Global Corporate Bond Markets and Local Monetary Policy Transmission

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Abstract

When tight monetary policy curtails domestic supply of credit and raises domestic borrowing costs, some firms can mitigate higher local borrowing costs by tapping global bond markets. This paper investigates whether this prediction holds for non-financial companies in the euro area. I first show that euro area firms exploit borrowing cost differentials between USD and EUR by issuing corporate bonds in USD when swap-adjusted U.S. dollar funding costs fall below euro rates. Using proxies for such opportunistic borrowing behavior, I then find that firms capable of seizing these opportunities in global corporate bond markets do not reduce their fixed capital investment to the same extent as other firms in response to monetary tightening. Further findings reveal that this differential investment response is not explained by differences in financial constraints or investment opportunities; instead, it reflects the ability to switch to lower cost offshore bond finance. Overall, the results underscore heterogeneity in the real effects of monetary policy and suggest that capital market openness can attenuate the domestic investment channel when global conditions allow lower cost funding abroad.

Keywords: monetary policy, firm heterogeneity, international financial markets, corporate bonds
JEL Classification: E22, E44, E52, F34, F62, G12, G15

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1 Introduction

Monetary tightening typically reduces credit supply and raises domestic borrowing rates, both of which depress firm borrowing. Consequently, firms cut back externally funded investment. This description corresponds to the conventional investment channel of monetary policy transmission in its most stripped-down form and implicitly assumes a closed economy setting. With rising global funding opportunities, however, the closed economy approach overlooks some important aspects of how monetary policy transmission works in an open economy. Notably, firms that tap foreign debt markets can shield themselves from contractionary impacts of local monetary tightening when foreign markets offer cheaper funding opportunities. In doing so, they may not reduce their investment as much as other firms without access to these markets, leading to heterogeneous and possibly impaired monetary policy transmission. In this paper, I test whether this hypothesis holds for euro area (EA) non-financial companies (NFCs) by examining their borrowing activity in global corporate bond markets.

Two facts motivate the analysis. Firstly, as Figure 1 illustrates, the debt mix of EA NFCs has tilted toward bond finance over the past two decades[1]. In quantitative terms, the bond to loan ratio has risen from 13 percent to above 30 percent, highlighting the growing role of bond finance in the EA financial system[2]. Secondly, while the growing share of bond financing in the EA has recently gained attention from scholars and policymakers (Schnabel, 2021; European Central Bank, 2021), there is a neglected aspect of this trend: the international finance dimension. Figure 2 demonstrates that U.S. Dollar (USD) denominated bonds issued by EA NFCs constitute a substantial part of bond financing. In my bond issuance sample, 1,039 out of 1,073 USD-denominated tranches are issued outside the euro area – with a large fraction issued in the United States (676 tranches). Thus, there is a significant international finance dimension of expanding corporate bond markets in the EA which has been overlooked so far. In this sense, this paper uniquely contributes to the literature by addressing this international aspect and studying its implications for monetary policy transmission.

I ask two main questions. First, do EA NFCs exploit currency specific borrowing

[1] From 2001 onward, outstanding amounts of bonds issued by EA private sector more than doubled, reaching €17 trillion in 2020.

[2] See Darmouni and Papoutsi (2022) for a detailed exposition of the rising corporate bond market in the EA with a special focus on changing issuer and investor composition.

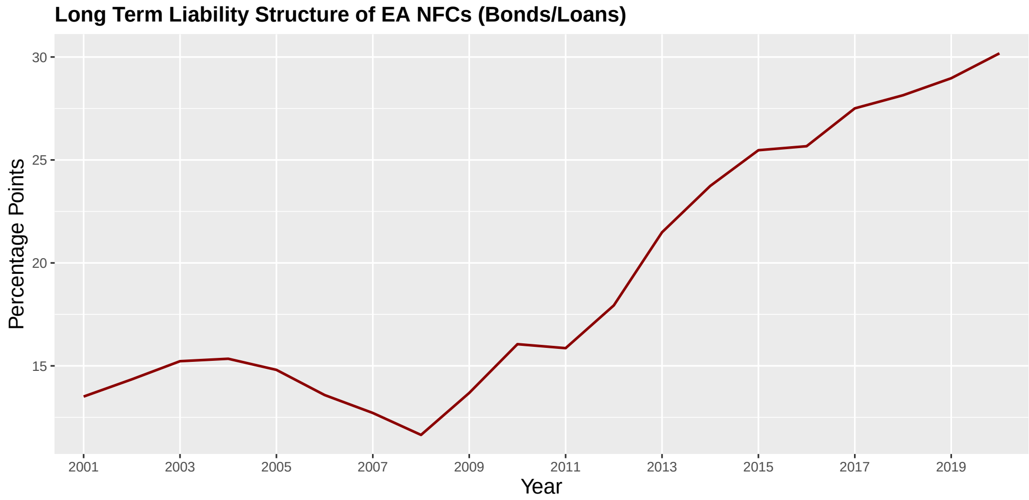

FIGURE 1
Long-Term Liability Structure of Euro Area NFCs. Source: ECB

cost differentials by issuing in USD when swap-adjusted dollar funding costs fall below euro rates? Second, conditional on such behavior, do investment responses to monetary tightening differ systematically between firms that can switch to offshore bond funding and otherwise similar firms that cannot?

To answer the first question, I calculate an FX-hedged borrowing cost differential measure, also referred to as the "corporate basis" by Liao (2020). The corporate basis measures the difference in borrowing costs between issuing in EUR and issuing in USD once exchange rate risk is hedged (via FX swaps) – i.e., the cost difference between direct and synthetic EUR borrowing. Utilizing the corporate basis in panel non-linear binary outcome and panel censored regression models, I show that the answer is affirmative. As corporate basis increases by one standard deviation, the probability of a given firm issuing in USD increases by 0.5%, a highly sizable effect given the unconditional probability of 2.1%.

Following McBrady and Schill (2007), I call this behavior as "opportunistic borrowing"[3]. This finding, *per se*, is of limited value as other studies also provide evidence for this behavior in different contexts (Liao, 2020; McBrady and Schill, 2007; Galvez et al., 2021; Caramichael et al., 2021). I validate the existence of this behavior for EA NFCs using more recent matched bond-firm level data. This discovery serves as a bridge to understand differing reactions of firms to monetary policy leading to the second question which constitutes the core contribution of this pa-

[3]I use "opportunistic borrowing" purely as a neutral descriptor of cost-sensitive issuance, i.e., when FX-hedged USD-EUR borrowing costs diverge and firms temporarily issue in the lower-cost currency. The term carries no normative or moral connotation.

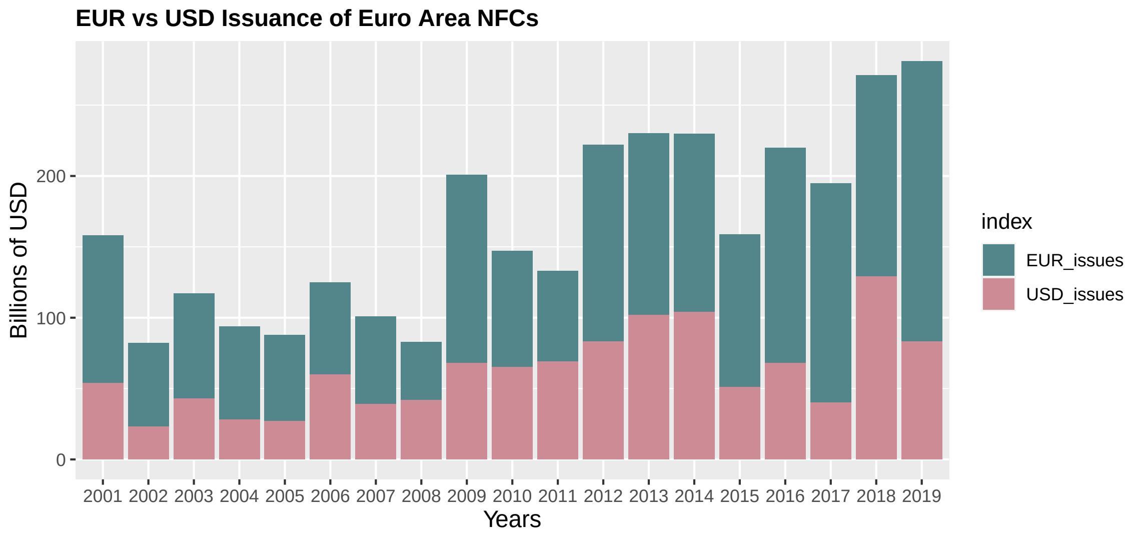

FIGURE 2

EUR vs USD Bond Issuance of Euro Area NFCs. Source: Refinitiv Eikon.

per.

Having established opportunistic borrowing behavior of EA NFCs, I then turn to the second question: whether access to offshore bond funding shapes real responses to monetary tightening. Opportunistic borrowing lets firms lower funding costs by switching markets or currencies[4]. When domestic credit tightens and local rates rise, firms with access to foreign bond markets can obtain cheaper financing abroad, limiting increases in their effective cost of capital and their investment responses as a result. A panel local projection analysis à la Jordà (2005) coupled with high-frequency identification of monetary policy surprises confirms that these firms indeed reduce their fixed capital investment to a lesser extent in response to monetary tightening compared to firms that only borrow in the local bond market.

An important threat to identification arises if opportunistically borrowing firms react less to monetary policy since they could be less financially constrained compared to their peers due, for instance, to their higher credibility. If this is the case, then heterogeneous firm reaction to monetary tightening can also be driven by differential financial constraints firms face that are independent of their access to global corporate bond markets. However, the observed heterogeneity persists even after controlling for differences in financial constraints across firms. Moreover,

[4]In this paper, I generally use offshore issuance and foreign currency issuance interchangeably. Even though these two concepts can describe fundamentally different phenomena in certain contexts, they are very close substitutes in the case of EA NFCs. For instance, the bulk of USD denominated bonds are issued outside the Eurozone. See Table 1 in Section 4.1 for more details.

if information asymmetries -unrelated to firms' access to global corporate bond markets- drive the heterogeneity in firm responses, then that heterogeneity should be independent of FX-hedged borrowing cost differentials across currencies (the corporate basis). Nevertheless, a placebo test reveals that the heterogeneity disappears when FX-hedged foreign currency issuance is more expensive than local currency issuance, suggesting that cost-saving opportunities in global bond markets, instead of other types of asymmetric financial constraints, are at the heart of this heterogeneous response.

After addressing this identification concern, I also assess external validity in equity markets using stock market's reaction to monetary policy surprises. I find that firms that can borrow opportunistically in global bond markets experience smaller return declines after monetary tightening than firms limited to local markets, corroborating the baseline investment results. In an ongoing companion work, we also document similar heterogeneity for U.S. firms: a difference-in-differences analysis around the 2014-2016 Fed-ECB policy divergence episode shows that U.S. firms with access to European bond markets increase their EUR issuance and reduce investment significantly less than firms restricted to domestic bond markets.

This paper makes at least two contributions. First, to my best knowledge, it is the first paper that studies the implications of global corporate bond markets for monetary policy transmission. Prior work has enriched our understanding of monetary policy transmission by showing that firms' financing choices -e.g., the loan-bond mix- interact with monetary policy (Crouzet, 2021, 2018; Bolton and Freixas, 2006; Darmouni et al., 2020; Chen, 2025). Adopting closed economy models, however, these studies remain silent on firms' bond financing opportunities in international markets. As global debt markets deepen, firms with access to these markets can partially offset domestic tightening by issuing abroad when FX-hedged foreign-currency funding is cheaper. I show that this mechanism operates for EA NFCs. This result is also policy relevant for financially open economies since cheaper global funding can attenuate monetary transmission implying that achieving a given stance may require stronger tightening than in a closed economy. The importance of this channel is likely to grow as global corporate bond markets expand and access widens[5].

[5]One can argue that since the number of firms tapping global corporate bond markets remains limited (122 on consolidated and 408 on unconsolidated basis), the effect of these markets will be negligible. These firms, however, are typically large firms with a mean asset of €35.8 billion and collectively span roughly 40,000 subsidiaries and affiliates. In the spirit of Gabaix (2011), they can be considered as granular firms whose investment dynamics likely impact aggregate investment patterns. In fact, some back of the envelope calculations indicate that they account for around 12-16%

Second, I introduce access to global corporate bond markets as a novel source of firm-level heterogeneity in exposure to monetary policy. Prior work on heterogeneous firm responses emphasizes financial frictions proxied by leverage, liquidity, size, age, or local bond market access. For instance, seminal work by Kashyap et al. (1994) and Gertler and Gilchrist (1994) find that bank-dependent and small firms are more exposed. I instead focus on non-bank-dependent issuers as the firm sample comprises firms that have issued at least one corporate bond during the sample period. Accordingly, this lets me ask whether exposure differs even among firms that all access domestic bond markets. I show that the relevant margin is international: firms with access to global corporate bond markets are less exposed than otherwise similar firms without such access.

Related Literature. This paper contributes to four strands of the literature. First, it relates to a small stream of corporate finance literature studying how firms' choice between bond and loan financing interacts with monetary policy transmission. A key premise of the popular bank lending channel view is the imperfect substitutability of bank loans and bonds. According to this view, should bonds be perfect substitutes of bank loans, the only effect of monetary tightening would materialize via the standard interest rate channel as firms could easily switch from bank loans to bond financing in response to a reduction in loan supply. Consistent with this view, Crouzet (2018), Altavilla, Pariès and Nicoletti (2019) and Alder et al. (2023) find that corporate bond issuance increases in response to a negative bank loan supply shock but this shift is not enough to compensate the reduction in bank lending. As a result, aggregate borrowing and investment declines. Likewise, Crouzet (2021) documents evidence suggesting that bank-dependent firms reduce their investment more compared to bond-financed firms in response to monetary shocks[6]. By adopting closed economy models, these papers largely abstract from the international finance dimension of corporate debt structure. By incorporating

of Euro Area private investment even without considering any network effects. Thus, studying their differential responses to monetary policy offers valuable insights into understanding how monetary policy propagates within the corporate sector.

[6]Related work also highlights alternative and sometimes seemingly conflicting patterns. Darmouni et al. (2020) find that bond-reliant firms can be more sensitive to monetary policy shocks and attribute this to the relative flexibility of bank lending as bond-reliant firms are likely to be more prudent in financially stressful episodes. Chen (2025) documents that monetary tightening is associated with an expansion in bank business loan origination driven by higher loan demand among large, financially unconstrained firms, which substitute bank loans for bonds when the bond-loan spread widens.

neglected global funding opportunities, I seek to deepen our understanding of how the corporate sector's debt composition shapes monetary policy transmission in open economies.

Another strand of the literature to which this paper is affiliated concerns the determinants of offshore bond issuance. While this literature counts many reasons behind the offshore issuance of a firm such as deeper foreign markets, desire to hedge foreign currency cash flows, funding diversification and signaling (Allayannis et al., 2003; Munro et al., 2011; Black and Munro, 2010; Mota and Siani, 2023), my paper is closest to studies emphasizing the importance of borrowing cost differentials across markets/currencies (McBrady and Schill, 2007; McBrady et al., 2010; Liao, 2020; Galvez et al., 2021; Bruno and Shin, 2017; Salomao and Varela, 2022; Huang et al., 2024)[7]. That said, this literature does not establish a link between opportunistic borrowing and monetary policy transmission. My paper extends it by characterizing that link and examining associated firm-level effects.

Third, a recently emerging literature documents reduced effectiveness of monetary policy due to various international leakage channels. Barajas et al. (2018) find that remittance inflows weaken monetary policy efficacy. Ongena et al. (2021) conclude that banks' foreign currency lending is less sensitive to domestic policy than their local currency lending, diluting transmission for multi-currency lenders. Using bank-level data from Norway, Cao and Dinger (2022) show that favorable global financial conditions insulate banks from local monetary policy. Finally, Fendoglu et al. (2019) argue that ample global liquidity reduces effectiveness of monetary policy tightening in Turkey as banks substitute toward international wholesale funding when domestic conditions tighten. My paper adds to this strand by proposing another potential impairment channel that operates through NFCs' issuance in global bond markets.

Finally, a growing literature studies heterogeneity in firms' investment responses to monetary policy. Ottonello and Winberry (2020) find that firms with low default risk are more responsive to monetary shocks whereas Jeenas (2024) conclude that firms with less balance sheet liquidity react more. On the other hand, Cloyne et al. (2023) demonstrate that young and no dividend paying firms adjust their fixed capital expenditure more compared to older and dividend paying firms. Ippolito

[7] Most notably, Graham and Harvey (2001) document that 44% of the firms in their survey respond that lower foreign rates are important/very important drivers of their decision to incur FX debt. Along similar lines, Gozzi et al. (2015) demonstrate that bonds issued abroad tend to have lower yields compared to bonds issued at the home country.

et al. (2018) examine multiple firm outcomes and conclude that firms (especially financially constrained ones) with more unhedged loans on their liability side react more to monetary policy owing to the floating rate nature of most loan payments. More recently, complementary evidence on heterogeneity in monetary transmission comes from the equity side. Almeida et al. (2025) show that financing constraints of equity-focused firms amplify the effects of monetary tightening on firm valuations, equity issuance, R&D spending, and capital expenditure. I contribute to this literature by introducing a new form of heterogeneity: firms' access to global corporate bond markets. As elaborated in the next section, such access can attenuate the investment response to monetary policy under certain conditions.

The rest of the paper is organized as follows. Section 2 outlines the mechanism through which the leakage effect can arise and generate substantial heterogeneity in firms' responses to monetary tightening. Section 3 computes the FX-hedged borrowing cost differential between EUR and USD for the euro area corporate sector which serves as a key input for the subsequent analysis. Section 4 examines, with a focus on the corporate basis, the determinants of firms' foreign currency issuance decisions. Section 5 analyzes firms' heterogeneous investment responses to monetary policy surprises. Section 6 presents robustness checks and evidence on external validity. Section 7 concludes and discusses avenues for future research.

2 The Leakage Mechanism and the Eurozone

Standard bank lending and interest rate channels of monetary policy transmission predict that monetary tightening leads to a contraction in loan supply and an increase in bank lending rates. In turn, credit squeeze and higher borrowing costs would induce firms to cut back externally funded investment. There is, however, another way out for firms in need of external finance. If they have the sufficient means, they can resort to market finance (e.g. issue bonds) to substitute for curtailed and costlier loan financing. To the degree that they offset reduction in bank loans and the rise in lending rates in this way, they can maintain their investment at desired levels.

