The Demand for Government Debt*

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Abstract

We document that the sectoral composition and marginal buyers of government debt differ notably across jurisdictions and have evolved significantly over time. Focusing on the United States, we estimate the yield elasticity of demand across sectors using instrumental variables constructed from monetary policy surprises. Our estimates point to a 14% increase in the demand by non-central-bank players for a 1 percentage point increase in long-term yields. Hence, a counterfactual reduction in the central bank balance sheet through quantitative tightening of around $266 billion increases long-term yields by 10 basis points. We find commercial banks, foreign private investors, pension funds, investment funds, and insurance companies to be the sectors whose demand is most sensitive to changes in long-term yields, but to varying degrees. The foreign official sector by contrast has price-inelastic demand. Our results imply compositional shifts towards more price elastic players as central banks normalize balance sheets with important implications for fiscal policy and financial stability.

Keywords: government debt, demand, yield elasticity, quantitative easing, quantitative tightening
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"...it will be like watching paint dry..."

Janet Yellen (2017) about balance sheet normalization

"...I would just stress how uncertain the effect is of shrinking the balance sheet..."

Jerome Powell (2022)

1 Introduction

Many central banks embarked on quantitative easing (QE) policies since the Global Financial Crisis (GFC) with the goal of lowering long-term interest rates as short-term interest rates had fallen to their lower bound. As central banks eventually normalize their balance sheets, a process also known as quantitative tightening (QT), the impact of these policies on government debt markets is frought with uncertainty. In order to understand the impact of these policies, we focus on the role of quantities and study compositional shifts in the investors that hold the government debt.

Focusing on the United States, we estimate that a 1 percentage point increase in long-term yields lead to a 14% increase in the demand by non-central-bank players. We arrive at these estimates by taking a demand-system perspective to bond holdings and using instrumental variables constructed from monetary policy surprises. Based on estimates of the yield-elasticity of different sectors, we infer the market-clearing yields under different quantitative tightening scenarios. For example, we find that, all else constant, a hypothetical reduction in the central bank balance sheet of around $266 billion leads to an increase in long-term bond yields by 10 basis points in the United States. Our estimates are quantitatively close to those for the Euro area and also in a similar ballpark as estimates of the impact of quantitative easing obtained via other methodologies.

To set the stage, in the first part of our analysis we study the evolution of holdings by different sectors of government debt in the United States, Euro area, Japan and United Kingdom using publicly available data.[1] In particular, we run regressions that measure the marginal absorption by different investor groups as the total amount of government debt outstanding changes following the methodology laid out in Fang, Hardy and Lewis (2022). According to the data from the

[1] For brevity, we report the results of the analysis for the United States and only discuss the main results for other jurisdictions in Section 2. We relegate the analysis for other jurisdictions to Appendix A.

Bank for International Settlements, these jurisdictions comprise around two thirds of total global government debt outstanding. We document that the evolution of the holdings of government debt by different sectors has varied significantly across jurisdictions. Even within jurisdictions, the marginal responses of different sectors to changes in government debt have changed significantly over time.

Our results shed light on how the role of the central banks in government debt markets has differed across jurisdictions and over time. Following rounds of QE, the average market shares of central banks increased to 22% in the United States, 16% in the Euro area, 44% in Japan, and 31% in the United Kingdom. Our estimates suggest that for every one unit change in the government debt outstanding only 7% was absorbed by the Federal Reserve between the first QE program and the Covid-19 crisis. This marginal response substantially increased to 43% after the Covid-19 crisis. In the Euro area, since the start of the QE program, the central bank has absorbed around two thirds of every unit increase in government debt. In Japan, following the start of the qualitative and quantitative easing program (QQE), the Bank of Japan absorbed 185% of every unit increase in government debt, which later declined to 33% after the Covid-19 crisis. Finally, the marginal response of the Bank of England was 38% after the QE program following the GFC, which increased to more than 50% after March 2020.

At the same time, our results show notable shifts in the holdings of non-central-bank sectors. In the United States, the marginal role of money market funds has increased since the Covid-19 crisis, with these players absorbing around one third of every unit of increase in government debt since March 2020.[2] The role of foreign investors rose after 2009 and declined again since the Covid-19 crisis. In the Euro area, all non-central-bank players, especially foreign investors, experienced a decline in marginal absorption (see Koijen, Koulischer, Nguyen and Yogo, 2021, for a similar finding using a confidential security-level dataset). In Japan, central bank purchases coincided with a decrease in bank holdings, indicating banks as major sellers, while foreign investors gained market share. In the United Kingdom, insurance companies and pension funds witnessed the most

[2] Money market funds are short-term investment vehicles and their holdings are also tilted towards shorter-maturity Treasury bills. In this paper, our main focus is on longer-term government bonds. For the impact of money market funds in the Treasury bill market, we refer the interested readers to Doerr, Eren and Malamud (2023).

significant decline in holdings, implying they were major sellers during quantitative easing, and foreign investors' market share has risen since 2009.

As many central banks embark on quantitative tightening (QT) to normalize balance sheets, a key determining factor of how different sectors would absorb government debt is how sensitive their demand curves are to changes in the yields (or prices) of government bonds. It is not possible to consistently estimate the yield elasticity of demand using ordinary least squares, since both prices and quantities demands are determined endogenously in equilibrium. One way to overcome the endogeneity is to find instrumental variables for government bond yields which do not directly affect the holdings of each sector, but only indirectly affect them through their effect on yields.

In the second part of the analysis, we quantify the yield elasticity of demand of different investor groups. To do so we take a demand system approach and focus on the United States given the data availability. To account for any endogeneity between latent demand by various investor groups and bond yields, we rely on monetary policy surprises as instruments for government bond yields in two-stage least squares regressions. In the first stage, we distil information contained in different measures of monetary policy surprises through a principal components analysis and take the first principal component as our instrument for yields.[3] This instrument has desirable properties. First, it is relevant as monetary policy surprises do have a direct impact on long-term yields. Moreover, we reject the null hypothesis of weak instruments. Second, since these are unanticipated by construction, we argue that they do not affect holdings through other channels than changes in yields and present further checks to alleviate any potential concerns.

Our second-stage regressions indicate that most sectors have downward-sloping demand curves with respect to prices. Since prices and yields move in opposite directions, this means these investors demand more when the yield goes up, and vice versa. Taking the weighted average of the elasticity estimates of each sector, we find a 1 percentage point increase in the 8-year yield[4] to increase demand by all non-central-bank sectors by around 14%. Commercial banks, foreign

[3] In particular, we use monetary policy surprises constructed by Jarociński and Karadi (2020); Swanson (2021); Bu, Rogers and Wu (2021); Kearns, Schrimpf and Xia (2022).

[4] Throughout the paper, we mainly focus on the 8-year yield which roughly corresponds the average duration of the Fed's SOMA portfolio as well as the average duration of other US Treasury investors (Tabova and Warnock, 2022). We provide results for different maturities in the appendix, which are qualitatively similar.

private investors, pension funds and investment funds are the sectors with the most elastic demand functions, but to varying degrees. Importantly, and in in line with intuition, we also find that the demand functions of the foreign official sector and households are not sensitive to prices. Using our elasticity estimates, we run scenario analyses for different hypothetical QT configurations. We estimate that a hypothetical $266 billion reduction in the central bank balance sheet increases the 8-year yield by 10 basis points. This estimate is roughly in line with estimates of the effect of QE using different methodologies (see Borio and Zabai, 2018, for a review).

These results have policy implications. As central banks have added balance sheet policies to their toolkit, a natural question is how to use and sequence interest rate and balance sheet policies for monetary tightening. Our analysis can be helpful for central banks in judging the possible impact of balance sheet tightening policies on long-term yields. During the Covid-19 crisis, market functioning issues came to the fore in government bond markets with several non-bank financial institutions contributing to the selling pressure in an environment where dealers' absorption capacity was constrained (e.g. Duffie, 2020; Schrimpf, Shin and Sushko, 2020; Vissing-Jorgensen, 2021; Eren and Wooldridge, 2021; He, Nagel and Song, 2022). Heterogeneous elasticity estimates across sectors imply a change in the composition of government debt holders as central banks normalize balance sheets. In an environment with higher yields, the role that foreign private investors, pension funds, commercial banks and investment funds play is set to increase further. Focusing on commercial banks, we show that commercial banks substitute between central bank reserves and government bonds to comply with the liquidity coverage ratio (LCR) and supervisory liquidity and resolution stress testing. A smaller central bank balance sheet (and hence a decline in reserves) might therefore lead to an even greater increase in banks' demand for bonds than our estimates indicate. This would imply that banks would resume again a more important role in government bond markets compared to their more subdued role in much of the post-GFC environment. Finally, a greater role for non-bank financial institutions similarly makes it more important to address vulnerabilities in this sector (see, e.g. FSB, 2022).

Related literature. Our paper contributes to the literature that emphasizes the role of quantities and investor demand in driving financial markets and macroeconomic phenomena. Recent pioneering work includes the inelastic markets hypothesis by Gabaix and Koijen (2021) who trace asset price movements to the impact of flows and Koijen and Yogo (2019) who propose a new demand system methodology based on market clearing conditions. Related to our work on bond markets, Fang, Hardy and Lewis (2022) study how different types of investors absorb debt supply in a broad panel of sovereign bond markets. They emphasize the role of non-bank investors, especially in emerging market economies. Zhou (2023) shows that accounting for foreign investor base differences helps explain the heterogeneous influence of the Global Financial Cycle on sovereign borrowing of emerging market economies. Choi, Kirpalani and Perez (2023) study the macroeconomic implications when the US government internalizes the downward sloping demand curve for its demand and exploits its market power when issuing debt. We contribute to this literature by studying how heterogeneous groups of investors differ in the yield-sensitivity of their demand for government and how this shapes the way the government bond market adjusts in case of imbalances.[5]

Our paper also relates to the literature that has studied the re-balancing in investor portfolios in response to central bank balance sheet policies. Koijen, Koulischer, Nguyen and Yogo (2021), in particular, show that the main counterparties to the ECB's asset purchase programmes since 2015 have been investors residing outside of the euro area. They also gauge the price impact of asset purchases through an estimated demand system setting. Saito and Hogen (2014), in turn, study how investors re-balanced their portfolios in response to the Bank of Japan's Quantitative and Qualitative Easing (QQE) policy. They find foreign entities to have responded the most via asset sales to the Bank of Japan, followed by domestic banks. Carpenter, Demiralp, Ihrig and Klee (2015) show that households – a heterogeneous group that also includes hedge funds – were the primary seller of securities to the Fed in the early phases of QE. They provide evidence suggestive of a re-balancing towards riskier assets such as corporate bonds—in line with a portfolio balance channel.

