Decomposing the Investment Channel of Monetary Policy*

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Abstract

Monetary policy announcements simultaneously change borrowing costs and firms' expectations about future economic conditions. Conventional estimates of the investment response to monetary policy therefore conflate price effects with belief revisions and fail to identify the structural sensitivity of firm investment to financing conditions. We exploit quasi-experimental variation from the ECB's 2016 Corporate Sector Purchase Programme (CSPP), which created persistent cross-sectional differences in firms' financing costs through bond market segmentation, to isolate the causal effect of borrowing costs on firm investment. We find that investment elasticities estimated from aggregate monetary policy shocks understate the effect of a pure cost-of-debt shock by 38 percent over four quarters. Our results imply that the structural sensitivity of firm investment to borrowing costs is larger than existing estimates indicate. The gap reflects a sizable firm-level information effect. A simple model of firm investment with an information effect rationalizes this attenuation. Our findings carry direct implications for heterogeneous real transmission of monetary policy and the calibration of quantitative macroeconomic models.

Keywords: Investment, Information Effect, Monetary Policy, Monetary Non-Neutrality

JEL-Codes: E52, E44, G12, G32

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1 Introduction

The extent of monetary non-neutrality is a classical question in macroeconomics (Friedman, 1968). Much of the resulting work has focused on households' intertemporal substitution of consumption, giving rise to the canonical New Keynesian framework (Clarida et al., 1999). Far less is settled about how monetary policy affects firm investment, even though investment accounts for a large share of aggregate demand and is one of the most interest-sensitive components of GDP. Empirically, firms' investment sensitivity to interest rates is typically measured using investment responses to high-frequency monetary policy shocks (see e.g. Ottonello and Winberry, 2020).[1] Yet such shocks affect firms' borrowing costs while simultaneously altering beliefs about future economic fundamentals – an effect commonly referred to as the *information effect* of monetary policy surprises (Romer and Romer, 2000; Nakamura and Steinsson, 2018). As a result, observed aggregate investment responses conflate price effects with belief revisions and fail to isolate firms' underlying sensitivity to financing conditions.

Disentangling the borrowing-cost channel from information effects using conventional high-frequency monetary policy shock series in practice requires additional – and potentially restrictive – identifying assumptions, such as sign restrictions or orthagonalization (Jarociński and Karadi, 2020; Miranda-Agrippino and Ricco, 2021; Bauer and Swanson, 2023). In this paper, we propose a novel quasi-experimental estimation setup which exploits cross-sectional variation generated by the ECB's Corporate Sector Purchase Programme (CSPP) to isolate the borrowing-cost channel. We find that the CSPP announcement created a yield wedge between eligible and ineligible bonds, while firms differed in their exposure to this wedge due to persistent pre-CSPP financing structures. The resulting cross-sectional variation in firms' financing costs is plausibly exogenous, orthogonal to belief revisions at the time of the announcement, and unavailable in conventional monetary policy shocks.

This paper provides evidence that standard monetary policy shocks identify a composite

[1] High-frequency monetary policy shocks capture the surprise component of how market beliefs about future policy rates change in a tight event window around a monetary policy communication (see e.g. Kuttner, 2001; Gürkaynak et al., 2005).

object that substantially understates firms' true sensitivity to financing costs. We find that aggregate investment elasticities over four quarters are 38 percent smaller than the response implied by a pure cost-of-debt shock. This dampening reflects belief revisions connected to monetary policy announcements and contributes new evidence to the ongoing debate on the quantitative importance of information effects in monetary policy (Cochrane and Piazzesi, 2002; Faust et al., 2004; Bauer and Swanson, 2023).

Beyond the debate on information effects, the distinction matters because macroeconomic models rely on investment elasticities as behavioral primitives. By isolating the borrowing-cost channel, our results suggest that calibrations based on conventional high-frequency estimates may systematically understate the true responsiveness of firm investment to changes in borrowing costs. We show that in standard investment models, the calibrated adjustment-cost parameter is approximately inversely related to the targeted investment elasticity, implying that a downward-biased elasticity leads to an upward-biased estimate of capital adjustment frictions, which play a central role in quantitative models of firm investment. The findings also matter for understanding the heterogeneous transmission of monetary policy shocks with varying information contents and highlight the importance of monetary policy communication for transmission.

To guide intuition, we introduce a simple model of firm investment with an information channel. Monetary policy surprises mechanically affect the user cost of capital but simultaneous convey signals about other economic fundamentals, with implications for expected long-run prospects and project profitability. Following a monetary policy surprise, firms adjust their investment decision for projects close to the investment threshold. When a sufficiently large mass of projects lies near the threshold, the model can replicate attenuation of the borrowing-cost channel, consistent with our empirical estimates. The framework also clarifies how conventional monetary policy shocks and the CSPP-based design load on different sources of variation.

Related Literature. Our paper relates to two strands of literature. First, it relates to the literature on the role of central bank information. Following the seminal contribution by Romer and Romer (2000), a growing body of empirical work like Campbell et al. (2012) and Nakamura and Steinsson (2018) argues that policy announcements convey information about macro fundamentals beyond their mechanical effect on interest rates.[2] They find that monetary policy tightening is associated with a significant downward revision of blue-chip forecasts of unemployment and upward revisions of real-GDP respectively, which goes contrary to what standard theory would suggest. They argue that this is due to "delphic forward guidance" where FOMC statements convey information to the private sector about the future of the economy. As a consequence, monetary policy shocks may conflate rate changes with belief revisions and fail to identify the structural real transmission of monetary policy. The idea that monetary policy surprises carry information, however, is not uncontested (Cochrane and Piazzesi, 2002; Faust et al., 2004; Bauer and Swanson, 2023). By exploiting cross-sectional variation in CSPP-exposure, this paper adds to this debate by providing novel firm level evidence in favor of an information effect.

Empirically, there also exist complementary approaches to ours that aim to correct for the information effect in series of high frequency monetary policy shocks (Jarociński and Karadi, 2020; Bauer and Swanson, 2023; Miranda-Agrippino and Ricco, 2021). Jarociński and Karadi (2020) for example decompose monetary policy surprises into pure "monetary policy" shocks and "central bank information" shocks depending on whether equity prices move in the opposite or the same direction as interest rates, respectively. One caveat of such sign restriction based approaches (also see Cieslak and Schrimpf, 2019) is that the decomposition is set identified, not point identified: there are infinitely many valid decompositions of monetary policy shocks into the two structural components. Jarociński and Karadi (2020) address this by using a uniform prior over all permissible decompositions. While different priors deliver compar-

[2]Besides the empirical contributions, a large theoretical literature discussing central bank signaling through its actions (e.g. changes in the federal funds rate) or communication (FOMC statements). Early work includes Cukierman and Meltzer (1986) and Ellingsen and Söderström (2001). More recent contributions include Berkelmans (2011), Melosi (2016), and Andrade et al. (2019).

able results for some aggregate outcomes, the choice of prior matters for firm investment. When studying the robustness of investment impulse response functions to other priors on rotations following Giacomini and Kitagawa (2021), we show in appendix section A that the choice of prior matters a lot for how quantitatively meaningful central bank information is for firm investment. Miranda-Agrippino and Ricco (2021) and Bauer and Swanson (2023) similarly provide corrections of high-frequency monetary policy shock series by controlling for the Fed's internal Greenbook forecasts, and publicly available macroeconomic and financial news, respectively. Their approach, however, has to explicitly assume the source of central bank information and identify an accurate empirical proxy thereof. In appendix section A, we compare our findings to results implied by all three aforementioned off-the-shelf shock series. Our decomposition is in line with predictions of these empirical corrections for the information effect. However, by exploiting quasi-experimental firm-level variation in borrowing costs, our investment-specific decomposition delivers tighter confidence bands relative to the statistically insignificant results from high frequency approaches and further sidesteps additional - and potentially restrictive assumptions on the prior choice or the empirical proxy for the information effect.

The second strand of the literature this paper connects to is on monetary transmission through financing conditions. At least since the introduction of the *financial accelerator* mechanism by Bernanke et al. (1999), models of monetary transmission have emphasized how borrowing costs and financial constraints shape firms' investment decisions (Ottonello and Winberry, 2020). Specifically, our findings speak to the quantitative investment literature in which firm-level nonconvexities, irreversibility, and heterogeneity shape aggregate dynamics. Early work emphasizes that irreversible or lumpy capital adjustment can generate inaction regions and extensive-margin investment responses at the micro level (Bertola and Caballero, 1994; Caballero and Engel, 1999; Thomas, 2002). A central question in this literature is whether these microeconomic nonlinearities survive aggregation. Thomas (2002) and Khan and Thomas (2008) show that general-equilibrium price movements can substantially neutralize the aggregate relevance of plant-level lumpiness. In contrast, Bachmann et al. (2013), and Winberry (2021) show that once models are disciplined by empirically plausible investment price elasticities and interest-rate dynamics, the distribution of firms across adjustment regions becomes quantitatively important for aggregate investment and for state-dependent policy transmission. Our contribution is complementary: rather than asking whether the distribution of firms changes the effect of a given financing shock, we show that the financing shock itself is mismeasured when conventional monetary policy surprises bundle borrowing-cost movements with information-induced revisions in expected returns. We show that possible mismeasurement in such investment responsiveness can lead to overstating the role of capital adjustment frictions that underpin many aggregate dynamics in the literature.

Outline. The rest of the paper is organized as follows. Section 2 presents a simple framework that illustrates how monetary policy surprises generate an information effect in firms' investment decisions and clarifies why conventional monetary policy shocks identify a composite object, conflating borrowing-cost and belief channels. Section 3 lays out the institutional background for the CSPP that enables our empirical decomposition, as well as the different data sources and their descriptive statistics. Section 4 provides illustrative evidence that the CSPP has driven a wedge between the bond prices of eligible- and ineligible bonds to which firms are persistently exposed. Section 5 lays out the empirical decomposition approach by using firms' CSPP pre-exposure as an instrument for their change in borrowing cost to estimate investment responses from a pure cost-of-debt shock as well as from aggregate monetary policy shocks. Section 6 discusses the empirical results and their robustness. Section 7 discusses the implications of our findings for macro models that use investment elasticities as behavioral primitives. Section 8 concludes.

2 A Simple Model of Firm Investment

To guide intuition, we develop a two-period model that illustrates how information effects shape firms' investment responses to monetary policy announcements. There is a set of firms, each with access to a continuum of investment projects. Each firm weighs the expected return of every available project against its borrowing rate and invests in the subset whose returns exceed it. When the central bank announces a surprise change in the policy rate, this affects borrowing costs but simultaneously transmits a signal about other economic fundamentals which can make projects more or less profitable in the future. We show that the investment response of each firm can be decomposed into a cost-of-debt channel – the response only due to the changed user cost of capital – and an offsetting information effect related to firms updating their beliefs about return prospects. Aggregating these individual decisions across a large number of firms yields the economy-wide investment response.

2.1 Environment

There are two periods, $t \in \{1, 2\}$. One can think of the periods as the short- and the long-run. The long-run state of the economy is stochastic and described by θ. In the long-run, economic fundamentals can be either good ($\theta = G$) or bad ($\theta = B$), but there is uncertainty about the realization of the state when firms make investment decisions in period one. In the following, we describe a single firm's investment decision over a continuum of investment opportunities, which we call projects henceforth.

Each firm is endowed with a continuum of projects. Each project k has a state-depend return, $R_k(\theta)$. The user cost of capital for firms is $i + r$, where i is the common risk-free rate and r is a firm-specific idiosyncratic component. The return of an investment into project i is a function of the state of the economy θ:

$$\pi_k(\theta) = R_k(\theta) - (i + r) \tag{1}$$

The heterogeneity of projects is reflected in the return component. The user cost of capital is firm-specific and, thus, the same across projects within firms.

The policy rate i is governed by a central bank. At the start of period one, the central bank receives a noisy private signal $s \in \{H, L\}$ about the likelihood of the good state in the long run:[3]

$$p_s \ := \ P(\,\theta = G \mid s\,), \tag{2}$$

where $p_H > p_L$. After receiving the signal, the central bank announces a monetary policy surprise: a change in the economy-wide interest rate from the expected i_e to the realized i. The direction of the interest surprise is a perfect signal about the central bank's previously private information: a surprise-cut signals $s = L$ associated with belief-updating to p_L, while a surprise-hike signals $s = H$ associated with p_H. Before the monetary policy surprise, firms hold ex-ante beliefs that $P(s = H) = \zeta$, so firms' unconditional expectation before observing the monetary policy shock can be written as:

$$p := \zeta p_H + (1 - \zeta) p_L$$

After observing the monetary policy surprise, firms update their beliefs about the future of the economy to p_H or p_L, depending on the direction of the surprise before choosing whether or not to invest in projects.

2.2 The Investment Decision

Firms in period one make a decision which projects to invest in. After observing the central bank signal, they choose to invest considering their borrowing costs and updated beliefs about project returns. Conditional on some belief $\tilde{p} \in \{p_H, p_L\}$, a firm will invest in a project if and only if the expected return on investment is positive. The expected gross return of

[3] It is well documented that the private sector updates its expectations around monetary policy surprises (Nakamura and Steinsson, 2018; Jarociński and Karadi, 2020; Andrade and Ferroni, 2021). There are two complementary ways to rationalize this with private information held by the central bank. First, the *delphic* view proclaims that the central bank has an information advantage over the private sector through superior forecasts about inflation and growth (Romer and Romer, 2000). Another *odyssian* explanation is that the central bank has private information about its own policy path, which determines long-run growth. For example, the central bank plausibly has an information advantage on whether it places a large value on inflation rate-stabilization over closing the output gap in the short-run, which has been shown to have a big effect on long-run growth (see e.g. Rogoff, 1985; Kydland and Prescott, 1977; Barro and Gordon, 1983; Walsh, 1995; Fischer, 1993). In this paper, we remain agnostic about the precise weight of the convex combination of the two channels. What we assume is that the central bank reveals some private information previously unbeknownst to firms via monetary policy surprises that affect expectations for long-run growth.

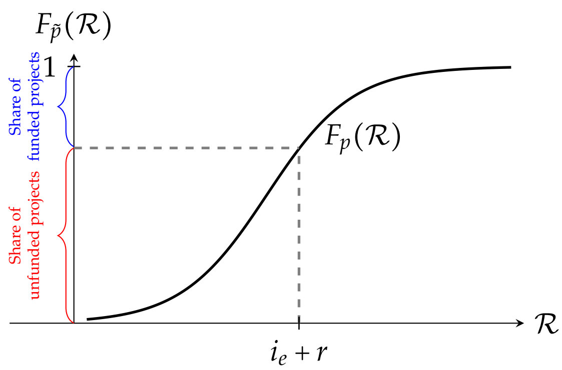
investing in project k under beliefs $\tilde{p}$ equals:

$$\mathcal{R}_k(\tilde{p}) \;:=\; \tilde{p}R_k(G) + (1-\tilde{p})R_k(B) \tag{3}$$

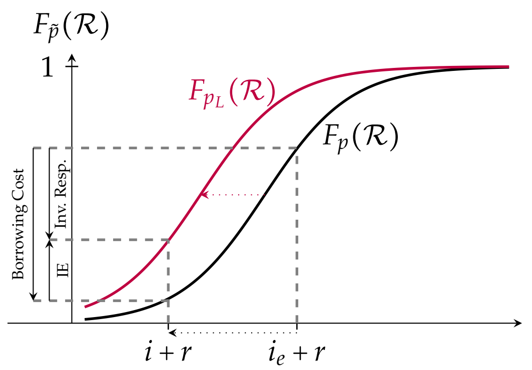
Given beliefs $\tilde{p}$, we denote the CDF according to which projects' expected gross return are distributed by $F_{\tilde{p}}(\mathcal{R})$. Ultimately, firms invest in those projects where the gross return exceeds the user cost of capital $(i+r)$. Figure 1a visualizes the static trade-off between the cost of capital and expected return absent a monetary policy surprise. Only projects whose expected gross return exceeds the user-cost-of-capital get funded. Thus, under belief p, the share of funded projects equals $1 - F_p(i_e + r)$.

Figure 1b visualizes the effect of a monetary policy surprise on the share of projects that get funded. When the monetary policy shock contains a signal on economic fundamentals that updates beliefs about the long-run state of the economy, the distribution of expected returns shifts simultaneously with the user-cost-of-capital. Take the case where a surprise cut in the safe rate i signals that the central bank received the private signal $s = L$. This causes firms to update their beliefs to p_L. The illustrative model allows us to cleanly decompose

FIGURE 1: THE FIRM INVESTMENT RESPONSE



(A) Static Investment Choice



(B) A Rate Cut

Note: The figures depict two stylized distributions for $F_p(\mathcal{R})$ and $F_{p_L}(\mathcal{R})$. The left figure shows what share of projects – under beliefs p – exhibit gross returns of more than the user-cost-of-capital which consists of the common risk-free rate i_e and the firm-specific component r. The right figure depicts the increase in the share of funded projects after a monetary surprise cut to i that updates beliefs to p_L.

the investment response: a lower real cost of capital entices investment, but at the same time firms adjust their beliefs about economic fundamentals, which shifts the distribution of gross returns and partially offsets the borrowing cost channel.[4] Proposition 1 formalizes this.

Proposition 1 (Firm-Level Decomposition)**.** *If the central bank cuts rates from i_{old} to $i < i_{old}$, the change in investment of firm j can be decomposed into an information effect and a borrowing cost channel:*

$$\Delta_{i_{old}\to i}\left(\mathcal{I}_j\right) = \underbrace{F_p(r+i_{old}) - F_p(r+i)}_{\textit{Borrowing Cost Channel}} - \underbrace{\left[F_{p_L}(r+i) - F_p(r+i)\right]}_{\textit{Information Effect}} \tag{4}$$

Proof. The difference between investment with and without the interest rate shock is

$$\Delta_{i_{old}\to i}\left(\mathcal{I}_j\right) = F_p(r+i_{old}) - F_{p_L}(r+i)$$

Using the definition of the conditional CDF F_p, we can split the second part into the shift in the interest rate at a constant return distribution and a component that reflects the shift in the CDF. □

Now suppose there is a mass of firms with the idiosyncratic borrowing cost component r distributed according to some distribution $\mu(r)$. Then, if r is the only source of heterogeneity, the aggregate investment response in the economy is

$$\Delta_{i_{old}\to i}\left(\mathcal{I}\right) = \int_r \left\{F_p(r+i_{old}) - F_p(r+i)\right\} d\mu(r) - \int_r \left\{F_{p_L}(r+i) - F_p(r+i)\right\} d\mu(r) \tag{5}$$

The information channel is large when, for a sufficient mass of firms, (i) the monetary policy surprise is informative, i.e. $|p-\tilde{p}|$ is large and (ii) projects near the threshold of investment are state-sensitive, i.e. the mass difference in F_p and F_{p_L} (or F_p and F_{p_H} for a rate hike) is sizable.