The shift away from bank loans toward market-based debt in response to monetary tightening has been widely studied. In particular, Kashyap et al. (1993) document substitution toward commercial paper, while Holm-Hadulla and Thürwächter (2021) study substitution toward corporate bonds; this pattern is interpreted as evidence for the bank lending channel of monetary policy transmission (Becker and Ivashina, 2014). That said, this substitution is imperfect and the local bond market may not serve as a "spare tire" even for firms which have access to market finance (i.e.firms that are not bank-dependent). A simple partial equilibrium model of investment developed by Crouzet (2021) implies that a monetary policy tightening shock steepens both types of credit supply curves but the effect is milder for loan supply. The reason is related to different natures of loan-financing and bond-financing with the former providing more flexibility due to the possibility of renegotiating the terms of the loan contract with the borrower's bank to avoid liquidation of the borrower in times of financial distress [8]. As a result, the model predicts that bank-financed firms reduce their borrowing less compared to bond-financed firms which have a dispersed base of lenders, diminishing the flexibility of their financing structures.

The idea that substitution effect is limited since monetary policy affects not only loan supply but also credit supply in the bond market can also be found in policy oriented work. For instance, International Monetary Fund (2016) discusses that monetary policy affects investor behavior in the domestic bond market as well by moving the risk premia, leading to reduced risk appetite during tightening episodes. This would reduce credit supply and drive up the cost of credit in the local bond market. Moreover, Schnabel (2021) and European Central Bank (2021) both argue that the rise of non-banks in the euro area, in fact, strengthened monetary policy transmission due to higher responsiveness of non-banks' (compared to banks') balance sheets to policy changes that primarily affect the long end of the yield curve. Then, given the high share of debt securities in non-banks' asset portfolio (around 40% in the euro area), rising domestic corporate bond markets, if anything, might have fostered the impact of monetary policy on corporate sector especially when policy change aims long term rates. Hence, the euro area evidence suggests that domestic bond market acts as a complement to rather than as a substitute for loan financing. If so, domestic bond market may fail to offer a resort for firms in need of external finance and remain unable to attenuate the effectiveness of the bank lending channel.

[8] A large stream of corporate finance literature studies limited substitutability between bond vs loan financing and implications of the debt structure for firm-level outcomes. A common theme in these studies is that the flexibility provided by bank loans may prove to be quite valuable in times of financial distress. See Darmouni et al. (2020), De Fiore and Uhlig (2011), Bolton and Freixas (2006), Crouzet (2018), Rajan (1992), Diamond (1991) and Bolton and Scharfstein (1996).

Since the investor base in global corporate bond markets is likely to be much less affected by local monetary policy changes, however, the complementary relation between loan finance and bond finance should exist only in the case of the local bond market. Global corporate bond markets could well emerge as an alternative and cheaper funding source for firms especially when local credit supply contracts and local funding becomes costlier. In fact, the literature on the determinants of firms' offshore bond issuance decisions demonstrate that firms borrow in foreign debt markets with lower cost of borrowing motives. Moreover, a recent study by Cortina et al. (2021) shows that firms switch internationally across markets in times of crisis and change the currency composition of their debt. By moving away from crisis-hit markets, they compensate, even if partly, the decline in borrowing in these markets and maintain the maturity of their debt.

These two observations tell us that a certain set of firms actively seek the best conditions in global debt markets by switching across markets/currencies. Under monetary tightening, such active debt management would prompt them to seek for alternative markets/currencies through which they can secure cheaper funding. Consequently, they would be, even if partially, isolated from tightened domestic funding conditions and might not reduce their investment as much as other firms[9].

Figure 3 illustrates this leakage channel working through firms' activity in global debt markets. There are two major credit related aspects of monetary tightening: quantity and price effects. Quantity effect works through curtailed supply of credit in loan and bond markets. This effect is generally conceptualized under the umbrella of the "credit channel" of monetary policy (Bernanke and Blinder, 1988, 1992; Bernanke and Gertler, 1995). Price effect on the other hand works through borrowing costs and thus affects the user cost of capital and in turn firm-level investment[10].

[9]This mechanism can be reinforced if local monetary tightening renders borrowing in foreign currency cheaper compared to borrowing in local currency. The results reported in Appendix C verify this prediction by showing that monetary policy differential measured by the difference between ECB and Fed controlled rates is a significant determinant of currency-induced borrowing cost differential between EUR and USD.

[10]In this paper, I focus on bond market activity of firms since I study their investment response to monetary policy and longer-term rates matter most for investment decisions. Another leakage effect that is not considered in this paper might be working through short-term borrowing needs of firms. Firms frequently borrow in short-term debt markets to fund their working capital needs (Barth III and Ramey, 2001; Gaiotti and Secchi, 2006; Christiano et al., 1997). A tighter monetary policy increases production costs by raising cost of external borrowing and curtails available short-term credit. In response, firms might tap foreign markets to issue commercial paper in an effort to reduce their borrowing costs especially when the market expects that monetary tightening will be followed by other tight policy actions. I leave this aspect of the leakage channel for future research.

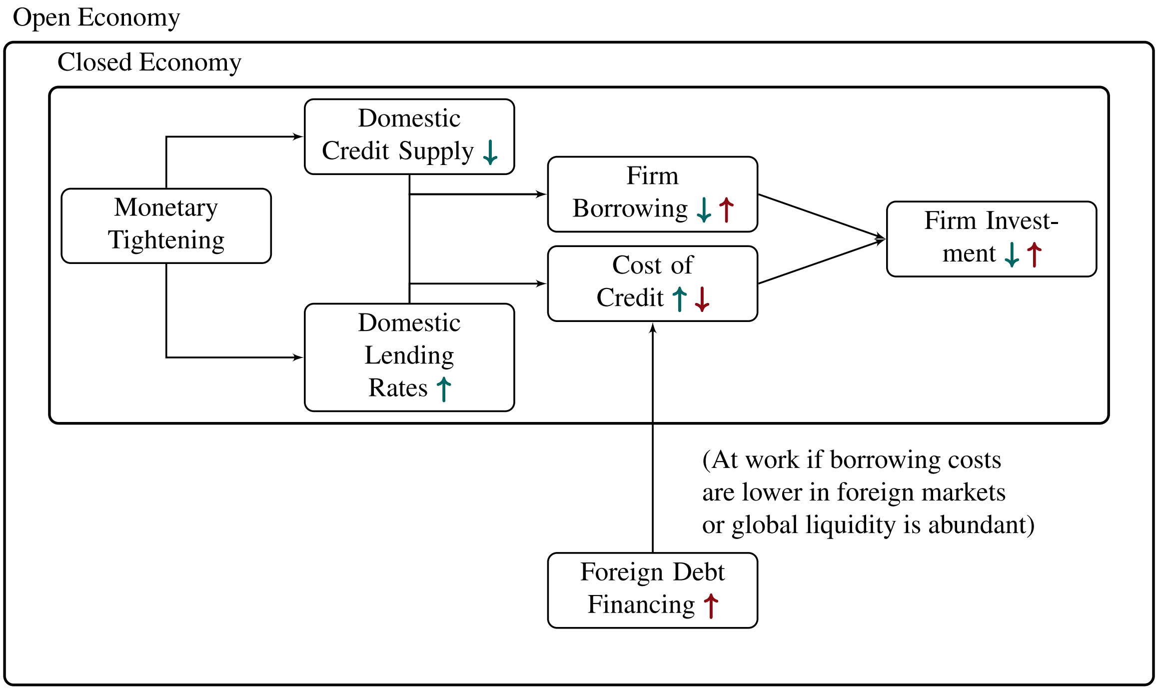

FIGURE 3
The Exposition of How Global Corporate Bond Markets Can Impair Monetary Policy Transmission

These effects are demonstrated by the green arrows in Figure 3. Firms with access to global markets can compensate for the decline in their domestic borrowing from foreign financial markets if global liquidity is abundant. That would impair the credit channel. They can also reduce their borrowing costs if foreign debt markets offer cheaper credit, impairing the cost of borrowing channel. These international substitution effects are depicted by the red arrows in Figure 3 and work in the opposite direction of local monetary policy transmission[11].

In this paper, I specifically focus on borrowing cost differentials between domestic and foreign currencies since measuring quantity effects pose considerable practical challenges. To test whether the credit channel is also impaired would require identifying the episodes during which local credit supply is tight and global liquidity is abundant. This is a notoriously difficult task that can easily lead to incorrect conclusions since it requires a considerable level of subjective assessment of prevailing credit conditions. On the other hand, comparing borrowing cost dif-

[11]This paper focuses on firms' activity in the bond market but a similar substitution of local credit with foreign credit can also happen in the loan market as well. This issue too is left for future research.

ferentials is largely free from these problems as comparison relies completely on a quantitative framework (a bond pricing model is introduced in Section 3). That said, it is highly conceivable that there is a correlation between these effects in the sense that when local credit conditions are tight and global liquidity is abundant, tapping global markets can reduce borrowing costs relative to borrowing domestically.

This paper focuses on EA NFCs, however the mechanism laid out here is likely to exist in other countries as well. Studying the Eurozone, however, brings forth several advantages promoting the robustness of the analysis. First, the Eurozone is largely free from problems associated with bond market incompleteness. In small economies with insufficient levels of bond market depth, issuers are likely to have difficulty in issuing sophisticated debt securities. Instead, they could issue offshore where they could meet a much larger investor base that matches their interests. Consequently, some firms may have an inherent propensity to issue offshore irrespective of borrowing cost considerations, complicating the empirical analysis: issuance abroad may simply proxy shallow domestic markets. In the Eurozone, this problem is much less severe thanks to well-developed corporate bond markets. Second, the way I define opportunistic borrowing allows firms to hedge their FX borrowing operations. The most natural way for a firm to hedge its FX exposure is to enter into a swap agreement. Yet, this requires the availability of swap counterparties. For less frequently traded currency pairs, lack of swap counterparties could prevent firms from engaging in hedged opportunistic borrowing. A large currency swap market between EUR and USD removes this problem. Finally, the Eurozone's rapidly expanding corporate bond market provides a compelling setting and heightens the policy relevance of the analysis.

3 Corporate Basis

There is one condition to be satisfied for firms to be able to borrow opportunistically in global corporate bond markets: borrowing in the foreign currency should be cheaper compared to domestic currency. There are several ways to measure borrowing cost differentials between currencies. First, the simplest method is to compare nominal interest rates, such as money market rates as in Bruno and Shin

(2017)[12]. This could prove to be a good indicator only if the majority of firms engage in unhedged FX borrowing as in the case of many emerging market economies[13]. Second, deviation from covered interest parity (CIP) in benchmark rates is another proxy that measures borrowing cost differential between two currencies assuming that borrowers hedge their open FX positions. However, since firms can face different credit spreads in different currencies, CIP deviation based on benchmark rates might not reflect the true long-term borrowing conditions of the corporate sector.

Another measure introduced lately by Liao (2020) is corporate basis which focuses on currency related differences in borrowing costs in corporate bond markets. Corporate basis remains largely free from problems associated with other approaches. First, its construction entails a bottom-up approach through the use of bond-level data. Thus, unlike other proxies, it is designed specifically for corporate sector's borrowing conditions. Second, it allows firms to hedge their FX debt. Third, it controls for bond-level and issuer-level characteristics that might affect borrowing cost differential between currencies, thereby providing us a more refined currency-induced borrowing cost differential. For all these reasons, corporate basis arguably stands out as the best proxy for borrowing cost differential between currencies in corporate bond markets. In this section, I calculate the corporate basis between EUR and USD for EA firms[14].

3.1 Calculation of Corporate Basis

Calculation of corporate basis is based on Liao (2020). In simplest terms, corporate basis is defined as follows:

$$CB_t = (rb_t - rb_t^{\$}) + (f_t - s_t) \tag{1}$$

[12]Gutierrez et al. (2023) provide a more sophisticated approach by measuring the difference between interest rates for loans denominated in USD and in domestic currency in a regression framework. This way, they are able to control for loan-level and firm-level characteristics and purge their interest rate difference measure from effects that are not related to the currency in which the loan is denominated.

[13]Even so, it might still fail to be a good proxy unless expected exchange rate movements between domestic currency and USD are of negligible nature. In this vein, Gutierrez et al. (2023) provide an interest rate difference measure that is adjusted for uncovered interest parity.

[14]It is important to make this calculation exclusively for the EA firms since corporate basis between the two currencies could be significantly different for firms of different countries. For instance, Liao (2020) shows that borrowing cost of US firms when issuing in USD is significantly lower compared to borrowing costs faced by other countries firms when issuing in USD.

where rb_t is the risky bond yield in EUR, $rb_t^{\$}$ is the risky bond yield in USD and $f_t - s_t$ is the forward premium. In words, corporate basis measures how much a EA firm can expect to gain by issuing in USD instead of in EUR and then swap USD into EUR, i.e. cost saving resulting from synthetic local currency (EUR) borrowing[15]. If we add and subtract risk-free yields (rf_t and $rf_t^{\$}$) to CB_t, we get:

$$CB_t = [(rb_t - rf_t) - (rb_t^{\$} - rf_t^{\$})] + [(rf_t - rf_t^{\$}) + (f_t - s_t)] \tag{2}$$

where the first term is the credit spread differential (CSD) between EUR and USD and the second term is the deviation from the CIP condition based on risk-free rates. Simply put, we have:

$$CB_{\$t} = CSD_{\$t} + CIPdev_{\$t} \tag{3}$$

Corporate basis, defined this way, implies that risk is priced differently depending on the currency of the bond issued. This, in turn, results from the segmentation of credit market along currency lines (Liao, 2020) which is mostly a post GFC phenomenon. I will exploit this segmentation of credit market to identify episodes when borrowing in USD provides cost-saving opportunities to EA firms.

Appendix A explains the details of how credit spread differential is calculated using bond-level data. In brief, the credit spread differential is the estimated currency-of-issuance component in yield spreads after controlling for standard bond characteristics, including bond age, rating, remaining maturity, amount issued, and issue size. The estimated credit spread differential is presented in Figure 4 along with its 95% confidence interval. The values below zero imply that credit spread of EUR denominated bonds is less than that of USD denominated bonds. Figure 4 shows that credit spread differential falls sharply around 2008-2009 which matches the turmoil in US financial markets when bond spreads soared in the US. After the launch of ECBs asset purchase program in 2014, credit spread differential decreases again significantly.

Figure 5, on the other hand, depicts credit spread differential, CIP deviation and corporate basis on the same graph. CIP deviation, proxied by the negative of 5-year cross-currency basis, rises substantially around the GFC when dollar shortage became a major problem for European banks and then moves upward again around the Eurozone sovereign crisis of 2011-2012.

[15]It is possible to calculate corporate basis for currencies other than USD. However, the overwhelming majority of corporate bonds issued by EA firms are denominated either in USD or in EUR. Hence, I restrict my analysis to EUR-USD pair.

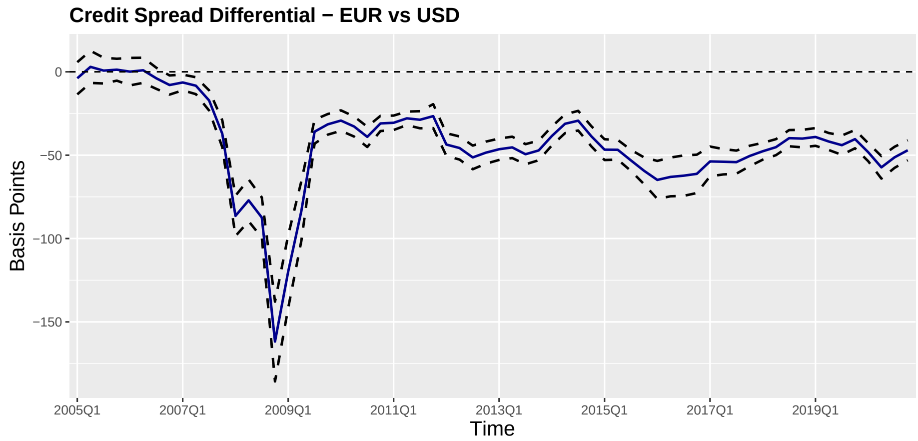

FIGURE 4
Credit Spread Differential - EUR vs USD. *Source:* Authors calculations, Refinitiv Eikon and Datastream.

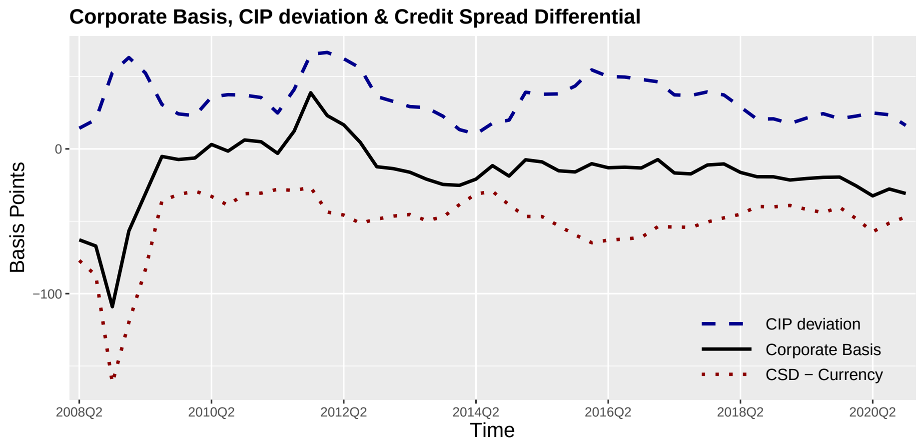

FIGURE 5
Corporate Basis, CIP Deviations & Credit Spread Differential
Source: Authors calculations, Refinitiv Eikon and Datastream.

4 The Choice of Foreign Currency Issuance

The main purpose of this section is to examine whether corporate basis drives foreign currency issuance decisions of EA NFCs. If it does, this implies that firms resort to global corporate bond markets to reduce their borrowing costs. In turn, this information will be used when studying heterogeneous firm responses to monetary policy surprises.

TABLE 1
Bond Issuances in USD vs Offshore Issuances by EA NFCs

USD Issuance vs Offshore Issuance

	USD denom.	Issued in US
Total Tranches	1,073	839
USD denom.	1,073	676
Issued in US	676	839
Issued in Euro Area	34	-
Issued by Parent	320	272
Issued by Subsidiary	753	567

Source: Refinitiv Eikon.