Our work also contributes to the broad literature about the impact of central banks' balance

[5]Aside from government bond markets, the perspective on quantities and investor demands has also been fruitfully applied, inter alia, to study effects in stock markets (e.g. Gabaix and Koijen, 2021), FX (e.g. Koijen and Yogo, 2020; Aldunate et al., 2022; Camanho et al., 2022; Jiang et al., 2022) and corporate bond markets (e.g. Coppola, 2021; Bretscher et al., 2022).

sheet size on government bond yields.[6] A large body of literature finds that purchases of government bonds from central banks lowers government yields as the net duration supply to the public falls (see Borio and Zabai (2018) and CGFS (2019) for reviews on the impact of various asset purchases programs implemented by central banks since the GFC). Our estimates for the impact of QT are roughly similar to the estimates for QE in the literature reviewed by Borio and Zabai (2018). For example, during QE2, in which the central bank purchased $667 billion in longer term US Treasuries (more than 6 years) and sold the same amount of short-term Treasuries, estimates for the impact on the 10-year yield range between 16 and 45 bps. Our estimate of the yield-impact of purchases of such quantity would be 25 bps (closer to estimates by Hamilton and Wu (2012) and Swanson (2011)). In the context of QT, Wright (2022) argued that the effects of reducing the Fed's Treasury holdings should be equivalent to the Treasury increasing the duration of their issuance, and found that Fed's QT is likely to have small effects on term premia and bond yields – 10 basis points higher in 10-year term premia. Based on projections by the Federal Reserve (Anderson, Marks, Na, Schlusche and Senyuz, 2022), SOMA holdings are expected to decline from $8.5 trillion in 2022 to $6.3 trillion in 2024. Our estimates thus suggest a total impact of QT on long-term yields of around 45 basis points.

Our work further contributes to the literature that emphasises market segmentation and the role of preferred habitat investors. Important contributions in this literature include Greenwood and Vayanos (2014); Greenwood and Vissing-Jorgensen (2018); Vayanos and Vila (2021) who show how such segmentation can have a bearing on asset prices and notably the yield curve. Jansen (2023) studies how changes in regulatory discount rates for Dutch insurers generated a demand shift affecting other players and had aggregate implications for the yield curve. Using an administrative dataset, Tabova and Warnock (2022) document the preferred habitats of different investors in the US Treasury market. Our paper also relates to the work that ascribes a special status to US Treasuries given their liquidity and safety attributes (see, e.g. Krishnamurthy and Vissing-Jorgensen, 2012; Greenwood, Hanson and Stein, 2015; Nagel, 2016; d'Avernas and Vandeweyer,

[6] Another related literature examines the impact of flows by the foreign official sector. See, for example, Bernanke, Reinhart and Sack (2004); Warnock and Warnock (2009); Beltran, Kretchmer, Marquez and Thomas (2013); Ahmed and Rebucci (2022).

2023; Doerr, Eren and Malamud, 2023; Krishnamurthy and Li, 2023; Acharya and Laarits, 2023). We contribute to this literature by providing new stylised facts on compositional shifts among various investor groups in absorbing US government debt and by estimating their respective yield elasticities.

2 The evolution of government debt holdings and marginal buyers

The supply of government debt has increased considerably since the GFC, boosted further by the massive fiscal expansion in response to the Covid-19 crisis. Between 2008 and 2021, total outstanding government debt has risen around five-fold in the United States and the United Kingdom and it roughly doubled in the Euro area and Japan. Even taking economic growth into account, debt-to-GDP ratios have increased from less than 40% to around 120% in the United States, from 70% to 95% in the Euro area, from 120% to more than 220% in Japan, and from around 30% to around 100% in the United Kingdom as shown in dashed lines in Figure 1(a), 5(a), 6(a) and 7(a).

In this section, we document how government debt holdings by different sectors have evolved over time across the different jurisdictions. In particular, we use a simple accounting framework to estimate how changes in each sector's government debt holdings co-move with changes in the total outstanding government debt across different time periods. This exercise allows us to quantify how the footprint of central banks and that by other key sectors has changed since the early 2000s and in response to quantitative easing programmes.

To assess how the absorption of debt supply by each sector has changed over time, we ask how much each sector absorbs of a one-unit increase in the supply of total government debt, and estimate the marginal response of different sectors to such changes. Our approach to estimate the marginal absorption by different types of investors follows the methodology laid out in Fang, Hardy and Lewis (2022). Specifically, for each jurisdiction j we regress separately the change in government debt held by each investor group (normalized by lagged total debt) on the growth of total debt:

$$\frac{H_t^{s,j} - H_{t-1}^{s,j}}{D_{t-1}^j} = \alpha^{s,j} + \beta^{s,j} \frac{D_t^j - D_{t-1}^j}{D_{t-1}^j} + \varepsilon_t^{s,j}, \tag{1}$$

where D_t^j represents the total outstanding government debt at time t of jurisdiction j; $H_t^{s,j}$ denotes the holdings of government debt by sector s of jurisdiction j at time t; $\alpha^{s,j}$ is a constant. The estimated coefficients $\beta^{s,j}$ can be interpreted as the marginal holding response of sector s to variations in the total outstanding government debt as the sum of the coefficients for different sectors will sum to 1, i.e. $\sum_s \beta^{s,j} = 1$ for each jurisdiction j.[7]

As we are interested in the marginal responses over time, we introduce interaction terms with time-dummies to Equation (1). Specifically, we run the following regression for each sector s in jurisdiction j:

$$\frac{H_t^{s,j} - H_{t-1}^{s,j}}{D_{t-1}^j} = \alpha^{s,j} + (\beta_{\text{pre-QE}}^{s,j} \times \mathbf{1}_{\text{pre-QE}}^j + \beta_{\text{QE}}^{s,j} \times \mathbf{1}_{\text{QE}}^j + \beta_{\text{post-Covid}}^{s,j} \times \mathbf{1}_{\text{post-Covid}}) \frac{D_t^j - D_{t-1}^j}{D_{t-1}^j} + \varepsilon_t^{s,j}, \tag{2}$$

where $\mathbf{1}_{\text{pre-QE}}^j$, $\mathbf{1}_{\text{QE}}^j$ and $\mathbf{1}_{\text{post-Covid}}$ are dummy variables that take the value 1 before QE, between QE and the Covid-19 crisis and after the Covid-19 crisis respectively, and 0 otherwise. $\sum_s \beta_{pre-QE}^{s,j} = \sum_s \beta_{QE}^{s,j} = \sum_s \beta_{post-Covid}^{s,j} = 1$.[8]

We interpret the coefficients $\beta_{pre-QE}^{s,j}$, $\beta_{QE}^{s,j}$ and $\beta_{post-Covid}^{s,j}$ as the marginal response of sector s in jurisdiction j during the respective time period (pre-QE, QE and post-Covid).

It is important to note that our dataset has two limitations. First, we only observe the holdings and not the maturity composition of holdings by sector. Second, total holdings correspond to market values and not face values. Therefore, changes in holdings might reflect in part valuation effects. To the extent that maturities of holdings across sectors are similar this concern would be

[7]Note that this does not impose non-negativity on these coefficients as long as their sum is 1. This allows for trade between different players. For example, if a sector has a negative coefficient, it suggests that the sector was a net seller to others.

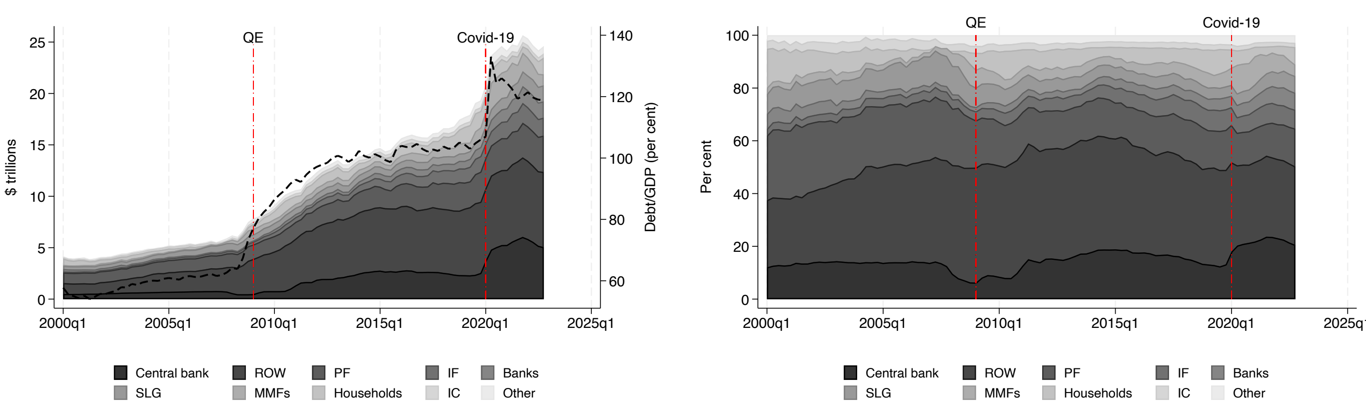
[8]The starting time of QE differs across jurisdictions: Q1 2009 for the United States and the United Kingdom, Q1 2015 for Euro area and Q2 2013 for Japan. The time dummy for post-Covid is one starting from Q1 2020 for all jurisdictions. Note that in the case of euro area, as the data is only available at annual frequency, we only consider pre-QE and QE periods.

alleviated, but holdings of different maturities coupled with heterogeneous changes in yields across the yield curve may still bias our estimates. We address the implications of this data limitation and describe ways to remedy them wherever applicable in the remainder of the paper.