[4]The figure plots an inward shift of the CDF as firms become more pessimistic about future economic fundamentals. This results in a dampening of the borrowing cost channel. The model would predict an amplification of the borrowing cost channel if a more pessimistic outlook improves profitability of projects. While this may be the case for some counter-cyclical firms, we depict the more standard case that is likely to hold in the aggregate.

Such behavior is in line with the idea of an information effect for the overall economy as proposed by Romer and Romer (2000) and Nakamura and Steinsson (2018). Yet whether this effect is quantitatively meaningful for firm investment decisions remains ambiguous from the model, as the size of the information effect is governed by how much information monetary policy surprises transmit and what measure of firms lies in the respective interval. In the remainder of this paper, we provide empirical evidence that the offsetting information channel exists in practice and is quantitatively meaningful.

2.3 Identification

The simple model clarifies why conventional monetary policy shocks do not identify the structural effect of borrowing costs on investment. Variation in the user cost of capital, $i + r$, driven by changes in the policy rate i, is correlated with revisions in expected fundamentals. Because these information effects are unobservable and enter the error term of a standard investment regression, the borrowing-cost variation induced by i is endogenous. As a result, the estimated elasticity reflects a composite object that conflates price effects with belief revisions. The magnitude of this bias increases the more information is transmitted through the monetary policy surprise.

This paper instead exploits cross-sectional variation in the firm-specific component r induced by firms' heterogeneous, quasi-random exposure to the ECB's Corporate Sector Purchase Programme (CSPP). The shock shifts borrowing costs in a manner orthogonal to firms' revisions in expected fundamentals, allowing us to isolate a pure cost-of-debt channel.

The macroeconomic relevance of the information effect remains contested in the literature. We show that it is quantitatively important for aggregate dynamics: the firm investment response to a conventional monetary policy shock is 38 percent smaller than the response to an isolated cost-of-debt shock. Our approach is agnostic about the precise mechanism through which information transmits, which may differ across industries and firms. Instead, we identify the common borrowing-cost channel directly and recover the information effect – a potentially heterogeneous bundle of mechanisms – as a residual.

3 Institutional Background and Data

The Corporate Sector Purchase Programme (CSPP) belongs to a series of quantitative easing efforts launched by the ECB since 2012. These efforts comprise the ECB's Public Sector Purchase Programme (PSPP), its Asset-Backed Securities Purchase Programme (ABSPP), three rounds of Covered Bond Purchase Programme (CBPP), and the CSPP. While the PSPP focused on purchases of sovereign bonds issued by euro-area governments, and the ABSPP, CBPP1, CBPP2, and CBPP3 chiefly targeted bonds backed or covered by pools of loans such as residential mortgages, auto loans, or public sector loans, the CSPP was the first program allowing the ECB to purchase large quantities of corporate bonds on primary and secondary markets. Launched on 10 March 2016, the CSPP was expected to last at least until March 2017:

> "Third, we decided to include investment-grade euro-denominated bonds issued by non-bank corporations established in the Euro Area in the list of assets that are eligible for regular purchases under a new corporate sector purchase programme. This will further strengthen the pass-through of our asset purchases to the financing conditions of the real economy. Purchases under the new programme will start towards the end of the second quarter of this year."
>
> — ECB Monetary policy decisions, 10 March 2016, 14:30 CET

Notably, the decision to include investment-grade non-bank corporate bonds as part of the ECB's asset purchases caught the market by surprise (Todorov, 2020). The CSPP was a novel measure of QE, as the ECB had so far almost exclusively purchased government bonds. The only other central bank to have purchased corporate bonds at scale as a means of unconventional monetary policy was the Bank of Japan. The CSPP's novelty, however, is not its only attractive feature for our analysis: the program was also substantial in size. Asset holdings under the CSPP peaked at €345 billion, which equaled about a third of the market capitalization of all outstanding eligible European corporate bonds at the time. A final attractive feature of studying bonds around the CSPP announcement is that 2016 and following years were a period of relative market calm, as the main goal was not to provide new monetary stimulus, but rather to strengthen pass-through to financing conditions. Unlike many other QE programs set up in response to the Great Financial Crisis 2008/09 or COVID-19, the CSPP announcement did not coincide with a major global financial crisis. Thus, the CSPP announcement constitutes a large shock to the European corporate bond market that was largely unanticipated and occurred in a period with few other coinciding trends distorting investment decisions.

The core eligibility criteria for corporate bonds were disclosed on the initial announcement on March 10. For a corporate bond to be eligible it had to satisfy the following five conditions: (i) be denominated in euro, (ii) have a minimum rating of BBB- or equivalent, (iii) have a remaining maturity of at least six months and at most 30 years, (iv) be issued by a firm incorporated in the Euro Area, (v) be issued by a non-financial firm. Purchases were scheduled to begin at the start of June 2016. A further announcement on 21 April 2016 addressed the implementation aspects of the CSPP. Yet the core eligibility requirements were released in March.

3.1 Data

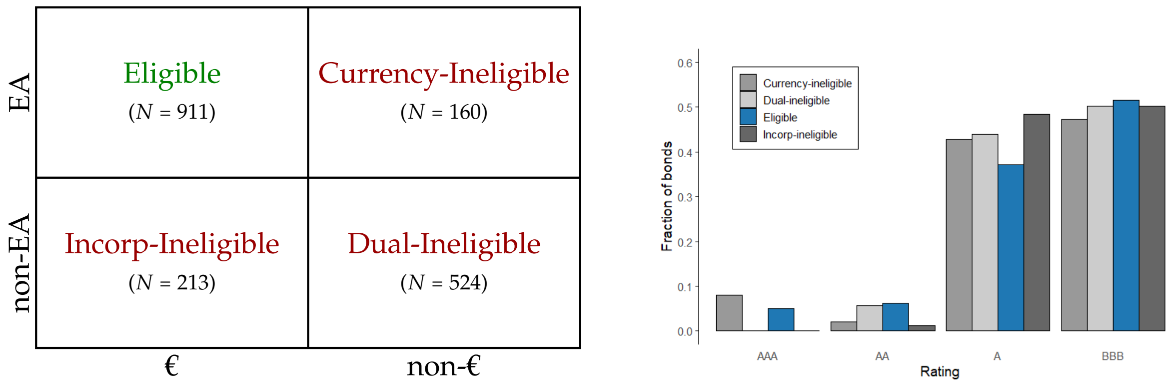
Our empirical analysis draws on three types of data. First, to study financial market dynamics, we use daily financial data on corporate bond yields at the ISIN level sourced from Bloomberg and Refinitiv Eikon. Second, at the firm level, we source annual balance sheet data from Moody's Orbis database. Lastly, we compare estimates of our new cost-of-debt shock to estimates from conventional investment elasticity approaches, so we also leverage an existing series on high-frequency monetary policy shocks.

Bond Data. To document the relevance of the CSPP for bond borrowing costs, we study the reaction of corporate bond yields in a 6-month window surrounding the CSPP announcement on 10 March 2016. Specifically, we collect data from January to June of that year. We only collect data on bonds that were eligible under the CSPP, as well as other European bonds ineligible due to their currency of denomination or their country of incorporation.[5] Our analysis is based on the universe of bonds satisfying the following five criteria: (i) be investment grade with a rating of at least BBB-, (ii) be issued by a non-financial firm, (iii) have a maturity of between 6 months and 30 years, (iv) be denominated in the currency of a European country, and (v) be issued by a firm incorporated in a European country. According to Bloomberg, as of 10 March 2016, there were 1,956 active corporate bonds satisfying these criteria, of which 1,808 had daily price and yield data available on Bloomberg or Refinitiv Eikon. Of these, 911 were eligible bonds and 897 were ineligible bonds.

Throughout the rest of the paper, we will refer to a particular partition of our bond sample as laid out in figure 2a. Namely, we partition the sample along two dimensions of eligibility into four groups: (i) *Eligible* bonds, (ii) *Incorp-ineligible* bonds that are ineligible solely because the issuing firm is incorporated in a European country that is not in the Euro Area, (iii) *Currency-ineligible* bonds that are ineligible only due to their denomination in a European currency that is not the Euro, and (iv) *Dual-ineligible* bonds that are ineligible due to both denomination and incorporation.

Figure 2b reports the rating distribution for the final sample of corporate bonds used in the

FIGURE 2: BOND DATA



(A) Partition

(B) Bond Ratings

[5]Other cutoffs are conceivable, such as eligibility around a BBB- credit-rating. We do not pursue this direction because prior work documents large spillovers from corporate sector purchase programs to ineligible lower-rated corporate bonds (De Santis and Zaghini, 2021), leaving no borrowing-cost wedge between exposed and unexposed firms.

analysis across this partition. Additionally, table 1 presents summary statistics on maturity, coupon rates, and par values. Bonds across the four segments appear similar in average maturity, coupon rate, par value, and bond rating. To ensure that our results are not driven by differences in maturity or bond rating, we also rerun our regressions within maturity and bond rating buckets as a robustness check in appendix section B.

For every bond, we collect two daily yield measures. First, we collect the bond's raw yield to maturity from Bloomberg or Refinitiv Eikon. Second, we construct daily government-yield spreads between the corporate bond's yield and the yield of the currency- and maturity-matched government bond. Our sample includes bonds denominated in euro, Swiss franc, Danish krone, Swedish krona, Norwegian krone, and British pound. Thus, we obtain daily estimates for each currency's yield curve from the corresponding central bank. In the case of bonds denominated in euro, we use the ECB estimates for the Euro Area yield curve.[6]

TABLE 1: SUMMARY STATISTICS OF CORPORATE BONDS

					Deciles				
		Mean	Std	Min	20%	40%	60%	80%	max
Maturity, Years	Eligible	5.40	4.28	0.52	1.89	3.56	5.51	8.07	29.60
	Incorp-ineligible	6.00	4.44	0.50	2.26	4.17	5.98	9.03	29.0
	Currency-ineligible	5.33	5.62	0.52	1.45	2.37	4.26	8.30	26.50
	Dual-ineligibility	10.70	7.92	0.52	3.10	6.53	11.3	19.10	30.00
Coupon, %	Eligible	2.79	1.78	0.00	1.12	2.12	3.30	4.50	8.62
	Incorp-ineligible	2.59	1.66	0.00	1.12	1.88	2.75	4.12	7.38
	Currency-ineligible	3.43	2.30	0.00	1.23	2.38	3.85	5.80	10.0
	Dual-ineligibility	3.97	2.29	0.00	1.65	3.12	5.00	6.00	10.0
Par, mil $	Eligible	769	572	5	284	634	812	1,091	4,180
	Incorp-ineligible	723	425	22	388	645	782	1,004	2,264
	Currency-ineligible	531	487	38	134	305	472	818	2,485
	Dual-ineligibility	407	325	2	134	281	414	607	2,009

Note: The table shows summary statistics of outstanding corporate bonds active on the CSPP announcement date. Data are shown for bonds that can be purchased under the CSPP ('Eligible') as well as bonds ineligible due to their country of incorporation only ('Incorp-ineligible'), bonds ineligible due to their currency denomination only ('Currency-ineligible'), or bonds ineligible due to both incorporation and denomination ('Dual-ineligibility').

[6] The ECB publishes yield curve estimates based on the full universe of outstanding euro-denominated government bonds, as well as a version based only on 'AAA'-rated Euro Area government bonds. We use the latter.

Firm Data. We draw firm-level variables from the annual Orbis database, a panel of publicly and privately listed firms with global coverage. The scope of our analysis is large European non-financial companies that issue investment-grade bonds. To identify the universe of such companies, we first collect all distinct issuer names from Bloomberg for firms that appear as issuers in our previously defined sample of 1,956 European investment-grade bonds. This yields a sample of 402 distinct issuer firms. We use Orbis data on firm ownership structures to map the issuers to unique global ultimate corporate owners. This is necessary as many issuing firms are financial subsidiaries of their larger manufacturing parent companies.[7] We identify 292 unique global ultimate owner companies. While all of these companies issue investment-grade bonds in Europe, not all of them are incorporated in the European Single Market. For example, if a non-European firm establishes a subsidiary in the Euro Area to issue bonds, or issues bonds in a European currency, the firm features in our sample. To ensure comparability across firms in our sample, we exclude all global ultimate owners that are not incorporated in the European Single Market. This leaves us with 183 distinct firms. This is only a small fraction of the over 40,000 public and private companies incorporated in the European Single Market. Yet because firm size is heavily skewed, the 183 firms account for over 40 percent of aggregate revenue, making our sample relevant for aggregate dynamics.[8]

For the set of 183 firms, we collect annual data on firm characteristics from Moody's Orbis database between 2000 and 2023. We collect data on firm investment, the cost of debt, and several other firm characteristics. For investment, our main specification relies on growth in total tangible assets. In appendix section B, we also report robustness results for (i) growth in tangible fixed assets, and (ii) the change in capital expenditure (CapEx) as a share of total assets, where we calculate CapEx from the change in tangible fixed assets plus depreciation. To proxy for firms' effective cost of debt, we construct the ratio of total interest paid each year to end-of-year total debt, captured by the sum of current and non-current liabilities. We

[7] For example, Volkswagen AG issues its bonds via six subsidiaries incorporated in Australia, the Netherlands, Germany, Japan, the US, and Canada with bonds denominated in 13 different currencies.

[8] Data on the universe of public and private companies in the European Single Market is drawn from Moody's Orbis database.

TABLE 2: SUMMARY STATISTICS OF 2015 FIRM CHARACTERISTICS

	Mean	Median	Std	Min	Max	# Obs
Total Assets (m EUR)	41,060	17,058	62,374	726	415,812	150
Total Debt (m EUR)	27,935	10,015	47,841	432	319,713	150
Interest-to-Debt-ratio (ppt)	2.17	2.02	1.35	0.05	12.40	141
Age (years)	53.9	40.0	43.79	6.0	180	145
Profitability (ppt)	10.91	10.51	5.32	-1.99	31.70	144
Liquidity (ppt)	9.76	7.75	9.87	0.00	92.06	150
Leverage (ppt)	64.77	64.69	14.61	23.61	134.75	150

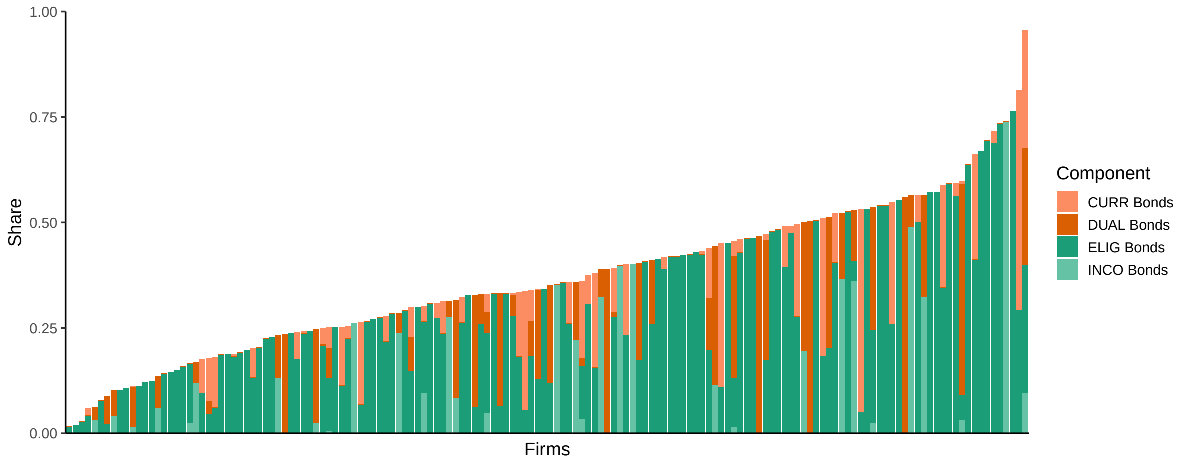
Note: The table shows summary statistics for firms in our sample. Summary statistics are on the cross-section of firm data for 2015, which serves as a benchmark in our main specification. Data are based on a total of 183 firms.

further collect data on (i) firm age based on the year of incorporation, (ii) firm size proxied by the log of total assets, (iii) profitability, as the ratio between earnings-before-interest-taxes-depreciation-and-amortization (EBITDA) and total assets, (iv) liquidity, as the ratio between firms' total cash or cash equivalent holdings and total assets, and (v) leverage, as the ratio of total debt to total assets. Table 2 presents summary statistics for key firm characteristics in our final sample.

Construction of CSPP Exposure. Our empirical identification leverages the fact that otherwise similar, large European firms differed in the extent to which their borrowing consisted of corporate bonds eligible under the CSPP. This is informative as the CSPP drove a wedge between the borrowing cost in CSPP-eligible and ineligible bonds as we show later in section 4. To capture this, we construct firms' pre-CSPP exposure to borrowing under CSPP-eligible bonds by merging firms' balance sheet data with Bloomberg data on their outstanding bond series. In addition to the price data for our sample of 1,956 bonds, we further collect the nominal outstanding amount of all bonds issued by the 183 firms we study, as well as their subsidiaries' outstanding bonds. Here we collect data on the whole universe of bonds issued by these firms, not just the ones placed in a European country and in a European currency. We construct a firm's total pre-CSPP exposure as the ratio of a firm's Q4 2015 borrowing in CSPP-eligible corporate bonds to its Q4 2015 total debt balance.

Figure 3 visualizes the distribution of firms' CSPP exposure in our final sample. For

FIGURE 3: FIRM CSPP-EXPOSURE



Note: This figure shows the debt portfolio composition before the CSPP in Q4 2015 for all 152 firms that have Orbis data on their total debt amount and that issued investment-grade bonds in Europe and whose global parent company is incorporated in a European country. Total debt is measured using Orbis data (current plus non-current liabilities), while the share of different corporate bond segments is calculated using Bloomberg data on the outstanding amount of the global universe of active bonds for each such firm. ELIG bonds, CURR bonds, INCO bonds, and DUAL bonds are defined as in previous sections.

each firm, we plot the share of debt raised via corporate bond borrowing, broken down into the four segments defined above: eligible, currency-ineligible, incorp-ineligible, and dual-ineligible. Firm reliance on non-bank borrowing via bond issuance exhibits rich heterogeneity, with most firms falling between 10 and 65 percent. There is also considerable heterogeneity in how reliant firms are on eligible bond issuance, conditional on their overall bond reliance. This allows us to compare the investment response of firms that are more exposed to the CSPP-induced wedge in bond yields, but also lets us compare firms that differ in this exposure while otherwise similar in their reliance on corporate bond financing.