4.1 Data and Methodology

After applying several filters to the raw bond dataset obtained from Refinitiv Eikon[16] and consolidating the bonds at the ultimate parent level, I end up with 5,375 corporate bonds (4,302 EUR + 1,073 USD) issued by 1,199 distinct EA private NFCs in consolidated basis [17] between 2008Q2 and 2019Q4[18]. The details of the filtering procedure along with the summary statistics of the resulting bond dataset are presented in the Data Appendix B.2. There, I also show that Refinitiv Eikon's bond dataset is fairly representative of overall market trends by comparing it with the ECB's aggregate corporate bond issuance data. Table 1, on the other hand, summarizes the relationship between offshore issuance and issuances in USD. In this paper, I generally use these two different concepts interchangeably. The reason I do this is that the vast majority of USD issuances take place outside the Eurozone border and mostly via subsidiaries. Similarly, bonds issued in the US are typically USD-denominated and issued by subsidiaries of European firms.

Concerning the empirical investigation, I consider four main specifications. The first introduces a binary dependent variable taking 1 if firm i issues a USD denominated bond at quarter t and 0 otherwise as in equation (4). In this case, I estimate a panel Probit model with the following explanatory variables: firm size proxied

[16] As confirmed by an Eikon representative, this database includes all bond data available in SDC Platinum which has been heavily used by the earlier literature on corporate bond market.

[17] Before consolidation, the number of firms that issue these bonds is 2,463.

[18] I focus on 2008Q2-2019Q4 because, post-GFC, deviations from covered interest parity and thus the EUR-USD "corporate basis" became economically meaningful and the necessary micro-market data are consistently available. The window spans both tightening and easing yet ends before the Covid-19 shock to avoid the bond market turmoil and associated extraordinary interventions that would confound normal monetary transmission and funding substitution behavior. The sample period also coincides with the rapid expansion of euro area corporate bond markets, ensuring sufficient issuance volume and cross-sectional heterogeneity.

by the logarithm of firms total assets; leverage defined as the total debt of the firm divided by its total assets; balance sheet liquidity proxied by the sum of cash and short-term investments of the firm divided by its total assets; sales growth given by the quarterly change in net sales; cash flow over total assets where cash flow is calculated as the sum of net income before extraordinary items, depreciation and amortization; short term debt over total assets and finally the corporate basis. Summary statistics of the firm balance sheet, income statement and cash flow statement variables are presented in Data Appendix B.3. All explanatory variables except corporate basis are lagged by one quarter to reduce endogeneity concerns and winsorized at 1st and 99th percentile. All explanatory variables including corporate basis are standardized.

$$USD^1_{it} = \begin{cases} 1, & \text{if } USDiss_{it} > 0 \\ 0, & \text{otherwise} \end{cases} \tag{4}$$

The second specification mimics the same Probit exercise with the same independent variables but with a slightly different dependent variable. This time, I treat the value of the dependent variable in no bond issuance quarters as missing. Mathematically, the dependent variable takes the form of equation (5):

$$USD^2_{it} = \begin{cases} 1, & \text{if } USDiss_{it} > 0 \\ 0, & \text{if } USDiss_{it} = 0 \; \& \; EURiss_{it} > 0 \\ NA, & \text{if } USDiss_{it} = 0 \; \& \; EURiss_{it} = 0 \end{cases} \tag{5}$$

The regression form of the Probit model is given by equation 6 where $G(.)$ is the cumulative distribution function of the standard normal distribution, CB_t is the corporate basis, w is one of the firm-level covariates described above and $k \in \{1, 2\}$. α_s, β_q and γ_c represent sector, quarter and country fixed effects, respectively. Sector fixed effects are at the two-digit level using Thomson Reuters Business Classification codes.

$$P(USD^k_{it} = 1 | CB_t, w_{i,t-1}) = G(\alpha_s + \beta_q + \gamma_c + \theta CB_t + \sum_{w \in W} \delta_w w_{i,t-1} + \varepsilon_{i,t}) \tag{6}$$

The dependent variable in the third specification is the amount of USD issuances of a given firm to its total issuances at each quarter as in equation (7). This allows the dependent variable to take values between 0 and 1. In this case, I estimate a two-limit panel Tobit model[19] with the same explanatory variables as in the Probit specification[20]. In the last specification, I repeat the Tobit exercise but treat the values of the dependent variable as missing if firm i did not issue a bond in EUR or USD at quarter t. In mathematical terms, the dependent variable in this case is given by equation (8).

$$USD^3_{it} = \begin{cases} \frac{USDiss_{it}}{USDiss_{it}+EURiss_{it}}, & \text{if } USDiss_{it} + EURiss_{it} > 0 \\ 0, & \text{otherwise} \end{cases} \tag{7}$$

$$USD^4_{it} = \begin{cases} \frac{USDiss_{it}}{USDiss_{it}+EURiss_{it}}, & \text{if } USDiss_{it} + EURiss_{it} > 0 \\ NA, & \text{otherwise} \end{cases} \tag{8}$$

The regression form of the two-limit Tobit model is given by equations 9 and 10 where $l \in \{3, 4\}$.

$$y^*_{i,t} = \alpha_s + \beta_q + \gamma_c + \theta CB_t + \sum_{w \in W} \delta_w w_{i,t-1} + \varepsilon_{i,t} \tag{9}$$

Here, $y^*_{i,t}$ is an unconstrained latent variable capturing the propensity to tilt issuance toward USD (and thus can take any real value), while the observed issuance share is censored to lie in $[0, 1]$ as shown in equation (10).

$$USD^l_{it} = \begin{cases} 0, & \text{if } y^*_{i,t} \leq 0 \\ y^*_{i,t}, & \text{if } 0 < y^*_{i,t} < 1 \\ 1, & \text{if } y^*_{i,t} \geq 1 \end{cases} \tag{10}$$

4.2 Results

The results of the currency choice model regressions are given in Table 2. Column 1 presents the results for the Probit case where the dependent variable is given by (4). Size, leverage, cash flow and corporate basis are statistically significant at conventional levels with expected signs. We observe that as firm size increases, the probability of the firm issuing in USD increases. This is consistent with the notion

[19] Bruno and Shin (2017) also use a Tobit model where the dependent variable is the ratio of USD-denominated bond proceeds to total bond proceeds in a year.

[20] Theoretically, fixed effects Tobit/Probit model suffers from incidental parameters problem leading to inconsistent coefficient estimates. However, bias approaches zero for large T. Moreover, using a Monte-Carlo analysis, Greene (2004) shows that slope coefficients can be estimated consistently even for small T in the case of Tobit model.

that large firms are tapping global markets more frequently than others. The same is true for the leverage: more leveraged firms have higher propensity to tap foreign markets. On the other hand, firms with abundant cash flow are less likely to issue in USD. Finally, and most importantly for this paper, corporate basis is a significant determinant of a firms USD issuance decision. As corporate basis increases -in other words, as swap-hedged USD issuance becomes cheaper compared to issuing in EUR-, the probability that a given firm issues a corporate bond in USD increases.

Column 2 repeats the same Probit exercise with the dependent variable given by equation (5). In this case, size and corporate basis continue to be statistically significant whereas leverage and cash flow cease to be significant predictors of firms' USD issuance decision. Columns 3 and 4 present the results for the Tobit case with dependent variables given by equations (7) and (8), respectively. The results are in accordance with the Probit case with size and corporate basis being significant determinants of USD issuance decision of EA NFCs.

In terms of economics significance, reported average marginal effects indicate that the impact of corporate basis on USD issuance decision is substantial. In the case of the first model, a one standard deviation increase in corporate basis leads to a 0.5 percentage point (pp) higher probability of issuing in USD, almost a quarter of the unconditional probability that a given firm issues in USD in any quarter (2.1 pp). The marginal effect of corporate basis rises to 3.7 pp in the second model in which no bond issuance quarters are removed from the dataset[21]. Tobit models yield similar results.

Columns 5-8 report the results of the same analysis done in columns 1-4 with a new firm sample where firms operating in the energy sector are excluded. As discussed previously, one of the main reasons behind a firms offshore issuance choice is to hedge foreign exchange cash flows. As firms in the energy sector typically have high levels of foreign exchange cash flows, they might issue USD-denominated bonds in order to hedge those cash flows rather than to exploit borrowing cost differentials. By removing firms in the energy sector, I intend to address this concern to some extent by having a more homogeneous firm sample in terms of offshore issuance decisions. The results with the reduced firm sample are qualitatively similar to columns 1-4 with size and corporate basis remaining significant predictors

[21] If we limit the sample to firms that issued a USD denominated bond at least once in the sample period, the marginal effect of one standard deviation change in corporate basis rises to 2.1 pp and 4.7 pp in models 1 and 2, respectively. These results are not reported in the paper to save from space but are available upon request.

of firms' USD issuance decision in all cases.

Finally, columns 9 and 10 present the results of the Probit analysis for the sub-periods 2008Q2-2013Q4 and 2014Q1-2019Q4, respectively. This breakdown shows us that corporate basis remains to be statistically significant during the 2008-2013 sub-period and ceases to be so in the 2014-2019 sub-period. This difference hints us that firms may be ignoring changes in corporate basis when the basis is in the negative territory as was the case after 2013 (see Figure 5). After all, from a EA firm's perspective, a negative corporate basis implies that issuing in EUR is cheaper compared to issuing in USD and movements within the negative territory do not provide any incentives to switch to USD issuance.

TABLE 2
Regression Results of Firms' Currency Choice Model

	1	2	3	4	5	6	7	8	9	10
Corporate Basis	0.132***	0.203***	1.579***	1.142***	0.132***	0.218***	1.591**	1.122**	0.144**	−0.001
	(0.043)	(0.078)	(0.577)	(0.437)	(0.045)	(0.084)	(0.620)	(0.467)	(0.057)	(0.049)
Size	0.684***	0.492***	8.073***	2.383***	0.693***	0.533***	8.228***	2.592***	0.816***	0.599***
	(0.057)	(0.092)	(1.550)	(0.597)	(0.060)	(0.096)	(1.682)	(0.668)	(0.086)	(0.081)
Leverage	0.075*	0.069	0.932*	0.199	0.066	0.012	0.843	0.208	0.050	0.130**
	(0.040)	(0.073)	(0.491)	(0.350)	(0.043)	(0.076)	(0.535)	(0.381)	(0.061)	(0.055)
Bal. Sheet Liq.	0.016	0.160*	0.203	0.631	0.037	0.163*	0.456	1.007**	0.067	−0.066
	(0.047)	(0.089)	(0.564)	(0.431)	(0.051)	(0.092)	(0.613)	(0.485)	(0.066)	(0.075)
Sales Growth	−0.022	−0.121	−0.297	−0.829*	−0.010	−0.107	−0.162	−0.736	−0.023	−0.013
	(0.048)	(0.099)	(0.594)	(0.494)	(0.050)	(0.101)	(0.632)	(0.536)	(0.077)	(0.065)
Cash Flow	−0.143***	−0.081	−1.751***	−1.119***	−0.075	−0.114	−0.922	−0.534	−0.116*	−0.179***
	(0.041)	(0.077)	(0.566)	(0.402)	(0.051)	(0.086)	(0.637)	(0.452)	(0.064)	(0.055)
ST Debt	0.053	0.035	0.622	−0.019	0.043	0.008	0.508	0.117	0.043	0.042
	(0.044)	(0.080)	(0.541)	(0.421)	(0.052)	(0.099)	(0.638)	(0.510)	(0.067)	(0.062)
Intercept	−2.945***	−2.378***	−35.15***	−10.48***	−2.889***	−2.103***	−35.49***	−10.55***	−3.192***	−2.824***
	(0.196)	(0.289)	(6.542)	(2.212)	(0.205)	(0.314)	(7.04)	(2.46)	(0.301)	(0.268)
Mean (Y)	0.021	0.231	0.019	0.214	0.020	0.234	0.019	0.216	0.024	0.018
Marginal Effect of										
Corporate Basis	0.005***	0.037***	0.005***	0.039***	0.005***	0.041***	0.005***	0.038***	0.006**	−0.000
	(0.002)	(0.014)	(0.002)	(0.014)	(0.002)	(0.016)	(0.002)	(0.014)	(0.002)	(0.002)
Observations	10,782	963	10,782	963	9,850	845	9,850	845	5,298	5,484
Industry FE	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Quarter FE	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Country FE	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Note:	*p<0.1; **p<0.05; ***p<0.01									

Notes: The table provides coefficient estimates from regressing dependent variables in (4), (5), (7) and (8) on firm characteristics and corporate basis. Columns 1-4, 9 and 10 use the whole firm sample whereas firms in the energy sector are excluded in columns 5-8. Columns 1-8 are based on the whole sample period while columns 9 and 10 use 2008Q2 - 2013Q4 and 2014Q1 - 2019Q4 sub-periods. The dependent variables in columns 1-4 are (4), (5), (7) and (8), respectively. Similarly, dependent variables in columns 5-8 are (4), (5), (7) and (8), respectively. Finally, dependent variable in columns 9 and 10 is given by (4). All models include sector, quarter and country fixed effects.

5 Heterogeneous Investment Responses: Identification from Monetary Policy Surprises

Section 4 demonstrates that reducing borrowing costs is a driving factor behind EA NFCs' USD denominated bond issuances. Thus, we know that these firms actively seek for the best terms for their borrowing operations. The next question is then, whether firms that have access to global corporate bond markets use this access to insulate themselves from local monetary tightening. In this section, I study investment reactions of EA NFCs to monetary policy to answer this question. This requires a careful identification of exogenous monetary shocks which I address by following the high-frequency identification approach popularized by Gürkaynak et al. (2005) and Bernanke and Kuttner (2005).

5.1 High Frequency Identification of Monetary Policy Surprises and the Information Effect

In a nutshell, high-frequency identification (HFI) of monetary policy surprises involves an event-study analysis through which changes in prices of specific asset types such as stock prices, government bond yields of various maturities or interest rate futures are measured around a short time interval (typically intraday movements) surrounding monetary policy announcements. Provided that there is no other major event that would move these assets prices within such a short period, we can safely argue that changes in asset prices are mainly driven by monetary policy announcements. Since the expected component of monetary policy changes is most likely to be priced in before the announcement in forward-looking asset markets, such HFI amounts to measuring solely the surprise component of monetary policy announcements[22].

[22]The major advantage of HFI of monetary policy surprises is that it largely eliminates the endogeneity problem associated with the omitted variable and simultaneity biases which would likely exist in lower frequency analysis. For instance, using monthly or even weekly frequency, it is not easy to establish a causal relationship between monetary policy announcements and asset prices. It is quite possible that the central bank and asset prices are both responding to some other external shock in which case measuring the impact of monetary policy suffers from an omitted variable bias problem. Alternatively, the central bank may also be responding to abrupt movements in asset prices to calm financial markets in which case the simultaneity related bias would lead to inconsistent estimates. HFI removes these concerns to a great extent by narrowing the time interval during which asset price changes are measured so that they can exclusively be attributed to monetary policy surprises.

In this paper, I use the recently published, regularly updated and publicly available Euro Area Monetary Policy Event Database (EAMPD) à la Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019). EAMPD allows us to observe movements in the yield curves of German, French, Italian and Spanish government bonds and of Overnight Index Swap (OIS) rates. I choose working with the OIS rates as its term structure is typically the best proxy of the risk-free yield curve in the EA (European Central Bank, 2014)[23]. Given that surprise data for OIS maturities greater than three years is not available before 2011, I use OIS rates with 1-month, 3-month, 6-month, 1-year, 2-year and 3-year maturities. This choice also allows us to capture the impact of conventional monetary policy target rate changes along with the impact of forward guidance and quantitative easing[24]. Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019) present OIS rate changes for three time intervals, namely the press release window, the press conference window and monetary event window that comprises the first two windows[25]. I use monetary event window in my analysis to study the impact of both target rate changes and unconventional policies.

To purge the monetary policy surprise series from the information effect that they carry, I apply "poor man's sign restrictions" as suggested by Jarociński and Karadi (2020)[26]. This approach involves keeping the level of monetary policy sur-

[23]An overnight index swap is an interest rate swap whereby swap counterparties exchange fixed-rate cash flows with floating-rate cash flows with the floating leg being tied to the geometric average of an overnight interest rate, EONIA in the case of euro area. Being quoted in the fixed rate, these swaps reflect markets expectations about future EONIA rates. As EONIA follows ECB's monetary policy rate very closely, OIS rates also provide valuable information about expectations of ECB's future policy stance.

[24]Studies on the impact of monetary policy focusing on pre-GFC period typically use changes in short-term rates such as 1-month Fed fund futures as proxy for monetary policy surprises. After hitting the zero lower bound, however, central banks expanded their policy toolkit to affect long-term rates. Thus, high frequency changes in short-term rates may not capture the true monetary policy stance post-GFC. In line with this, Wright (2012) uses US Treasury bond futures of 2,5,10,30-year maturity whereas Gertler and Karadi (2015) use 1-year and 2-year government bond rates as their policy indicators. Besides, Gürkaynak et al. (2022) show that the surprise effect of monetary policy materialized mostly through forward guidance both before and after the zero lower bound period.

[25]ECB's monetary policy announcements have two distinct phases. In the first phase, a press release is delivered stating the policy decision without further explanation. It is followed by the second phase when a press conference is held communicating the rationales behind the decisions taken which also shapes expectations regarding the future path of monetary policy. See Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019) for the detailed characteristics of ECB's monetary policy announcements and a chronological exposition of each monetary policy announcement event.

[26]In recent years, a growing number of studies emphasize the need to purge monetary policy surprises from the information shocks that they carry when constructing true monetary policy sur-

prise same if it is of the opposite sign with the stock markets reaction around the event window and restricting it to zero otherwise. When applying this restriction, I compare the signs of 2-year maturity OIS surprises and changes in EURO STOXX 50 index around monetary policy announcement events as drawn in Figure 6[27]. If their signs are the same, then I set the OIS surprise value for each maturity to zero. After applying the restrictions where necessary, I aggregate OIS surprises to quarterly frequency for each maturity by summing OIS surprise changes that happen at the same quarter. Finally, I take the first principal component of these restricted and aggregated surprise series as my measure of "true" monetary policy surprises which I call as OISPRCT.