2.1 United States

There has been a compositional shift in the holders of government debt in the United States since early 2000s and in particular since the GFC as shown in Figure 1 and Table 1 Panel (a), with the share of holdings by the Federal Reserve rising strongly from around 14% to 22%. Similarly, the shares of money market funds (MMFs), investment funds and commercial banks have also picked up since 2020. While the share held by foreign investors has drifted up from around 18% to 41% since the beginning of the sample until the Covid-19 crisis, it has since then declined again to about 30%. This decline was especially pronounced for foreign official investors (e.g. reserve managers). The share of holdings of pension funds, another large investor group, has also declined from around 15% to around 13%.

Figure 1: Total holdings and shares of different sectors - United States



(a) Total holdings **(b)** Shares

Note: Panels 1(a) and 1(b) show the total market value of the government debt holdings and market shares of each sector in the United States, respectively, between 2000Q1 and 2022Q3 (quarterly data). Central bank refers to the holdings of the Federal Reserve. ROW refers to foreign investors (official and private). PF refers to pension funds, IF refers to investment funds (open-ended mutual funds, exchange-traded funds and closed-end funds). Banks refer to commercial banks. SLG refers to state and local governments. MMFs refers to money market funds. Households refer to the direct and indirect holdings (e.g. through hedge funds) of households. IC refers to insurance companies and Other refers to all other sectors combined. QE starts at 2009Q1 and Covid-19 is at 2020Q1. *Source: Federal Reserve.*

Our accounting framework sheds further light on the evolution of holdings. We report the regression results of Equation (2) for the United States in Table 1 Panel (b). The rows show the OLS estimates of $\beta^{s,j}_{pre-QE}$, $\beta^{s,j}_{QE}$, and $\beta^{s,j}_{post-Covid}$, respectively. Each column shows the results for sector s denoted in the column heading. Before the launch of QE, for every additional change in government debt, 23% was absorbed by foreign investors, 20% was absorbed by the domestic household sector, 15% was absorbed by commercial banks, and 11% was absorbed by pension funds.[9] This changed notably with the advent of QE policies: between 2009Q1 and 2020Q1, the central bank increased its marginal absorption to 7%, foreign investors increased their marginal role from 23% to 31% and pension funds increased their role from 11% to 14%. Since the Covid-19 crisis, however, only two sectors, the central bank and MMFs, jointly account for the absorption of 75% of a one unit increase government debt supplied. The marginal role of almost all other sectors, by contrast, has declined. This change also illustrates key differences between the asset purchases conducted following the GFC compared to the purchases following the Covid-19 crisis, especially in relation to the role played by the central bank.

MMFs are an important source of demand for Treasury securities. While their outright purchases are tilted towards short-maturity instruments, such as Treasury bills, recently they have also come to play a crucial *indirect* role in the Treasury market through the reverse repos at the Federal Reserve's overnight reverse repurchase agreement (ON RRP) facility. In particular, the usage of the ON RRP facility has skyrocketed to around \$2.5 trillion in 2022 (see Doerr, Eren and Malamud, 2023, for the role of MMFs in the Treasury bill market and the ON RRP facility).

2.2 Euro area, Japan and the United Kingdom

A similar analysis for the Euro area, Japan and the United Kingdom highlights various patterns of the evolution of holdings in these jurisdictions.[10]

Similar to the United States, central banks in these jurisdictions increased their footprints in their respective sovereign bond markets after rounds of QE. In the Euro area, the central bank has

[9] Other sectors in column (10) combine the holdings of government-sponsored enterprises, broker-dealers and other financial companies that are not captured in other columns and non-financial corporates.

[10] In this section, we report the summary of the results. We refer the interested reader to Appendix A for the detailed analysis and discussion.

Table 1: Average shares and marginal response by sectors in the United States

Panel (a): Average shares by sector over different periods

Avg. Share	CB	ROW	PF	IF	Banks	SLG	MMF	HH	IC	Other
Pre-QE	13.5	18.6	18.5	2	12.4	9.8	1.4	17.3	3.6	2.9
QE	14.6	40.8	14.8	5.6	3.1	5	4.3	6.6	2.3	2.9
Post-Covid	21.7	30.4	13.1	6.9	5.7	5.3	7.8	4.5	1.6	3

Panel (b): Marginal holdings by sector over different periods

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	CB	ROW	PF	IF	Banks	SLG	MMF	HH	IC	Other
Pre-QE * Pct. Ch. Gov. Debt	-0.01	0.23***	0.11***	0.01	0.15***	0.04	0.10***	0.20***	0.04***	0.12***
	(0.02)	(0.07)	(0.04)	(0.01)	(0.06)	(0.03)	(0.04)	(0.03)	(0.01)	(0.01)
QE * Pct. Ch. Gov. Debt	0.07**	0.31***	0.14**	0.07***	0.09***	-0.02	0.07	0.21***	0.02***	0.05***
	(0.04)	(0.06)	(0.06)	(0.01)	(0.02)	(0.02)	(0.05)	(0.06)	(0.01)	(0.01)
Post-Covid * Pct. Ch. Gov. Debt	0.43***	0.07***	0.02	-0.00	0.07***	0.05***	0.32***	-0.05***	0.01***	0.08***
	(0.04)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.02)	(0.01)	(0.00)	(0.01)
Observations	227	227	227	227	227	227	227	227	227	227
R-squared	0.43	0.33	0.16	0.14	0.21	0.06	0.40	0.18	0.29	0.16

Note: Panel (a) shows the average share of each investor group over different time periods. Panel (b) reports the coefficients of the OLS regression of Equation (2) for the United States. CB refers to the Federal Reserve. ROW refers to foreign investors (official and private). PF refers to pension funds, IF refers to investment funds (open-ended mutual funds, exchange-traded funds and closed-end funds). Banks refer to commercial banks. SLG refers to state and local governments. MMF refers to money market funds. HH refer to the direct and indirect holdings (e.g. through hedge funds) of households. IC refers to insurance companies. Other refers to all other sectors combined. Data are quarterly. Pre-QE is between 2000Q1 and 2008Q4. QE is between 2009Q1 and 2019Q4. Post-Covid is between 2020Q1 and 2022Q3. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. *Source: Federal Reserve.*

absorbed two thirds of every unit of new government debt since its PSPP, raising its market share from less than 3% to 16%. In the United Kingdom, the central bank has taken down about 40-50% of new debt, leading to an average of one third market share. The expansion is most aggressive in Japan. Between QQE and the Covid outbreak, the central bank bought 1.85 units for every unit increase in the outstanding amount of government bonds. While the pace slowed after Covid, the central bank accounted for almost almost half market share in the post-Covid period.

The compositional shift among non central bank holders is more heterogeneous across regions. In the Euro area, the marginal absorption by all non-central-bank players declined, with the most pronounced declines observed for foreign investors (see Koijen, Koulischer, Nguyen and Yogo, 2021, for a similar finding using a confidential security-level dataset). In Japan, most of central bank purchases of government debt are mirrored by declines in bank holdings – which suggests that banks were the major sellers; at the same time, the market share of foreign investors has increased during our sample period. In the United Kingdom, insurance companies and pension funds saw the largest decline in holdings and marginal response, suggesting that these players were the sellers to the central bank during QE. Like in Japan, the average market share of foreign investors has also increased since 2009 in the UK debt markets.

3 The yield elasticity of demand for different types of investors

As central banks embarked on quantitative easing programs to impact long-term yields, the sensitivity of different sectors to changes in government bond yields was a key factor influencing the changes in the composition of non-central bank holders of government debt. In this section, we estimate the elasticity of demand by different groups of investors to government bond yields building on insights from the demand system approach to asset pricing (Koijen and Yogo, 2019). This quantification is crucial for assessing how bond markets will shift to a new equilibrium as central banks normalize their balance sheets through QT – a question we shed light on in Section 4 using our estimates from this section. Due to data availability, we focus on the United States in this section. We also compare our estimates to those for the Euro area in Koijen, Koulischer, Nguyen and Yogo (2021).

3.1 The demand system approach

Taking a demand system approach to asset pricing, Koijen and Yogo (2019) derive weights of different assets in a portfolio based on three assumptions: (i) investor preferences are such that the optimal portfolio is a mean-variance portfolio, (ii) returns have a factor structure, (iii) and expected returns and factor loadings depend only on an asset's own prices and characteristics. Under these conditions, the portfolio weight of an asset can be expressed as a logit function of its own price p_t (or yield Y_t in the case of government bonds) and a vector of the asset's characteristics $\mathbf{X_t}$:

$$\delta_t^s = exp(\alpha^s + \beta_1^s\ Y_t + \beta_2^{s'}\ \mathbf{X_t})\ \epsilon_t^s, \tag{3}$$

where $\delta_t^s \equiv \frac{H_t^s}{H(0)_t^s}$ is the ratio of holdings of an asset at time t by sector s (H_t^s) to the outside asset ($H(0)_t^s$), α^s is a constant pertaining to a particular investor sector, β_1^s is the responsiveness of each sector s to changes in the yield of an asset Y_t, $\mathbf{X_t}$ is a vector of characteristics of the asset and $\beta_2^{s'}$ capture the sensitivity of each investor group's demand with respect to these characteristics. Finally, ϵ_t^s is the latent demand which represents characteristics that are unobserved to the econometrician where $\eta_t^s \equiv log(\epsilon_t^s)$ is assumed to be normally distributed. The latent demand might capture the differences in beliefs about expected returns and risk aversion across sectors (among other unobserved investor traits).

While the demand system approach generates empirically tractable portfolio weights, two challenges remain in practice. First, the latent demand is jointly endogenous with asset prices, therefore estimates using OLS would yield biased and inconsistent estimates. Therefore, a valid instrument is needed to consistently estimate the main parameters of interest, β_1^s. Second, the outside asset in a logit demand system is defined as the asset investors would hold more of if the prices of all "inside assets", those that are in the investor choice set, went up. Without detailed visibility into the entire portfolios held by different groups of investors and due to its rather abstract nature, in most cases, the holdings of the outside asset are not observable. In this case, one approach would be to include time fixed effects, but it is not feasible since elasticities are estimated from time-series variation in yields. Faced with this challenge, empirical researchers often work with a parametric specification of the outside asset which aims to capture other investment opportunities investors have (see, for example, Koijen, Koulischer, Nguyen and Yogo, 2021; Fang, Hardy and Lewis, 2022).