Monetary Policy Shocks. Our novel identification strategy isolates investment elasticities to a cost-of-debt shock by leveraging heterogeneous firm exposure to the CSPP. We complement it with a conventional estimation of firm elasticities on the same sample of firms and a comparable time period, using a series of high-frequency monetary policy shocks. We follow the event-study approach pioneered by Kuttner (2001), and use the high-frequency shock series constructed for the Euro Area by Altavilla et al. (2019), available as part of the Euro Area

TABLE 3: SUMMARY STATISTICS OF MONETARY POLICY SHOCKS

	High Frequency	Sum
Mean	0.06	2.1
Median	0.0	0.5
S.D.	2.1	7.2
Min	-9.4	-19.4
Max	13.9	16.9
Observations	151	27

Note: Summary Statistics of monetary policy shocks from 01/01/2000 to 31/12/2023. 'High-frequency' shocks are estimated based on equation 6 and taken from Jarociński and Karadi (2020). 'Sum' shocks are aggregated to a yearly frequency. Surprises in the EONIA rate are in basis points.

Monetary Policy Event-Study Database (EA-MPD). Shocks ε_t^m are constructed as:

$$\varepsilon_t^m = (\text{EONIA}_{t+\Delta+} - \text{EONIA}_{t-\Delta-}) \tag{6}$$

where t is the time of the monetary announcement, and EONIA is the Overnight Index Swap (OIS) rate based on the Euro OverNight Index Average (EONIA). Altavilla et al. (2019) use swap rates for the Euro Area rather than futures contracts because European swap markets are more liquid and have a longer history than their futures counterparts. We focus on changes in the 3-month swap rate. Δ_+ and Δ_- determine the event window and are set to 10 minutes before and 20 minutes after the monetary policy announcement, yielding a 30-minute window. Our shock series runs from 2000 to 2023, in line with our firm data. We aggregate the high-frequency shocks to annual frequency to merge them with our firm-level data. Table 3 reports summary statistics for the high-frequency and aggregated series.

4 Evidence on CSPP Exposure

The cornerstone of the analysis is the relevance of firms' CSPP exposure for their cost of debt. In the following, we perform several exercises showing how the CSPP persistently changed borrowing conditions for European firms. First, in line with the preferred habitat view (see e.g. Culbertson, 1957; Modigliani and Sutch, 1966; Vayanos and Vila, 2021), we show that the

CSPP lowered interest rates for borrowing in CSPP-eligible corporate bonds. We show this both empirically and in a stylized model of a segmented corporate bond market. Second, we show that firm's pre-exposure borrowing mix is sticky: firms do not shift towards cheaper CSPP-eligible corporate bond borrowing after the announcement. The CSPP therefore induced a wedge between different types of borrowing to which firms remained persistently exposed.

Empirical Framework. To document the effect of the CSPP announcement on bond yields, we compare CSPP-eligible and ineligible corporate bonds before and after the announcement date:

$$y_{it} = \alpha_i + \lambda_t + \gamma_{Treat \times Post} \times Elig_i \times Post_t + \varepsilon_{it}. \tag{7}$$

Here, y_{it} is either the corporate bond's yield to maturity or its spread over the maturity- and currency-matched government bond. α_i is an individual bond fixed effect and λ_t is a time-fixed effect. The dummies $Elig_i$ and $Post_t$ indicate whether a bond is eligible for purchase under the CSPP and whether an observation falls on or after the CSPP announcement, respectively. The sample period analyzed in this study runs from January 2016 to June 2016, which incorporates 127 trading days. The methodology is akin to Todorov (2020). To the best of our knowledge, no other major central-bank announcements or macroeconomically relevant event coincided with the CSPP announcement on March 10, 2016.

Table 4 reports the results from estimating the baseline regression (7) for two yield measures, as well as for four different sets of control groups: (i) all ineligible bonds, (ii) only currency-ineligible bonds, (iii) incorp-ineligible bonds, and (iv) dual-ineligible bonds. We report estimates for the coefficient $\hat{\gamma}$, the effect of the CSPP on eligible corporate bonds relative to the respective control group of ineligible bonds. In line with the preferred habitat view as well as previous work on corporate sector purchasing programs (Todorov, 2020; De Santis and Zaghini, 2021), the CSPP reduced yields of eligible bonds relative to ineligible ones by 8.6

basis points.[9]

Two novel findings stand out. First, while the CSPP has reduced bond yields of eligible bonds relative to ineligible bonds by 8.6 basis points, this effect on the yield of eligible bonds varies across control groups and is largest when comparing eligible bonds to dual-ineligible bonds (-17.5 basis points). Second, most of the total effect comes not from the CSPP's effect on the underlying yield curve but from its compression of government yield spreads. When studying the cleanest control group, the CSPP announcement reduced yields of eligible bonds relative to dual-ineligible bonds by 17.5 basis points, 13.1 of which are driven by a reduction in yield-spreads, not the yield curve.

We attribute the former finding to varying spillovers from heterogeneous portfolio rebalancing effects. Figure 4 makes the point descriptively, plotting average yield spreads around the CSPP announcement for the four bond market segments. Eligible bonds' government-yield spreads dropped sharply on the CSPP announcement date, by 17.9 basis points – equivalent to 6.0 standard deviations of their typical daily changes. Incorp-ineligible bonds fell

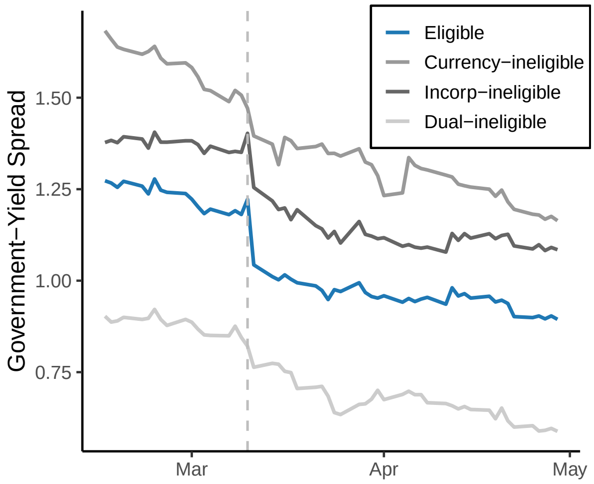
TABLE 4: IMPACT OF THE CSPP ON BOND YIELDS

Dep. Variable	Yields				Yield Spreads			
Control Group	ALL	INCO	CURR	DUAL	ALL	INCO	CURR	DUAL
Post × Treat	-8.6*	7.5	-2.8	-17.5***	-4.8	6.2	5.0	-13.1***
	(4.2)	(15.2)	(4.8)	(2.0)	(4.1)	(15.1)	(4.7)	(1.9)
# Observations								
(Bond-Time)	221k	137k	133k	173k	221k	137k	133k	173k
Bond FEs	YES	YES	YES	YES	YES	YES	YES	YES
Time FEs	YES	YES	YES	YES	YES	YES	YES	YES
F-Statistic	67.9	12.6	26.9	2974.2	21.3	8.7	86.0	1698.9

Note: Columns one to four estimate equation (7) for corporate bond yields as the dependent variable. Columns five through eight estimate equation (7) for corporate government-yield spreads as the dependent variable. Regressions are estimated based on eligible bonds as well as a varying group of ineligible bonds. Columns indicated by ALL are based on all eligible and all ineligible bonds, while columns indexed by INCO, CURR, and DUAL are based on eligible bonds and incorp-ineligible bonds, currency-ineligible bonds, and dual-ineligible bonds only, respectively. We report the coefficient of $Treat \times Post$. Standard errors as shown in parentheses are heteroskedasticity-robust and double-clustered by bond and time. Statistical significance is indicated by *, **, and *** at the 5%, 1%, and 0.1% level. Units for bond yields and government-yield spreads are in basis points.

[9] This finding is consistent with Krishnamurthy et al. (2011), who similarly document a disproportionately large price effect for asset classes targeted by the Fed's QE1 and QE2.

FIGURE 4: GOVERNMENT YIELD SPREADS AROUND THE CSPP ANNOUNCEMENT



Note: The figures show average trajectories of our corporate bond sample around the CSPP announcement on 10 March 2016 (dashed vertical line). We partition our sample of bonds into bonds that could be purchased under the CSPP ('Eligible') as well as bonds ineligible due to their country of incorporation only ('Incorp-ineligible'), bonds ineligible due to currency denomination only ('Currency-ineligible'), or bonds ineligible due to both incorporation and denomination ('Dual-ineligibility').

similarly, by 14.8 bps (4.9 standard deviations of their typical daily changes). Dual-ineligible bonds' government-yield spreads, meanwhile, dropped by only 5.7 (3.1 standard deviations), indicating weaker spillovers. These high-frequency estimates on yield-spreads match our findings reported in Table 4.

In appendix section B, we further provide estimates for the coefficient of the interaction dummy estimated for a particular maturity-rating segment of bonds in table 9. Results hold at this more granular level. As in our main specification (Table 4), results broken down by maturity- and rating buckets confirm that the CSPP induced a sizable yield wedge between eligible and ineligible bonds. The effect is consistently larger when comparing eligible bonds only to dual-ineligible bonds, which appear to be the cleanest control group experiencing the least spillover. The wedge is driven predominantly by an effect on yield-spreads.

A Stylized Model of Imperfect Arbitrage. Empirically, the CSPP's spillover to different ineligible bond market segments varies throughout our observation window. This is in line with the preferred habitat view (see e.g. Culbertson, 1957; Modigliani and Sutch, 1966; Vayanos and Vila, 2021). One explanation for this heterogeneity is differing degrees of market segmentation, as in Greenwood et al. (2018). To support this hypothesis, we show that a stylized model of a segmented corporate bond market can closely replicate the observed yield dynamics following the CSPP announcement, for varying degrees of market segmentation. Here, market segmentation refers to how frictionlessly capital can flow between segments. We further show that, in this model, such effects are persistent.

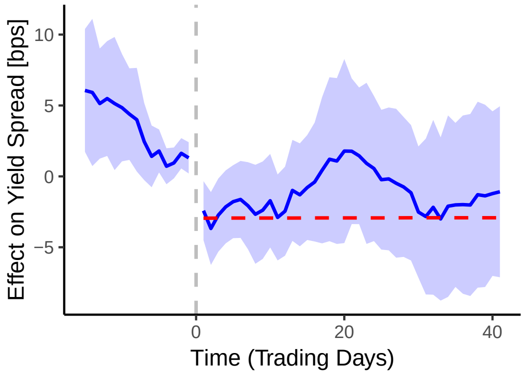
Supplementary materials C lay out the model structure of a partially segmented corporate bond market in detail. It describes a stylized model for pricing two long-term defaultable bonds A and B, representing CSPP-eligible bonds and a segment of the ineligible corporate bond market, respectively. We extend the model of Greenwood et al. (2018) by allowing for default risk in both bond market segments to study market segmentation within the corporate bond market. The model features two key asset-pricing frictions: partial market segmentation and slow-moving capital across markets. Markets are segmented in the sense that there are three types of traders: specialists A and B, who can only trade in bonds A and B, respectively, as well as generalist traders, who can trade both bonds. Capital is slow-moving in the sense that generalists can only adjust their portfolio every k periods, akin to a Calvo (1983) friction. Bond supply in the model is stochastic. We model the CSPP announcement as a shock to the supply of eligible bonds A under different degrees of market segmentation, which we proxy by the measure of generalist traders q_G in the model.

For each of the three eligible-ineligible bond pairs (currency-ineligible, incorp-ineligible, and dual-ineligible), we structurally estimate the measure of generalist traders q_G that minimizes the distance between (i) the model-implied yield spread between bonds A and B following the CSPP shock, and (ii) the following empirical local projection à la Jordà (2005):

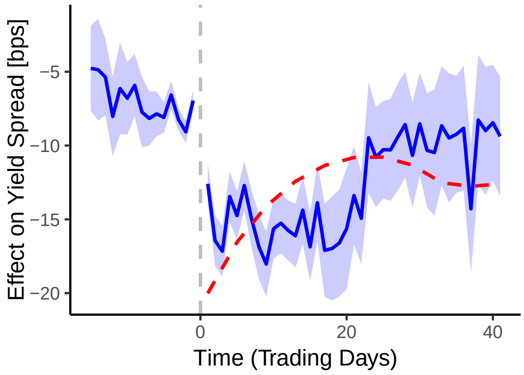
$$y_{i,t_{CSPP}+h} - y_{i,t_{CSPP}} = \alpha + \gamma_h \times D_{i,t} + \varepsilon_i. \tag{8}$$

Here, $D_{i,t}$ is a dummy variable that is one if bond i is eligible and t equals the announcement date of the CSPP and is zero otherwise. The simulated method of moments minimizes the dif-

FIGURE 5: TRANSITIVE EFFECT OF THE CSPP ($\hat{\gamma}_k$)



(A) Eligible spread to Incorp-ineligible



(B) Eligible spread to Dual-ineligible

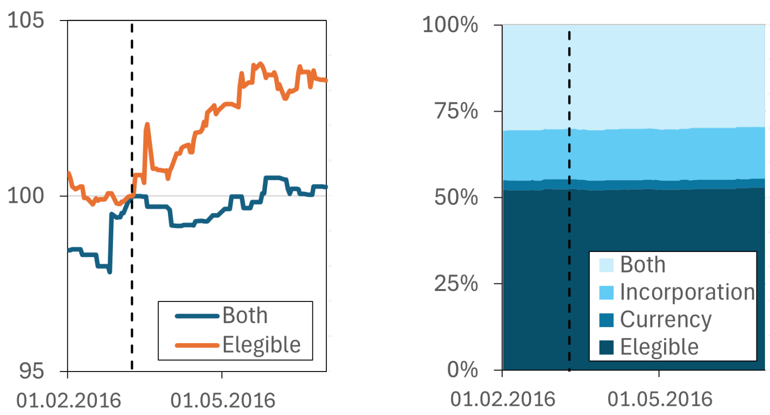
Note: The two figures compare empirical estimates (solid blue line) of the transitive effect of the CSPP to model outcomes from matched moments thereof (dashed red line). We do so for the relative effect of eligible bonds relative to incorp-ineligible bonds (left) and dual-ineligible bonds (right). Empirical estimates are based on parameter γ_h from equation (8), where the confidence bands reflect heteroskedasticity-consistent standard errors. Model predictions reflect the difference in yield spread $y_{A,t} - y_{B,t}$, h periods after the supply contraction.

ference between $\hat{\gamma}_t$ and the model-implied yield difference between bonds A and B following the CSPP announcement in the model.

Figure 5 shows the local-projection results ($\hat{\gamma}_t$) alongside the model-implied yield spread for the degree of market segmentation that most closely replicates them. We present results for eligible versus incorp-ineligible bonds, which exhibit the largest spillovers, and eligible versus dual-ineligible bonds, which exhibit the smallest. We closely replicate the former under no market segmentation, where all traders are generalists. We replicate the latter under partial market segmentation, where only 40 percent of traders are generalists. The effect on yield spreads is also persistent in the latter case: ineligible-bond yields stay depressed ten basis points below eligible yields.

Sticky Firm Exposure. Lastly, we document stickiness in firms' CSPP exposure. Pre-CSPP exposure to CSPP-eligible bond borrowing would not be indicative of actual borrowing costs if firms could switch frictionlessly between different credit lines. However, similar to the notion of relationship lending in bank-based borrowing (see e.g. Boot, 2000; Ongena and Smith, 2000; Kysucky and Norden, 2016), we find substantial frictions and fixed costs associated

FIGURE 6: DAILY EVOLUTION OF OUTSTANDING CORPORATE BONDS



(A) Total Volume Growth (B) Within-Firm Shares

Note: Figures are based on the daily universe of bonds that are investment-grade, denominated in a European currency, whose issuer is a non-financial firm incorporated in a European country, and whose outstanding maturity is between 6 months and 30 years. Figure (A) plots the trajectory of total outstanding eligible bond volume over time relative to total dual-ineligible bond volume. Figure (B) plots the trajectory of within-firm reliance on the four partitions of the European corporate bond market. Here, firms are clustered based on their Bloomberg ticker, which exhibits a nearly one-to-one mapping with the global ultimate owners of issuing companies identified via Orbis ownership structures. The dashed line reflects the CSPP announcement on 10 March 2016.

with switching from bank borrowing to corporate bond issuance, and in switching market segments within the corporate bond market. In particular, we find that the CSPP led to some increased issuance of eligible bonds on the intensive margin, but not on the extensive margin.

To show this, we study the universe of European bonds as defined earlier that were active between 1 February 2016 and 30 June 2016. Figure 6b shows that the average within-firm reliance on bonds eligible under the CSPP remains stable both before the CSPP and up to three months after the CSPP announcement. This indicates that firms do not shift to borrowing under comparatively cheaper CSPP-eligible bonds. We further confirm this by analyzing the bond issuance behaviour of firms that have a 100 percent reliance on dual-ineligible bonds. Of these firms – which had issued only dual-ineligible bonds before the announcement – none issued bonds from other bond market segments in the three months that followed the CSPP announcement. This indicates that while firms appear to increase issuance of eligible bonds on the intensive margin, firms do not shift towards eligible bonds on the extensive margin (Figure 6a).

Such persistence of firms' debt portfolios is also in line with the reality of corporate bond placements, which often include substantial fixed costs, chiefly driven by large underwriter fees. According to Dalal (2018), the average underwriting fee received by intermediary banks for the placement of new corporate bonds averaged around 0.7 percent (70 bps) of the par value for investment-grade bonds in the United States around the announcement of the CSPP. In addition to underwriting fees, firms have to pay legal fees, marketing and road-show costs, as well as regulatory filing fees, especially if they are planning on setting up a new subsidiary in a European Economic Area country. Relative to our estimated 17-basis-point reduction in bond yields, such fixed costs are a severe deterrent to switching bond markets.