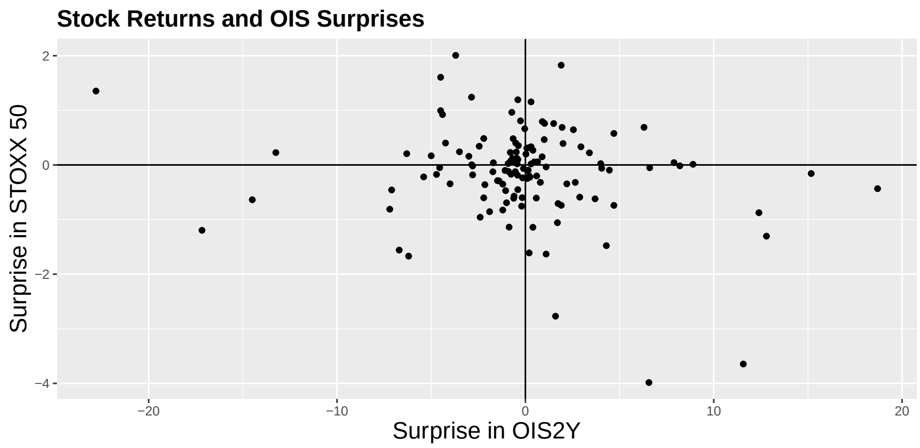

FIGURE 6
Surprises in STOXX50 and OIS2Y. *Source:* Euro Area Monetary Policy Event Study Database.

prises (Jarociński and Karadi, 2020; Nakamura and Steinsson, 2018; Miranda-Agrippino and Ricco, 2021). Information shocks are at work when monetary policy announcement implicitly reveals the central bank's assessment of the state of the business cycle. For instance, a surprise policy rate hike could induce lower stock prices and lower investment through a genuine monetary shock effect while it could also be suggestive of a stronger economic outlook than what is perceived before by market participants leading to a strong information effect. If the information effect dominates the genuine effect, then it is possible that the market responds to monetary policy changes in ways that contradict the standard theory. In this vein, Jarociński and Karadi (2020) show that positive interest rate changes that are accompanied by positive stock returns -indicative of a strong information shock- around monetary policy announcements lead to higher real activity and higher price level. This concern is particularly important for the Eurozone given ECB's highly transparent monetary policy implementation.

[27] I use the 2-year rate due mainly to two reasons. First, the 2-year rate is likely to represent the stance of monetary policy the best since it has the highest correlation with the first principal component of various maturities. Second, while it is widely used in the literature since it captures the impact of unconventional monetary policy, the 2-year rate is also largely free from the zero lower bound as shown by Swanson and Williams (2014) for the U.S. Using 1-year rate as the benchmark instead of 2-year rate produces very similar results.

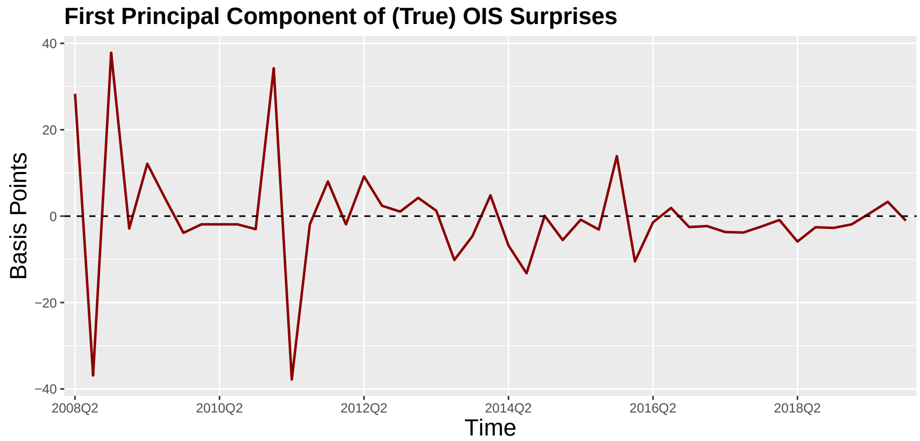

FIGURE 7
First Principal Component of (True) OIS Surprises. *Source:* Authors calculations based on Euro Area Monetary Policy Event Study Database.

Figure 7 depicts time series of OISPRCT. The correlations between quarterly aggregated surprise changes in OIS rates of various maturities including OISPRCT are given in Figure B.3 in the Data Appendix B.4. Table B.5 in the Appendix presents summary statistics of monetary policy surprises.

5.2 Methodology

Since investment is a slowly moving variable, monetary policy affects it with some lag. Following the recent literature (Jeenas, 2024; Ottonello and Winberry, 2020; Cloyne et al., 2023; Crouzet, 2021), I adopt the panel version of local projections approach pioneered by Jordà (2005). More specifically, I consider the following model for each horizon $h = 0, 1, , 16$.

$$\begin{aligned}\Delta_h log(k_{i.t+h}) = log(k_{i,t+h}) - log(k_{i,t-1}) =& f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}\eta_t \\ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1}\eta_t + \varepsilon_{i,t+h}\end{aligned} \tag{11}$$

where $k_{i,t}$ is the capital stock of firm i, f_i^h represents firm-fixed effects that control for firm specific time-invariant factors, $\lambda_{s,t}^h$ and $\psi_{c,t}^h$ are sector-time and country-time fixed effects controlling for time-varying sector-level and country-level heterogeneity within the euro area[28], η_t stands for the information effect corrected

[28]Some studies show that some industries (e.g. consumer durables sector) are affected more by monetary shocks due to a higher interest rate elasticity of demand (Peersman and Smets, 2005;

monetary policy surprise OISPRCT as described in Section 5.1, and W is the set of the quarterly-reported firm characteristics sourced from Refinitiv Eikon. Here, $OB_{i,t}$ is a binary indicator for opportunistic borrowing, equal to one for firms with prior USD bond issuance experience when the corporate basis is positive (so that FX-hedged USD borrowing is cheaper than EUR borrowing), and zero otherwise (see equation 12 below).

Equation (11) is symmetric in the sense that monetary easing and tightening episodes are treated equally. Given that the leakage effect that I mention in Section 2 is likely to be active during monetary tightening episodes[29], the baseline regression model is a slightly modified version of (11) where I simply replace η_t with η_t^+ which is equal to the interaction of η_t with a monetary tightening dummy in the spirit of Bernanke and Kuttner (2005), Dao et al. (2021) and Canofari et al. (2025). Accordingly, θ^h measures the differential dynamic response of investment to monetary tightening for firms that have the means to borrow opportunistically. A positive θ^h implies that these firms do not reduce their investment as much as others in response to monetary tightening. Hence, if the mechanism in Section 2 holds, we expect a significantly positive θ^h.

The firm sample consists of EA NFCs that issued a corporate bond at least once between 2008Q2 and 2019Q4[30]. The firm-level covariates that I include in the investment dynamics model are quite standard in the literature and include size, leverage, balance sheet liquidity, sales growth, cash flow over total assets and short term debt over total assets. Data Appendix B.3 provide more information about firm-level data. All firm-level covariates are winsorized at 1% and 99% level to reduce the impact of outlier observations. Following the standard practice in firm-level analyses of growth rates, I also trim the growth rate in the dependent variable separately for each horizon, h.

An important variable in this section is a proxy variable indicating whether a firm borrows opportunistically or not. Unfortunately, it is practically impossible to gauge whether a given firm borrows in foreign markets due to opportunistic motives as there can be other reasons behind a firm's offshore issuance decision. Nev-

Durante et al., 2022).

[29]During monetary tightening, domestic credit supply contracts and local borrowing costs rise, so firms with access to global bond markets can substitute into cheaper FX-hedged funding when the corporate basis is favorable, dampening transmission. In easing phases, domestic credit is already abundant and cheap -and the basis often compresses or reverses- so issuing abroad offers little incremental advantage, leaving the "leakage" margin largely inactive.

[30]See Appendix B.2 and Table B.2 for information about the bond sample.

ertheless, since corporate basis is a significant determinant of firms' USD issuance as shown in Section 4, there is sufficient ground to be confident that opportunistic borrowing is one of these reasons[31].

In the baseline case, the binary opportunistic borrowing variable $OB_{i,t}$ takes 1 if firm i has issued at least one USD-denominated bond until quarter $t-1$ and if corporate basis is positive. It takes 0 otherwise. Bond issuance condition implicitly assumes that if a given firm issued a USD-denominated bond in the past, it has the means to do so should the need arises given large fixed costs of accessing global corporate bond markets. Thus, without any further condition imposed, it rather would serve as an access to global corporate bond markets dummy. To take cost reduction motives into account, I further impose the condition that corporate basis is positive. The combination of the two conditions allows us to identify firms that are able to borrow opportunistically given by their access to global corporate bond markets when borrowing in USD is cheaper compared to borrowing in EUR. In fact, when the positive corporate basis condition is not imposed, as we shall see shortly, response heterogeneity disappears, suggesting that access to global markets alone is not sufficient to drive heterogeneous firm behavior. Rather, firms react heterogeneously to monetary tightening only when borrowing in USD is cheaper than borrowing in EUR, highlighting the importance of favorable borrowing conditions in global markets[32].

$$OB^1_{it} = \begin{cases} 1, & \text{if } USDiss_i \text{ until } t-1 > 0 \ \& \ CB_t > 0 \\ 0, & \text{otherwise} \end{cases} \tag{12}$$

5.3 Results

Before studying heterogeneous firm reactions, I first estimate average effects of monetary policy by removing the sector-time and country-time fixed effects and interaction terms from the model. This leads to equation (13) where monetary pol-

[31] A related question is how bond proceeds are used. Colla and Nagler (2025) show that bond issuance tends to raise firms' cash buffers -i.e., a portion of proceeds is saved rather than spent immediately, unlike loans. De Gregorio and Jara (2024), on the other hand, show that this saving is largely a save-to-invest strategy: offshore issuance elevates cash when conditions are favorable, and those balances later unwind as firms deploy funds into real investment. Taken together, the evidence implies that bond proceeds often pass through cash first but are ultimately aimed at financing investment.

[32] In Section 6, I also consider a slightly modified version of equation (12) where access is limited to firms which issued in USD in the last 5 years.

icy tightening variable is included as a standalone regressor. ξ_q^h represents quarter fixed effects which control for seasonality effects for firm investment.

$$log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \xi_q^h + \gamma^h \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \varepsilon_{i,t+h} \tag{13}$$

Estimated impulse response coefficients of the monetary tightening variable, γ^h, are drawn in Figure 8. The coefficients in the figure are scaled so that they represent the change in fixed capital expenditure following a one standard deviation increase in η_t^+. The same scaling will be held throughout the rest of the analysis. I double cluster standard errors at the firm and time (quarter-year) level. Figure 8 indicates that γ^h is negative as expected for each horizon. It reaches its minimum around nine quarters after the monetary policy tightening surprise and the effect of monetary policy diminishes thereafter. In economic terms, a one standard deviation increase in monetary tightening leads to a 2.0%-2.4% cumulative reduction in fixed capital expenditure around 7-12 quarters following the policy tightening surprise[33]. This is a sizable effect since the mean investment in the firm sample over 10 quarters is 11% as can be seen from Table B.4. Moreover, the magnitude aligns with studies focused on monetary tightening. Using Compustat data, Perez-Orive and Timmer (2023) show that a one standard deviation increase in their tightening surprise measure yields a cumulative 2.8% decline in firm investment over eight quarters.

Since my focus is on monetary tightening, I present only the results for equations with η_t^+ in the main text and leave the results for the η_t (symmetric case) to the Appendix. The results for the average effect of monetary policy in the symmetric case are qualitatively similar as shown in Figure D.1. The main difference is that, in the symmetric case, γ loses statistical significance after four quarters and transmission is more muted since the peak response to a one standard deviation monetary policy surprise is a 1.1% decline in fixed capital expenditure. This accords with evidence that documents stronger effects of monetary tightening compared to easing (Tenreyro and Thwaites, 2016; Canofari et al., 2025). In a recent study, Perez-Orive et al. (2024) argue that when firms face multiple financing constraints, monetary policy transmission is asymmetric, with easing dampened because the least responsive constraint continues to bind whereas tightening is amplified because the most responsive constraint binds.

[33]Notice that the vertical axis represents accumulated (log) change in physical capital and not the change in the investment rate between quarter h and $h-1$.

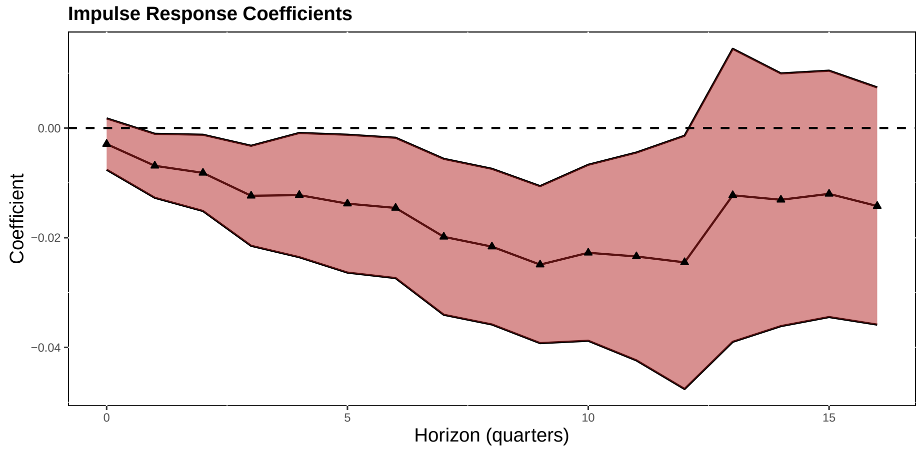

FIGURE 8

The Average Effect of Monetary Tightening Surprises on Firms Fixed Capital Expenditure

Notes: The figure depicts impulse response coefficients, γ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \xi_q^h + \gamma^h \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text. The coefficient γ^h is scaled so that it represents the change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

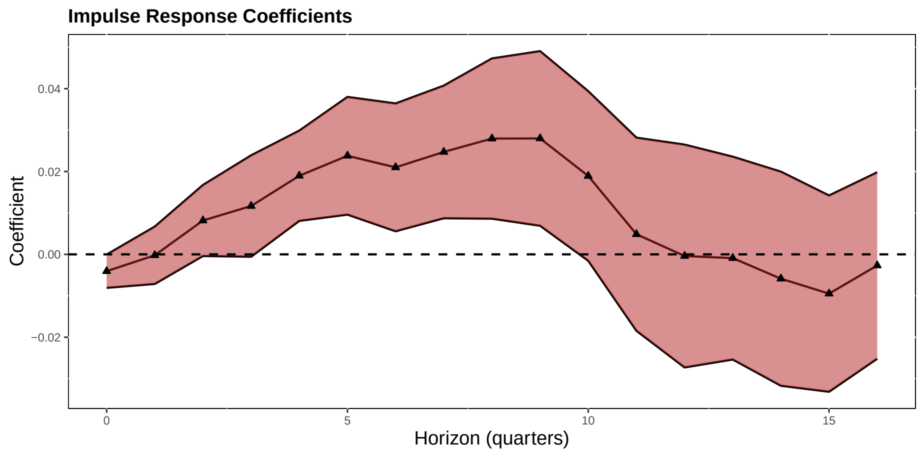

FIGURE 9

The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text and $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

Next, I study differential investment responses to monetary policy (equation 11). Here, the coefficient of interest is the interaction term, θ^h which provides differential responses of firms that can borrow opportunistically to monetary tightening. Figure 9 depicts estimated θ^h for each horizon h using OB^1_{it} given by (12). In Figure 9, θ^h becomes significantly positive after around four quarters and ceases to be so after ten quarters. In terms of magnitude, it is very close to the average effect coefficient, implying that firms able to borrow opportunistically nearly neutralize the average contraction in investment following monetary tightening.

This finding suggests that the impact of monetary policy on firms investment decisions is heterogeneous and depends on whether a firm is able to tap global corporate bond markets when issuing in foreign currency provides cost-saving opportunities. In other words, firms that are able to reduce their borrowing costs do not decrease their investment as much as other firms in response to monetary tightening.

The estimated coefficients from the symmetric case with η_t is reported in Figure D.2 in Appendix D for comparison. The symmetric case is both qualitatively and quantitatively similar to the asymmetric case (tightening episodes).

A caveat is that this heterogeneity pertains to the pre-2013 period. Following the 2013 tapering news alongside Federal Reserve's subsequent tightening cycle and ECB's quantitative easing efforts, currency hedged borrowing cost of issuing in USD never became lower than the cost of issuing in EUR for EA NFCs evidenced by the negative corporate basis for this period. Consequently, after 2013, EA NFCs could not simultaneously hedge currency risk and lower borrowing costs via global bond markets[34]. By contrast, the post-2013 environment offered cost-saving opportunities to U.S. firms that have access to European bond markets, shaping their investment responses to Fed policy as I elaborate in Section 6.

In interpreting this result, one should be cautious since heterogeneous firm response may reflect asymmetric financial constraints rather than the mechanism of interest. If firms classified as having global bond market access are intrinsically less financially constrained for reasons unrelated to cross-border access than firms issuing only locally, they may react less to tightening regardless of that access. Differences in investment opportunities could likewise generate differential responses. The next section details these concerns and presents additional robustness checks.

[34] Admittedly, because the corporate basis is an aggregate, estimated measure, borrowing in USD may still offer cost-saving opportunities to certain firms post-2013. My analysis does not attempt to identify such opportunities, which are difficult to detect in reliable ways.

6 Robustness Checks and Additional Results

This section presents some further analyses that show whether the results obtained in Section 5 are robust to changes in methodological details. I also conduct a placebo test to investigate whether heterogeneous investment response survives when borrowing in global corporate bond markets do not offer cost-saving opportunities. Finally, I demonstrate that the baseline results have external validity as evidenced by the stock market's reactions to monetary policy.

6.1 Different Interest Rate Measures

Monetary policy changes are at the heart of the analysis conducted in this paper. Therefore, it is important that results are not very sensitive, at least qualitatively, to the choice of how we measure monetary policy. In this section, I consider four alternatives to the baseline surprise series.

