-0.005	
									(-0.63)	
Payout										3.80***
										(5.27)
Observations	2457	2417	2390	2459	2363	2158	993	2063	1011	2118
Constant	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: This table presents pooled probit regression estimates of the impact of different plan-level (rows (1)-(4)) and sponsor-level characteristics (rows (5)-(9)) on the probability that a sponsor is tax exposed. The dependent variable is Tax Exposure, a dummy variable equal to 1 if (i) a firm has a positive ex-contributions tax bill and (ii) the firm sponsors at least one plan with funding ratio below 150%. z-statistics obtained using robust standard errors in parentheses.

Table 2.4: Pension Contributions, Tax-Based Incentives and the Tax Cuts & Jobs Act of 2017

	2014		2015		2016		2017		2018	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Tax Exp.	0.020***	0.024***	0.018***	0.019*	0.023***	0.023**	0.029***	0.037**	0.008***	0.008**
	(6.00)	(4.75)	(6.26)	(2.47)	(6.77)	(3.00)	(6.59)	(3.23)	(5.45)	(2.89)
Obs.	530	150	503	140	483	128	471	142	430	134
R^2	0.04	0.25	0.05	0.11	0.04	0.15	0.04	0.13	0.04	0.08
Controls	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes

Notes: This table presents regression estimates of the response of voluntary pension contributions to fiscal incentives in each year between 2014-2018. The dependent variable is voluntary pension contributions deducted from 2014 tax returns in Columns (1) and (2); from 2015 returns in Columns (3) and (4); from 2016 returns in Columns (5) and (6); from 2017 returns in Columns (7) and (8); and from 2018 returns in Columns (9) and (10). Tax Exposure is a dummy-variable equal to 1 if (i) a firm has a positive ex-contributions tax bill and (ii) the firm sponsors at least one plan with funding ratio below 150%. The Tax Cuts & Jobs Act of 2017 reduced the federal corporate tax rate from 35% to 21% beginning in 2018. As a result, contributions counted towards the 2017 corporate tax return could be deducted at 35%, while contributions counted towards 2018 returns at 21%. Columns (2), (4), (5), (6) and (8) include the following plan-level controls: Funding Ratio, PBGC Premium, Return on Investment, Discount Rate. They also include the following sponsor-level controls: Altman's Z-score, Cash Flows, CAPEX, Tobin's Q, Payout and DB Plans Share. t-statistics obtained using robust standard errors in parentheses.

Table 2.5: Tax Exposure and Plan- and Sponsor-Level Characteristics: Robustness

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Funding Ratio	-1.67***									
	(-8.34)									
PBGC Premium		7293.38***								
		(3.48)								
Investment Return			0.77							
			(1.58)							
Discount Rate				0.03						
				(0.29)						
DB Plans Significance					0.01					
					(0.19)					
CAPEX						1.29				
						(1.22)				
Tobin's Q							-0.01			
							(-0.70)			
Non-Pension Cash-Flows								1.91***		
								(3.55)		
Altman's Z-Score									-0.007	
									(-0.87)	
Payout										0.72
										(1.00)
Observations	2457	2417	2390	2459	2363	2158	993	2063	1011	2118
Constant	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: This table presents pooled probit regression estimates of the impact of different plan-level (rows (1)-(4)) and sponsor-level characteristics (rows (5)-(9)) on the probability that a sponsor is tax exposed. The dependent variable is Tax Exposure, a dummy variable equal to 1 if the firm sponsors at least one plan with funding ratio below 150%. z-statistics obtained using robust standard errors in parentheses.