5 Econometric Approach

Thus far, we have shown the three key ingredients that enable our analysis. First, the CSPP has induced a systematic shift in borrowing costs: borrowing in bonds eligible for purchase under the CSPP has become cheaper relative to other forms of borrowing. Second, we show that this wedge in borrowing costs persists. Third, the extent to which firms are exposed to this shift varies, as shown in Figure (3). Even conditional on a firm's reliance on borrowing in corporate bonds, the CSPP's eligibility cutoffs along lines of incorporation and denomination induce quasi-experimental variation in how exposed otherwise comparable multinational companies from the European Single Market are to the wedge. Extending this logic, we use CSPP pre-exposure as a source of variation in the cost of debt that is orthogonal to other forces that co-move with the monetary policy shock and usually cannot be partialled out. We discuss this assumption in more detail in subsection 5.1 on threats to identification.

Despite its many desirable properties, pre-CSPP exposure is only an imperfect proxy for a cost-of-debt shock. Thus, we do not use it as a direct shock measure, but rather as an instrument for the observed change in the firm-level interest-to-debt ratio. This is in the spirit of the

Bartik (1991)-style shift-share IV literature. Namely, we run the following two regressions:

$$\Delta_h r_i = \alpha Z_i + \gamma X_{i0} + \varepsilon_i, \tag{9}$$

$$\Delta_h Y_i = \delta \Delta_h \hat{r}_i + \tilde{\gamma} X_{i0} + \tilde{\varepsilon}_i. \tag{10}$$

Here, Z_i is a firm's CSPP exposure share, defined as the share of a firm's 2015 debt raised in CSPP-eligible corporate bonds, $\Delta_h r_i$ is the percentage change in firm i's effective interest rate (interest-to-debt ratio) over the h-year period following 2015, and X_{i0} is a vector of controls that includes firms' 2015 bond reliance, proxied by the share of 2015 debt raised via corporate bonds.[10] Lastly, $\Delta_h Y_i$ denotes the percentage change in our investment measure over the h-year period following the benchmark year 2015.

Choice of Instrument. Our econometric approach uses the share of bond financing in eligible bonds as the instrument. An advantage of this design is that it lets us remain agnostic about the relevant time window for yield changes. An alternative instrument incorporating yield changes directly, such as a convex combination of mean yield changes by bond category, would require taking a stance on the appropriate time window over which to compute those changes. Such a choice has no obviously correct answer. This exposes the instrument to specification-searching concerns and would require extensive robustness checks across alternative windows.

5.1 Robustness

Our identification strategy relies on the assumption that firms that differ in how much of their bond debt is raised via CSPP-eligible bonds do not meaningfully differ across other dimen-

[10] In appendix section B table 11, we provide estimates for investment elasticities from a larger set of controls, including firm size (log of total assets), profitability (EBITDA-to-asset ratio), leverage (debt-to-asset ratio), liquidity (cash-to-asset-ratio), firm age in years, and firms' procyclicality with the Euro Area, in the first and second stage. Doing so does not materially change estimated effect sizes, or levels of statistical significance. However, including all controls simultaneously reduces the number of observations our estimates are based on, which is why our main specification controls only for bond reliance.

sions that affect their investment elasticity, their responsiveness to the information effect, or other forces co-moving with monetary-policy-induced cost-of-debt shocks. The core rationale for why this holds originates in considering only large, predominantly multinational firms, all incorporated in the European Single Market, that issue investment-grade corporate bonds. Firms' bond-issuing choices in non-euro-area currencies or countries make some more or less exposed to the CSPP. Yet under the European Single Market, their common incorporation lets all firms sell goods and services frictionlessly within the single market, access capital markets across borders, and hire workers from any member country. In other words, because the single market functions like "one country" for firms in many respects, they should be similarly exposed to macro and monetary policy shocks in the euro area.

We formally test this assumption in two ways. First, we provide evidence that firm exposure to the CSPP is not meaningfully correlated with other relevant firm characteristics, and that firms more exposed to borrowing in CSPP-eligible bonds are not more pro-cyclical with Euro Area GDP. Second, we adapt the Donaldson (2026) test to our setting. Here, we test whether firms that differ in CSPP exposure exhibited different investment responses to ECB monetary policy shocks prior to 2016. We do not find this to be the case.

Descriptive Evidence. To show that firm exposure to the CSPP is not meaningfully correlated with other relevant firm characteristics, we first regress firm CSPP exposure on firms' overall bond-exposure in 2015:

$$Z_i = \alpha + \beta Bond_i + e_i. \tag{11}$$

Here, Z_i is the firm's share of debt raised via CSPP-eligible corporate bonds, while $Bond_i$ is the firm's share of debt raised via corporate bonds. We estimate this regression on our sample of firms to obtain the residual $\hat{e}_i$, then test whether this residual – the share of CSPP exposure not explained by bond exposure – is meaningfully correlated with other firm characteristics.

Table 5 summarizes the correlation between $\hat{e}_i$ and several firm characteristics. As shown, no firm characteristic exhibits a statistically significant correlation with the share of CSPP exposure that cannot be explained by bond reliance ($\hat{e}_i$). In addition to standard firm characteristics, summarized in the data section of the paper, we construct a firm-level measure of procyclicality with euro-area output. Namely, we calculate firm-level procyclicality as the correlation between the HP-filtered cyclical component of each firm's revenue growth and the cyclical component of the European Economic Area's aggregate GDP growth. Importantly, this measure also exhibits no correlation with $\hat{e}_i$, indicating that more CSPP-exposed firms have no greater incentives to respond to ECB monetary policy shocks because of greater reliance on the EEA market. To some extent, this reflects the fact that firms in our sample are very large and raise most of their revenue globally. While some of these firms appear to have a home bias in where they raise their bonds, their revenue generation appears globally diversified.

Donaldson (2026) Test. In a recent contribution, Donaldson (2026) shows that cross-sectional exposure designs can fail to identify partial-equilibrium elasticities if the exposure shifter is correlated with heterogeneous sensitivity to aggregate variables that respond in general equilibrium. Applied to our setting, the relevant concern is that firms with greater pre-CSPP exposure may also be systematically more sensitive to aggregate monetary conditions. In that case, an aggregate monetary-policy response coinciding with the CSPP announcement could generate differential investment responses correlated with CSPP exposure, contaminating the cross-sectional estimate. We therefore adapt the diagnostic proposed by Donaldson (2026) and ask whether CSPP exposure predicts differential investment responses to high-

TABLE 5: CORRELATION BETWEEN CSPP EXPOSURE AND FIRM CHARACTERISTICS

	Procyclicality	Size	Leverage	Liquidity	Profitability	Age
Correlation with $\hat{e}_i$	0.018	-0.155	-0.148	0.100	-0.008	-0.038
	(0.082)	(0.078)	(0.079)	(0.080)	(0.084)	(0.083)
P-Value	0.832	0.056	0.071	0.227	0.922	0.6476
Observations	148	149	148	148	141	144

Note: This table summarizes how different firm characteristics correlate with the share of firms' CSPP exposure that can't be explained via firms' bond-exposure. We report Pearson correlation coefficients with standard errors in brackets below. Statistical significance is based on a two-tailed test, and p-values are reported. Procyclicality refers to a firm's revenue procyclicality with EEA GDP since the incorporation of the firm. Size, leverage, liquidity, profitability, and age are defined in the data section. Statistical significance is indicated by *,**, and *** at 5%, 1%, and 0.1%.

frequency ECB monetary-policy surprises in the pre-CSPP period. We estimate the following specification using data from 2005 to 2016:

$$\Delta Y_{f,t} = \mu_f + \lambda_t + \delta\left(MP_t \times Z_i\right) + \gamma X_{f,t-1} + \varepsilon_{f,t}. \tag{12}$$

Here, λ_t and μ_f are time and firm fixed effects. MP_t reflects standard aggregated high-frequency monetary-policy surprises for the Euro Area, following Altavilla et al. (2019), as well as aggregated high-frequency "central-bank information" shocks, following Jarociński and Karadi (2020). The null of interest is $H_0 : \delta = 0$.

Table 6 summarizes the results of the adapted Donaldson (2026) test. It reports the coefficient of the interaction term once for a conventional high-frequency monetary-policy shock, as in Altavilla et al. (2019), and once for a central-bank information shock as per Jarociński and Karadi (2020).

TABLE 6: DONALDSON TEST

	Investment Response	
	Altavilla	JK (CBI)
$MP \times Z$	1.36	-0.41
	(0.969)	(0.716)
Obs	488	488

Note: The table reports the coefficient of the interaction term for $(MP_t \times Z_i)$ in equation (12). The left column uses standard high frequency monetary policy shocks aggregated to the annual frequency for the Euro area from Altavilla et al. (2019), while the right column uses aggregated Central Bank Information (CBI) shocks as reported in Jarociński and Karadi (2020). Statistical significance is indicated by *, **, and *** at the 5%, 1%, and 0.1% level. Standard errors, reported in parentheses, are heteroskedasticity-robust and clustered at the year level, using the Arellano covariance estimator with a finite-sample correction.

We find no statistically significant differential response associated with CSPP exposure. This suggests firms with greater pre-CSPP exposure were not systematically more sensitive to ECB monetary-policy shocks before 2016. As an additional diagnostic, we obtain a similar null result using Jarociński–Karadi central-bank-information shocks, suggesting that CSPP exposure does not systematically predict differential sensitivity to the information component of ECB announcements. These results are consistent with the identifying assumption that

CSPP exposure is unrelated to firms' sensitivity to aggregate forces that could move alongside the CSPP announcement. However, given the limited precision of some estimates, we interpret these findings as supportive rather than definitive evidence. Together with the balance results in Table 5, they reduce the concern that heterogeneous general-equilibrium sensitivities contaminate our estimated borrowing-cost elasticity.

5.2 High-Frequency Shock Comparison

The previously described setup allows us to estimate the investment elasticity in response to an isolated cost-of-debt shock. Yet, decomposing the aggregate investment elasticity into a cost-of-debt component and a residual still requires information about the aggregate response. To this end, we estimate an alternative first-stage regression using a series of high-frequency monetary-policy shocks as the instrument for the change in the interest-to-debt ratio:

$$\Delta r_{f,t} = \alpha MP_t + \gamma X_{f,t-1} + \varepsilon_{f,t}, \tag{13}$$

$$\Delta Y_{f,t} = \delta \Delta \hat{r}_{f,t} + \tilde{\gamma} X_{f,t-1} + \tilde{\varepsilon}_{f,t}. \tag{14}$$

where MP_t reflects aggregated high-frequency changes in the 3-month EONIA swap rate around ECB monetary policy announcements, following Altavilla et al. (2019), t indexes the year at an annual frequency, and f indexes the firm. We consider the same sample of 183 firms for which we previously calculated pre-CSPP exposure. We consider the time period between 2005 and 2019, which most closely resembles the time following the CSPP announcement studied earlier. We study only one-year changes over the year in which the monetary-policy shocks to the EONIA swap rate materialized. Other variables are defined as for equations (9) and (10).

Importantly, using conventional monetary policy shocks as an instrument allows us to estimate the aggregate investment elasticity to a monetary policy shock, not just the cost-of-debt channel. Consider, for example, the case where the rate change in the EONIA around a monetary policy announcement correlates with revisions of the economic outlook. Then in the second stage $\tilde{\varepsilon}_{f,t}$ would include a changed economic outlook that would matter for the investment decision, but would also be correlated with $\Delta\hat{r}_{f,t}$: years in which many rate cuts led to a reduced cost of debt would also be years in which the economic outlook improved. Thus, here δ is not the pure investment elasticity due to a pure cost-of-debt shock, but rather also incorporates all other correlated channels of the monetary policy announcement and should be interpreted as the aggregate investment elasticity induced by a monetary policy shock. Comparing these investment elasticities with our previous estimate – which leverages cost-of-debt variation orthogonal to such co-moving channels – lets us implicitly estimate the drag of the information effect on the investment elasticity.

6 Results

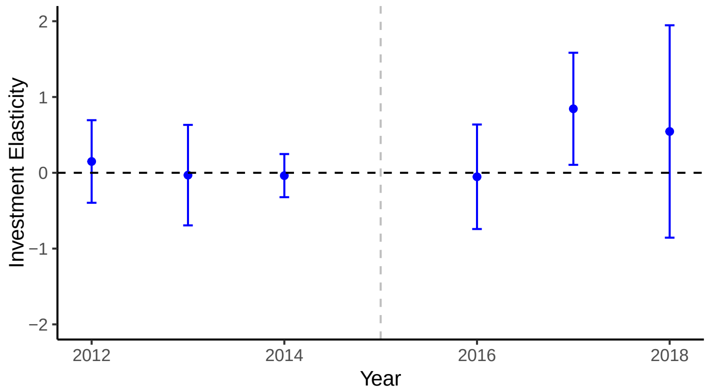
Our headline finding is summarized in Table 7, which compares the investment elasticity from a pure cost-of-debt shock with the aggregate elasticity from the series of monetary policy announcements. The borrowing cost specification estimates equations (9) and (10) using CSPP exposure as an instrument, while the monetary policy shock specification estimates equations (13) and (14) using aggregated high-frequency monetary policy shocks as an instrument. We present estimates for the investment elasticities $\hat{\delta}$. The first stages indicate

TABLE 7: DECOMPOSING THE INVESTMENT ELASTICITY

	Borrowing Cost Shock		MP Shock	
	2SLS	Fuller	2SLS	Fuller
Total Asset Growth	0.98**	0.83**	0.30**	0.27**
	(0.47)	(0.38)	(0.14)	(0.12)
First stage F	5.05		6.94	
Obs	138		998	

Note: The table reports the Two-Stage Least Squares (2SLS) estimates for δ. The borrowing cost shock specification estimates equations (9) and (10) over the two-year 2015-2017 window. The MP Shock specification, on the other hand, estimates equations (13) and (14). In addition to the 2SLS estimate, we also report the Fuller estimates, which are more robust to weak instruments. To further account for the possibility of a weak instrument, we compute Conditional Likelihood Ratio (CLR)–based confidence intervals for both Two-Stage Least Squares (2SLS) and Fuller estimators. Statistical significance is indicated by *, **, and *** at the 5%, 1%, and 0.1% level.

FIGURE 7: INVESTMENT RESPONSE



Note: The figure plots Fuller (1977) estimates for the investment elasticity $\hat{\delta}$ from equation (10) in response to a one percent reduction in the interest-to-debt ratio induced by a loosening of borrowing costs. The time axis corresponds to the window over which investment elasticities are estimated relative to the benchmark year 2015. Years below 2015 correspond to a negative h. Confidence bands Conditional Likelihood-Ratio based and reflect the 95% confidence interval.

a weak instrument in both cases, evidenced by an F-statistic below ten (Staiger and Stock, 1997; Stock et al., 2002). To account for this, we report estimates based on Fuller (1977) in our main specification, which reduce the small-sample bias of limited-information maximum-likelihood estimations, while maintaining low variance – a more robust estimator under a weak instrument. Additionally, we report confidence intervals from the conditional likelihood ratio (CLR) approach (Moreira, 2003), which is robust to weak instruments.

In the borrowing-cost shock specification of Table 7, results are based on differences over the two years following the benchmark year 2015. We focus on the two-year specification, as the first stage is strongest. In other words, pre-CSPP exposure is most relevant for the 2015-2017 change in firms' interest-to-debt ratios. Table 10 in supplementary materials B provides full regression results over one-, two-, and three-year windows following 2015 for our main investment measure, as well as the two other proxies: tangible fixed-asset growth and total capital-expenditure growth. Figure 7 further plots investment elasticities for our main investment measure over different time windows. Similarly, results are not statistically significant over one- and three-year windows. There are two possible explanations for this finding. First, because much of firms' debt is locked in fixed-rate debt contracts, it takes some time before lower borrowing costs in secondary markets affect the average interest rate of a firm's debt portfolio. In our sample, the average maturity of outstanding corporate bonds is 7.05 years, which could explain why the effect of the CSPP announcement on actual interest payments was muted in 2016. Second, while we document stickiness in firms' CSPP exposure via their bond portfolios, the longer the time window, the more likely some firms are to issue bonds in different market segments, which weakens the link between pre-CSPP exposure and firms' effective cost of borrowing. The two-year window appears optimal: long enough for effects to pass through to debt portfolios, but not so long that debt portfolios have time to rebalance. Figure 7 summarizes these considerations by showing estimates for the investment elasticity in the borrowing-cost shock specification based on an h-year window following the benchmark year 2015, for varying h. Here, we further consider negative h, which represents a retrospective application of our methodology to years before 2015. Finding no effect here suggests that firms with different pre-CSPP exposures did not differ in their pre-trends for borrowing costs and investment.

When annualizing investment elasticities from Table 7, we find that the elasticity from a cost-of-debt shock is much larger than the investment elasticity induced by a monetary policy shock. The former estimate of a 0.83 investment elasticity over two years corresponds to a conservative, annualized investment elasticity of 0.41, which exceeds the investment elasticity to an MP shock of 0.27 by 51.8 percent.[11] This interpretation is consistent with the existence of an information effect. If a monetary policy announcement also changes firms' economic outlook in the opposite direction from the rate cut's effect on investment, this should mute the investment elasticity relative to a raw cost-of-debt shock unaccompanied by a correlated revision to the economic outlook.

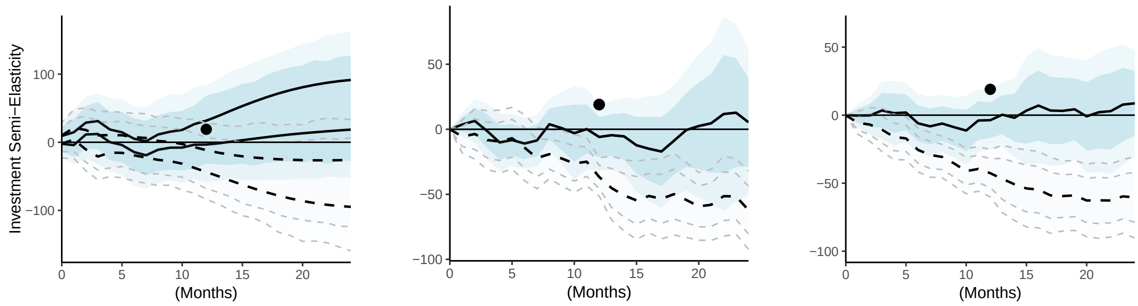
Comparison to High-Frequency Corrections. We compare our finding to semi-elasticities of firm investment to a monetary policy loosening, based on existing corrections for the information effect in high-frequency monetary-policy shock series. Figure 8 plots our results

[11] We annualize the elasticity over the two-year window as follows: $(1+\hat{\delta})^{0.5}$. This is conservative in that it assumes the reduction in borrowing costs has persisted equally over the whole two-year period, although the weak first stage over the one-year period suggests this is likely not the case.

against the investment responses to the shocks in Jarociński and Karadi (2020), Miranda-Agrippino and Ricco (2021), and Bauer and Swanson (2023). Appendix section A details how we replicate investment elasticities using the structural monetary policy shock in Jarociński and Karadi (2020) and the orthogonalized monetary policy shocks in Miranda-Agrippino and Ricco (2021) and Bauer and Swanson (2023). We mark our point estimate for the investment semi-elasticity at the twelve-month mark in all three figures.