If the outside asset were observable, one could use the following identity:

$$log(H_t^s) = log(\delta_t^s) + log(H(0)_t^s). \tag{4}$$

Using Equation (3) and Equation (4) together would then lead to the following linear equation:

$$log(H_t^s) - log(H(0)_t^s) = \alpha^s + \beta_1^s\ Y_t + \beta_2^{s'}\ \mathbf{X_t} + \eta_t^s. \tag{5}$$

In most cases, when the outside asset is not observable, the previous literature parametrically specifies the outside asset to capture investment opportunities outside of investors' likely choice sets (e.g. Koijen, Koulischer, Nguyen and Yogo, 2021; Fang, Hardy and Lewis, 2022; Jansen, 2023). This takes a linear form with $log(H(0)_t^s) = \phi^s + \gamma^{s'}\ \mathbf{W_t}$, where $\mathbf{W_t}$ is a vector that is meant to capture factors that affect the holdings of the outside asset, and $\gamma^{s'}$ is the sensitivity of each sector's demand for the outside asset to these factors. In our setup, however, we leverage the information contained in the Financial Accounts of the United States (also known as Flow of Funds) about total financial assets which is available for some sectors. That said, information on the total assets of some key investor groups such as foreign investors is unfortunately missing from this dataset. Hence, for foreign investors we cannot use an empirical specification based on Equation (5).

3.2 Baseline empirical specification to estimate demand

We proceed with our estimation strategy as follows. To start, similar to the previous literature, we parametrically specify the outside asset in a parsimonious way to get comparable estimates for all sectors. In particular, we estimate Equation (6) for *all sectors* with two-stage least squares:

$$log(H_t^s) = (\alpha^s + \phi^s) + \beta_1^s\ Y_t^8 + \beta_2^{s'}\ \mathbf{X_t} + \gamma^{s'}\ \mathbf{W_t} + t + t^2 + \eta_t^s. \tag{6}$$

In addition, for the subset of sectors included in the Flow of Funds dataset, we make use of our knowledge of the total financial assets to proxy for the holdings of the outside asset, $H(0)_t^s$ by subtracting the total holdings of US Treasury securities from the total assets. We then estimate Equation (7) for these sectors using two-stage least squares:

$$log(H_t^s) - log(H(0)_t^s) = \alpha^s + \beta_1^s \ Y_t^8 + \beta_2^{s'} \ \mathbf{X_t} + t + t^2 + \eta_t^s. \tag{7}$$

In both specifications, our main coefficient of interest is β_1^s which measures the per cent change in the holdings of a sector in response to a 1 percentage point change in the 8-year (zero coupon) US government bond yield. We choose to study the sensitivity to 8-year yields as that maturity roughly corresponds to the duration of assets held by the Federal Reserve.[11] We include a (quadratic) trend in order to get identification from the deviations from the trend.[12]

We use instrumental variables and estimate β_1^s using 2SLS regressions in order to address endogeneity between the latent demand and yields. For example, if investor beliefs are behind the latent demand which affects yields and holdings jointly, the estimates using OLS would be inconsistent. We use cleanly identified high-frequency monetary policy surprises as instruments for government bond yields which are measured within a short window around monetary policy announcements. We describe our identification strategy in detail in the next section.

In our baseline estimations, $\mathbf{X_t}$ includes log GDP, GDP growth, core inflation and the broad dollar index (in logs) to capture core characteristics of US Treasury securities. While the first three variables control for macroeconomic factors, we include the spot exchange rate in order to parsimoniously control for the convenience yield of US Treasuries (see Jiang, Krishnamurthy and Lustig, 2021, for a theory linking the spot exchange rate to convenience yields).

In the baseline $\mathbf{W_t}$ that captures factors influencing demand of the outside asset, we include the zero coupon rate on 5-year German government bonds and the VIX (in logs). The former is similar to Koijen, Koulischer, Nguyen and Yogo (2021) as they include the yield on US Treasury bonds when they parameterize the outside asset for European investors. Including the VIX is

[11] According to data from the Federal Reserve Bank of New York, the weighted average maturity of Treasury holdings in the SOMA portfolio was 7.91 years on January 4th, 2023. To assess the impact of QT, in which we focus on how private sector will absorb the lack of purchases by the central bank, we focus on the elasticity of demand to 8-year zero-coupon yields. In the Internet Appendix, we estimate Equation (6) using different yields and get similar results.

[12] Using a linear trend yields similar results, but we use the quadratic trend as it is a better specification to account for trends in holdings.

meant to capture risk-on and risk-off episodes which would affect the holdings of safe assets (e.g. flight-to-safety episodes).

Our results do not vary much depending on the exact choice of variables in $\mathbf{X_t}$ and $\mathbf{W_t}$. In the Internet Appendix, we report numerous robustness checks where we vary the variables in both $\mathbf{X_t}$ and $\mathbf{W_t}$. These results show that our estimates are to a large extent qualitatively and quantitatively unaffected by the control variables in the regression.[13]

Finally, we make valuation adjustments to address a data limitation which might impact our estimates. Since our dataset reflects the *market value* of holdings by sectors, the change in holdings we observe might reflect both the change in demand and valuation effects, which would bias our estimates.[14] For our baseline results, we assume 8 years as the average modified duration for each sector.[15] We calculate the percentage change in the price of their holdings as:

$$\%\Delta Price^{8y} \approx -\Delta 8\text{y yield} \times \text{Modified Duration}.$$

Using the total holdings by each sector, we translate these percentage changes into those effects that are only due to valuation effects. In the second stage regressions, we subtract these values from the market value of holdings for each sector and use the adjusted holdings as our dependent variable.

3.3 Monetary policy surprises as instruments

We use high-frequency monetary policy surprises as instruments for government bond yields. These surprises are typically measured within a short window around monetary policy announcements.

[13]See Section 3.5 for a discussion of the robustness checks and the Internet Appendix.

[14]To see why, assume that holdings remain unchanged from quarter-to-quarter, but yields decline. This would amount to a positive valuation effect and in our data appear as an increase in holdings. As a result, this is likely lead to a downward bias in our elasticity estimate, making the demand curve appear upward sloping with respect to prices.

[15]Here, we choose 8 years to be consistent with our use of the 8-year zero coupon yield, but our results are qualitatively similar if we use alternative assumptions for the average modified duration. Using an administrative data, Tabova and Warnock (2022) show that the duration of US Treasuries in portfolios varied around 7 and 8 years for major investors. The duration of foreigners' holdings are around 5 years, which is accounted for in our robustness checks.

Since our other data are quarterly, we aggregate the estimated monetary policy surprises within short windows over the quarter.

This high-frequency identification approach can cleanly address any endogeneity concerns. It also delivers a relevant instrument given the information content of the identified surprises for bond yields (see, e.g. Kuttner, 2001; Cieslak, 2018; Bauer and Swanson, 2022). However, this clean identification comes at the cost of reduced statistical power as the estimated effects tend to be small in the order of a few basis points (Nakamura and Steinsson, 2018).

Our identification argument is that monetary policy surprises are uncorrelated with the latent demand of each investor group for government bonds. One might think of them as "supply shifters" as they affect the borrowing decisions of the Treasury which also takes these surprises as given. We estimate demand elasticities by tracing out the demand curves following these surprises. In this case, a potential threat to identification would be simultaneity if those investors whose demand elasticity we are interested in are precisely those that respond to monetary policy surprises within the windows in which the surprises are estimated. We believe this is unlikely as the main sectors we are interested in typically rebalance their portfolios in a slow manner. The estimated surprises, by contrast reflect reactions of market makers and fast money investors such as high-frequency traders which are not the main players of interest for us. Another potential threat to identification would be if the central bank were to take demand by individual sectors into account in its interest rate decision and tried to surprise the market accordingly. However, we also believe this case to be unlikely.

There are multiple measures of monetary policy surprises that have been proposed in the literature. We take an agnostic approach on how to construct our instrument. In particular, we run the first stage regression separately for each set of monetary policy surprises. We also conduct a principal component analysis in order to make use of information contained in all of them while keeping the dimensionality in check since we have a relatively short sample period. We report the first stage results for all and select our baseline for the second stage according to the results of weak instruments tests.

We rely on monetary policy surprises from four sources. The first source is Swanson (2021) who uses a factor analysis to separately identify surprise changes in the federal funds rate, forward guidance, and large-scale asset purchases (LSAPs) for each FOMC announcement through their different impact on various asset prices including interest rates of different maturities, equities and exchange rates. The second one is Kearns, Schrimpf and Xia (2022) who share a similar goal as Swanson but take a simpler approach and construct target rate, path, and long-rate surprises. The third is Bu, Rogers and Wu (2021) who develop a heteroscedasticity-based partial least squares approach, combined with Fama-MacBeth style regressions, to identify a common US monetary policy surprise reflecting both conventional and unconventional monetary policy news. Finally, Jarociński and Karadi (2020) decompose monetary policy surprises into monetary policy and information shocks using high frequency co-movement between interest rates and stock prices.[16]

In addition to using the respective monetary policy surprises individually, we also generate a time series of shocks by taking the first principal component of these surprises. Given that the sample periods of the different studies are not the same, we consider three different combinations when constructing the first principal component.[17] We construct three principal components, one using the three series that start early, another using all four of them during the sample period in which they overlap, and finally use Bu, Rogers and Wu (2021) and Kearns, Schrimpf and Xia (2022) to cover the Covid-19 period.