6.1.1 True monetary policy surprises vs original surprises

In Section 5, I used monetary policy surprise series that is purged of the information effect that it carries by imposing restrictions elaborated in Section 5.1. In this section, I do not impose any restrictions and use the original surprise series instead. Each OIS surprise variable (with maturities: 1M, 3M, 6M, 1Y, 2Y, 3Y) is aggregated into quarterly frequency and their first principal component is computed. Some descriptive statistics and time series plot of the resultant series are presented in Table B.6 and Figure B.4 in the Appendix.

6.1.2 Grouping policy changes

In this part, I group each monetary policy surprise observation in three bins that represent easing, tightening and no action with values -1, 1 and 0, respectively. Cutoffs for no action is taken as -10 bps and 10 bps. The resulting surprise series is called OISPRCbins. To illustrate, if the value of the surprise is -12 bps, OISPRCbins takes -1; if surprise is 13 bps, OISPRCbins takes 1; and if surprise is -5 bps or 4 bps, OISPRCbins takes 0. Descriptive statistics and time series plot of OISPRCbins are presented in Table B.6 and Figure B.5.

6.1.3 A wider term structure

While the current analysis uses up to 3-year OIS rates due to high-frequency data unavailability for longer term rates before 2011, it could be important to consider a wider term structure incorporating rates of 4-10, 20 and 30 years which would reflect better the impact of post-crisis QE and forward guidance efforts[35]. I achieve this by including German bond yields of these longer maturities in my calculation of the first principal component of monetary policy surprises. The resulting series is called widerPRC. Descriptive statistics and time series plot of widerPRC are presented in Table B.6 and Figure B.6.

6.1.4 Nominal interest rates

I also use levels of nominal interest rates instead of high-frequency monetary policy surprises in line with Ippolito et al. (2018). While stock market is forward looking and responds only to unanticipated changes in monetary policy, investment is likely to respond to expected interest rate changes as well through the latter's impact on cost of capital and consumer demand. For this purpose, I aggregate daily OIS rates into quarterly frequency by taking their quarterly average. The underlying rate of swaps is EONIA. Again, I calculate the first principal component of OIS rates of different maturities. Descriptive statistics and time series plot of the resulting series are presented in Table B.6 and Figure B.7.

6.1.5 Results

Each of the four alternative monetary policy variables replaces η_t in equation (11) and I interact them with a monetary tightening dummy. The results with new monetary policy variables are given by Figures D.3-D.6 in Appendix D. Overall, the results are in line with the baseline and suggestive of heterogeneous monetary policy transmission with positive and statistically significant coefficients between 4-10 quarters after monetary tightening.

[35] A wider term structure can also help overcome the information effect problem better (see Section 5.1) as shown by Bu et al. (2021).

6.2 Threats to Identification I - Asymmetries of Financial Constraints between USD-issuers and EUR-only Issuers

It is possible that the results suggesting a heterogeneous monetary policy transmission of Section 5.3 are driven by asymmetries of financial constraints between USD-issuers and EUR-only issuers[36] in the sense that firms that suffer less from financial frictions may be the ones that are able to issue in global corporate bond markets. Thus, the reason they react less to monetary policy could be the fact that they are less financially constrained anyway independent of whether they tap foreign debt markets. This is an important concern which I address in four ways in this paper.

First, all firms in my sample are bond-issuers. This provides a natural control for financial frictions since all firms have access to at least the local bond market, therefore they are not completely bank-dependent. Second, I have $size \times \eta_t$, $BSL \times \eta_t$ and $cashflow \times \eta_t$ in my baseline regressions which already control for the differential effect of monetary policy for larger and more liquid firms. Third, I also add an additional control for financial frictions: Standard & Poors Long Term Issuer Rating. I create three dummies standing for Not Rated, Non-Investment Grade and Investment Grade firms. Interactions of these dummies with monetary policy surprises are included in the model to control for remaining financial frictions. The results are presented in Figure D.7.

Finally, if heterogeneous responses are driven by underlying asymmetries of financial constraints between USD-issuers and EUR-only issuers rather than by USD-issuers' cost-saving opportunities in global debt markets, we would expect to see a positive θ^h independent of the level of the corporate basis. To test whether this prediction holds, I modify opportunistic borrowing variable as follows[37]:

$$OB^2_{it} = \begin{cases} 1, & \text{if } USDiss_i \text{ until } t-1 > 0 \text{ \& } CB_t < 0 \\ 0, & \text{otherwise} \end{cases} \tag{14}$$

If θ^h is not significantly positive under this scenario, it would imply that heterogeneous reaction depends on the borrowing cost differential between USD and

[36] "EUR-only issuers" refers to firms that issued at least one bond denominated in EUR but never issued a USD-denominated bond throughout the sample period.

[37] Naming this new variable as "opportunistic borrowing" is at odds with the definition of opportunistic borrowing as explained in the introduction. However, I keep its name as it is in order to ease comparison with the baseline.

EUR, hence is unlikely to be driven by underlying asymmetries of financial constraints between firms that are not related to their access to global markets. The results with this placebo test are given in Figure D.8. The insignificant coefficients indicate that heterogeneous responses emerge only when USD funding is cheaper than EUR; when the cost ranking reverses, the differential response disappears.

Overall, results do not support the idea that financial frictions that are not related to firms' access to global corporate bond markets drive heterogeneous firm reactions to monetary tightening. Rather, it is the ability of certain firms to tap global corporate bond markets when issuing in foreign currency provides cost-saving opportunities that leads to heterogeneous firm reactions.

6.3 Threats to Identification II - The Role of the Equity Constraint Channel

In Section 6.2, I focus on financial constraints in a broad sense. However, a growing literature emphasizes that financing frictions can be dimension specific in the sense that some firms are primarily constrained in their access to debt, whereas others face tighter constraints on raising external equity (Hoberg and Maksimovic, 2015; Linn and Weagley, 2024). These distinctions matter for monetary policy transmission.

A key premise in Almeida et al. (2025) is that the firms commonly labeled as "financially constrained" may often rely on equity issuance rather than debt to obtain liquidity. Moreover, the pecking order theory and its empirical support imply that external equity is the least preferred source of funding. This suggests that firms constrained in equity finance are likely to be more severely constrained than debt-focused constrained firms and therefore more sensitive to contractionary monetary policy shocks. Consistent with this view, Hoberg and Maksimovic (2015) show that equity-focused constrained firms appear more constrained overall and are particularly affected by large negative shocks. Accordingly, Almeida et al. (2025) argue that an "equity constraint channel" amplifies the effects of monetary tightening because equity-dependent firms experience a sharper deterioration in financing conditions and consequently cut equity issuance and real expenditures, including investment and R&D, by more.

A natural concern is whether the opportunistic borrowing channel I document could be confounded with an equity financing constraint channel of the type emphasized by Almeida et al. (2025). Measuring equity-focused constraints in their setting relies on the U.S. institutional infrastructure of 10-K filings and text-based classifications that identify firms reporting delayed investment due to liquidity conditions and a reliance on equity financing (Hoberg and Maksimovic, 2015), with Linn and Weagley (2024) extending coverage via a random-forest mapping from accounting variables to those text-based measures. While this approach is powerful in the U.S., it is difficult to implement in a comparable way for European firms because there is no single, standardized and readily usable counterpart to the 10-K across jurisdictions and languages. Thus, implementing an HM/LW-style text-based approach would require collecting and harmonizing firm disclosures across multiple countries and building a new model, which is outside the scope of this paper.

To assess whether the opportunistic borrowing channel is distinct from an equity financing constraint channel, I introduce a transparent equity-constraint proxy based on the firm characteristics that Linn and Weagley (2024) and Almeida et al. (2025) identify as distinguishing equity-focused constrained firms from both financially unconstrained firms and debt-focused constrained firms: firm size (smaller), cash flow (lower), balance-sheet liquidity (higher), and Tobin's Q (higher; proxied by price-to-book). These sign restrictions have a straightforward economic interpretation[38]. In Linn and Weagley (2024) and Almeida et al. (2025), equity-focused constrained firms look like "early-stage, growth" firms: they are younger and smaller, have lower current cash flow, have higher Tobin's Q and their balance sheets are more liquid.

To construct the index, I standardize each variable within quarter: size (log assets), cash flow/total assets, balance-sheet liquidity (as defined in Section 4.1), and Tobin's Q (proxied by price-to-book ratio). I then construct a simple Equity-Constrained Index: $ECI_{i,t} = \frac{1}{4}(-z_{Size_{i,t}} - z_{CF_{i,t}} + z_{Liquidity_{i,t}} + z_{Q_{i,t}})$[39]. I then define

[38] Smaller firms typically have a thinner public information environment, with less analyst coverage and less investor attention, which exacerbates information frictions. This raises the marginal cost of issuing equity relative to internal funds. Lower cash flow captures limited internal financing capacity, so these firms are more likely to need external financing when it is most costly. Higher Q typically indicates stronger growth opportunities and a greater importance of intangible investment, which further limits the ability to finance through collateralized borrowing and pushes financing needs toward equity. Finally, higher balance-sheet liquidity reflects precautionary cash management: when access to external financing is fragile, constrained firms have incentives to hold liquid assets and to be reluctant to run them down.

[39] I also consider an alternative version of the index that includes firm age, reflecting the fact that equity-focused constrained firms are typically younger (Linn and Weagley, 2024; Almeida et al., 2025). The results are qualitatively similar. I focus on the four-variable version in the main analysis

an indicator variable that equals one for the least equity constrained firms, defined as those in the bottom quintile of $ECI_{i,t}$ in a given quarter, and zero otherwise.[40].

In the baseline regression, I add this indicator and its interaction with the monetary policy shock, and I drop the individual firm controls and the associated interaction terms. The newly introduced interaction term captures the differential effect of monetary policy on the least constrained firms. The interaction between the opportunistic borrowing dummy and the monetary policy shock then captures the remaining heterogeneity in monetary transmission that is orthogonal to equity constraints. This specification yields estimates of the opportunistic borrowing channel that are very similar in magnitude and statistical significance to the baseline, as shown in Figure D.9. Hence, the main results are not driven by equity financing constraints as captured by this proxy.

6.4 Threats to Identification III - Profitable Investment Opportunities

It is also possible that firms that have access to global corporate bond markets have more profitable investment opportunities in comparison with other firms. A more profitable firm with ample investment opportunities can be expected to reduce its investment less relative to other firms when monetary policy tightens. If so, heterogeneous firm reaction may emerge due to different investment opportunities these firms have and not because of cost-saving opportunities of borrowing in global corporate bond markets.

To isolate my analysis from such effects, I control for investment opportunities. In the baseline case, the econometric model already incorporates sales growth which is frequently used as a proxy for investment opportunities firms have. In this section, I also add Tobin's Q along with its interaction with monetary policy surprises. Tobin's Q is another frequently used proxy for profitable investment opportunities in the literature. Q is itself proxied by price-to-book ratio which is also obtained from Refinitive Eikon for each firm in my sample. The results with the modified firm characteristics set are given in Figure D.10 and are largely in line with baseline results.

because incorporating age leads to a substantial loss of observations, as age information is missing for a sizeable share of firms in the sample.

[40]Defining the indicator using the bottom quartile or the bottom tercile yields qualitatively similar results.

Finally, it is also possible to use equation (14) in this section as well. If profitable investment opportunities that firms with access to global corporate bond markets have is the main reason why these firms are less reactive to monetary tightening, we would expect this relation to be independent of the level of corporate basis. However, Figure D.8 demonstrates that heterogeneous response disappears when borrowing in USD is more expensive borrowing in EUR.

6.5 Other Robustness Checks

I further do the following. First, I tighten the constraint when constructing the opportunistic borrowing variable in a way that it takes 1 only if the firm issued at least one USD-denominated bond in the last five years (instead of anytime until $t-1$). This choice aims to remove the concern that a firm might not be able to borrow in global markets anymore even though it did so in the distant past. Hence, a firm is assumed to have access to global corporate bond markets only if it issued a foreign currency bond within the last five years. This specification leads to (15):

$$OB_{it}^{3} = \begin{cases} 1, & \text{if } USDiss_i \text{ in the last five years} > 0 \ \& \ CB_t > 0 \\ 0, & \text{otherwise} \end{cases} \tag{15}$$

Results from the alternative specification are presented in Figure D.11 in Appendix D and are broadly consistent with the baseline, corroborating heterogeneous firm responses to monetary-policy surprises.

Second, I also adapt the baseline empirical model to inventory investment in order to study heterogeneous inventory investment response of firms to monetary tightening. Appendix E presents the results that reinforce the findings from the study of fixed capital investment response of firms.

6.6 External Validity Check I - Stock Returns and Monetary Policy

So far, the analysis has examined firms' financing choices under shifting financial conditions and their investment responses to monetary policy. An equally interesting aspect of heterogeneous monetary policy transmission is stock market participants' view of how monetary policy affects each firm[41]. Studying stock market's

[41] Darmouni et al. (2020) show that bond-reliant firms' stock prices react more to monetary shocks compared to bank-dependent firms. Gorodnichenko and Weber (2016) report that conditional volatility of stock returns of firms with sticky prices are higher compared to those with more

reaction to monetary policy changes also helps assess the external validity of the results on heterogeneous investment responses. In the spirit of Bernanke and Kuttner (2005), I conduct an event study of individual stock returns around monetary policy announcement windows, asking whether pricing differs with a firm's ability to borrow opportunistically in global corporate bond markets. For this aim, I estimate variants of the following specification.:

$$\Delta log(p_{i,t}) = f_i + \psi_t + (\alpha + \theta\eta_t)OB_{i,t} + (\beta' + \gamma'\eta_t)w_{i,t-1} + \varepsilon_{i,t} \tag{16}$$

where $\Delta log(p_{i,t})$ is the log change (in p.p.) in closing quote of the stock price of firm i between the day after the monetary policy announcement and the day before the announcement. The time subscript t stands for one of the 120 monetary policy announcement events that happened between 2008 and 2019. f_i and ψ_t capture firm fixed effects and event fixed effects, respectively. η_t is the monetary policy surprise variable that is described in section 5 without quarterly aggregation and $OB_{i,t}$ is as in equation 12. $w \in W$ represents firm characteristics and include firm size, firm leverage, balance sheet liquidity, short-term debt over total assets and Q as described in sections 4, 5 and 6. The firm sample includes 594 listed ultimate parent EA NFCs which issued a corporate bond on a consolidated basis in EUR or in USD during the sample period.

In equation 16, θ captures the differential stock return response to monetary policy for firms with access to cheaper offshore bond finance relative to firms without such access. A positive θ implies that, when USD issuance offers cost-saving opportunities, investors expect access firms to fare better under a monetary tightening -showing higher (less negative) announcement window returns- than other firms.

I use a two-day window for stock returns as in Gürkaynak et al. (2022). In my case, this choice aims to address two concerns. First, the window should be narrow enough so that the impact of news releases other than the monetary policy announcement on stock returns are minimized. Second, it should also be wide enough so that there is enough time for individual stocks to be exchanged with significant volumes and price movements do not reflect only a handful of trades. The results presented below, however, are robust to a more conservative one-day

flexible prices following monetary shocks. In their seminal paper, Ehrmann and Fratzscher (2004) find that firms with high Tobin's q, low debt, low cash flows and small size respond more to monetary policy announcements. Using the S&P 500 sample, Gürkaynak et al. (2022) show that stock returns of firms with more cash flow exposure are affected more by monetary policy surprises.

window. Some summary statistics for stock returns are presented in Table B.7.

TABLE 3
Stock Return Regression Results

	(1)	(2)	(3)	(4)	(5)	(6)
η_t	-0.378^{***}					
	(0.048)					
$OB \times \eta_t$		0.441^{**}	0.539^{**}	0.467^{**}	0.479^{**}	0.598^{**}
		(0.207)	(0.220)	(0.232)	(0.233)	(0.298)
Firm Controls	N	Y	Y	Y	Y	Y
Firm Controls $\times$ η_t	N	Y	Y	Y	Y	Y
Firm FE	Y	Y	Y	Y	N	N
Event FE	N	Y	N	N	N	N
Sector $\times$ Event FE	N	N	Y	Y	Y	N
Country $\times$ Event FE	N	N	N	Y	Y	N
Sector $\times$ Country $\times$ Event FE	N	N	N	N	N	Y
Adj. R^2	0.002	0.226	0.243	0.267	0.264	0.278
N	55,493	26,386	26,386	26,386	26,386	26,386

*p<0.1; **p<0.05; ***p<0.01

Notes: The table provides coefficient estimates from variants of $\Delta log(p_{i,t}) = f_i + \psi_t + (\alpha + \theta\eta_t)OB_{i,t} + (\beta' + \gamma'\eta_t)w_{i,t-1} + \varepsilon_{i,t}$. The dependent variable, $\Delta log(p_{i,t})$ is the log change (in pp) in closing quote of the stock price of firm i between the day after the monetary policy announcement at time t and the day before announcement. η_t is monetary policy surprise series purged from information effect as explained in Section 5. Firm controls include size, leverage, balance sheet liquidity, short-term debt over total assets and Tobin's Q. All firm-level covariates and stock returns are winsorized at 1% and 99% level. Firm controls are lagged by one year prior to the monetary policy announcement event. Standard errors are clustered at the firm level.

Table 3 shows estimation results of equation 16 with different sets of fixed effects. Monetary policy surprises are rescaled so that a one unit increase represents a 25 bps increase in η_t which corresponds to roughly two standard deviations. The first column measures the average stock return response. In this regression, there is no firm-level control or an interaction term but only monetary policy surprise as a standalone regressor and firm fixed effects. The average effect is negative implying that a 25 bps increase in the monetary policy surprise variable leads to an approximately 0.4% decrease in stock returns on average.

Column 2 drops η_t from Column 1 and includes event fixed effects ψ_t. There, and in the rest of the columns, I also include firm characteristics and their interactions with monetary policy surprises. Column 3 includes sector-event fixed effects that control for event-varying industry-level heterogeneity that captures industry-specific effects of monetary policy announcement events. Column 4 saturates the model even further by adding country-event fixed effects that capture any time-varying country-level heterogeneity in stock returns. Columns 3 and 4 explicitly control for differential effect of monetary policy for firms in different sectors and different constituent countries of the Eurozone. In Column 5, I drop firm fixed effects from Column 4. In column 6, I again drop firm fixed effects and include only sector-country-event fixed effects. This sharpens identification and enables me to exploit only cross-sectional variation among firms (in particular, USD issuers vs EUR-only issuers) within each sector-country-event cell and absorb any common shocks (including monetary policy) that affect the firms in a given sector of a given country around a given monetary policy announcement.