Two findings stand out. First, all three corrections support our finding that the investment response is offset by a central bank information channel. For all three shocks, the corrected investment response is stronger and more in line with standard theory than the shock series that plausibly contains central bank information – evident in the larger investment response to a monetary policy loosening. Second, the existing approaches are also in line with the magnitude of the estimate presented in this paper. Our semi-elasticity of 19 percent after twelve months is in line with the set identification of Jarociński and Karadi (2020) and also falls within the 90 percent confidence bands of Miranda-Agrippino and Ricco (2021) and Bauer

FIGURE 8: INVESTMENT RESPONSE TO HF-MP SHOCKS



(A) Jarociński & Karadi (B) Miranda-Agrippino & Ricco (C) Bauer & Swanson

Note: The three figures plot the investment semi-elasticity for aggregate total private nonresidential construction spending in the United States to a one percentage point surprise loosening in the policy rate. Figures depict responses to the Jarociński and Karadi (2020) decomposition (left), the Miranda-Agrippino and Ricco (2021) shock series (middle), as well as the Bauer and Swanson (2023) series (right). In the case of Jarociński and Karadi (2020), the figure compares the set identified impulse response function to a monetary policy shock (solid) to a central bank information shock (dashed). Robustness for set identification is based on Giacomini and Kitagawa (2021). For Miranda-Agrippino and Ricco (2021); Bauer and Swanson (2023), the figures compare the investment response to orthogonalized high frequency monetyary policy shocks (solid) against the non-orthogonality shocks (dashed). Shaded areas and dashed gray lines reflect 68 and 90 percent confidence bands. Estimates on the left follow the VAR approach outlined in Jarociński and Karadi (2020). Estimates for Miranda-Agrippino and Ricco (2021); Bauer and Swanson (2023) are based on an OLS local projections with Newey–West standard errors, using horizon-dependent bandwidth. The point at 12 months and a semi-elasticity of 19 percent corresponds to the finding of our paper for the structural semi-elasticity of investment to a change in borrowing cost.

and Swanson (2023). Unlike these three approaches, our paper identifies a significant positive investment response, consistent with standard theory. Estimates from Jarociński and Karadi (2020) suffer from wide confidence bands because their structural decomposition of monetary policy shocks is not point-identified. Miranda-Agrippino and Ricco (2021) and Bauer and Swanson (2023) have point-identified estimates. However, their approach relies on finding an empirical proxy that spans the universe of all potential sources of central bank information. While they seem to capture some of this, as indicated by the non-negative investment response of the orthogonalized shock series, the fact that they cannot reject a zero investment response indicates that some information contamination remains in their shock series. Our approach suffers from neither limitation. Our estimate is set identified, and does not require assumptions about empirical controls for everything that might cause a central bank information effect. Our estimation also delivers tighter confidence bands, while all high frequency estimations fail to identify a statistically significant investment response.

Reduced Form Evidence. A separate concern with weak instruments is that a weak first stage can inflate the IV estimate even when the reduced-form relationship between instrument and outcome is itself weak. In the just-identified case, the estimand is the ratio β/α, where α is the coefficient from the first-stage regression in Equation (9) and β is the coefficient from the reduced-form regression:

$$\Delta_h Y_i = \beta Z_i + \gamma X_{i0} + \varepsilon_i. \tag{15}$$

When the first-stage coefficient α is close to zero, the ratio can become large even if β is small. Following Angrist and Pischke (2009), a common sanity check is to report the reduced-form regression in Equation (15) directly and confirm that the effect β is sizable. We report the reduced-form regression results in Table 8. For 2017 and 2018, the coefficient is large and significantly different from zero. This is not the case in 2016, consistent with the muted first-stage over the one-year window documented in Figure 7. This suggests our IV estimates are not simply an artifact of a small first stage inflating a small reduced-form effect.

TABLE 8: REDUCED-FORM REGRESSION

Investment change ($\Delta_h Y_i$)	Year 2016 (h = 1)	2017 (h = 2)	2018 (h = 3)
Share of eligible bonds (Z_i)	4.92	29.65**	29.88*
	(9.26)	(10.90)	(12.78)
R^2	0.008	0.051	0.036
Obs	149	150	148

Note: The table reports the estimates for the coefficient β from the reduced-form regression in Equation (15). The estimated equation features the same vector of controls as above and an intercept, which are not reported in the table. Statistical significance is indicated by *, **, and *** at the 5%, 1%, and 0.1% level.

Moreover, just-identified IV estimators are median-unbiased under conventional conditions (Angrist and Pischke, 2009). Our estimate is thus as likely to lie above the true value as below it, offering reassurance against systematic bias. Our findings on the reduced-form regression are also in line with Grimm et al. (2021), who document that firms eligible for the ECB's CSPP saw a significant increase in corporate research and Research and Development (R&D) spending relative to comparable ineligible firms.

7 A Calibration Exercise

Our finding that firms' structural sensitivity to borrowing costs is larger than previously measured has first-order implications for calibrating quantitative macro models. For illustration, consider a textbook dynamic investment model in the tradition of Hayashi (1982). Although simple, variants of this dynamic capital-adjustment framework, featuring both convex and nonconvex adjustment frictions, remain common building blocks in larger, more sophisticated models used in recent research and policy-shaping applied work (e.g., Caballero and Engel, 1999; Khan and Thomas, 2008; Bachmann et al., 2013; Lanteri, 2018; Ottonello and Winberry, 2020; Winberry, 2021; Gnewuch and Zhang, 2025).

Under the model, time t is discrete. There is a (representative) firm that owns its capital stock and has productivity θ, which can be stochastic. The firm produces with constant returns to scale using capital K as its only input. In period t, the firm receives revenue $\theta_t K_t$ and chooses investment I_t. The capital stock evolves according to:

$$K_{t+1} = (1-\delta)K_t + I_t \tag{16}$$

where δ is the depreciation rate and investment is non-negative. Capital is subject to a quadratic adjustment cost:

$$C(I_t, K_t) = \frac{\varphi}{2}\left(\frac{I_t}{K_t}\right)^2 K_t \tag{17}$$

where φ is positive but finite and governs the strength of the adjustment friction. The firm-specific interest rate used for discounting is $i_t + r_t$, where i is the common risk-free rate and r is a firm-specific component.

The firm value at time t given its productivity and capital stock equals:

$$V_t(\theta_t, K_t) = \max_{I_t}\left\{\theta_t K_t - I_t - C(I_t, K_t) + \frac{\mathbb{E}_t V_{t+1}(\theta_{t+1}, K_{t+1})}{1 + i_t + r_t}\right\} \tag{18}$$

The first-order condition yields the following Euler Equation of investment:

$$(1+\varphi x_t)(1+i_t+r_t) = \mathbb{E}_t\theta_{t+1} + (1-\delta)\mathbb{E}_t(1+\varphi x_{t+1}) + \frac{\varphi}{2}\mathbb{E}_t(x_{t+1}^2) \tag{19}$$

where $x_t \equiv I_t/K_t$ is the investment rate and $1 + \varphi x_t$ is the price of installing an additional unit of capital today, which is usually defined as q_t in this literature.[12]

A common calibration approach determines the adjustment cost parameter φ by targeting a specific steady-state investment elasticity with respect to the interest rate. Consider the

[12]The expectation operators in the Euler Equation can be replaced with the realized values plus an error term. This yields a testable moment condition: the error term at $t+1$ must be orthogonal to any information available at t, or the equation is rejected by the data. We can extend the dynamic investment model to show how the Euler Equation suggests an alternative test for the information channel, beyond the scope of this paper. We can explicitly introduce an information effect akin to the stylized firm investment model in section 2 to the dynamic model. Then, the expectation of θ, which can be interpreted as a demand shifter, enters the error term and result in an empirical violation of the moment condition from the Euler Equation if the information channel is strong.

model economy in a deterministic steady state with constant productivity, investment, and interest rates. Up to first-order, Equation (19) reduces to

$$(1 + \varphi x)(\delta + i + r) \approx \theta \tag{20}$$

This approximation implies the following semi-elasticity of investment:

$$\varepsilon \equiv \frac{d\log(x)}{d(i+r)} = -\frac{1 + \varphi x}{\varphi x(\delta + i + r)} \tag{21}$$

In calibration, Equation (21) embeds an inverse relationship between the chosen adjustment cost parameter φ and the targeted semi-elasticity ε. As a result, targeting a downward-biased elasticity leads researchers to overestimate the adjustment friction in calibration. Quantitative models calibrated to a lower target elasticity, in turn, underestimate the impact of policies and shocks that do not move beliefs.

For steady states with small adjustment costs, i.e. where φx is close to zero, the ratio of two elasticity targets is approximately inversely proportional to the ratio of their corresponding calibrated adjustment-cost parameters, $\varepsilon_2/\varepsilon_1 = \varphi_1/\varphi_2$. Using our estimates of the investment elasticities in Table 7, the calibration would overstate the firm-level adjustment costs by up to 58 percent. The same holds true when comparing our estimated investment elasticity to the commonly used estimate from Zwick and Mahon (2017), which puts the investment elasticity at only 10.2. More generally, our results suggest that calibrating capital-adjustment frictions to reduced-form interest-rate sensitivities that confound borrowing-cost and information effects can substantially understate firms' structural responsiveness to monetary and financial shocks. The precise quantitative mapping is model-specific, but the direction of the bias applies broadly to quantitative models in which investment adjustment frictions are disciplined by firms' observed responsiveness to the cost of capital.

8 Conclusion

This paper provides firm-level evidence that the investment channel of monetary policy is materially attenuated by an offsetting information effect. Using quasi-experimental variation from the ECB's Corporate Sector Purchase Programme (CSPP), we isolate changes in firms' financing costs that are plausibly orthogonal to contemporaneous belief revisions. The central empirical finding is that the investment elasticity typically inferred from aggregate monetary policy shocks is substantially smaller than the elasticity to a *pure* cost-of-debt shock: over four quarters, the observable aggregate investment response is dampened by 38 percent relative to the borrowing-cost channel alone.

We present a simple model of firm investment under central bank information to interpret these results. Firms invest only in the share of available projects whose expected return exceeds the user cost of capital. If a monetary policy surprise cut simultaneously lowers the cost of capital and makes firms more pessimistic about the future, the shock should, on the one hand, boost investment via cheaper borrowing costs, but, on the other hand, could disincentivize investment as expected project returns fall. In particular, the model implies that the strength of the information effect – and therefore the net potency of monetary policy – depends on the distribution of firm-level credit spreads relative to investment thresholds.

Our findings carry two main implications. First, calibrations of investment elasticities based on standard high-frequency identification of monetary policy shocks may understate the true sensitivity of firm investment to financing conditions, because they embed the net effect of borrowing-cost changes and belief revisions. We show that this can have first-order implications for upward-biased estimates of capital-adjustment frictions in quantitative models of firm investment. Second, changing firm beliefs around monetary policy surprises appear to matter substantially for firms' investment decisions. This means that understanding how different monetary shocks affect firm beliefs is of first-order importance for understanding how these shocks transmit to the real economy. An average monetary policy shock already offsets the investment response by 38 percent. Shocks that affect firm beliefs more than average could have a larger offsetting effect, while shocks that do not materially affect beliefs could be much closer to the structural elasticity. Specifically, the distribution of credit spreads emerges as a central object for how strongly information effects can offset the borrowing-cost channel, suggesting a parsimonious way to incorporate state-dependent transmission into macro-finance models.

Several directions for future work follow naturally. Empirically, richer measurement of expectation revisions at the firm level – through analyst forecasts, earnings guidance, or text-based measures of managerial sentiment – could sharpen the mapping from the residual component of monetary policy shocks to belief updating. Relatedly, extending the decomposition to other real outcomes (employment, R&D, or intangibles) could clarify whether information effects are particularly salient for irreversible, lumpy expenditures. On the policy side, the results motivate closer attention to how central bank communication interacts with unconventional policies that directly affect private financing costs. If real transmission depends on where firms sit in the cross-section of spreads and investment thresholds, then the same policy action may have different real effects across episodes even when aggregate conditions look similar.

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A Comparison of Findings to Other Identification Approaches

Jarociński and Karadi. We replicate the decomposition of monetary policy surprises into a Monetary Policy (MP) shock component and a Central Bank Information (CBI) shock component as proposed by Jarociński and Karadi (2020). We then study what role the choice of a uniform prior over all permissible decompositions plays for the study of firm investment following the robustness check proposed by Giacomini and Kitagawa (2021) and how our results relate to investment semi-elasticities derived from their decomposition.

We take the off-the-shelf series of high frequency monetary policy shocks presented in Jarociński and Karadi (2020). They record surprises in the three-month federal funds rate futures, as well as high frequency changes in the S&P 500 stock market index. The interest rate surprise is denoted by i^{Total} while the stock market surprise is denoted by s. These surprises are then decomposed into a series of interest rate surprises by leveraging the direction of stock price surprises:

$$i^{\text{Total}} = i^{MP} + i^{CBI}, \tag{22}$$

where i^{MP} is negatively correlated with s and i^{CBI} is positively correlated with s. In line with standard theory, we think of the structural component i^{MP} as a monetary policy shock and we think of i^{CBI} as a central bank information effect. Concretely, the decomposition satisfies

$$M = UC, \quad \text{with} \quad U'U = \text{diagonal matrix} \quad \text{and} \quad C = \begin{pmatrix} 1 & c_{MP} < 0 \\ 1 & c_{CBI} > 0 \end{pmatrix} \tag{23}$$

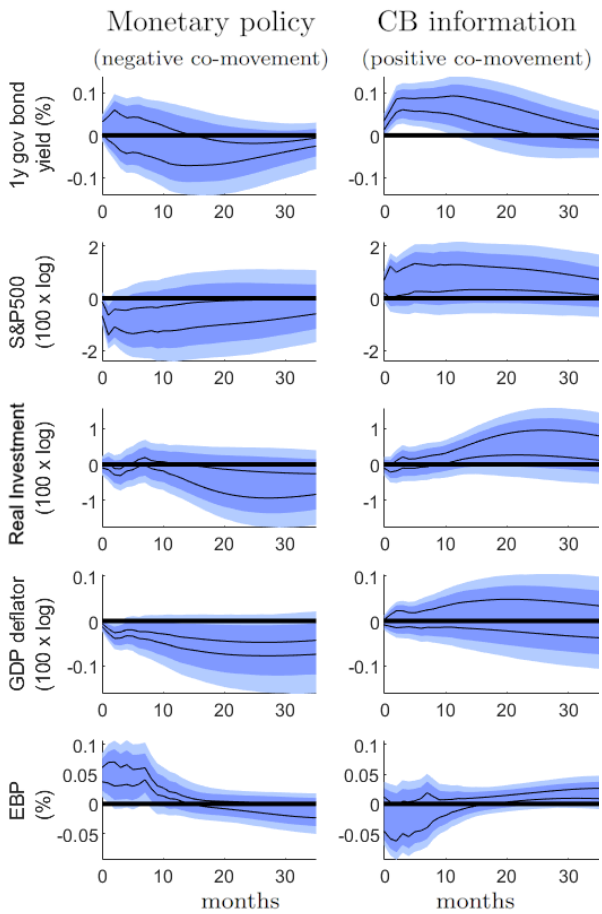
where $M = (i^{\text{total}}, s)$ is a $T \times 2$ matrix with i^{Total} in the first column and s in the second column, $U = (i^{MP}, i^{CBI})$ is a $T \times 2$ matrix with i^{MP} in the first column and i^{CBI} in the second column, T is the number of central bank announcements, i^{MP} and i^{CBI} are mutually orthogonal, and matrix C captures how i^{MP} and i^{CBI} translate into financial market surprises. The 1's in the first column of C reflect equation 22. The second column of C contains the elasticities of stock prices to i^{MP} and i^{CBI}, $c_{MP} < 0$ and $c_{CBI} > 0$.

The decomposition in 23 is not unique. If we rotate U and C with an orthogonal matrix, the new shocks $\tilde{U}$ will still be orthogonal and for many rotation angles the sign restrictions $c_{MP} < 0$ and $c_{CBI} > 0$ will still be satisfied. Smaller rotation angles yield decompositions with less negative c_{MP} and larger c_{CBI}, so monetary policy shocks account for more variation of the interest rates and central bank information accounts for more variation of the stock prices. Larger rotation angles yield decompositions with more negative c_{MP} and smaller c_{CBI}, so monetary policy shocks account for more variation of the stock prices and central bank information account for more variation of the interest rates. To obtain a point estimate, Jarociński and Karadi (2020) compute the median of the admissible rotation angles.

To derive an elasticity of firm investment to a monetary policy shock that affects stock markets as standard theory would suggest, we follow Jarociński and Karadi (2020) in estimating the investment response to a MP shocks in a baseline VAR that includes seven variables: two high frequency variables (surprises to the three-month federal funds futures, and the S&P 500 stock market index), and five low-frequency variables consisting of monthly interest rates, a stock price index, an indicator of real activity, the price level, and financial conditions. We follow Jarociński and Karadi (2020) in their choice of data-sources, but deviate from their setup in that we use real firm investment as a proxy for the real economy instead of real GDP to study the transmission of monetary policy shocks to firm investment. We use the monthly series of total private, non-residential construction spending from the FRED database, which is the only time series proxying firm investment available at the monthly level. The VAR has 12 lags, and the sample period is from January 1993 until December 2016. Firm investment at the monthly level is unavailable before January 1993. Results are reported based on 2,000 draws from the Gibbs sampler akin to Jarociński and Karadi (2020).

Figure 9 presents the impulse response functions to the monetary policy and the central bank information shock. A monetary policy shock is associated with a 4.5 points increase of the three months federal funds futures rate and a 37.5 basis point drop in the S&P 500 index in the 30 minute window. Focusing on the response to firm investment, we see that a contractionary monetary policy surprise appears to lower firm investment, while a contractionary central bank information shock appears to raise output.