Table 2.6: Contributions, Tax-Based Incentives and the Tax Cuts & Jobs Act of 2017: Robustness

	2014		2015		2016		2017		2018	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Tax Exp.	0.021***	0.029***	0.020***	0.023***	0.026***	0.025**	0.034***	0.040***	0.010***	0.011***
	(5.46)	(4.93)	(11.71)	(4.97)	(9.61)	(3.21)	(11.07)	(3.33)	(9.53)	(3.98)
Obs.	530	150	503	140	483	128	471	142	430	134
R^2	0.02	0.24	0.04	0.10	0.03	0.14	0.03	0.12	0.04	0.07
Controls	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes

Notes: This table presents regression estimates of the response of voluntary pension contributions to fiscal incentives in each year between 2014-2018. The dependent variable is voluntary pension contributions deducted from 2014 tax returns in Columns (1) and (2); from 2015 returns in Columns (3) and (4); from 2016 returns in Columns (5) and (6); from 2017 returns in Columns (7) and (8); and from 2018 returns in Columns (9) and (10). Tax Exposure is a dummy-variable equal to 1 if the firm sponsors at least one plan with funding ratio below 150%. The Tax Cuts & Jobs Act of 2017 reduced the federal corporate tax rate from 35% to 21% beginning in 2018. As a result, contributions counted towards the 2017 corporate tax return could be deducted at 35%, while contributions counted towards 2018 returns at 21%. Columns (2), (4), (5), (6) and (8) include the following plan-level controls: Funding Ratio, PBGC Premium, Return on Investment, Discount Rate. They also include the following sponsor-level controls: Altman's Z-score, Cash Flows, CAPEX, Tobin's Q, Payout and DB Plans Share. t-statistics obtained using robust standard errors in parentheses.

Table 2.7: Contributions, Tax-Based Incentives and the Tax Cuts & Jobs Act of 2017: Robustness

	2014		2015		2016		2017		2018	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Tax Exp.	0.025***	0.34***	0.025***	0.22**	0.026*	0.30**	0.048**	0.53***	0.010**	0.15**
	(3.72)	(4.76)	(3.88)	(2.47)	(2.71)	(2.95)	(2.79)	(3.39)	(2.94)	(3.21)
Obs.	150	150	140	140	128	128	142	141	134	135
R^2	0.31	0.33	0.20	0.16	0.27	0.21	0.20	0.16	0.13	0.09
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sector FE	Yes	No	Yes	No	Yes	No	Yes	No	No	No

Notes: This table presents regression estimates of the response of pension contributions to fiscal incentives in each year between 2014-2018. The dependent variable is *voluntary* pension contributions deducted from 2014 returns in Column (1); from 2015 returns in Column (3); from 2016 returns in Column (5); from 2017 returns in Column (7); and from 2018 returns in Column (9). The dependent variable is *total* pension contributions deducted from 2014 returns in Columns (2); from 2015 returns in Column (4); from 2016 returns in Column (6); from 2017 returns in Column (8); and from 2018 returns in Column (10). Tax Exposure is a dummy-variable equal to 1 if (i) a firm has a positive ex-contributions tax bill and (ii) the firm sponsors at least one plan with funding ratio below 150%. The Tax Cuts & Jobs Act of 2017 reduced the federal corporate tax rate from 35% to 21% beginning in 2018. As a result, contributions counted towards the 2017 corporate tax return could be deducted at 35%, while contributions counted towards 2018 returns at 21%. All columns include the following plan-level controls: Funding Ratio, PBGC Premium, Investment Return, Discount Rate. They also include the following sponsor-level controls: : Altman's Z-score, Cash Flows, CAPEX, Tobin's Q and DB Plans Share. t-statistics obtained using robust standard errors in parentheses.

Table 2.8: Funding Ratios and the Tax Cuts & Jobs Act of 2017

	ΔFR 16-17		ΔFR 17-18		ΔFR 16-18	
	(1)	(2)	(3)	(4)	(5)	(6)
Tax Exposure in 2017	3.37***	2.49**	-1.11	-2.00*	1.47	-0.22
	(4.53)	(3.28)	(-1.19)	(-2.29)	(1.31)	(-0.20)
Observations	457	425	307	284	311	248
R^2	0.02	0.06	0.00	0.17	0.16	0.12
Controls	No	Yes	No	Yes	No	Yes