We report the results of the first stage regression for the estimation Equation (6) in Table 2 and relegate the results of the first stage with equation (7) to the Internet Appendix.[18] While in many instances the coefficients have the economically meaningful sign and are statistically significant,

[16] Jarociński and Karadi (2020) differentiate between these shocks as follows. The monetary policy shock moves interest rates and stock prices in opposite directions, while the central bank information shock lead to same directional change in interest rates and stock prices. There is an ongoing debate in the literature whether these represent private information that the Federal Reserve transmits to the market or they simply measure the response of the Fed to news (Bauer and Swanson, 2023). In the latter case, for which we believe the evidence is compelling, there is no threat to our exclusion restriction. However, if the former interpretation is right, it might threaten the exclusion restriction if the Fed's private information affects holdings also through the release of information about the economy. We present robustness checks in the Internet Appendix in which we exclude the estimated information shocks and our results remain similar.

[17] The series by Jarociński and Karadi (2020) and Swanson (2021) starts in 1990s, but ends in mid-2019; hence, it does not cover the Covid-19 crisis and its aftermath. The sample of Bu, Rogers and Wu (2021) starts in early 1990s and goes through the end of 2021. The measures constructed by Kearns, Schrimpf and Xia (2022) start 2004 and go through the end of 2022.

[18] In columns (1)-(4), each of the monetary policy surprise series is standardized to ease comparison. Therefore, for example, a coefficient of 0.0004 means a one standard deviation change in the surprise variable corresponds to a 4 basis point change in the 8-year zero-coupon yield.

in all these specifications, the effective F-statistics constructed using the methodology in Olea and Pflueger (2013) are low and below the critical values. Hence, we fail to reject that they are weak instruments.[19] This is problematic since weak instruments lead to biased estimates in small samples. These results may reflect that these series, when considered individually, are noisy measures of monetary policy surprises that affect long-term yields at quarterly frequency.

We proceed with the monetary policy surprises constructed using principal components analysis. Since each series covers different time periods, we report the results with first principal components as instruments in three alternative specifications. The first one (PCA 1 (BRW, Swanson, JK)) aims to have a long sample and is reported in column (5). It uses three of these four series and runs between 1994q2 and 2019q2. This yields a positive, but insignificant estimate with a low effective F-statistic. The second one (PCA 1 (All)) uses all series when they overlap. The sample period is between 2004q3 and 2019q2. The estimated coefficient is positive and significant and the effective F-statistic improves as we reject the null hypothesis of weak instruments at the critical values used in Olea and Pflueger (2013). In the third specification, in column (7), we report the results using the first principal component of the Bu, Rogers and Wu (2021) and Kearns, Schrimpf and Xia (2022) series (PCA 1 (BRW, KSX)), which covers a sample period between 2004q3 and 2021q4. The estimated coefficient is similar to the one in column (6) and it is statistically significant. We also reject the null hypothesis of weak instruments in this case. Given these results, we take the (PCA 1 (All)) as our baseline for the second stage. We report the results for various alternative specifications (using different instruments, restricting the sample periods to be the same, among others) in the Internet Appendix, but our main take-aways remain largely similar. We also present further details on the principal components analysis in the Internet Appendix.

Overall, these results suggest that distilling information contained in each series through dimension reduction improves the first stage results. In terms of magnitudes, the estimates suggest that a one standard deviation increase in the first principal component of the monetary policy surprises (i.e. tightening) corresponds to a 4-12 basis point increase in the 8-year zero-coupon yield.

[19] We also repeat this exercise to check whether specifying the government bond yield and other control variables in changes instead of levels since the monetary policy surprise might in principle capture more information about changes rather than levels. However, this also leads to a failure to reject the null hypothesis of weak instruments. Therefore, we stick to our specification in Equations (6) and (7).

Table 2: **First stage results with alternative specifications of monetary policy surprises**

VARIABLES	(1) 8Y Yield (ZC)	(2) 8Y Yield (ZC)	(3) 8Y Yield (ZC)	(4) 8Y Yield (ZC)	(5) 8Y Yield (ZC)	(6) 8Y Yield (ZC)	(7) 8Y Yield (ZC)
Swanson FFR	-0.0002						
	(0.0003)						
Swanson FG	0.0011***						
	(0.0004)						
Swanson LSAP	0.0001						
	(0.0001)						
BRW		0.0005*					
		(0.0003)					
KSX (3M)			0.0012***				
			(0.0004)				
KSX (2Y)			-0.0006				
			(0.0007)				
KSX (10Y)			0.0010**				
			(0.0004)				
JK MP				0.0000			
				(0.0003)			
JK CBI				0.0014***			
				(0.0003)			
PCA 1 (BRW, Swanson, JK)					0.0003		
					(0.0002)		
PCA 1 (All)						0.0005***	
						(0.0001)	
PCA 1 (BRW, KSX)							0.0008***
							(0.0001)
Observations	102	112	74	102	102	60	70
R-squared	0.9548	0.9427	0.9031	0.9549	0.9512	0.9207	0.9058
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Sample	1994q2-2019q2	1994q2-2021q4	2004q3-2022q4	1994q2-2019q2	1994q2-2019q2	2004q3-2019q2	2004q3-2021q4
Effective F-stat	4.80	3.36	3.22	5.68	1.36	43.62	27.54
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	18.89	23.11	19.46	9.96	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation (6). The sample period varies depending on the availability of data across different monetary policy surprises. In columns (1)-(4), each of the monetary policy surprise series is standardized. Effective F-stat is calculated using the methodology in Olea and Pflueger (2013). The final row reports the critical values of a test of weak instruments with a 5% confidence level and a 10% worst-case bias. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

3.4 Second-stage results

We replace the government bond yields by the fitted values recovered from the first stage to consistently estimate β_1^s in the second stage. For inference, we report standard errors that are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. We also report the p-values of the Anderson-Rubin Wald test and the p-values of the underidentification LM test.

We report the results of the estimation of Equation (6) in the first panel of Table 3. We order the columns based on the market share of each sector in 2022q3 (as reported in Section 4).[20] We find that the yield elasticity of demand is positive and statistically significant for commercial banks (31.3), foreign private investors (24.3), pension funds (23.2), investment funds (21.3), insurance companies (16.3), and state and local governments (6.7), with the elasticity estimates indicated in parentheses. This indicates that these groups of investors have a downward-sloping demand curve (with respect to prices). The results are similar in the Panel (b) of the same table, which reports estimates of Equation (7) for the subset of sectors for which we have information on the total financial assets.

A higher elasticity means that the demand from a particular investor group is more responsive to changes in bond yields. Quantitatively, if the 8-year yield increases by one percentage point, the demand by sector s increases by the elasticity estimates reported, i.e. $\widehat{\beta_1^s}$ %. Interestingly, for the foreign official sector, we find the elasticity estimate to be statistically indistinguishable from zero. This finding in turn points to an inelastic demand curve by these players, which means they are not sensitive to changes in the yield of government debt. All other investors not included in these categorizations, which make up for less than 3% of total holdings, exhibit a negative coefficient (significant at 10% level) which suggests an upward sloping demand curve for some investors included in there.

Importantly, we obtain a weighted average elasticity estimate of 14.2 across investor groups, when applying weights according to the latest holdings shares of the sectors. The magnitude suggests that for a one percentage point increase in the 8-year yield, the demand of non-central-bank players on aggregate goes up by 14.2%. These estimates are similar in aggregate to those estimated for Euro area government bonds by Koijen, Koulischer, Nguyen and Yogo (2021). The weighted average yield elasticity estimate for the Euro area is around 12, which is remarkably

[20] We exclude money market funds since they are only allowed by regulation to hold short-term securities.

Table 3: The yield elasticity of demand across different sectors in the United States

Panel (a): Using a parametric specification for the outside asset (Equation (6))

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
8Y Yield (ZC)	1.49	24.34***	23.28***	21.36***	2.98	31.35***	6.73*	16.39***	-29.75*
	(2.58)	(7.78)	(8.72)	(7.26)	(43.32)	(11.45)	(3.84)	(4.72)	(16.36)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.56	0.00	0.00	0.00	0.95	0.00	0.04	0.00	0.06
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07

Panel (b): Using portfolio information to control for the outside asset (Equation (7))

VARIABLES	(1) log(PF)-log(OA_{PF})	(2) log(IF)-log(OA_{IF})	(3) log(HH)-log(OA_{HH})	(4) log(Banks)-log(OA_{Banks})	(5) log(SLG)-log(OA_{SLG})	(6) log(IC)-log(OA_{IC})
8Y Yield (ZC)	20.52**	18.06**	-8.34	27.34*	5.20	15.44**
	(9.49)	(7.41)	(49.83)	(14.25)	(4.15)	(6.34)
Observations	60	60	57	60	60	60
Controls	**X**	**X**	**X**	**X**	**X**	**X**
Anderson-Rubin Wald test (p-val)	0.002	0.000	0.867	0.003	0.247	0.000
Underidentification LM stat (p-val)	0.094	0.102	0.106	0.089	0.102	0.102

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) in Panel (a) and Equation (7) in Panel (b) using PCA 1 (All) as an instrument for yields. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

similar to our average elasticity estimate for the United States.[21] Using data from the Netherlands, Jansen (2023) also finds that banks are the most price elastic sector, similar to our findings.

[21] In the euro area, the estimated yield elasticities of demand for government bonds are 7.3 for commercial banks, 13.1 for mutual funds, -34.4 for pension funds and insurance companies, 1.4 for households, 42.8 for foreign investors, and 1.3 for all other investors. We use the shares of each sector reported in the Table 4 of Koijen, Koulischer, Nguyen and Yogo (2021) to compute the weighted averages.

3.5 Robustness checks and other considerations

We report the results of several alternative regression specifications in the Internet Appendix. In this section, we provide a high-level discussion of the key findings of these checks. Overall, our main take-aways remain similar.

We report the elasticities of different sectors with respect to variables used in $\mathbf{X_t}$ and $\mathbf{W_t}$ in Appendix B.1. Important take-aways are that the coefficient in the second stage for the 5-year German yield is negative, consistent with it being defined as the outside asset. In addition, for many sectors the coefficient for log GDP is negative suggesting lower holdings of Treasuries in boom periods, presumably as investors load up on riskier assets in booms.