FIGURE 9: IMPULSE RESPONSES WITH 'ROBUST ERROR BANDS'



Note: The figure plots the impulse response functions of the baseline VAR with 'robust'error bands of Giacomini and Kitagawa (2021). Posterior mean bounds (lines), 68% robustified region (darker band), and 90% robustified region (lighter band).

When accounting for the set-identification of the sign restriction by constructing 'robust' error bands of Giacomini and Kitagawa (2021), we find that while the set identified responses appear negative for a contractionary monetary policy shock and positive for a contractionary central bank information shock, depending on the prior chosen, the investment response is not significant over any time horizon. The set identification is also wide, with semi-elasticities of investment ranging from 0 to 7.17 after twelve months. This wide band of set-identified elasticities makes it hard to comment on the quantitative importance of the information channel for investment. However, estimates of investment semi-elasticities are consistent with our findings. Over twelve months, a semi-elasticity of 19 percentage points falls within the 90% robustious region of the confidence bands for the monetary policy shock. Since the central bank information shock exhibits an investment semi-elasticity of the opposite sign, this also supports our finding that the average monetary policy shock which consists of a weighted average between the structural monetary policy shock and the central bank information shock is partially offset by central bank information and transmits more weakly than a monetary policy shock. Due to the non-uniqueness of the structural shock decomposition and the wide range of impulse responses that are set-identified, commenting on the exact magnitude remains challenging.

Miranda-Agrippino and Ricco. We also replicate the decomposition by Miranda-Agrippino and Ricco (2021). Their starting point is the high-frequency surprise in the fourth federal funds futures contract (*FF4*), measured as the change in the implied interest rate within a 30-minute window around FOMC announcements. Event-level surprises are aggregated to monthly frequency by summing all surprises occurring within a given month. To remove predictable variation and the component of the surprise reflecting the Federal Reserve's private information about the macroeconomic outlook, the monthly *FF4* surprise is residualized with respect to its own lags and Greenbook forecasts and forecast revisions for output growth, inflation, and unemployment. Specifically, they estimate

$$FF4_t = \alpha + \sum_{j=1}^{p} \beta_j FF4_{t-j} + \Gamma' GB_t + \Delta' \Delta GB_t + \varepsilon_t, \tag{24}$$

where GB_t denotes Greenbook forecast levels and ΔGB_t denotes Greenbook forecast revisions. The residual $\widehat{\varepsilon}_t$ is orthogonal to both past high-frequency surprises and the Federal Reserve's internal macroeconomic forecasts. Miranda-Agrippino and Ricco (2021) use this residual as an external instrument for the structural monetary policy shock.

Figure 10a shows the impulse response of the investment semi-elasticity to a one percentage point surprise loosening of monetary policy. Our measure for firm investment is aggregate total private nonresidential construction spending in the United States. We report the impulse response function for investment semi-elasticities to a one percentage point drop in the fourth federal funds future contract (dashed), as well as to a one percentage point drop in the orthogonalized shock series by Miranda-Agrippino and Ricco (2021) (solid). We find that the investment response to a non-orthogonalized shock leads to a counterintuitive investment response: a loosening leads to a drop in investment. When studying the orthogonalized series, this effect disappears, and a change in monetary policy has no real effects. Confidence bands remain wide though, and our positive estimate for the semi-elasticity of investment to a change in borrowing costs of 19 percentage points lies in the 90 percent confidence band after twelve months. The difference between the impulse response of the orthogonalized shock series and the non-orthogonalized shock series supports our finding for information offsetting of the investment channel of monetary policy at the firm level. However, the residualization relies on reliable observable proxies of central bank information and does not account for the possibility of odyssian central bank information. This may not purge the entirety of the information effect.

Bauer and Swanson. Finally, we also use the orthogonalized high-frequency monetary policy surprise constructed by Bauer and Swanson (2023). Their baseline surprise is measured using changes in the first four quarterly Eurodollar futures rates (*ED1-ED4*) within a 30-minute window beginning 10 minutes before and ending 20 minutes after an FOMC announcement. The raw monetary policy surprise, denoted by MPS_t, is the first principal component of these four rate changes, rescaled such that a one-unit increase corresponds to a one-percentage-point increase in the *ED4* rate. The measure therefore captures news about both the current policy rate and the expected future path of monetary policy. Bauer and Swanson (2023) show that the raw surprise is systematically correlated with macroeconomic and financial information available before the FOMC announcement. They therefore estimate the event-level regression

$$MPS_t = \alpha + \boldsymbol{\beta}'\mathbf{X}_{t^-} + u_t, \tag{25}$$

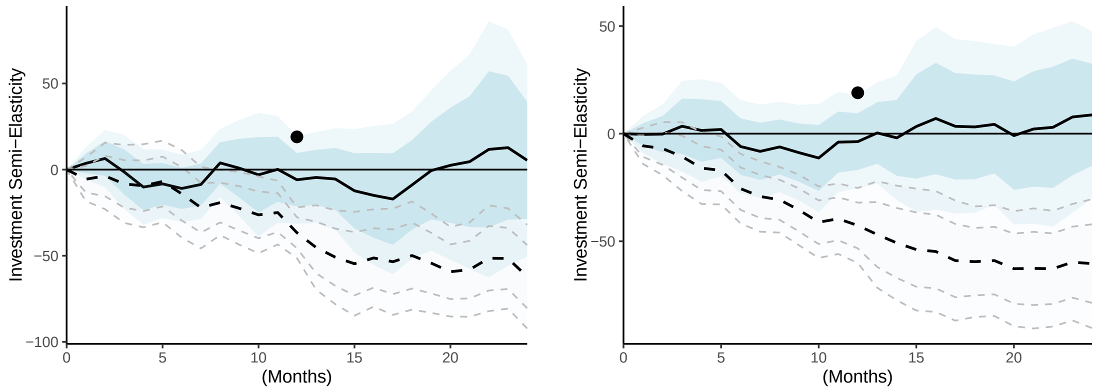
where $\mathbf{X}_{t^-}$ contains six variables observed prior to the announcement: the most recent nonfarm-payroll surprise, year-on-year employment growth, the three-month change in the log S&P 500 index, the three-month change in the slope of the Treasury yield curve, the three-month change in the log Bloomberg Commodity Spot Price Index, and the option-implied skewness of the 10-year Treasury yield. The orthogonalized monetary policy surprise is the estimated residual

$$MPS_t^{\perp} = MPS_t - \widehat{\alpha} - \widehat{\boldsymbol{\beta}}' \mathbf{X}_{t^-}. \tag{26}$$

By construction, $MPS_t^{\perp}$ is orthogonal to these pre-announcement macroeconomic and financial variables. Bauer and Swanson argue that this adjustment removes variation arising from the Federal Reserve's systematic response to recent economic news and makes the resulting series more suitable as an external instrument for structural monetary policy shocks.

Figure 10b similarly shows the impulse response of the investment semi-elasticity to a one percentage point surprise loosening of monetary policy. We report the impulse response function for investment semi-elasticities to a one percentage point drop in the first principal

FIGURE 10: INVESTMENT RESPONSE TO ORTHOGONALIZED HF-MP SHOCKS



(A) Miranda-Agrippino and Ricco (2021) (B) Bauer and Swanson (2023)

Note: Both figures plot the investment semi-elasticity for aggregate total private nonresidential construction spending in the United States to a one percentage point surprise loosening. The figures compares the investment response to orthogonalized high frequency monetyary policy shocks (solid) against the non-orthogonality shocks (dashed). The left figure reports the investment response to the Miranda-Agrippino and Ricco (2021) shock series, while the right reports the impulse response for the Bauer and Swanson (2023) shock series. Shaded areas and dashed gray lines reflect 68 and 90 percent confidence bands. Estimates are based on an OLS local projections with Newey–West standard errors, using horizon-dependent bandwidth. The point at 12 months and a semi-elasticity of 19 percent corresponds to the finding of our paper for the structural semi-elasticity of investment to a change in borrowing cost.

component of these four rate changes, rescaled such that a one-unit increase corresponds to a one-percentage-point increase in the *ED4* rate (solid). We also report the impulse response of investment to a non-orthogonalized surprise loosening in the ED4 rate (dashed). Results are similar to those for our replication of Miranda-Agrippino and Ricco (2021). This similarly supports our finding for information offsetting of the investment channel of monetary policy at the firm level. Also similar to the structural investment elasticities from the monetary shocks from Jarociński and Karadi (2020) and Miranda-Agrippino and Ricco (2021), our estimate lies within the confidence bands of the investment elasticity from the orthogoinalized shocks.

B Supplementary Materials

This section provides robustness checks for some of the empirical excercises described in the main body of the paper. First, table 9 explores the wedge between eligible and ineligible bonds after the CSPP announcement over a range of maturity-rating buckets. It does so for bond yields, as well as for yield spreads. Note that "AA", "A", and "BBB" refer to all subratings in these categories including "AA+", "AA-", "A+", "A-", "BBB+", "BBB+".

Panel A provides estimates for equation (7) for corporate bond yields as the dependent variable across specific maturity-rating segments of our bond sample. Panel B similarly provides estimates of equation (7) for corporate government-yield spreads as the dependent variable. Regressions are estimated based on eligible bonds as well as a varying group of ineligible bonds. Columns indicated by ALL are based on all eligible and all ineligible bonds, while columns indexed DUAL are based on eligible bonds and dual-ineligible bonds only. We report the coefficient of $Treat \times Post$.

It highlights that the CSPP announcement has driven a wedge between the bond yields of eligible bonds and ineligible bonds. This is particularly pronounced for eligible bonds vis-a-vis dual-ineligible bonds. Furthermore, much of this wedge can be explained by the reaction in the yield spreads.

Furthermore, table 10 and 11 show robustness checks for our estimate of investment elast-

TABLE 9: IMPACT OF CSPP ANNOUNCEMENT BY RATING-MATURITY SEGMENT

Panel A: Yields

	AA		A		BBB	
	ALL	DUAL	ALL	DUAL	ALL	DUAL
6 months - 2 years	NA	NA	-0.052**	-0.067*	0.03	0.008
	NA	NA	(0.026)	(0.040)	(0.085)	(0.085)
2 - 5 years	-0.136***	-0.136***	0.000	-0.059	-0.154**	-0.280***
	(0.025)	(0.025)	(0.025)	(0.036)	(0.070)	(0.053)
5 - 10 years	-0.107**	-0.167***	-0.077***	-0.156***	-0.265***	-0.337***
	(0.048)	(0.056)	(0.023)	(0.029)	(0.056)	(0.051)
10+ years	-0.148***	-0.148***	0.015	-0.241***	-0.235***	-0.264***
	(0.041)	(0.041)	(0.209)	(0.030)	(0.066)	(0.067)

Panel B: Yield Spreads

	AA		A		BBB	
	ALL	DUAL	ALL	DUAL	ALL	DUAL
6 months - 2 years	NA	NA	0.023	0.058*	0.088	0.088
	NA	NA	(0.025)	(0.034)	(0.085)	(0.085)
2 - 5 years	-0.015	-0.015	0.008	-0.035	-0.141**	-0.268***
	(0.023)	(0.023)	(0.022)	(0.030)	(0.070)	(0.052)
5 - 10 years	-0.058	-0.092*	-0.030	-0.078***	-0.231***	-0.286***
	(0.040)	(0.051)	(0.020)	(0.024)	(0.056)	(0.050)
10+ years	-0.005***	-0.005***	0.146	-0.088***	-0.090	-0.110
	(0.036)	(0.036)	(0.208)	(0.030)	(0.067)	(0.068)

Note: Panel A provides estimates for equation (7) for corporate bond yields as the dependent variable across specific maturity-rating segments of our bond sample. Panel B similarly provides estimates of equation (7) for corporate government-yield spreads as the dependent variable. Regressions are estimated based on eligible bonds as well as a varying group of ineligible bonds. Columns indicated by ALL are based on all eligible and all ineligible bonds, while columns indexed DUAL are based on eligible bonds and dual-ineligible bonds only. We report the coefficient of $Treat \times Post$. Standard errors as shown in parentheses are heteroskedasticity-robust and double-clustered by bond and time. Statistical significance is indicated by *, **, and *** at the 5%, 1%, and 0.1% level. Units for bond yields and government-yield spreads are in basis points. Regressions on fewer than 1,000 bond-day observations are excluded from the analysis.

icities. Table 10 shows investment elasticities for different investment measures and over different time windows. Similar to total asset growth, total capital expenditure likewise exhibits a positive investment elasticity to our instrumented reduction in borrowing costs. Tangible fixed asset growth also exhibits a positive such elasticity, however it is not statistically significant at the ten percent level. Furthermore, we investigate investment elasticities over different time windows. For all investment measures, the two year window between 2015 and 2017 appears to be the sweet spot. In part, because price dynamics in secondary financial markets

TABLE 10: ROBUSTNESS TO DIFFERENT INVESTMENT MEASURES

	$\Delta_1 Y_i$		$\Delta_2 Y_i$		$\Delta_3 Y_i$	
	2SLS	Fuller	2SLS	Fuller	2SLS	Fuller
Total Asset Growth	-0.26	0.05	0.98**	0.83**	3.67	0.57
	(0.75)	(0.33)	(0.47)	(0.38)	(8.83)	(0.77)
Tangible Fixed Asset Growth	0.26	0.12	1.27	1.09	4.29	0.84
	(0.43)	(0.22)	(0.80)	(0.69)	(9.29)	(1.09)
Total Capital Expenditure	-0.25	0.07	0.36*	0.31**	0.46	0.09
	(0.45)	(0.14)	(0.19)	(0.15)	(1.21)	(0.13)
First Stage F	0.43		5.05		0.17	
N (Firms)	138		138		140	

NOTE: The table reports the Two-Stage Least Squares (2SLS) estimates for δ from equation 10 for three measures of investment (rows) and three different time-windows (columns). We estimate the percentage change of investment for a one percentage point drop in borrowing costs. In addition to the 2SLS estimate, we also report the Fuller estimates which are more robust to weak instruments. To further account for the possibility of a weak instrument, we compute Conditional Likelihood Ratio (CLR)–based confidence intervals for both Two-Stage Least Squares (2SLS) and Fuller estimators. This approach allows us to mitigate small-sample bias and obtain valid inference even when instrument strength is limited. Statistical significance is indicated by *, **, and *** at the 5%, 1% and 0.1% level.

TABLE 11: ROBUSTNESS TO DIFFERENT CONTROLS

	$\Delta_2 Y_i$						
	Bond Reliance	Size	Leverage	Liquidity	Prof-itability	Procycl. EA	All
Total Asset Growth	0.74*	0.85**	0.84**	0.84**	0.72**	0.85**	0.61
	(0.37)	(0.43)	(0.37)	(0.40)	(0.34)	(0.39)	(0.37)
First Stage F	4.057	3.895	5.191	4.482	5.086	4.848	3.512
N (Firms)	137	137	137	136	134	130	127

Note: The table reports Fuller estimates from equation (10) for investment elasticities to a loosening of monetary policy. Every column reflects the inclusion of size, leverage, liquidity, profitability, procyclicality with the Euro Area, and all of the aforementioned characteristics as controls in the first and second stage of the 2SLS.

have had enough time to affect actual firm borrowing costs and firm investment decisions, unlike for the one-year elasticity. On the other hand, the window appears to be sufficiently short for there not to be too much noise, a weakening of the CSPP induced wedge over time due to financial markets overcoming market segmentation gradually, or firms readjusting their debt portfolios.

Table 11 reports Fuller estimates from equation (10) for investment elasticities to a loosening of monetary policy. Every column reflects the inclusion of size, leverage, liquidity, profitability, procyclicality with the Euro Area, and all of the aforementioned characteristics as controls in the first and second stage of the 2SLS.

The excercise shows that results are robust to pairwise inclusion of various firm characteristics. When including all characteristics, the sample size of available firms drops substantially relative to our main specification.

C A Model of Market Segmentation

In section 4, we established that the CSPP reduced the yields of eligible bonds primarily by lowering government-yield spreads. The degree to which this effect spilled over to different market segments of ineligible bonds, however, differed. We argue that this can be explained by different degrees of market segmentation between eligible bonds and the different segments of ineligible corporate bonds. To test this hypothesis, in this section, we build on Greenwood et al. (2018), to develop a model of a segmented corporate bond market. For different degrees of market segmentation and under a reasonable model calibration, we can indeed replicate the empirical findings presented in section 4 and account for the heterogeneity observed in the yield response across different market segments.

The model presented in this section is a stylized model for pricing two long-term defaultable bonds A and B that reflect bonds eligible under the CSPP, as well as some control groups of ineligible bonds. We extend the model from Greenwood et al. (2018) by allowing for default risk in both bond market segments to study market segmentation within the corporate bond market, instead of studying market segmentation between sovereign and corporate bond markets, which is the focus of Greenwood et al. (2018). The model features two key asset pricing frictions: partial market segmentation and slow-moving capital across markets. Bond Supply in our model is stochastic and we model the announcement of the CSPP as a shock to the supply of eligible bonds A. This allows us to analyze the counterfactual effect of a drop in the active bond supply on the yield spread between bonds under different degrees of market segmentation.

C.1 Bonds

We consider two types of bonds in our model: A and B. Both bonds are exposed to interest risk and bond-idiosyncratic default risk. We will think of bond A as bonds eligible under the CSPP, while we will think of bond B as bonds ineligible for the CSPP.

For $i \in \{A, B\}$, the i perpetuities pay a coupon C_i each period, but a random fraction $h_{i,t+1}$ of the bonds default at $t+1$ and are worth $(1 - L_{i,t-1})(P_{L,t+1} + C_i)$, where $0 \leq 1 - L_{i,t+1} \leq 1$ is the recovery rate of the bond. As a consequence, in the next period, $Z_{i,t+1} \equiv h_{i,t+1} L_{i,t+1}$ of the bond value is lost due to default. The return of bond i is

$$1 + R_{i,t+1} = \frac{(1 - Z_{t+1})(P_{i,t+1} + C_i)}{P_{i,t}} \tag{27}$$

Using the Campbell and Shiller (1988) log-linear approximation, we find that the excess return on the i perpetuity can be written as

$$rx_{i,t+1} \equiv ln(1 + R_{i,t+1}) - r_t \approx \frac{1}{1-\theta_i} y_{i,t} - \frac{\theta_i}{1-\theta_i} y_{i,t+1} - z_{i,t+1} - r_t \tag{28}$$

where $\theta_i \equiv 1/(1 + C_A) < 1$, and $y_{i,t}$ is the log yield-to-maturity on bond i at time t, and $z_{i,t} \equiv -ln(1 - Z_t)$.