Notes: This table presents regression estimates relating changes in funding ratios to sponsor tax exposure in 2017. The dependent variable in Columns (1) and (2) is the change in plan funding ratios between end-2016 and end-2017. The dependent variable in Columns (3) and (4) is the change in plan funding ratios between end-2017 and end-2018, for plans sponsored by firms that were tax-exposed in both 2017 and 2018. The dependent variable in Columns (5) and (6) is the change in plan funding ratios between end-2016 and end-2018, for plans sponsored by firms that were tax-exposed in both 2017 and 2018. Funding ratio changes are reported in percentage points. Columns (2) and (4) include the following controls: Funding Ratio in 2016 (respectively, 2017), change in the Discount Rate between 2016 and 2017 (2017 and 2018), and actual 2017 (2018) Return on Investment. Column (6) includes the change in Discount Rate between 2016 and 2018 (average of 2017 change and 2018 change) and actual Return on Investment (average of 2017 return and 2018 return) as controls. Controls winsorized at the bottom and top 1%. t-statistics obtained using robust standard errors in parentheses.

Table 2.9: Variable Definitions (Plan-Sponsor Level)

Variable	Definition
Total Contributions (Plan)	Total sponsor contributions reported on tax return
Required Contributions I	Contributions allocated towards unpaid MRC from prior years
Required Contributions II	Contributions allocated towards MRC for the current year
Special Contributions	Contributions made to avoid restrictions on benefits
Mandatory Contributions	Required Contributions (I+II) + Special Contributions
Voluntary Contributions (Plan)	Total- Mandatory Contributions
Credit Balances	Funding Standard Carryover Balance + Pre-Funding Balance
Assets	Market value of plan assets at year end. Contributions not included
Safe Assets	Sum of investment grade bonds, insurance contract and cash
Safe Assets Share (Plan)	Safe Assets/Assets
Liabilities	Present value of plan benefits accumulated to year end
Return on Investment (Plan)	Investment Income/(L1.Assets - Total Contributions)
Discount Rate (Plan)	Interest rate used to compute liabilities
Vested Benefits	The share of liabilities that employees will receive regardless of their continued participation in the sponsor's pension plan
Participants	Number of plan participants
PBGC Premium (Plan)	Variable-rate benefits insurance premium
	$\max[0, R(\text{Vested Benefits-Assets})/1000]$

Table 2.10: Variable Definitions (Sponsor-Level)

Variable	Definition
Aggregates of plan-level variables (IRS 5500 Filings)	
Voluntary Contributions	Sum of Voluntary Contributions (Plan) over sponsored plans
Total Contributions	Sum of Total Contributions (Plan) over sponsored plans
Funding Ratio	Sum of (Assets + Total Contributions - Credit Balances) over sponsored plans/Sum of Liabilities over sponsored plans
Return on Investment	Assets-weighted average of Return on Investment (Plan), over sponsored plans
Discount Rate	Liabilities-weighted average of Discount Rates (Plan), over sponsored plans
PBGC Premium	Sum of PBGC Premium (Plan) over sposored plans (Compustat)
Other sponsor-level variables	
Net Tax	Federal corporate income tax expense
Gross Tax	Net Tax + $\tau\times$sum of Total Contributions over sponsored plans. $\tau = 35\%$ until 2017, 21% after
Tax Exposure	A dummy variable = 1 if Gross Tax > 0 and Funding Ratio < 150% for at least one firm pension plan
Net Income	Net income
Depreciation	Depreciation and amortization
Pensions Expense	The sum of the service cost and an interest cost (the change in the present discounted value of the pension obligations arising from the approach of the time when these obligations come due) minus an assumed return on pension plan assets (see Bergstresser et al., 2006)
Cash Flows	Net Income + Depreciation + Pensions Expense + sum of Total Contributions over sponsored plans
CAPEX	Capital expenditures
Altman's Z	$\left(3.3\times \text{EBIT}+\text{Sales}+1.4\times \text{Retained Earnings}+1.2\times \text{Net Working Capital}\right)/\text{Operating Assets} +$ Market Value of Equity / Total Liabilities
Tobin's Q	$\left(\text{Assets (book)}+\text{Equity (market)}-\text{Common Equity (book)}-\text{Deferred taxes}\right)/\text{Assets (book)}$
Employees	Current number of employees
DB Plans Share	Sum of Participants over sponsored plans/Employees
Payout	Purchase of Common and Preferred Stock + Dividends for Common Stock + Dividends for Preferred Stock

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