Throughout columns 2-6, the interaction coefficient θ ranges from 0.4% to 0.6% and is statistically significant in all cases at 95% confidence level. These estimates imply that the stocks of firms able to borrow opportunistically in global bond markets fall less on tightening days than those of otherwise similar firms without such access. This offers external validation of the investment results: equity investors discount monetary shocks less for firms with foreign funding options.

6.7 External Validity Check II - The Case of the US Firms

To assess whether the heterogeneous investment responses documented above are unique to EA NFCs or reflect a broader phenomenon, it is informative to consider the case of U.S. corporations. A specific policy divergence episode between the Federal Reserve and the European Central Bank provides a natural setting for such a comparison. While the Fed began tapering its asset purchases in 2014 and initiated its first interest rate hike since the global financial crisis in December 2015, the ECB moved in the opposite direction -starting its quantitative easing program and pushing rates into negative territory. By the end of 2016, the shadow rate differential between the two monetary areas exceeded 600 basis points. In response to these divergent conditions, many U.S. firms increased their bond issuance in Europe, with EUR-denominated issuance by U.S. non-financial corporations surpassing 50 billion dollars per quarter during the 2014-2016 period.

In ongoing companion work (Benlialper and Ozturk, 2025), we investigate whether these firms exhibited heterogeneous investment responses to the Fed's tightening measures, depending on their access to European bond markets which is proxied by prior EUR denominated bond issuance. To this end, we match firm-level financial data from Compustat with bond issuance records from Refinitiv Eikon and estimate a difference-in-differences model centered around the divergence period. Firms with and without prior access to European markets display parallel investment trends before the divergence, but begin to diverge once policy paths split. The estimated responses suggest that firms with access to European funding sources exhibit a significantly more muted investment decline in response to Fed's tightening efforts.

These findings -while preliminary- suggest that the heterogeneity and leakage channels identified in the euro area context may also be active in the U.S., underscoring the external validity of the mechanism proposed in this paper.

7 Conclusion

This paper shows that access to global bond markets shapes the real effects of monetary tightening for a subset of EA NFCs. I document cost-driven USD issuance when the corporate basis favors dollar funding and show that firms that have the means to exploit these episodes experience smaller declines in fixed investment after monetary tightening than comparable non-issuers.

This finding confirms that there is a significant level of heterogeneity in firms' reactions to monetary tightening. This heterogeneity is not driven by asymmetric financial constraints faced by USD denominated bond issuers and firms that only issue in EUR. Nor is it driven by profitable investment opportunities that opportunistically borrowing firms might have. Furthermore, stock market participants' pricing behavior provides external validation of the heterogeneous investment responses and a similar pattern holds for U.S. NFCs.

From a corporate finance perspective, the results highlight how currency and venue choice in debt issuance alter the firms' effective cost of capital and, in turn, their investment responses to policy shocks. From a policy perspective, the findings imply that the investment channel of monetary transmission can be attenuated when global bond markets offer lower hedged funding costs; the magnitude of any attenuation depends on the prevalence of firms with reliable access to those markets.

The paper's findings have also indirect implications for the working of the bank lending channel of monetary policy transmission. As Sobrun and Turner (2015) discuss, as larger and more credible firms switch to foreign debt markets, domestic banks need to find other -less credible- domestic customers to extend loans. This will increase the risk taking of the domestic banking sector. At the same time, these market switching firms are likely to deposit the cash they raise offshore into their local bank accounts, easing the funding constraints of the domestic banks. Both indirect channels work against what the local central bank aims to achieve by monetary tightening, leading to further impairment. A quantitative investigation of these predictions would be an important contribution to the literature on the bank lending channel.

Two avenues for future research are particularly promising. First, banks' own cross-border issuance and its pass-through to loan pricing could interact with the mechanism studied here. Although this paper focuses on NFCs' foreign currency bond issuance, financial institutions are very active participants in corporate bond markets as well. Consistent with the leakage mechanism outlined in this paper, banks with superior access to foreign currency funding could be less exposed to domestic monetary tightening and contract loan supply less than other banks. Borrowers with relationships to such banks would, in turn, face softer funding constraints. Second, an analogous leakage channel likely operates through NFCs' activity in foreign currency loan markets as well: large foreign currency borrowings via (typically syndicated) bank loans can similarly insulate firms from local tightening, creating another impairment channel for monetary policy transmission.

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Online Appendix

Global Corporate Bond Markets and Local Monetary Policy Transmission

by Ahmet Benlialper

A Calculation of Corporate Basis

This Appendix elaborates the specific steps in the calculation of corporate basis. As equation (2) makes it clear, we need two terms to calculate the corporate basis. The term in the first bracket is the credit spread differential between EUR and USD. The second bracket is the CIP deviation which is proxied by the cross-currency basis of a given maturity. Corporate basis is, then, simply the sum of credit spread differential and the risk-free CIP deviation.

In this paper, I measure CIP deviations as the 5-year cross currency basis swap based on USD LIBOR and EURIBOR rates multiplied by minus one. For the credit spread differential, however, we do not have a clear-cut proxy. In the most ideal scenario, one can use the bond yield spreads of firms that issue two bonds at the same time, one in USD and the other in EUR, both of which have the same rating and maturity and so on, so that one can compare costs of issuing in USD and in EUR, ceteris paribus. Nevertheless, these cases being rare, it would be misleading to generalize such small number of occurrences.

Eventually, we need to come up with an estimation methodology and I do it by adopting the bottom-up approach using individual bond data pioneered by Gilchrist and Zakrajšek (2012) and estimating a bond pricing model along the lines of Liao (2020). Below, I explain the estimation procedure for credit spread differential between USD and EUR. The bond dataset that is used in estimating the credit spread differential is described in Data Appendix B.1.

$$S_i = \alpha + \beta D_i + \sum_{k \in \left\{r,m,a,ai\right\}} \sum_{j=2}^{3} \beta_{ji}^k D_{ji}^k + \sum_{j=2}^{F} \beta_{ji}^f D_{ji}^f + \varepsilon_i \tag{A.1}$$

$$CSD_{\$t} = \hat{\beta} \tag{A.2}$$

In (A.1), I regress the yield spread of bond i on a couple of bond characteristics such as the currency in which the bond is issued, amount issued, the remaining maturity, the age, and the rating of the bond. I estimate (A.1) at each quarter separately so there is no time subscript on variables. D_i is a currency dummy taking 1 if the bond is issued in EUR and 0 if in USD. Dummy variables for r, m, a, ai represent rating, remaining maturity, age divided by original maturity, and amount issued of the bond, respectively. When constructing these dummy variables, I put each bond into one of the three bins associated with the bond characteristic variable[42]. Then, each dummy is arranged so that it takes 1 if the bond is in the bin and 0 otherwise. Lastly, D^f_{ji} gives us the firm-fixed effect with F being the number of distinct firms at each quarter.

In (A.2), I define the credit spread differential between EUR and USD as the OLS estimate of β since it gives us the residual spread differential related to the currency of the bond after controlling for basic bond characteristics.

[42]Rating bins are: no rating, investment grade and high yield. Remaining maturity bins are: 1-5 years, 5-10 years, 10+ years. Age over initial maturity bins are: old (if the ratio is greater than 0.67), mature (if the ratio is between 0.34 and 0.67) and young (if the ratio is smaller than 0.34). Amount issued bins are: small (if the amount issued is less than $ 100 mil.), medium (if the amount issued is between $ 100 mil. and $ 500 mil.), large (if the amount issued is greater than $ 500 mil.)

B Data Appendix

B.1 Estimation of Credit Spread Differential

Estimating (A.1) requires data on bond characteristics which I obtain from Refinitiv Eikon for each bond. I match this data with the secondary market bond yield spread data obtained from Datastream using bond International Securities Identification Numbers (ISIN). Spread is calculated by subtracting the maturity-matched USD or EUR swap rate from bond i's yield. I winsorize bond spreads at 5% and 95% level to remove bonds with outlier prices.

Before estimating (A.1), I apply several filters to the raw bond dataset. First, I remove all bonds whose issuer's parent domicile is other than the EA, bonds with principal currency other than USD and EUR, bonds issued before 01.01.2001, bonds with maturity at issuance less than one year, and bonds without ISIN. Second, I apply liquidity related filters to ensure that bonds in my dataset are frequently traded so that they truly reflect pricing movements. To achieve this, I eliminate all bonds with face value less than $10 million notional and bonds with remaining maturity less than one year. Third, I apply homogeneity related filters to have a homogenous sample of bonds so that price comparison among them is meaningful. Accordingly, I exclude all floating rate coupon, convertible, asset based (covered), perpetual, callable and putable bonds from my dataset. This procedure leaves me some 61,802 bonds issued by 3,512 firms. 52,713 of these bonds are denominated in EUR while the remaining 9,089 are denominated in USD.

In addition to these filters, I also remove all bonds whose issuer does not have an outstanding bond in the other currency at the same quarter with the aim of improving the precision of the analysis. After this final filter, I merge this dataset with the bond spread data obtained from Datastream. Ultimately, 15,772 bonds out of 31,923 bonds are successfully merged, of which 12,957 is denominated in EUR and 2,815 in USD. The whole procedure leaves me with 2,825 observations on average per quarter. The summary statistics of this final dataset which is used in estimating (A.1) is given in Table B.1. One notable difference between USD-denominated and EUR-denominated bonds is that the formers mean (or median) amount is much larger than that of the latter while maturities of both types of bonds are similar.

Table B.1
Summary Statistics of Bonds in the Final Sample

Bond Summary

		All Bonds	USD-denom.	EUR-denom.
Number	Tranches	15,772	2,815	12,957
	Firms	213	213	213
Maturity (year)	Min	1	1	1
	Max	100.1	100.1	100
	Mean	6.61	6.85	6.55
	Median	5	5	5
	Sd	5.32	6.5	5.03
Amount(USD mil)	Min	10	10	10
	Max	12,218	7,000	12,218
	Mean	522	949.3	429.27
	Median	122	750	122.2
	Sd	1,048	1,071.3	1,020.22

Notes: Bonds whose issuers have no outstanding bond in the other currency and bonds for which spread data is not available in Datastream are excluded from the sample.
Source: Refinitiv Eikon, Datastream

B.2 Bonds Used in Estimating the Currency Choice Model

Again, I apply some filters to the raw bond dataset. The most important one is the exclusion of bonds issued by banks, other financial institutions, and state agencies so that I have a sample of bonds issued by EA non-financial private companies. This time, I restrict my bond sample to start from 2008 Q2 since corporate basis is very close to zero before the GFC. I also exclude bonds issued after 2019 Q4 in order to remove any external impact caused by the Covid-19 pandemic on the bond market. Next, I remove bonds whose maturity is less than one year and bonds with missing ISIN, currency, issuer, issue date or maturity information. Furthermore, I consolidate all bonds at the ultimate parent level. For instance, if a US subsidiary of a EA NFC issues a bond in the US, I consider it as the liability of the European ultimate parent company. I also remove all bonds whose ultimate parent domicile is other than EA countries and whose ultimate parent operates in financial sector or is owned by a state agency.

The summary statistics of the final sample which will be used both in this section and in the coming sections are presented in Table B.2. In the final sample, there are 5,375 bonds (4,302 EUR + 1,073 USD) issued by 1,199 distinct companies in consolidated basis. Again, a simple breakdown of the bond dataset along the currency lines shows that USD issuances are much larger in magnitude compared to EUR issuances. This time, average maturities are different too, with USD issuances having longer maturities.

TABLE B.2
Summary Statistics of Bonds Used in Estimating the Currency Choice Model

Bond Summary

		All Bonds	USD-denom.	EUR-denom.
Number	Tranches	5,375	1,073	4,302
	Firms	1,199	122	1,160
Maturity (year)	Min	1	1	1
	Max	100.1	60.54	100.1
	Mean	7.71	10.42	7.04
	Median	6.21	8	6
	Sd	6.94	9.19	6.07
Amount(USD mil)	Min	0.12	0.4	0.12
	Max	9,542.5	9,542.5	3,665.4
	Mean	442.3	840.3	343
	Median	254.8	584.8	146.7
	Sd	598.1	941.1	420.6

Source: Refinitiv Eikon.

In order to check for the representativeness of the bond-level data, I compare the corporate bond issuance data used in this paper with ECB's monthly gross corporate sector long-term debt security issuance data. I aggregate both datasets into annual frequency and depict their time series in Figure B.1. The correlation coefficient between the two series is 0.90 and Refinitiv Eikons bond data cover around 91% of ECB data on average[43]. This shows that Eikon's bond dataset sufficiently covers overall market trends.

[43] A certain portion of differences may result from the fact that ECB data is not consolidated at the ultimate parent level.

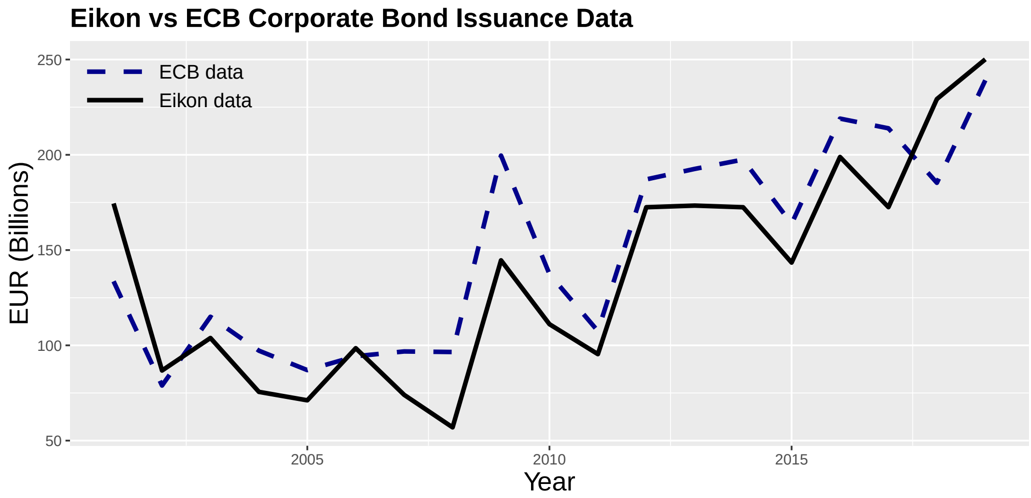

FIGURE B.1
Eikon vs ECB Corporate Bond Issuance Data. *Source:* Author's calculations, Refinitiv Eikon and ECB.

B.3 Firm-Level Characteristics

This section provides detailed information on firms' balance sheet and income statement variables used in the paper. Firm size is proxied by the logarithm of firms total assets; leverage is defined as the total debt of the firm divided by its total assets; balance sheet liquidity is taken as the sum of cash and short-term investments of the firm divided by its total assets; sales growth is given by the quarterly change in net sales; cash flow is calculated as the sum of net income before extraordinary items, depreciation and amortization and I divide it by total assets; short term debt is divided by total assets; Q is proxied by price-to-book ratio. Finally, fixed capital stock, $k_{i.t}$, is measured as the book value of a firms tangible capital stock (net property, plant and equipment).

Quarterly firm-level data is obtained from Refinitiv Eikon for the time period between 2008Q2 and 2019Q4. Figure B.2 presents the correlation matrix for firm characteristics while Table B.3 presents their summary statistics. The most notable difference between USD-issuers and firms which never issued in USD is that the former is significantly larger in size. Summary statistics of the quarterly growth rates of capital stock is reported in Table B.4. In general, the distribution of the investment rates in the EA NFCs sample is in line with other studies using Compustat data for the U.S. firms as in Bai et al. (2022), Crouzet (2021) and Ottonello and Winberry (2020).

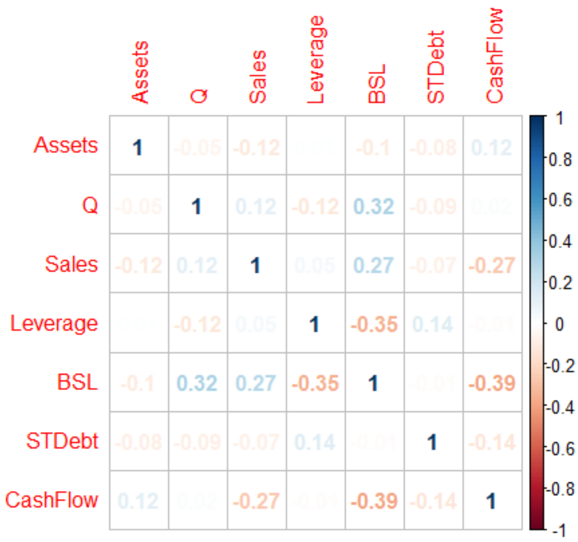

FIGURE B.2
Correlation Structure of Firm Characteristics. *Source:* Refinitiv Eikon and author's calculations.