C.2 Risk Factors

The model incorporates three different sources of risk: (i) short-term interest risk r_t, (ii) default risks $z_{i,t}$, and (iii) supply risks $s_{i,t}$, which all evolve randomly according to an AR(1) process as

$$r_{t+1} = \bar{r} + \rho_r (r_t - \bar{r}) + \varepsilon_{r,t+1} \tag{29}$$

$$z_{i,t+1} = \bar{z}_i + \rho_{z_i} (z_{i,t} - \bar{z}_i) + \varepsilon_{z_i,t+1} \tag{30}$$

$$s_{i,t+1} = \bar{s}_i + \rho_{s_i} (s_{i,t} - \bar{s}_i) + \varepsilon_{s_i,t+1}, \tag{31}$$

where $Var(\varepsilon_{r,t+1}) = \sigma_r^2$, $Var(\varepsilon_{z_i,t+1}) = \sigma_{z_i}^2$, $Var(\varepsilon_{s_i,t+1}) = \sigma_{s_i}^2$. This means that both bonds are in an exogenous time-varying supply, governed by independent shocks. At the same time, the log short-term interest rate and the default processes similarly follow an exogenous AR(1) process. At this point we do not assume the correlation structure of the underlying shocks $\varepsilon_{r,t}, \varepsilon_{z_i,t+1}, \varepsilon_{s_i,t+1}$, yet for simplicity, in the numerical illustration, we will solve for underlying orthogonal shocks.

C.3 Market Structure

The model features three types of investors: (i) A specialists, (ii) B specialists, and (iii) Generalists. A specialists can only trade bond A, while B specialists can only trade bone B. Generalists can trade both bonds, but can only adjust their portfolio every k periods due to the financial friction in the model.

For $i \in \{A, B\}$, there exists a measure q_i of i specialists, whose demand for bond i in period t is denoted by $b_{i,t}$. Fast moving i specialists have mean-variance preferences over one-period portfolio log returns

$$b_{i,t} = \tau \frac{E_t\left[rx_{i,t+1}\right]}{\mathrm{Var}_t\left[rx_{i,t+1}\right]}. \tag{32}$$

There also is a measure of $1 - q_A - q_B$ slow-moving generalist investors who can adjust their holdings of both A and B bonds, but only a fraction $1/k$ of them is active each period. When they are active, they need to maintain their portfolio for the next k periods. As a result, they have mean-variance preferences over their k period cumulative excess return. Defining $rx_{i,t\to t+k} \equiv \sum_{j=1}^{k} rx_{i,t+j}$ the excess return on a generalist's portfolio is

$$rx_{d_t,t\to t+k} \equiv d_{A,t} \times rx_{A,t\to t+k} + d_{B,t} \times rx_{B,t\to t+k}, \tag{33}$$

where $d_{i,t}$ denote the generalist's demand for bond i at time t. Generalists now choose their bond demand such that

$$\max_{d_{A,t},d_{B,t}} \left\{E_t\left[rx_{d_t,t\to t+k}\right] - (2\tau)^{-1}\left(\mathrm{Var}_t\left[rx_{d_t,t\to t+k}\right]\right)\right\}, \tag{34}$$

which implies that

$$\begin{bmatrix} d_{A,t} \\ d_{B,t} \end{bmatrix} = \tau \begin{bmatrix} \mathrm{Var}_t\left[rx_{A,t\to t+k}\right] & \mathrm{Cov}_t\left[rx_{A,t\to t+k}, rx_{B,t\to t+k}\right] \\ \mathrm{Cov}_t\left[rx_{A,t\to t+k}, rx_{B,t\to t+k}\right] & \mathrm{Var}_t\left[rx_{B,t\to t+k}\right] \end{bmatrix}^{-1}$$

$$\times \begin{bmatrix} E_t\left[rx_{A,t\to t+k}\right] \\ E_t\left[rx_{B,t\to t+k}\right] \end{bmatrix}. \tag{35}$$

C.4 Market Clearing

In equilibrium of the model, markets for both bonds must clear in a way that is consistent with the optimization of specialists A and B, generalists, and the law of motion for all exogenous and endogenous variables. Since the measure $q_G \equiv 1 - q_A - q_B$ of generalists can only adjust their portfolio every k periods, the active supply of bonds equals the total supply of bonds net what generalists have purchased over the past $k-1$ periods, which is now being kept fixed. The market clearing condition for bond $i \in \{A, B\}$ is

$$\underbrace{q_i b_{i,t}}_{\text{Specialist demand}} + \underbrace{q_G k^{-1} d_{i,t}}_{\text{Active generalist demand}} = \underbrace{s_{i,t}}_{\text{Total bond supply}} - \underbrace{(1 - q_A - q_B)\left(k^{-1}\sum_{i=1}^{k-1} d_{i,t-i}\right)}_{\text{Inactive generalist holdings}}. \tag{36}$$

Notably, in the model, the parameters k and q_G are the exogenous parameters that determine the degree of market segmentation in the model. They in down how many traders there are that operate in both market segments and at what speed they can adjust their portfolio.

C.5 Equilibrium Conjecture

Similar to Greenwood et al. (2018), we solve for a rational expectations equilibrium in which equilibrium yields and generalist demands are linear functions of a state vector x_t which includes steady state deviations of the short term interest rate, default realizations, the supply of A bonds, the supply of B bonds, and inactive generalist holdings of both bonds. Concretely, we assume that

$$y_{A,t} = \alpha_{A0} + \boldsymbol{\alpha}'_{A1}\boldsymbol{x}_t \tag{37}$$

$$y_{B,t} = \alpha_{B0} + \boldsymbol{\alpha}'_{B1}\boldsymbol{x}_t \tag{38}$$

$$d_{A,t} = \delta_{A0} + \boldsymbol{\delta}'_{A1}\boldsymbol{x}_t \tag{39}$$

$$d_{A,t} = \delta_{B0} + \boldsymbol{\delta}'_{B1}\boldsymbol{x}_t \tag{40}$$

where the $(5+2(k-1)) \times 1$ dimensional vector $\boldsymbol{x}_t$ is given by

$$\begin{aligned} \boldsymbol{x}_t = [&r_t - \bar{r}, z_{A,t} - \bar{z}_A, z_{B,t} - \bar{z}_B, s_{A,t} - \bar{s}_A, s_{B,t} - \bar{s}_B, \\ &d_{A,t-1} - \delta_{A0}, \ldots, d_{A,t-k} - \delta_{A0}, d_{B,t-1} - \delta_{B0}, \ldots, d_{B,t-k} - \delta_{B0}] \end{aligned} \tag{41}$$

which implies that the state vector $\boldsymbol{x}_t$ follows an AR(1) process as well:

$$\boldsymbol{x}_{t+1} = \boldsymbol{\Gamma}\boldsymbol{x}_t + \boldsymbol{\varepsilon}_{t+1} \tag{42}$$

where $\varepsilon_{t+1} \equiv [\varepsilon_{r,t+1}, \varepsilon_{z_A,t+1}, \varepsilon_{z_B,t+1}, \varepsilon_{s_A,t+1}, \varepsilon_{s_B,t+1}, 0, \ldots, 0]$, $Var(\boldsymbol{\varepsilon}_{t+1}) \equiv \boldsymbol{\Sigma}$, and the transition matrix $\boldsymbol{\Gamma}$ depends on generalist demand functions (δ_{A1}, and δ_{B1}) as described in detail in the appendix (equations 48 and 49).

Similar to insights from Greenwood et al. (2018) and as shown in appendix section D, a rational expectations equilibrium of the model is a fixed point of a specific operator which involves the price-impact coefficients $(\alpha_{A1}, \alpha_{B1})$ as well as generalists' demand impact coefficients $\delta_{A1}, \delta_{B1})$. Specifically, we let $\omega = (\alpha'_{A1}, \alpha'_{B1}, \delta'_{A1}, \delta'_{B1})$ and derive the operator $f(\omega_0)$ that gives the new price impact coefficients and demand impact coefficients that clear both bond markets when agents expect ω_0 to hold at all future dates. Thus we find our equilibrium as a fixed point of f where $f(\omega^*) = \omega^*$.

We follow Greenwood et al. (2018) in applying the Powell hynrid algorithm to solve for the fix point, which performs a quasi-Newton search for the fix-point from an initial guess. We account for possible equilibria-multiplicity by sampling over 10,000 initial guesses. While multiple equilibria can arise, we focus on the study of the unique stable equilibrium, defined by having all eigenvalues of the Jacobian $D_\omega f(\omega^*)$ less than one in magnitude.

C.6 An illustrative Comparison

After having set up the model and outlined how to solve for general equilibrium, in this section, we will model a change in the supply of bond A and how this affects bond yields in both market segments under given degree of market segmentation, as captured by a varying q_G. We do so to compare the heterogeneous effect of the CSPP on yield-spreads between eligible bonds and ineligible bonds from different market segments to the yield-spread dynamics arising from our model under different degrees of market segmentation. We summarize illustrative parameters for a numerical calibration of the model in Table 12.

We model the CSPP announcement as a supply contraction in A bonds. We match the size of the supply contraction to the observed total purchases under the CSPP as a share of outstanding eligible bonds - a contraction by $1/3$ of the market capitalization. We then derive

TABLE 12: ILLUSTRATIVE MODEL PARAMETERS FOR NUMERICAL CALIBRATION

Parameters	Description	Value
$\bar{r}$	Average annualized short-term interest rate	4%
σ_r	Volatility of quarterly shocks to annualized short-term riskless rate	0.64%
ρ_r	Quarterly persistence of short-term riskless rate	0.9
$\bar{z}_A, \bar{z}_B$	Expected annualized default losses	0.4%
$\sigma_{z_A}, \sigma_{z_B}$	Volatility of quarterly shocks to annualized default losses	0.4%
ρ_{z_A}, ρ_{z_B}	Quarterly persistence of default losses	0.96
$\bar{s}_A$, $\bar{s}_B$	Average asset supplies	5
σ_{s_A}, σ_{s_B}	Volatility of quarterly supply shocks	1
ρ_{s_A}, ρ_{s_B}	Quarterly persistence of supply shock	0.999
D_A, D_B	Macaulay duration in years (implies $\theta_A = \theta_B = 0.95$)	5 years
τ	Aggregate investor risk tolerance	1.75

NOTE: This table presents the illustrative model parameters that we use throughout our numerical exercises. One period corresponds to one week. We report annualized values for the mean and standard deviation of shocks to both the short rate and default losses.

model estimates for bond yields over time for both A and B bonds. To match our empirical findings on the transitive effect on eligible yields relative to different market segments of ineligible yields, we then take the difference $y_A - y_B$ and choose the degree of market segmentation in the model, such that we minimize the distance between our empirical estimates from Figure 5 and our model prediction $y_A - y_B$. To this end, we keep k fixed at 8 in line with Greenwood et al. (2018), but vary the measure of generalist traders q_G.

Figure 5 plots the empirical estimates of the transitive impact of the CSPP (solid lines) against our model estimates under varying degrees of market segmentation (dashed line). We find that we can match our empirical findings for the wedge between eligible and inco-ineligible bonds most closely under no market segmentation when $q_G = 1$. Likewise, we can match our empirical findings for the wedge between eligible and dual-ineligible bonds most closely by complete market segmentation ($q_G = 0$). For the wedge between eligible and currency-ineligible bonds, we find that a moderate degree of market segmentation ($q_G = 0.4$) replicates empirical findings most closely. This indicates, that dual-ineligible bonds are plausibly the most segmented ineligible bonds that experience the least amount of spillover and provide the best counterfactual in our empirical analysis from section 4. It further indicates some persistence of the wedge in borrowing costs on secondary markets following the CSPP, at least between eligible and dual-ineligible bonds that are not traded away by arbitrageurs instantaneously.

D Solution to the Baseline Segmented Market Model

In line with the solution method, we conjecture that the equilibrium yields in both markets at time t take the form

$$y_{A,t} = \alpha_{A0} + \boldsymbol{\alpha}'_{A1}\mathbf{x}_t, \tag{43}$$

$$y_{B,t} = \alpha_{B0} + \boldsymbol{\alpha}'_{B1}\mathbf{x}_t, \tag{44}$$

while generalists' demands for bonds A and B take the form

$$d_{A,t} = \delta_{A0} + \delta_{A1}' \mathrm{x}_t, \tag{45}$$

$$d_{B,t} = \delta_{B0} + \delta_{B1}' \mathrm{x}_t. \tag{46}$$

For concreteness and illustrative purposes, let us fix $k = 3$. To keep track of the inactive bond holdings of generalist traders, the state vector must include the past two lags of $d_{A,t}$ and $d_{B,t}$, as well as the fundamentals. In general x_t is the $2(k-1)+5$ dimensional state vector.

$$x_t = \begin{bmatrix} r_t - \bar{r} \\ z_{A,t} - \bar{z}_A \\ z_{B,t} - \bar{z}_B \\ s_{A,t} - \bar{s}_A \\ s_{B,t} - \bar{s}_B \\ d_{A,t-1} - \delta_{A0} \\ d_{A,t-2} - \delta_{A0} \\ d_{B,t-1} - \delta_{B0} \\ d_{B,t-2} - \delta_{B0} \end{bmatrix} \tag{47}$$

Based on the model assumptions as well as the conjecture for the generalists' demand, we can define the transition matrix Γ for the state vector as follows:

$$
\begin{aligned}
\mathbf{x}_{t+1} &= \Gamma(\delta)\mathbf{x}_t + \boldsymbol{\epsilon}_{t+1} \qquad (48) \\
&= \begin{bmatrix}
\rho_r & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & \rho_{z_A} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & \rho_{z_B} & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & \rho_{s_A} & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & \rho_{s_B} & 0 & 0 & 0 & 0 \\
\delta_{A1} & \delta_{A2} & \delta_{A3} & \delta_{A4} & \delta_{A5} & \delta_{A6} & \delta_{A7} & \delta_{A8} & \delta_{A9} \\
0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\
\delta_{B1} & \delta_{B2} & \delta_{B3} & \delta_{B4} & \delta_{B5} & \delta_{B6} & \delta_{B7} & \delta_{B8} & \delta_{B9} \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0
\end{bmatrix}
\begin{bmatrix}
r_t - \bar{r} \\
z_{A,t} - \bar{z}_A \\
z_{B,t} - \bar{z}_B \\
s_{A,t} - \bar{s}_A \\
s_{B,t} - \bar{s}_B \\
d_{A,t-1} - \delta_{A0} \\
d_{A,t-2} - \delta_{A0} \\
d_{B,t-1} - \delta_{B0} \\
d_{B,t-2} - \delta_{B0}
\end{bmatrix}
+
\begin{bmatrix}
\varepsilon_{r,t+1} \\
\varepsilon_{z_A,t+1} \\
\varepsilon_{z_B,t+1} \\
\varepsilon_{s_A,t+1} \\
\varepsilon_{s_B,t+1} \\
0 \\
0 \\
0 \\
0
\end{bmatrix}
\end{aligned}
$$

Under the assumption that the error terms are mutually orthogonal, we have

$$
\boldsymbol{\Sigma} \equiv \mathrm{Var}_t\left[\epsilon_{t+1}\right] =
\begin{bmatrix}
\sigma_r^2 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & \sigma_{z_A}^2 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & \sigma_{z_B}^2 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & \sigma_{s_A}^2 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & \sigma_{s_B}^2 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0
\end{bmatrix}. \qquad (49)
$$

We further adopt the following notation: Let $\mathbf{e}_r$ be the basis vector with a 1 corresponding to the first $r_t - \bar{r}$ column and zeros otherwise, let $\mathbf{e}_{z_A}$ be the basis vector with a 1 at the second row and zeros otherwise, and let $\mathbf{e}_{z_B}$ be the basis vector with a one in the third row and zeros otherwise. Finally, denote $C_{[t+i,t+j]} \equiv Cov(x_{t+i}, x_{t+j}|x_t)$ and note that

$$\mathbf{C}_{[t+i,t+j]} = \sum_{s=1}^{\min\{i,j\}} \left[\boldsymbol{\Gamma}^{i-s}\right] \boldsymbol{\Sigma} \left[\boldsymbol{\Gamma}^{j-s}\right]', \tag{50}$$

so $C_{[t+i,t+j]} = C'_{[t+i,t+j]}$.