We report two alternative first stage specifications in Appendix B.2. We get similar results when we restrict the sample period to be the same across specifications, suggesting that condensing the information content in more surprise series through the principal components analysis does indeed increase the signal in the estimates. We also check whether running the first stage in differences as opposed to levels. With first differences, we cannot reject the null hypothesis of weak instruments, therefore we keep levels in our baseline regressions.

We vary the instruments in these regressions in robustness exercises reported in Appendices B.3, B.4, and B.5. In Appendix B.3, we report the results using PCA 1 (BRW, KSX) as an instrument instead of PCA 1 (All). We find the results to be qualitatively similar, but with somewhat smaller point estimates of elasticities. In Appendix B.4, we drop the Jarociński and Karadi (2020) information shocks from the construction of the principal components. In Appendix B.5, we use the first two principal components (PCA 1 (All) and PCA 2 (All)) as instruments.

In Appendices B.6, we vary the control variables used in the regression Equation (6). In one specification, we use log GDP, GDP growth, inflation, log broad dollar index and the total face value of outstanding Treasuries in the asset characteristics vector X_t, and we use log VIX, the Option-Adjusted Spread (OAS) of the ICE BofA AAA US Corporate Index, 5-year (zero-coupon) German government bond yield, S&P 500 dividend yield for the outside asset specification W_t. In another specification, we control for a linear trend instead of a quadratic trend. Overall, the results we obtain for our parameter of interest are similar.

In Appendix B.7, we vary the valuation adjustments in the holdings of each sector. In our baseline results, we assume that the modified duration of each sector's US Treasury holdings is 8, in line with the weighted average duration of the Fed's SOMA portfolio. We show that our main results also qualitatively hold if we assume the modified duration to be 5 or 10. In Appendix B.8, we use 5-year and 10-year zero-coupon government bond yields instead of the 8-year zero-coupon yields in the baseline.

In Appendix C, we provide further details about the principal components analysis that we use to generate our monetary policy surprise instruments.

4 Implications for quantitative tightening

In this section, we use our estimates of yield elasticity to conduct scenario analyses for hypothetical QT configurations by the Federal Reserve. Specifically, we try to shed light on two key questions: How much will equilibrium yields need to move for non-central-bank participants to clear the market as quantitative tightening releases government bond supply to the market (either through the absence of central bank purchases or active central bank sales)? How will the market shares of different sectors evolve as the market finds its new equilibrium once central bank holdings shrink?

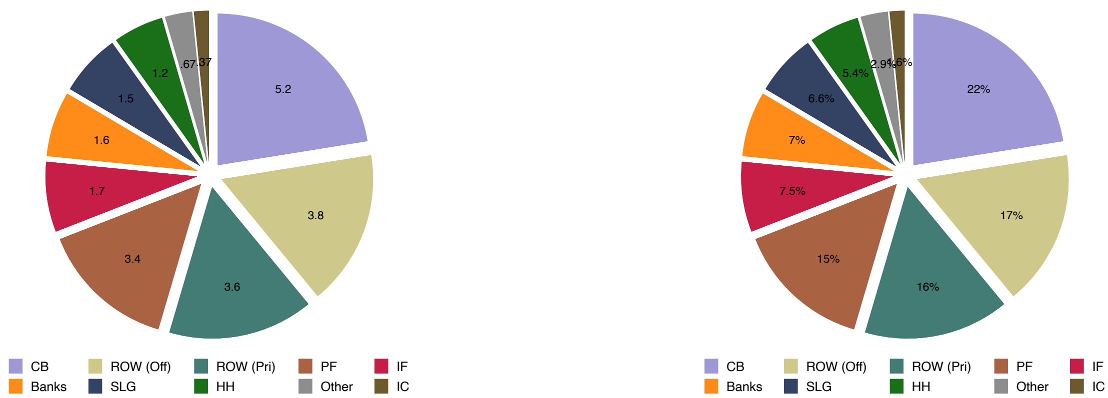
In these back-of-the-envelope calculations, we compute the increase in the holdings of non-central bank participants in response to hypothetical changes in the 8-year yield using the elasticity estimates from the previous section. As discussed above, we choose to focus on the 8-year yield since the average maturity of the SOMA portfolio Treasury holdings is around 8 years. We compute the hypothetical changes for the aggregate holdings of non-central bank participants using the elasticity estimates from Panel (a) of Table 3. Note that this scenario analysis rests on a ceteris paribus assumption – i.e. that everything else appearing in $\mathbf{X_t}$ and $\mathbf{W_t}$ is assumed to remain constant.

Based on elasticity estimates for each sector s reported in the Panel (a) of Table 3, we estimate the hypothetical change in demand for an array of changes in 8-year yield via the following formula:

$$\widehat{\Delta H}^{s}_{\Delta Y^8} = \Big(exp(\Delta Y^8 \times \widehat{\beta}^s) - 1\Big) \times H^s_{2022q3}, \tag{8}$$

where $\widehat{\beta}_1^s$ is the estimate of the 8-year yield elasticity of demand by sector s, H^s_{2022q3} is the total holdings of US Treasuries by each sector s in 2022q3 as reported in Figure 2, and the resulting $\widehat{\Delta H}^s_{\Delta Y_8}$ denotes the estimated change in demand by sector s for US Treasuries in response to the change in the 8-year yield denoted by ΔY^8. This equation naturally follows from the specification in Equation (6).

Figure 2: Total holdings by sector (exc. MMFs) at 2022 Q3



(a) Total holdings ($ trillions) **(b)** Percent of total held by sector

Note: Panel (a) of this figure shows the total market value of holdings by sector at 2022Q3 in the United States. Panel (b) shows the market shares of each sector. CB refers to the central bank. IF refers to investment funds (i.e. open-ended mutual funds, ETFs, closed-ended funds). Banks refer to commercial banks, PF refers to pension funds. IC refers to insurance companies. ROW (Pri) refers to foreign private investors. ROW (Off) refers to foreign official investors. SLG refers to state and local governments. HH refers to households. Other refers to all other sectors except money market funds.

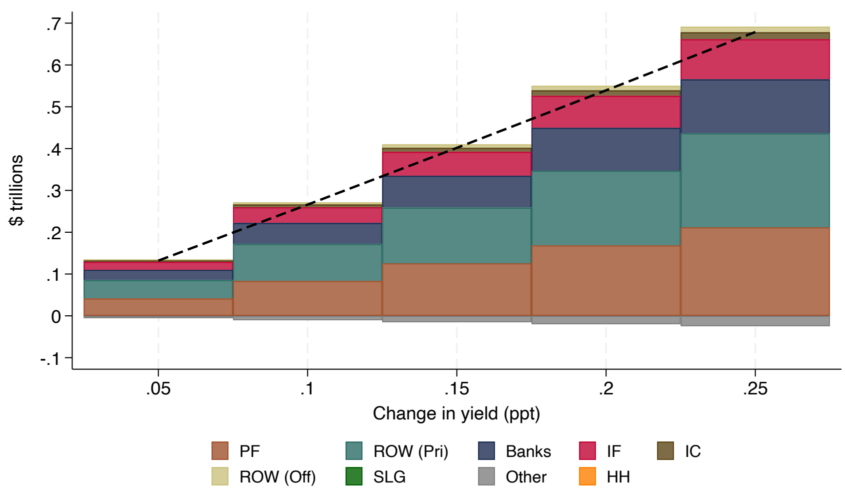
Using Equation (8), we are able to estimate how much each sector would change its demand for government debt in response to hypothetical changes in yield. Adding these dollar amounts up, we get a relationship between the size of QT and the market-clearing yields.[22]

Our scenario analysis maps how demand by each sector would change for a given change in yield as shown in Figure 3. We interpret the sum of these dollar amounts as the impact of the magnitude of QT on market-clearing yields. According to our estimates, non-central bank players would absorb a QT size of $266 billion at a market-clearing change of 10 basis points in the 8-year yield. The major increases in holdings would come from foreign private investors ($88 billion), pension funds

[22]Note that we conduct this scenario analysis only at the point estimates. Interested reader can compute the confidence intervals using information given in Table 3 and Figure 2(b).

($83 billion), commercial banks ($50 billion) and investment funds ($37 billion). These estimates imply that at the current QT rate of $60 billion per month ($180 billion per quarter), the quarterly market-clearing impact on long-term yields is around 7 basis points. Using estimates of Koijen, Koulischer, Nguyen and Yogo (2021), we do a similar analysis for the Euro area and find that a QT size of €188 billion would lead to a market-clearing change of 10 basis points in long-term yields, mostly driven by demand by foreign investors (see the Internet Appendix for details).

Figure 3: Implied absorption capacity by all sectors for different changes in yield in the United States

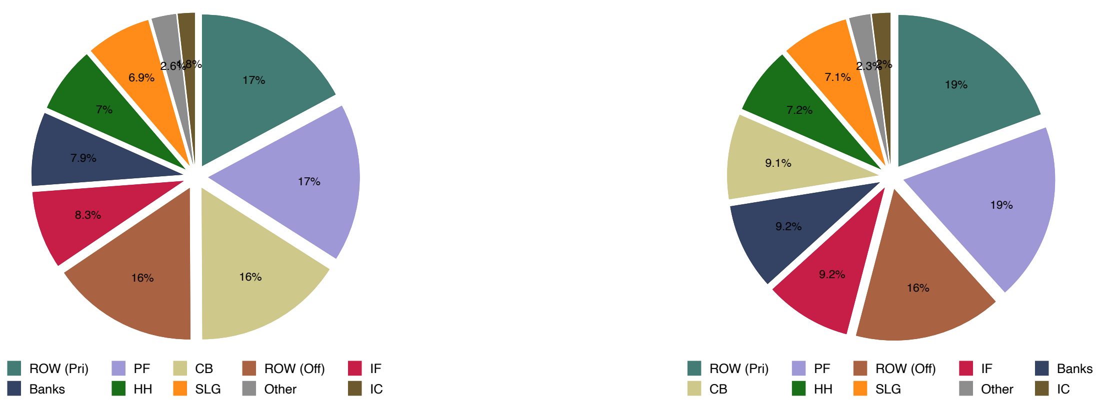


Note: This figure plots the relationship between $\widehat{\Delta H}^{s}_{\Delta Y^8}$ and ΔY^8 using Equation (8) based on the second stage estimates from Panel (a) of Table 3 where we make valuation adjustments to holdings of each sector assuming a modified duration of 8. The stacked bars correspond to how much demand by each sector would change in response to changes in yield using the elasticity estimates from Panel (a) of Table 3 and the latest holdings in Figure 2. The dashed line is the total change in demand (as we vary the yield along the horizontal axis), obtained by adding up all stacked bars.