TABLE B.3
Summary Statistics of Firm Characteristics

Sample		Assets (USD mil)	P/B Ratio	Sales Gr. (pp)	Leverage	BSL	ST Debt	Cash Flow
Whole	Mean	12,240	2.37	4.02	0.32	0.11	0.01	0.00
	Median	2,414	1.64	2.17	0.32	0.08	0.00	0.01
	Std	28,227	2.52	7.42	0.16	0.09	0.03	0.02
	5^{th} Perc.	31	0.24	-2.18	0.08	0.03	0.00	-0.03
	95^{th} Perc.	68,311	6.71	18.47	0.60	0.29	0.06	0.03
USD Issuers	Mean	35,837	2.57	2.47	0.32	0.10	0.01	0.01
	Median	14,298	1.93	1.45	0.30	0.10	0.00	0.01
	Std	47,078	2.68	3.62	0.15	0.07	0.01	0.01
	5^{th} Perc.	847	0.44	-0.49	0.13	0.03	0.00	-0.01
	95^{th} Perc.	132,262	7.06	6.91	0.54	0.26	0.03	0.02
EUR-only Issuers	Mean	6,108	2.31	4.41	0.33	0.11	0.01	0.00
	Median	1,567	1.57	2.25	0.32	0.08	0.00	0.00
	Std	15,880	2.48	8.04	0.16	0.10	0.04	0.03
	5^{th} Perc.	25	0.24	-2.79	0.07	0.03	0.00	-0.03
	95^{th} Perc.	29,449	6.70	21.29	0.60	0.33	0.08	0.03

Source: Refinitiv Eikon

Table B.4
Summary Statistics of Firms' Fixed Capital Expenditure

Summary Statistics of Firms' Fixed Capital Expenditure

$\Delta_h log(k_{i.t+h})$	Mean	Median	Std. Dev.	5^{th} Perc.	95^{th} Perc.
$h = 0$	0.013	0.002	0.113	-0.085	0.139
$h = 1$	0.026	0.009	0.188	-0.157	0.280
$h = 2$	0.039	0.015	0.248	-0.219	0.422
$h = 3$	0.053	0.022	0.301	-0.276	0.518
$h = 4$	0.062	0.029	0.348	-0.329	0.600
$h = 5$	0.070	0.037	0.390	-0.387	0.675
$h = 6$	0.080	0.043	0.430	-0.438	0.748
$h = 7$	0.090	0.050	0.467	-0.484	0.830
$h = 8$	0.097	0.057	0.510	-0.538	0.898
$h = 9$	0.106	0.068	0.543	-0.597	0.961
$h = 10$	0.113	0.076	0.579	-0.631	1.017
$h = 11$	0.122	0.086	0.606	-0.668	1.062
$h = 12$	0.131	0.096	0.638	-0.709	1.127
$h = 13$	0.141	0.105	0.656	-0.727	1.164
$h = 14$	0.151	0.114	0.677	-0.749	1.228
$h = 15$	0.160	0.122	0.700	-0.770	1.260
$h = 16$	0.167	0.127	0.721	-0.794	1.291

Source: Refinitiv Eikon

B.4 Monetary Policy Variables

B.4.1 Monetary Policy Surprises

This section presents summary statistics and figures for the monetary policy surprise series used in the paper. Between 2008Q2 and 2019Q4, 122 monetary policy announcement events occurred in total. The detailed information on how monetary policy surprise series is obtained and how I calculate OISPRCT is explained in Section 5.1.

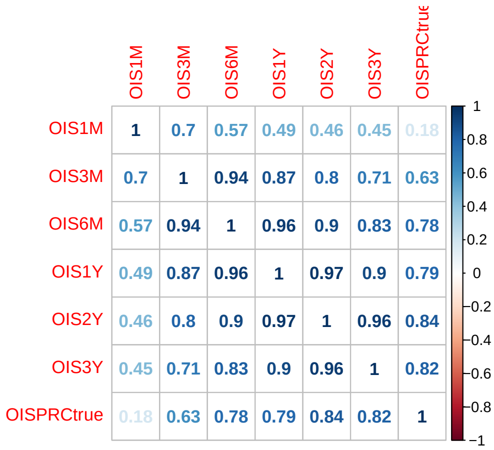

FIGURE B.3
Correlation Structure of OIS Rate Surprises of Different Maturities. *Source:* Author's calculations based on Euro Area Monetary Policy Event Study Database.

TABLE B.5
Summary Statistics of Monetary Policy Surprises

MP Summary

OIS Maturity	Min	Max	Mean	Median	Std. Dev.	Min. Date	Max. Date
OIS 1M	-6.60	8.24	0.04	0.00	2.87	2012 Q3	2019 Q3
OIS 3M	-9.65	10.25	0.07	0.00	4.06	2011 Q3	2008 Q2
OIS 6M	-14.00	15.00	0.09	0.21	5.72	2011 Q3	2008 Q2
OIS 1Y	-25.75	20.30	-0.09	-0.05	7.79	2008 Q3	2008 Q2
OIS 2Y	-37.50	20.38	-0.42	-0.25	9.23	2008 Q3	2011 Q1
OIS 3Y	-34.70	18.40	-0.84	-0.60	8.12	2008 Q3	2011 Q1
OISPRCT	-37.82	37.83	-0.27	-1.90	12.69	2011 Q2	2008 Q4

Notes: Monetary policy surprises are aggregated into quarterly frequency by summing monetary policy surprises that happen at the same quarter.
Source: Author's calculations on Euro Area Monetary Policy Event Study Database

B.4.2 Monetary Policy Variables Used for Robustness Checks

This appendix presents summary statistics and figures for the monetary policy variables used in robustness checks. I explain how I constructed each variable in Sections 6.1.1-6.1.4 in the main text.

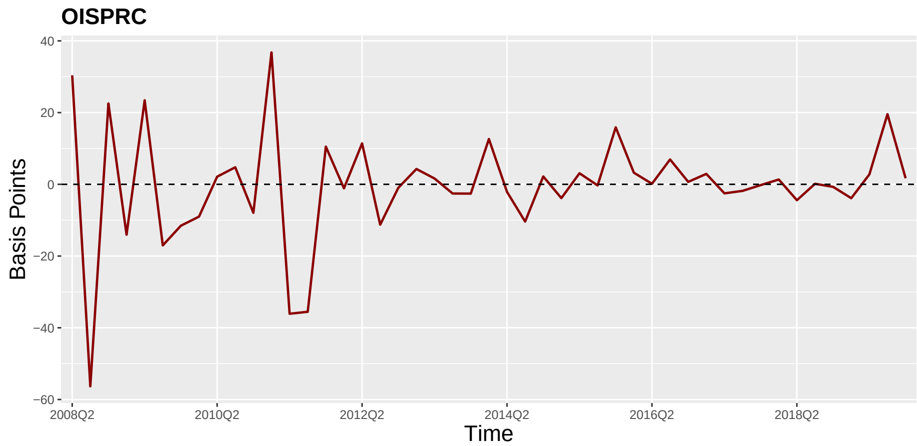

Figure B.4
OISPRC. *Source:* Author's calculations based on Euro Area Monetary Policy Event Study Database.

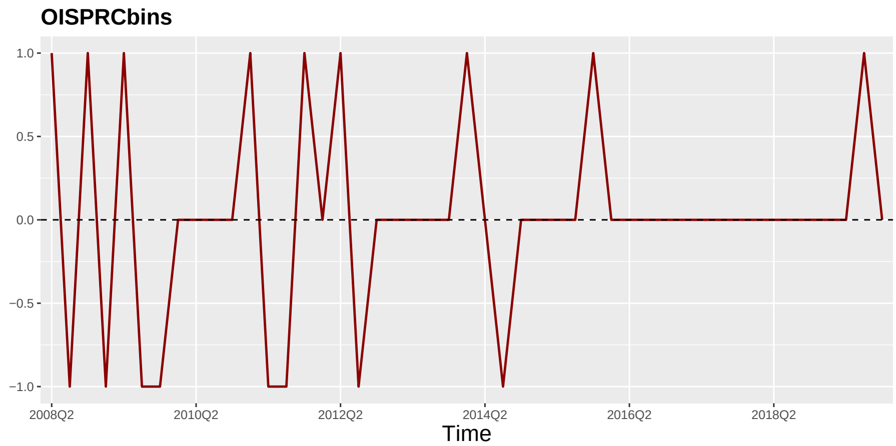

Figure B.5
OISPRCbins. *Source:* Author's calculations based on Euro Area Monetary Policy Event Study Database.

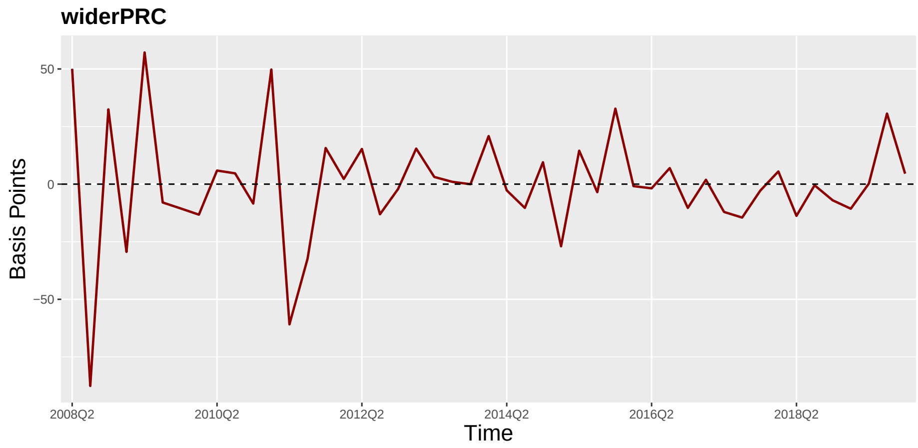

Figure B.6
widerPRC. *Source:* Author's calculations based on Euro Area Monetary Policy Event Study Database

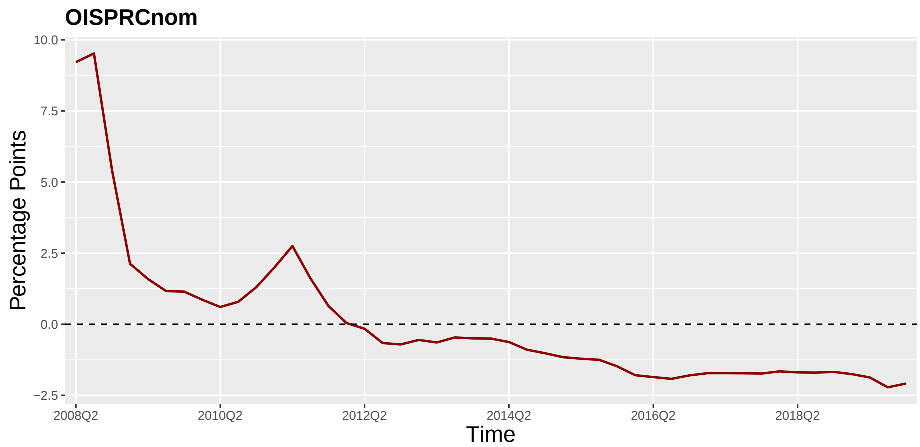

Figure B.7
OISPRCnom. *Source:* Author's calculations and Refinitive Eikon

Table B.6
Summary Statistics of Monetary Policy Variables

MP Summary

OIS	Min	Max	Mean	Median	Std. Dev.	Min. Date	Max. Date
OISPRC	-56.32	36.77	-0.32	0.14	15.69	2008 Q3	2011 Q1
OISPRCbins	0	1	0.02	0	0.61	-	-
widerPRC	-87.71	57.17	-0.07	-0.49	24.90	2008 Q3	2009 Q2
OISPRCnom	-2.22	9.52	0.00	-0.66	2.54	2019 Q3	2008 Q3

Source: Author's calculations on Euro Area Monetary Policy Event Study Database

B.5 Stock Returns

Table B.7 presents summary statistics for stock returns that I use in Section 6.5 using a two-day window around each ECB monetary policy announcement event between 2008 and 2019.

Table B.7
Summary Statistics of Stock Returns

	All Firms	**USD-Issuers**	**EUR-only Issuers**
Mean	-0.01	0.05	-0.03
Median	0	0.13	0
SD	3.71	3.45	3.78
5^{th} Perc.	-6.25	-5.86	-6.39
95^{th} Perc.	5.74	5.26	5.88
Obs.	$55,493$	$12,271$	$43,222$

C Corporate Basis and Monetary Policy

In this part, I study the effect of monetary policy on corporate basis. Theoretically, the impact of monetary policy on corporate basis is ambiguous. Consider equation (2). On one hand, a static interpretation reads an increase in domestic risk-free rate driven by monetary tightening (rf_t) as pulling the credit spread differential (first term on the right hand side) down. However, monetary tightening typically influences risky rates (rb_t) as well leading to higher credit spreads when financial conditions are tight. Similarly, prolonged interest rate reductions can squeeze credit spreads through higher risk appetite and search for yield efforts. Thus, changes in monetary policy can positively affect the first term of the right hand side of equation (2).

In a similar vein, a mechanical reading would suggest that a ECB controlled interest rate decline decreases the second term of the right hand side of equation (2) (CIP deviation). However, Du et al. (2018) show that monetary policy differential affects CIP deviation (measured as in equation (2)) negatively. As ECB-Fed differential decreases, higher demand for USD-denominated assets raise the cost of currency hedging in forward and swap markets leading to an increase in the second term of the right hand side of (2).

Due to counteracting forces at work, the direction of the impact of monetary policy on corporate basis needs to be empirically investigated. In Table C.1, I consider six specifications. As currency induced borrowing cost differential is affected by both local and foreign monetary policy, I calculate the difference between ECB controlled rate and Fed controlled rate. I then calculate the four quarter moving average of this differential and regress corporate basis on ECB-Fed differential. In the first three columns, I use monetary policy surprise series. I use OISPRCT for the ECB rate (see Section 5.1). For the Fed rate, I use monetary policy surprise series produced by Bu et al. (2021). The two series are compatible in that they both address the information effect problem.

In the last three columns, I use interest rate levels instead of surprises. I use OISPRCnom for the ECB rate (see Section 6.1.4). For the Fed rate, I obtain treasury yields of 1-month, 3-month, 6-month, 1-year, 2-year and 3-year maturities from St. Louis Fed's website and compute their first principal component. Sample period is from 2008 Q2 through 2019 Q4. Monetary policy differential seems to be a significant driver of corporate basis across all specifications albeit with differences in significance levels. Overall, results suggest that as monetary policy differential increases between ECB and Fed -indicative of a relative tightening of ECB's monetary policy-, issuing in USD becomes more favorable for EA NFCs in terms of FX-hedged borrowing costs.

TABLE C.1
Corporate Basis and Monetary Policy (Estimation Results)

	1	2	3	4	5	6
Dep. Variable	CB_t	ΔCB_t	ΔCB_t	CB_t	ΔCB_t	ΔCB_t
Intercept	-2.77^{*}	2.01	2.07	-2.97^{**}	2.13	2.21
	(1.50)	(1.75)	(1.77)	(1.33)	(1.74)	(1.76)
CB_{t-1}	0.62^{***}			0.59^{***}		
	(0.06)			(0.05)		
ΔCB_{t-1}			-0.06			-0.07
			(0.13)			(0.13)
ECB-FED (surp.)	3.19^{*}	5.81^{**}	5.92^{**}			
	(1.82)	(2.42)	(2.45)			
ECB-FED (nom.)				1.53^{***}	1.55^{**}	1.59^{**}
				(0.41)	(0.63)	(0.64)
Multiplier	8.35		5.57	3.76		1.58
R^2	0.70	0.12	0.13	0.76	0.13	0.13
Note:	*p<0.1; **p<0.05; ***p<0.01					

D Additional Results

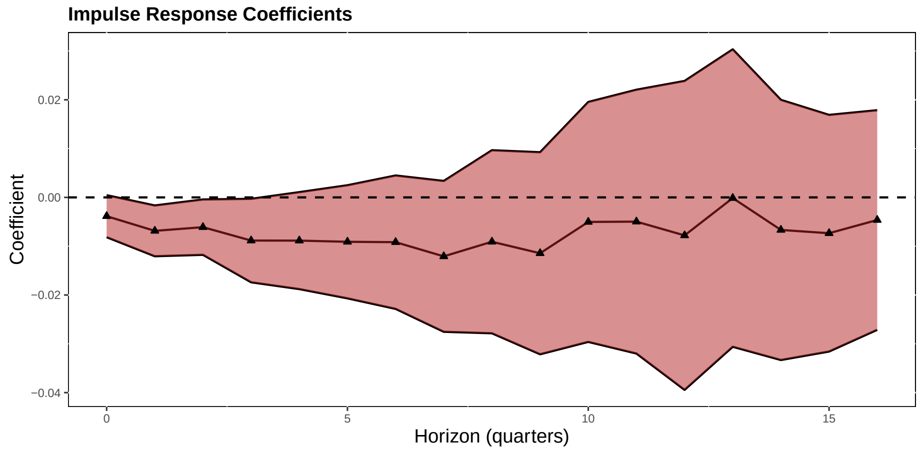

FIGURE D.1
The Average Effect of Monetary Policy Surprises on Firms Fixed Capital Expenditure

Notes: The figure depicts impulse response coefficients, γ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \xi_q^h + \gamma^h \eta_t + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \varepsilon_{i,t+h}$ where monetary policy variable, η_t, is defined within the text. The coefficient γ^h is scaled so that it represents the change in fixed capital expenditure following a one standard deviation increase in η_t. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

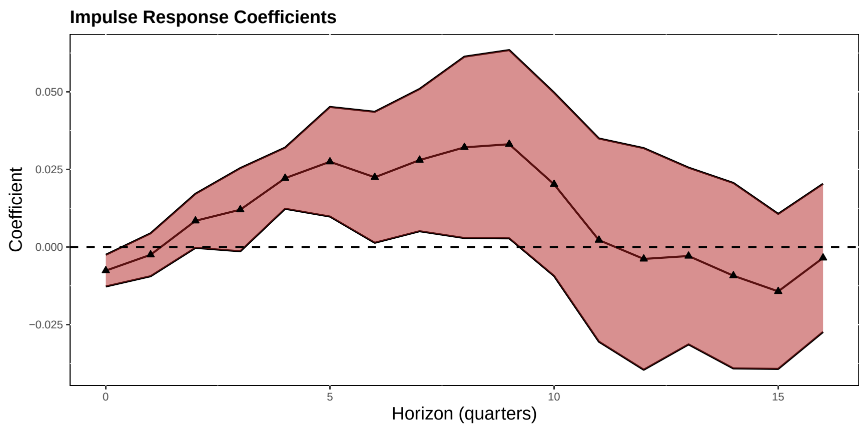

FIGURE D.2
The Differential Impact of Monetary Policy Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary policy variable, η_t, is defined within the text and $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

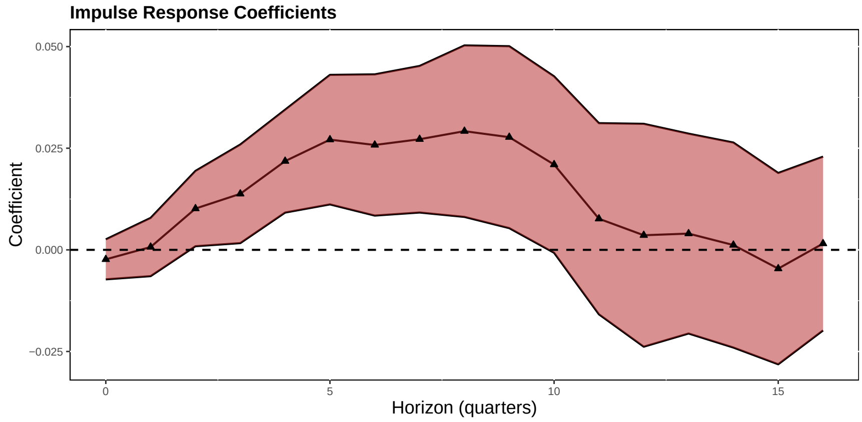

Figure D.3
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with OISPRC)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary policy variable, η_t, represents OISPRC as defined in Section 6.1.1 and $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