D.1 Specialist Demand

Given the conjectured forms of equilibrium yields from (43) and (44), we have

$$\begin{aligned} rx_{i,t+1} &= \frac{1}{1-\theta_i} y_{i,t} - \frac{\theta_i}{1-\theta_i} y_{i,t+1} - z_{i,t+1} - r_t \\ &= (\alpha_{i0} - \bar{r} - \bar{z}_i) + \left(\frac{1}{1-\theta_i}\boldsymbol{\alpha}_{i1} - \mathbf{e}_r\right)' \mathbf{x}_t - \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)' \mathbf{x}_{t+1} \end{aligned} \tag{51}$$

From this, we can find the expectation of excess return as well as its variance

$$E_t\left[rx_{i,t+1}\right] = (\alpha_{i0} - \bar{r} - \bar{z}_i) + \left(\frac{1}{1-\theta_i}\boldsymbol{\alpha}_{i1} - \mathbf{e}_r\right)' \mathbf{x}_t - \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)' \boldsymbol{\Gamma}\mathbf{x}_t \tag{52}$$

$$\mathrm{Var}_t\left[rx_{i,t+1}\right] = \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)' \boldsymbol{\Sigma} \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right) \tag{53}$$

Due to the mean-variance preferences of the specialist agent, demand for asset A under the conjecture for equilibrium yields is

$$\begin{aligned} b_{i,t} &= \tau \frac{E_t\left[rx_{i,t+1}\right]}{\mathrm{Var}_t\left[rx_{i,t+1}\right]} \\ &= \left[\tau \frac{\alpha_{i0} - \bar{r} - \bar{z}_i}{\left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)' \boldsymbol{\Sigma} \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)}\right] + \left[\tau \frac{\left(\frac{1}{1-\theta_i}\boldsymbol{\alpha}_{i1} - \mathbf{e}_r\right)' - \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)' \boldsymbol{\Gamma}}{\left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)' \boldsymbol{\Sigma} \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)}\right] \mathbf{x}_t \end{aligned} \tag{54}$$

D.2 Generalists' Demand

Generalists who can trade in both bonds, but are only allowed to adjust their portfolio every k periods solve

$$\max_{d_{A,t},d_{B,t}} \left\{ \begin{array}{c} d_{A,t}E_t\left[\sum_{i=1}^{k} rx_{A,t+i}\right] + d_{B,t}E_t\left[\sum_{i=1}^{k} rx_{B,t+i}\right] \\ -\frac{1}{2\tau}\left(\begin{array}{c} (d_{A,t})^2 \operatorname{Var}_t\left[\sum_{i=1}^{k} rx_{A,t+i}\right] + (d_{B,t})^2 \operatorname{Var}_t\left[\sum_{i=1}^{k} rx_{B,t+i}\right] \\ +2d_{A,t}d_{B,t} \operatorname{Cov}_t\left[\sum_{i=1}^{k} rx_{A,t+i}, \sum_{i=1}^{k} rx_{B,t+i}\right] \end{array} \right) \end{array} \right\}, \tag{55}$$

which implies

$$\begin{aligned}\begin{bmatrix} d_{A,t} \\ d_{B,t} \end{bmatrix} &= \tau \begin{bmatrix} V_A^{(k)} & C_{AB}^{(k)} \\ C_{AB,k} & V_B^{(k)} \end{bmatrix}^{-1} \begin{bmatrix} E_t\left[\sum_{i=1}^{k} rx_{A,t+i} \right. \\ E_t\left[\sum_{i=1}^{k} rx_{B,t+i}\right] \end{bmatrix} \\ &= \frac{\tau}{V_A^{(k)}V_B^{(k)} - \left(C_{AB}^{(k)}\right)^2} \begin{bmatrix} V_B^{(k)}E_t\left[\sum_{i=1}^{k} rx_{A,t+i}\right] - C_{AB}^{(k)}E_t\left[\sum_{i=1}^{k} rx_{B,t+i}\right] \\ V_A^{(k)}E_t\left[\sum_{i=1}^{k} rx_{B,t+i}\right] - C_{AB}^{(k)}E_t\left[\sum_{i=1}^{k} rx_{A,t+i}\right] \end{bmatrix},\end{aligned} \tag{56}$$

where $V_i^{(k)} \equiv Var_t\left[\sum_{j=1}^{k} rx_{i,t+j}\right]$ and $C_{AB}^{(k)} \equiv Cov_t\left[\sum_{j=1}^{k} rx_{A,t+j}, \sum_{j=1}^{k} rx_{B,t+j}\right]$. Under the conjectured forms for equilibrium yields, the realized k-period cumulative excess returns for both bonds are

$$\begin{aligned}\sum_{i=1}^{k} rx_{i,t+i} &= \sum_{i=0}^{k-1}\left(y_{i,t+i} - r_{t+i} - z_{i,t+i+1}\right) - \frac{\theta_i}{1-\theta_i}\left(y_{i,t+k} - y_{i,t}\right) \\ &= k\left(\alpha_{i0} - \bar{r} - \bar{z}_i\right) + \left(\boldsymbol{\alpha}_{i1} - \mathbf{e}_r\right)'\left(\sum_{j=0}^{k-1}\mathbf{x}_{t+j}\right) - \mathbf{e}_{z_i}'\left(\sum_{j=1}^{k}\mathbf{x}_{t+j}\right) - \frac{\theta_i}{1-\theta_i}\left(\boldsymbol{\alpha}_{i1}'\mathbf{x}_{t+k} - \boldsymbol{\alpha}_{i1}'\mathbf{x}_t\right).\end{aligned} \tag{57}$$

Given this formulation for excess returns, we find expressions for expected excess returns, the variance of excess returns, as well as the covariance between excess returns on A and B bonds:

$$E_t\left[\sum_{i=1}^{k} rx_{i,t+i}\right] = k\left(\alpha_{i0} - \bar{r} - \bar{z}_i\right) + \left(\left(\boldsymbol{\alpha}_{i1} - \mathbf{e}_r\right)'\left(\mathbf{I} - \boldsymbol{\Gamma}\right)^{-1} + \frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1}' - \mathbf{e}_{z_i}'\left(\mathbf{I} - \boldsymbol{\Gamma}\right)^{-1}\boldsymbol{\Gamma}\right)\left(\mathbf{I} - \boldsymbol{\Gamma}^k\right)\mathbf{x}_t. \tag{58}$$

$$V_i^{(k)} = \mathrm{Var}_t\left[(\boldsymbol{\alpha}_{i1} - \mathbf{e}_r - \mathbf{e}_{z_i})'\left(\sum_{j=1}^{k-1}\mathbf{x}_{t+j}\right) - \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)'\mathbf{x}_{t+k}\right] \tag{59}$$

$$= (\boldsymbol{\alpha}_{i1} - \mathbf{e}_r - \mathbf{e}_{z_i})'\left(\sum_{i=1}^{k-1}\sum_{j=1}^{k-1}\mathbf{C}_{[t+i,t+j]}\right)(\boldsymbol{\alpha}_{i1} - \mathbf{e}_r - \mathbf{e}_{z_i})$$

$$+ \left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)'\mathbf{C}_{[t+k,t+k]}\left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)$$

$$- 2\,(\boldsymbol{\alpha}_{i1} - \mathbf{e}_r - \mathbf{e}_{z_i})'\sum_{i=1}^{k-1}\mathbf{C}_{[t+i,t+k]}\left(\frac{\theta_i}{1-\theta_i}\boldsymbol{\alpha}_{i1} + \mathbf{e}_{z_i}\right)$$

$$C_{AB}^{(k)} = \mathrm{Cov}_t\left[\begin{array}{c}(\boldsymbol{\alpha}_{A1} - \mathbf{e}_r - \mathbf{e}_{z_A})'\left(\sum_{j=1}^{k-1}\mathbf{x}_{t+j}\right) - \left(\frac{\theta_A}{1-\theta_A}\boldsymbol{\alpha}_{A1} + \mathbf{e}_{z_A}\right)'\mathbf{x}_{t+k},\\ (\boldsymbol{\alpha}_{B1} - \mathbf{e}_r - \mathbf{e}_{z_B})'\left(\sum_{j=1}^{k-1}\mathbf{x}_{t+j}\right) - \left(\frac{\theta_B}{1-\theta_B}\boldsymbol{\alpha}_{B1} + \mathbf{e}_{z_B}\right)'\mathbf{x}_{t+k}\end{array}\right] \tag{60}$$

$$= (\boldsymbol{\alpha}_{A1} - \mathbf{e}_r - \mathbf{e}_{z_A})'\left(\sum_{i=1}^{k-1}\sum_{j=1}^{k-1}\mathbf{C}_{[t+i,t+j]}\right)(\boldsymbol{\alpha}_{B1} - \mathbf{e}_r - \mathbf{e}_{z_B})$$

$$- (\boldsymbol{\alpha}_{A1} - \mathbf{e}_r - \mathbf{e}_{z_A})'\sum_{i=1}^{k-1}\mathbf{C}_{[t+i,t+k]}\left(\frac{\theta_B}{1-\theta_B}\boldsymbol{\alpha}_{B1} + \mathbf{e}_{z_B}\right)$$

$$- (\boldsymbol{\alpha}_{B1} - \mathbf{e}_r - \mathbf{e}_{z_B})'\sum_{i=1}^{k-1}\mathbf{C}_{[t+i,t+k]}\left(\frac{\theta_A}{1-\theta_A}\boldsymbol{\alpha}_{A1} + \mathbf{e}_{z_A}\right)$$

$$+ \left(\frac{\theta_A}{1-\theta_A}\boldsymbol{\alpha}_{A1} + \mathbf{e}_{z_A}\right)'\mathbf{C}_{[t+k,t+k]}\left(\frac{\theta_B}{1-\theta_B}\boldsymbol{\alpha}_{B1} + \mathbf{e}_{z_B}\right)$$

Using our formulation for generalists bond demand from (56) and our closed form expressions for expectation, variance and covariance of excess returns, we can find a closed form for generalists bond demands and match these to the coefficients $\delta_{A1}, \delta_{A2}, \delta_{B1}, \delta_{B2}$ based on the conjectured form for bond demand:

$$\delta_{A0} = \tau\frac{V_B^{(k)}k\,(\alpha_{A0} - \bar{r} - \bar{z}) - C_{AB}^{(k)}k\,(\alpha_{B0} - \bar{r} - \bar{z})}{V_A^{(k)}V_B^{(k)} - \left(C_{AB}^{(k)}\right)^2} \tag{61}$$

$$\delta_{B0} = \tau \frac{V_A^{(k)} k \left(\alpha_{B0} - \bar{r} - \bar{z}\right) - C_{AB}^{(k)} k \left(\alpha_{A0} - \bar{r} - \bar{z}\right)}{V_A^{(k)} V_B^{(k)} - \left(C_{AB}^{(k)}\right)^2} \tag{62}$$

$$\delta'_{A1} = \tau \frac{\begin{pmatrix} V_B^{(k)} \left(\left(\alpha_{A1} - \mathbf{e}_r\right)' \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} + \frac{\theta_A}{1-\theta_A} \alpha'_{A1} - \mathbf{e}'_{z_A} \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} \mathbf{\Gamma} \right) \\ -C_{AB}^{(k)} \left(\left(\alpha_{B1} - \mathbf{e}_r\right)' \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} + \frac{\theta_B}{1-\theta_B} \alpha'_{B1} - \mathbf{e}'_{z_B} \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} \mathbf{\Gamma} \right) \end{pmatrix}}{V_A^{(k)} V_B^{(k)} - \left(C_{AB}^{(k)}\right)^2} \left(\mathbf{I} - \mathbf{\Gamma}^k\right) \tag{63}$$

$$\delta'_{B1} = \tau \frac{\begin{pmatrix} V_A^{(k)} \left(\left(\alpha_{B1} - \mathbf{e}_r\right)' \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} + \frac{\theta_B}{1-\theta_B} \alpha'_{B1} - \mathbf{e}'_{z_B} \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} \mathbf{\Gamma} \right) \\ -C_{AB}^{(k)} \left(\left(\alpha_{A1} - \mathbf{e}_r\right)' \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} + \frac{\theta_A}{1-\theta_A} \alpha'_{A1} - \mathbf{e}'_{z_A} \left(\mathbf{I} - \mathbf{\Gamma}\right)^{-1} \mathbf{\Gamma} \right) \end{pmatrix}}{V_A^{(k)} V_B^{(k)} - \left(C_{AB}^{(k)}\right)^2} \left(\mathbf{I} - \mathbf{\Gamma}^k\right) \tag{64}$$

D.3 Rational Expectations Equilibrium

In equilibrium, we need to clear markets for both bonds in a way that is consistent with the optimization of specialists A, and B as well as generalists. For both $i \in \{A, B\}$, the market clearing condition is

$$\left(1 - q_A - q_B\right) k^{-1} d_{i,t} + q_i b_{i,t} = s_{i,t} - \left(1 - q_A - q_B\right) \left(k^{-1} \sum_{i=1}^{k-1} d_{i,t-i}\right). \tag{65}$$

Plugging in $d_{i,t} = \delta_{i0} + \delta'_{i1} \mathrm{x}_t$ as well as our closed form expression for $b_{i,t}$ this becomes

$$\frac{\tau q_i}{V_i^{(1)}} \left(\alpha_{i0} - \bar{r} - \bar{z}\right) + q_i \left[\tau \frac{\left(\frac{1}{1-\theta_i} \alpha_{i1} - \mathbf{e}_r\right)' - \left(\frac{\theta_i}{1-\theta_i} \alpha_{i1} + \mathbf{e}_z\right)' \mathbf{\Gamma}}{V_i^{(1)}} \right] \mathbf{x}_t$$

$$= \left(\bar{s}_i - \left(1 - q_A - q_B\right) \delta_{i0}\right) + \left(\mathbf{e}_{s_i} - \left(1 - q_A - q_B\right) k^{-1} \left(\mathbf{1}_{(i)} + \delta_{i1}\right)\right)' \mathbf{x}_t, \tag{66}$$

where $\mathbf{1}_{(A)} \equiv \mathbf{e}_6 + \mathbf{e}_7$ and $\mathbf{1}_{(B)} \equiv \mathbf{e}_8 + \mathbf{e}_9$ in the case where $k = 3$. Matching constants for α_{i0} we have

$$\alpha_{i0} = \bar{r} + \bar{z} + \frac{V_i^{(1)}}{q_i \tau} \left(\bar{s}_i - \left(1 - q_A - q_B\right) \delta_{i0}\right) \tag{67}$$

Matching slopes for α_{i1} yields

$$\alpha_{i1} = \frac{1-\theta_i}{1-\rho_r\theta_i}\mathbf{e}_r + \frac{1-\theta_i}{1-\rho_{z_i}\theta_i}\rho_{z_i}\mathbf{e}_z \tag{68}$$

$$+ \frac{V_i^{(1)}}{\tau q_i}\left(\frac{1-\theta_i}{1-\theta_i\rho_{s_i}}\mathbf{e}_{s_i} - (1-\theta_i)(1-q_A-q_B)k^{-1}\left[\mathbf{I}-\theta_i\mathbf{\Gamma}'\right]^{-1}\left(\mathbf{1}_{(i)}+\delta_{i1}\right)\right) \tag{69}$$

Now note that α_{i0} and δ_{i0} are closed form expressions, and we express α_{i1} and δ_{i1} as closed form expressions that only depend on other exogenous closed form components and one-another. Namely, an equilibrium solves the following system of equations

$$\alpha_{A1} = \frac{1-\theta_A}{1-\rho_r\theta_A}\mathbf{e}_r + \frac{1-\theta_A}{1-\rho_{z_A}\theta_A}\rho_{z_A}\mathbf{e}_{z_1}$$

$$+ \frac{V_A^{(1)}(\alpha)}{\tau q_A}\left(\frac{1-\theta_A}{1-\theta_A\rho_{s_A}}\mathbf{e}_{s_A} - (1-\theta_A)(1-q_A-q_B)k^{-1}\left[\mathbf{I}-\theta_A\mathbf{\Gamma}(\delta)'\right]^{-1}\left(\mathbf{1}_{(A)}+\delta_{A1}\right)\right) \tag{70}$$

$$\alpha_{B1} = \frac{1-\theta_B}{1-\rho_r\theta_B}\mathbf{e}_r + \frac{1-\theta_B}{1-\rho_{z_B}\theta_B}\rho_{z_B}\mathbf{e}_{z_B}$$

$$+ \frac{V_B^{(1)}(\alpha)}{\tau q_B}\left(\frac{1-\theta_B}{1-\theta_B\rho_{s_B}}\mathbf{e}_{s_B} - (1-\theta_B)(1-q_A-q_B)k^{-1}\left[\mathbf{I}-\theta_B\mathbf{\Gamma}(\delta)'\right]^{-1}\left(\mathbf{1}_{(B)}+\delta_{B1}\right)\right) \tag{71}$$

$$\delta'_{A1} = \tau\frac{\begin{pmatrix} V_B^{(k)}(\alpha,\delta)\left((\alpha_{A1}-\mathbf{e}_r)'(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1} + \frac{\theta_A}{1-\theta_A}\alpha'_{A1} - \mathbf{e}'_{z_A}(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1}\mathbf{\Gamma}(\delta)\right) \\ -C_{AB}^{(k)}(\alpha,\delta)\left((\alpha_{B1}-\mathbf{e}_r)'(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1} + \frac{\theta_B}{1-\theta_B}\alpha'_{B1} - \mathbf{e}'_{z_B}(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1}\mathbf{\Gamma}(\delta)\right)\end{pmatrix}}{V_A^{(k)}(\alpha,\delta)V_B^{(k)}(\alpha,\delta) - \left(C_{AB}^{(k)}(\alpha,\delta)\right)^2}\left(\mathbf{I}-\mathbf{\Gamma}(\delta)^k\right) \tag{72}$$

$$\delta'_{B1} = \tau\frac{\begin{pmatrix} V_A^{(k)}\left((\alpha_{B1}-\mathbf{e}_r)'(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1} + \frac{\theta_B}{1-\theta_B}\alpha'_{B1} - \mathbf{e}'_{z_B}(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1}\mathbf{\Gamma}(\delta)\right) \\ -C_{AB}^{(k)}\left((\alpha_{A1}-\mathbf{e}_r)'(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1} + \frac{\theta_A}{1-\theta_A}\alpha'_{A1} - \mathbf{e}'_{z_A}(\mathbf{I}-\mathbf{\Gamma}(\delta))^{-1}\mathbf{\Gamma}(\delta)\right)\end{pmatrix}}{V_A^{(k)}V_B^{(k)} - \left(C_{AB}^{(k)}\right)^2}\left(\mathbf{I}-\mathbf{\Gamma}(\delta)^k\right). \tag{73}$$

Here, we write $V_A^{(1)}(\alpha)$, $V_B^{(1)}(\alpha)$ to emphizize their dependence on α_{A0}, α_{A1}, we write $\mathbf{\Gamma}(\delta)$ to emphizize $\mathbf{\Gamma}$'s dependence on δ_{A1} and δ_{B1}, we write $V_A^{(k)}(\alpha, \delta)$, $V_B^{(k)}(\alpha, \delta)$, and $C_{AB}^{(k)}(\alpha, \delta)$ to emphizize $V_i^{(k)}$'s and C_{AB}^k's dependence on α_{A1}, α_{B1}, δ_{A1}, and δ_{B1}. We can write the above equations compactly as

$$\alpha_{A1} = \mathrm{f}_{\alpha_{A1}}\left(\alpha_{A1}, \alpha_{B1}, \delta_{A1}, \delta_{B1}\right)$$

$$\alpha_{B1} = \mathrm{f}_{\alpha_{B1}}\left(\alpha_{A1}, \alpha_{B1}, \delta_{A1}, \delta_{B1}\right)$$

$$\delta_{A1} = \mathrm{f}_{\delta_{A1}}\left(\alpha_{A1}, \alpha_{B1}, \delta_{A1}, \delta_{B1}\right)$$

$$\delta_{A2} = \mathrm{f}_{\delta_{A2}}\left(\alpha_{A1}, \alpha_{B1}, \delta_{A1}, \delta_{B1}\right),$$

or simply as

$$\omega = f(\omega)$$

where $\omega = (\alpha'_{A1}, \alpha'_{B1}, \delta'_{A1}, \delta'_{B1})$. We solve this system of non-linear equations numerically in python by employing the Powell hybrid algorithm.