We show how market shares of all players would evolve under two different QT scenarios in Figure 4, extrapolating our estimates. In Panel (a) of Figure 4, we show that, all else constant, a QT of size $1.2 trillion, which is the baseline projection by the Federal Reserve (Anderson, Marks, Na, Schlusche and Senyuz, 2022), would increase the market share of foreign investors from 15% to 17%, that by pension funds from 14% to 17%, that by investment funds from 7% to 8%, and that by commercial banks from 7% to 8%. Along these shifts in relative holdings the equilibrium 8-year yield would be pressured up by 45 basis points in in total. In Panel (b) of Figure 4, we show that, all else constant, a larger QT of size $3 trillion, which would undo most of the post-Covid increase in the central bank balance sheet, would make foreign private investors and pension funds the largest investors with a market share of 19%. The market share of the central bank would go down to 9% from 22%. Investment funds and commercial banks would also see an increase in their market shares to around 9% each. Under both scenarios, the foreign official sector keeps its market share roughly constant.

These shifts could have implications for the pricing of Treasuries going forward. As the share of price-elastic investor groups increases, this has implications for how yields change in response to changes in the supply of government debt.

Figure 4: Counterfactual market shares for different QT configurations



(a) QT=$1.2 trillion & $\Delta 8Y = 0.45\%$ **(b)** QT=$3 trillion & $\Delta 8Y = 1\%$

Note: Panel (a) of the figure shows the hypothetical market shares of different sectors following a QT size of $1.2 trillion which corresponds to a 45 basis point change in the 8-year yields according to our estimates. Panel (b) of the figure shows the hypothetical market shares of different sectors following a QT size of $3 trillion which corresponds to a 1 percentage point change in the 8-year yields according to our estimates. CB refers to the central bank. ROW (Pri) refers to foreign private investors. ROW (Off) refers to foreign official investors. PF refers to pension funds. IF refers to investment funds (i.e. open-ended mutual funds, ETFs, closed-ended funds). Banks refer to commercial banks. HH refers to households. SLG refers to state and local governments. IC refers to insurance companies. Other refers to all other sectors except money market funds.

Finally, it is useful to compare our estimates quantitatively to those in the literature on the impact of QE on long-term yields. Borio and Zabai (2018) summarize the range of estimates in the literature of the impact of QE2, in which the Federal Reserve purchased around $660 billion long-term US Treasuries and sold shorter-term Treasuries, on the 10-year yield (Krishnamurthy and Vissing-Jorgensen, 2011; Swanson, 2011; Hamilton and Wu, 2012; d'Amico, English, López-Salido and Nelson, 2012). These estimates range from a 16 basis points to a 45 basis points decline in the 10-year yields. Since our sample period covers a longer period, our estimates are not directly comparable to these estimates in the literature. However, if we use our estimates to gauge the hypothetical impact of QT of roughly a similar amounts as the purchases during these programmes, i.e. sales of $667 billion government debt by the central bank, we find that the impact would correspond to a change in 8-year yields of 25 basis points. These numbers are roughly similar to the estimates of the impact of QE in the previous using different methodologies (mostly in the form of event studies). While in this paper, our main focus in the scenario analyses has been the unwinding of central bank balance sheets, our estimates can in principle be also be used to gauge the price impact of QE policies in a symmetric fashion.

5 Discussion and forward-looking considerations

Our elasticity estimates are based on historical relationships between yields and investor holdings. In this section, we discuss several forward-looking considerations that are not captured by our estimates, but which are likely to influence the demand for government bonds going forward. We also briefly discuss the policy implications of a shift in the composition of holders of government debt.

Foreign private and official investors. Foreign investors are currently the largest non-central bank holders of US Treasuries. Foreign official investors hold Treasuries to the tune of 16%, while foreign private investors hold 15% of outstanding Treasury securities. Our estimates suggest that the foreign official sector is largely insensitive to changes in yield, which is reasonable as reserve accumulation is likely influenced by autonomous factors. Foreign private investors, by contrast, have an elastic demand curve, which in turn suggests an ongoing important role for these investors in the US Treasury market.

Going forward, in addition to adjustment in yields, a couple of other factors may affect the Treasury demand from foreign investors. Foreign private investors often purchase US Treasuries on an FX-hedged basis. Their demand might therefore be influenced by how FX hedging costs change.[23] In addition, holdings of US Treasuries by foreign official investors might be influenced by geopolitical factors as well as FX reserve accumulation. The evolution of the latter depends on a number of factors, including the future of the dollar and other currencies as a reserve currency and the role of FX interventions in the central bank toolkit.

US commercial banks. Our analysis has shown that commercial banks are likely to play an important role in government debt markets as central bank balance sheets normalize. In this context, it is worth noting that commercial banks have already notably increased their demand for US Treasuries significantly since 2014, in part to comply with the liquidity coverage ratio (LCR) of Basel III regulations as well as for reasons of internal liquidity risk management. To comply with the LCR, banks need to hold sufficient amount of high-quality liquidity assets (HQLA) to cover cash outflows during a long period of liquidity stress lasting 30 days. High-quality liquidity assets are either central bank reserves or other assets that can be converted into cash quickly through sales (or by being pledged as collateral) with no significant loss of value. Central bank reserves and US Treasury securities are considered among the highest quality of these liquid assets, making them highly substitutable for banks to comply with LCR.

We show suggestive evidence in Table 4 using US call reports data that banks indeed respond to changes in the relative yields between government bonds and the interest on reserve balances, especially since the introduction of the LCR in 2015. To illustrate, using the bank (b) and quarter (t) panel dataset from call reports between 2009Q2 (following the first QE) and 2021Q2, we run the following regressions:

$$\begin{aligned} UST\ share\ in\ HQLA_{bt} = {} & \beta_1\ (Y^m - IOR)_t + \beta_2\ \mathbb{1}(LCR)_t + \beta_3\ (Y^m - IOR)_t \times \mathbb{1}(LCR)_t \\ & + \beta_4\ log(Size\ of\ CB\ BS)_t + \beta_5\ log(Assets)_{bt} + \beta_6\ Leverage_{bt} + \theta_b + \varepsilon_{bt}, \end{aligned} \tag{9}$$

[23] FX hedging depends on the cross-currency basis. As such, the relative slope of the yield curve between the US and other currency areas plays an important role in affecting foreign investor demand.

where $UST\ share\ in\ HQLA_{bt}$ denotes the share of US Treasury securities in Level-1 HQLA assets, namely US Treasuries and reserves, for a bank b at quarter t, $(Y^m - IOR)_t$ is the spread between the yield on m-year government debt and the interest on reserve balances paid by the Federal Reserve at t $\mathbb{1}(LCR)_t$ is a dummy variable which is one after the implementation of the LCR in January 2015, $log(Size\ of\ CB\ BS)_t$ is the log size of the balance sheet of the Federal Reserve, $log(Assets)_{bt}$ is the log total assets of a bank, $Leverage_{bt}$ is measured as the bank's liabilities over assets, and θ_b refers to bank fixed effects.

Table 4: US banks' portfolio allocation between different types of HQLA

	(1)	(2)	(3)
VARIABLES	UST share	UST share	UST share
1Y yield-IOR	-0.02		
	(0.04)		
(1Y yield-IOR) X LCR	0.08*		
	(0.05)		
5Y yield-IOR		0.02***	
		(0.01)	
(5Y yield-IOR) X LCR		0.03**	
		(0.01)	
10Y yield-IOR			0.01*
			(0.01)
(10Y yield-IOR) X LCR			0.02**
			(0.01)
LCR	0.00	-0.00	-0.00
	(0.02)	(0.01)	(0.02)
Observations	318,535	318,535	318,535
R-squared	0.79	0.79	0.79
Bank FE	✓	✓	✓
Controls	✓	✓	✓

Note: The Table reports the coefficients of the OLS regression of Equation (9) using a bank-quarter panel dataset between 2009Q2 and 2021Q2 with 8,587 banks. Standard errors are clustered at the bank level. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.
Source: US Call Reports

The results suggest that after the implementation of the LCR, commercial banks held more HQLA in the form of Treasuries as opposed to reserves when the spread between government bond yields and the interest on reserves (IOR) paid by the Federal Reserve was higher. The results are more pronounced for lower maturity government bonds. This result is consistent with banks substituting between liquid assets to comply with the LCR by responding to differences in return between government bonds and reserves. As central banks normalize balance sheets, a reduction in reserves might increase the demand for Treasuries by banks further than what has been captured by our estimates, resulting in an even more important role played by banks in the US Treasury market. At the same time, the recent events triggered by the collapse of the Silicon Valley Bank might also have implications for the demand for government debt by commercial banks as a form of HQLA.

An interesting counterfactual exercise using our elasticity estimates is to study the potential impact of banks not stepping in to purchase government bonds during QT due to worries about solvency following the SVB crisis. In this case, setting the elasticity coefficient to zero, we get that instead of our baseline estimates of $266 billion QT moving long-term yields by 10 basis points, we find that in the absence of bank purchases only $215 billion of QT suffices to increase long-term yields by the same amount.

US non-bank financial institutions. Our estimates suggest that pension funds and investment funds are two key players in government bond markets with an elastic demand. Pension funds are important players whose demand matters especially at the long end of the yield curve. Our estimates suggest that they are most yield-sensitive and are hence likely to play the most important role in adsorbing net supply of Treasuries.