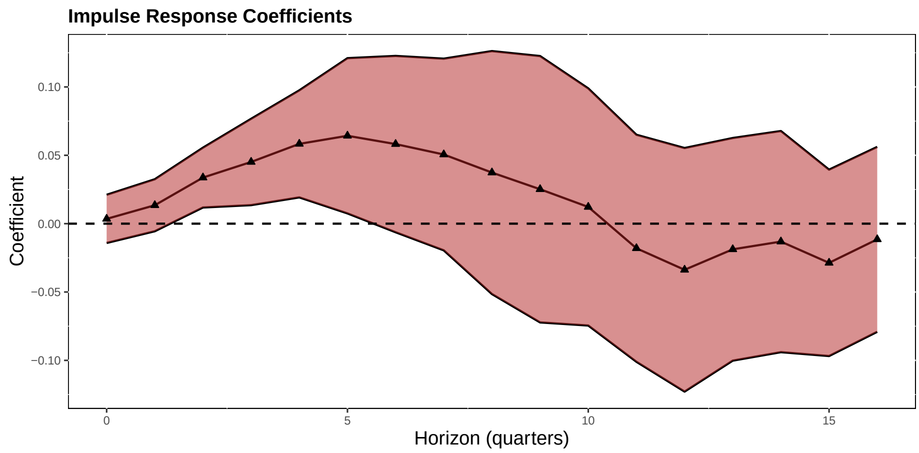

Figure D.4
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with OISPRCbins)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary policy variable, η_t, represents OISPRCbins as defined in Section 6.1.2 and $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

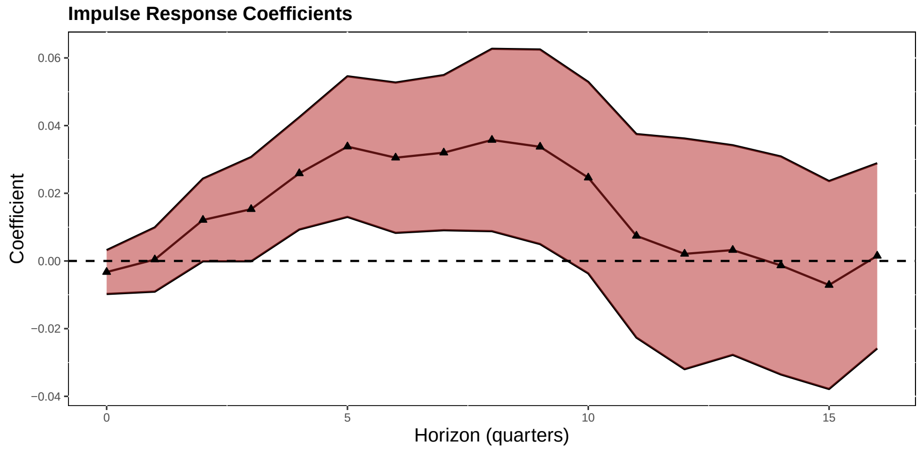

FIGURE D.5
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with widerPRC)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary policy variable, η_t, represents widerPRC as defined in Section 6.1.3 and $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

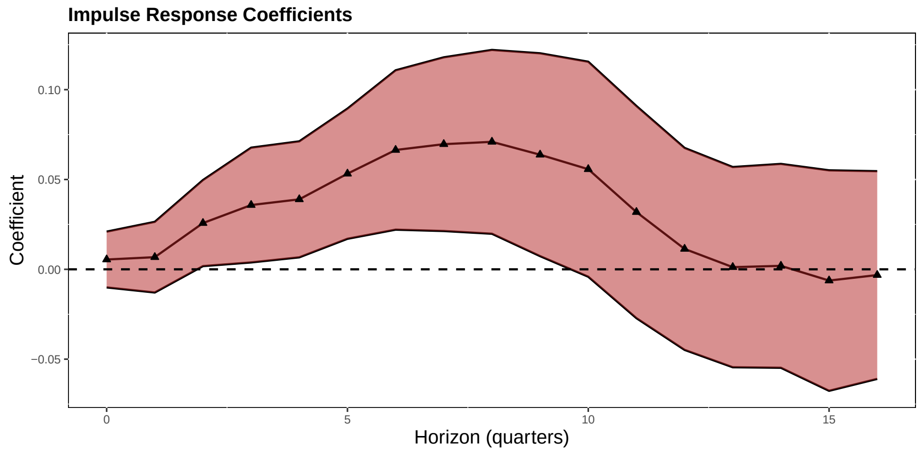

FIGURE D.6
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with OISPRCnom)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary policy variable, η_t, represents OISPRCnom as defined in Section 6.1.4 and $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

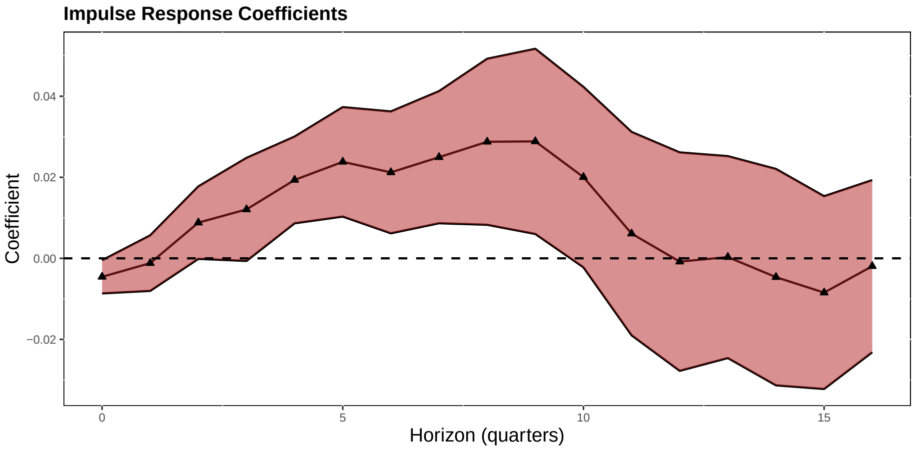

FIGURE D.7
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with Bond Ratings)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text and $OB_{i,t}^1$ is as described in equation 12. W includes bond rating dummies along with all the firm characteristics as elaborated in Section 5.2. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

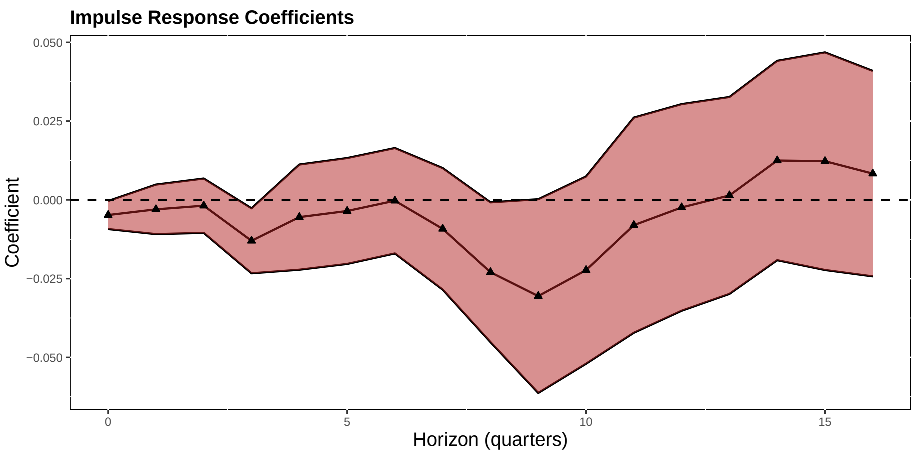

FIGURE D.8
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with OB_{it}^2)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^2 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text and $OB_{i,t}^2$ is as described in equation 14. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

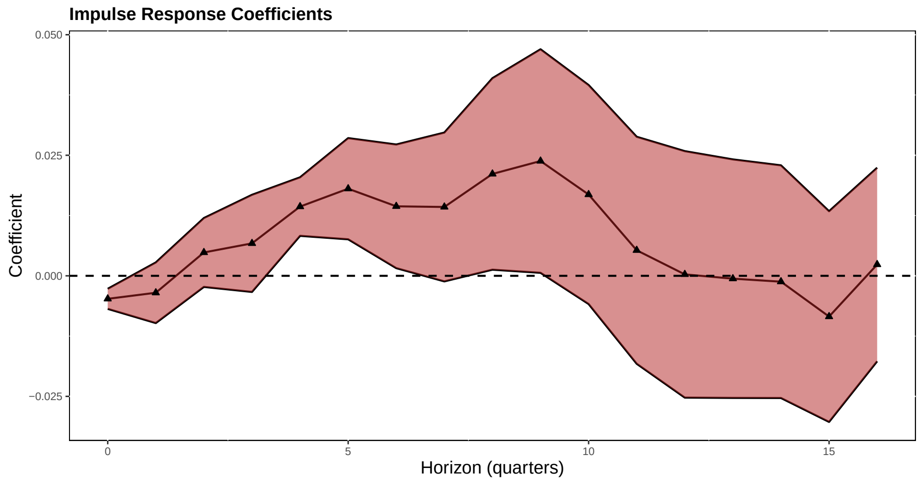

FIGURE D.9

The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (controlling for equity constraints)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \alpha^h I_{i,t} + \beta^h I_{i,t} \eta_t^+ + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, and the indicator variable representing the least constrained firms according to the equity-constrained index, $I_{i,t}$ is defined within the text. $OB_{i,t}^1$ is as described in equation 12. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

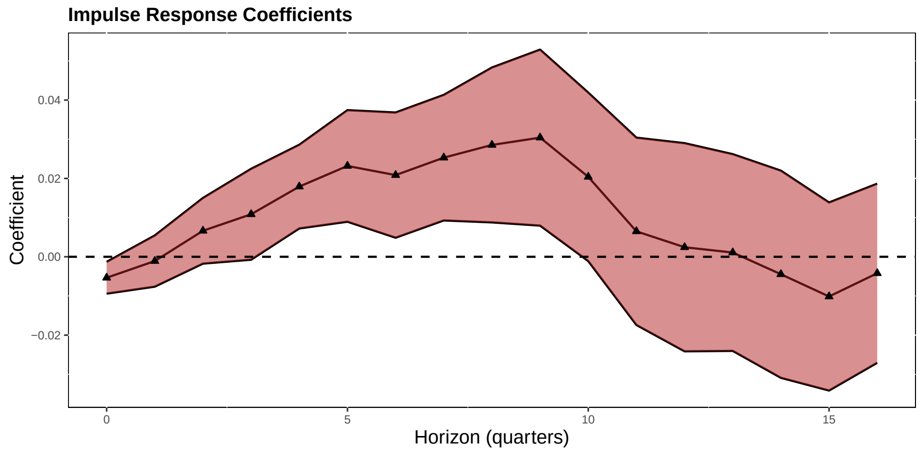

FIGURE D.10

The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with Tobin's Q)

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text and $OB_{i,t}^1$ is as described in equation 12. W includes Tobin's q proxied by price-to-book ratio along with all the firm characteristics as elaborated in Section 5.2. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

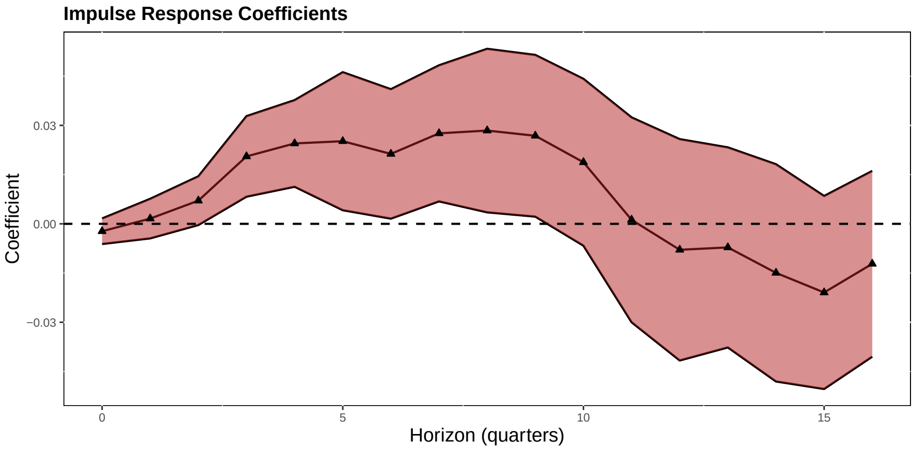

FIGURE D.11
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Fixed Capital Expenditure (with OB^3_{it})

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(k_{i,t+h}) - log(k_{i,t-1}) = f^h_i + \lambda^h_{s,t} + \psi^h_{c,t} + \theta^h OB^3_{i,t}\eta^+_t + \sum_{w\in W} \alpha^h_w w_{i,t-1} + \sum_{w\in W} \beta^h_w w_{i,t-1}\eta_t + \varepsilon_{i,t+h}$ where monetary tightening variable, η^+_t, is defined within the text and $OB^3_{i,t}$ is as described in equation 15. The coefficient θ^h is scaled so that it represents the differential change in fixed capital expenditure following a one standard deviation increase in η^+_t. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

E Heterogeneous Inventory Investment Responses to Monetary Tightening

In this section, I repeat the baseline exercise done in Section 5.3, this time for inventory investment response. I estimate equations (13) and (11) with inventories replacing the capital stock $k_{i,t}$.

Estimated coefficients are reported in Figure E.1 and Figure E.2. The average effect of tightening is akin to fixed capital investment case in terms of magnitude implying that monetary tightening dampens inventory investment in the firm sample. The difference is that the coefficient becomes statistically significant only seven quarters after surprise monetary tightening. The heterogeneous effect is also at a similar level to the baseline in terms of magnitude. However, it is not as strong as what I found for fixed capital investment in terms of statistical significance due possibly to more missing values for inventories in my sample compared to PPE, reducing the sample size for inventory analysis. The interaction coefficient is both positive and statistically significant 5-7 quarters after surprise tightening.

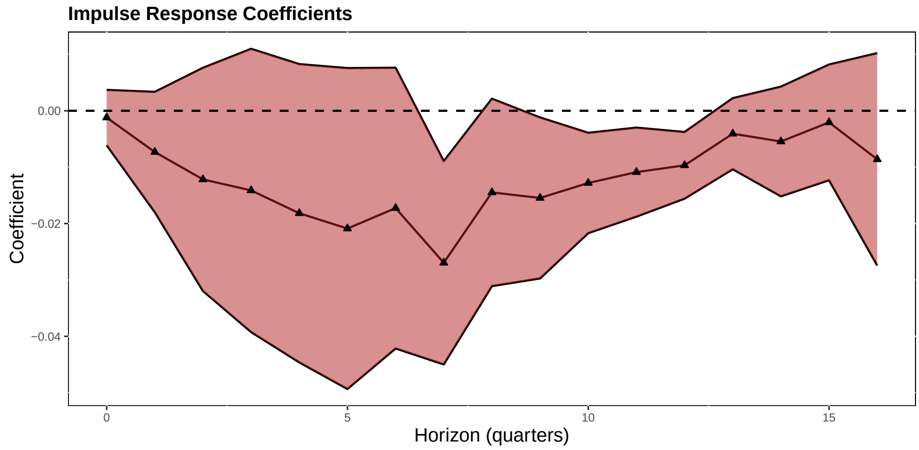

FIGURE E.1
The Average Effect of Monetary Tightening Surprises on Firms Inventory Investment

Notes: The figure depicts impulse response coefficients, γ^h estimated at each forecast horizon h, from the following regression: $log(inv_{i,t+h}) - log(inv_{i,t-1}) = f_i^h + \xi_q^h + \gamma^h \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text. The coefficient γ^h is scaled so that it represents the change in inventory investment following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.

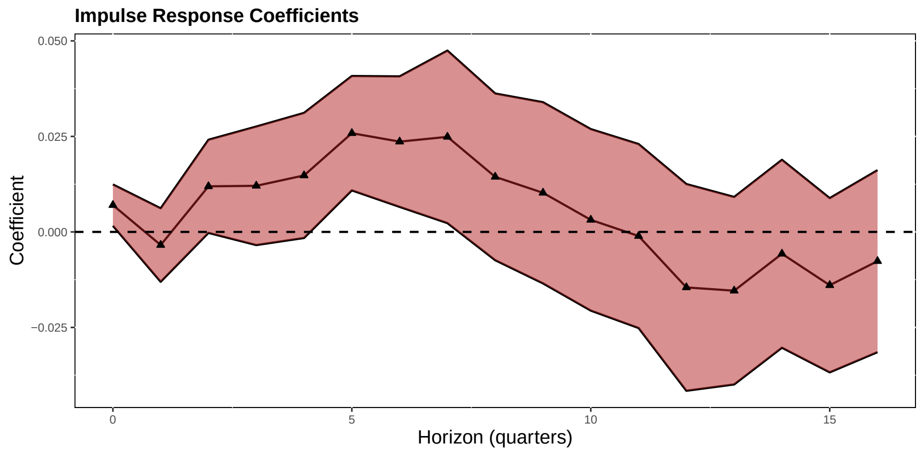

FIGURE E.2
The Differential Impact of Monetary Tightening Surprises on Opportunistically Borrowing Firms Inventory Investment

Notes: The figure depicts impulse response coefficients, θ^h estimated at each forecast horizon h, from the following regression: $log(inv_{i,t+h}) - log(inv_{i,t-1}) = f_i^h + \lambda_{s,t}^h + \psi_{c,t}^h + \theta^h OB_{i,t}^1 \eta_t^+ + \sum_{w \in W} \alpha_w^h w_{i,t-1} + \sum_{w \in W} \beta_w^h w_{i,t-1} \eta_t + \varepsilon_{i,t+h}$ where monetary tightening variable, η_t^+, is defined within the text and $OB_{i,t}^1$ is as described in equation 13. The coefficient θ^h is scaled so that it represents the differential change in inventory investment following a one standard deviation increase in η_t^+. The area between the two dashed lines represents the confidence interval at 90% level. Standard errors are double clustered at the firm and time (quarter-year) level.