While we group investment funds together for our analysis, there are substantial differences among them. It is important to differentiate them when we draw policy implications. In the Internet Appendix E, we show that open-ended mutual funds are the largest holder among all investment funds with the market values of their holdings exceeding $1.5 trillion at its peak. The value of holdings of exchange-traded funds is close to $500 billion, while the total holdings of closed-ended funds are lower than $50 billion. It is quite intuitive that the yield elasticity estimate we obtain for investment funds overall is primarily driven by open-ended mutual funds, whereas those for ETFs and closed-end funds are statistically indistinguishable from zero.

Open-ended mutual funds typically promise daily redemptions and operate with a certain degree of liquidity mismatch. These characteristics lead to a first mover advantage in redemptions and make these funds vulnerable to runs (Falato, Goldstein and Hortaçsu, 2021). If these funds hold a mix of illiquid assets combined with more liquid government bonds, they might sell more liquid bonds first to raise liquidity in the face of redemptions (Huang, Jiang, Liu and Liu, 2021). This might generate an externality and add to selling pressure in the aggregate. Indeed, these risks materialized and were on display during the March 2020 market turmoil (e.g. Vissing-Jorgensen, 2021). Therefore, a greater overall footprint of open-ended mutual funds as quantitative tightening progresses implies a greater urgency to address the externalities in this sector.

6 Conclusion

In this paper we study the demand for government debt in core markets— the United States, Euro area, Japan and United Kingdom. We start by illustrating the dramatic changes in the composition of government debt holders over the past decades, shedding light on the varying role of key investor groups. We also run regressions that measure the marginal response of each sector to changes in outstanding government debt. This helps us quantify the time-varying marginal role of each sector including the central banks in absorbing government debt supply across jurisdictions.

Next, focusing on the United States and taking a market-clearing perspective using monetary policy surprises as instrumental variables, we estimate the yield elasticity of demand of each sector for government debt. We find that most sectors have a downward-sloping demand curve (with respect to price) and some have inelastic demand. Among the different investor groups, pension funds, foreign private investors, commercial banks and investment funds have the most elastic demand. On the other hand, we find the demand functions of the foreign official sector and the household sector to be inelastic. We use our estimates to infer the market-clearing yields for assumed changes in the debt supply as the central bank shrinks its balance sheet. According to our estimates, a \$266 billion quarterly reduction in the central bank balance sheet results in a 10 basis point rise in long-term bond yields per quarter.

Our results suggest that as central banks' role in government debt markets shrinks, the role played by the most elastic sectors is set to increase. Our results thus have policy implications, not least as some of these investor groups played an important role amid the dash for cash during the Covid-19 crisis. Similarly, as commercial banks' role is set to increase, more research is needed to understand the incentives and constraints of commercial banks pertaining to government debt.

We view our analysis as a step to understand the role different sectors play in government debt markets and estimate how financial markets will adjust to a new equilibrium as central banks' footprint declines. To do so, we use publicly available quarterly data to conduct our analysis which, while comprehensive, also has limitations and necessitates certain assumptions. Future work could leverage more granular country-level data to run similar analyses overcoming some of the data limitations and to extend the analysis to different countries.

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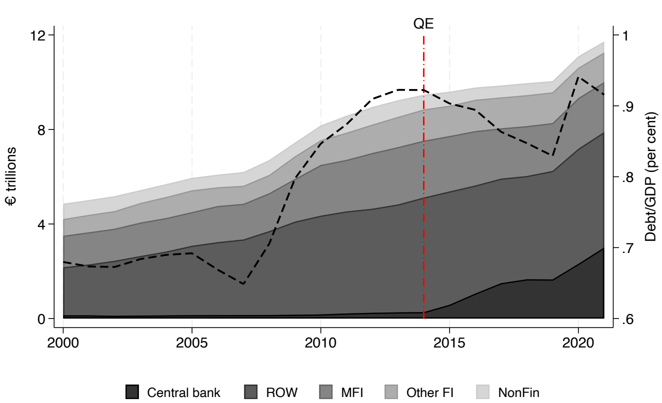
A The evolution of government debt holdings and marginal buyers for the Euro area, Japan and the United Kingdom

In this section, we repeat the same analysis as in Section 2 for the Euro area, Japan and the United Kingdom.

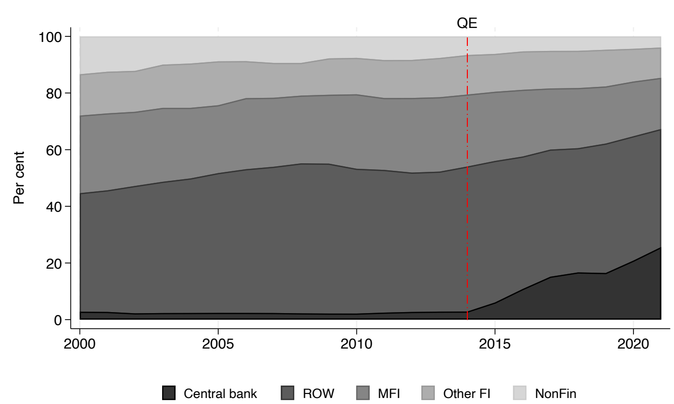
A.1 Euro area

In the Euro area, QE also brought about notable compositional shifts in the holders of government debt. We report the changes in the average shares of each sector in Table 5 Panel (a). The average share of the central bank in the holdings of government bonds has increased from 2.7% to 16%. The share of monetary financial institutions, which is comprised to a large extent of commercial banks, has fallen from 27% to 21%. The share of non-financial corporations has also dropped from 11.5% to 5%. While the share of foreign investors stayed roughly constant at 45%, that of other financial institutions (which include mutual funds, pension funds, insurance companies) has declined from 14% to 12.6%.

Figure 5: Total holdings and shares of different sectors - Euro Area



(a) Total holdings



(b) Shares

Note: Panels 5(a) and 5(b) show the total market value of the government debt holdings and market shares of each sector in the Euro area, respectively, between 2000 and 2021 (yearly data). Central bank refers to the holdings of the ECB/Eurosystem. ROW refers to foreign investors (official and private). MFI refers to monetary financial institutions, such as banks and money market funds. Other FI refers to non-monetary financial institutions, such as pension funds, insurance companies, mutual funds. QE starts in 2015. *Source: European Central Bank.*

The role of the central bank as a major player in European sovereign debt markets clearly stands out, as indicated by the regression results reported in Table 5 Panel (b). The central bank has absorbed 67% of every unit of new government debt since the beginning of the ECB/Eurosystem's public sector asset purchases programme in March 2015. This is in contrast with the estimate of -17% prior to QE (at a very low base as the central bank holdings of government debt averaged only 2.7% prior to the quantitative easing programme), likely owing to the fact that the Eurosystem had been shedding bond holdings at a time when total government debt supply increased pre-GFC.

Table 5: Average shares and marginal response by sectors in the Euro area

Panel (a): Average shares by sector over different periods

Avg. Share	CB	ROW	OtherFI	MFI	NonFin
Pre-QE	2.7	44.7	14	27.1	11.5
QE	16	45.2	12.6	21.2	5

Panel (b): Marginal holdings by sector over different periods

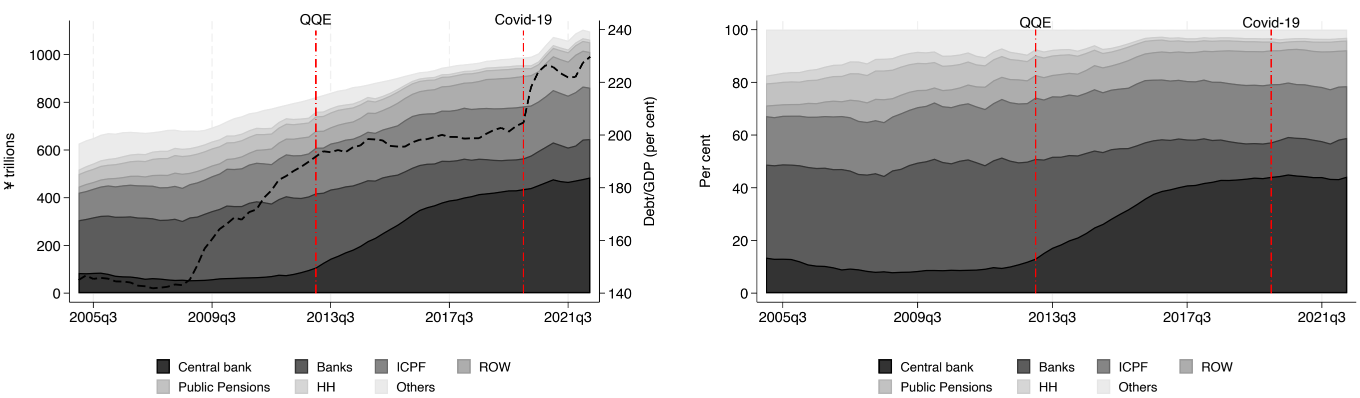
	(1)	(2)	(3)	(4)	(5)
VARIABLES	CB	ROW	OtherFI	MFI	NonFin
Pre-QE * Pct. Ch. Gov. Debt	-0.17*	0.46***	0.23***	0.42***	0.06***
	(0.10)	(0.12)	(0.04)	(0.02)	(0.02)
QE * Pct. Ch. Gov. Debt	0.67***	0.04	0.01	0.22***	0.06*
	(0.12)	(0.15)	(0.03)	(0.02)	(0.04)
Observations	26	26	26	26	26
R-squared	0.73	0.43	0.37	0.69	0.05

Note: Panel (a) shows the average share of each investor group over different time periods. Panel (b) reports the coefficients of the OLS regression of Equation (2) for the Euro area. CB refers to the holdings of the ECB/Eurosystem. ROW refers to foreign investors (official and private). MFI refers to monetary financial institutions, such as banks and money market funds. Other FI refers to non-monetary financial institutions, such as pension funds, insurance companies, mutual funds. Data are yearly. Pre-QE is between 2000 and 2014. QE is between 20015 and 2021. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** p<0.01, ** p<0.05, * p<0.1. *Source: European Central Bank*

The responsiveness of each sector to changes in government debt also uncovers an important story. According to our estimates, prior to QE, for an additional increase in government debt, the marginal purchase response of different sectors was 46% for foreign investors, 42% for monetary financial institutions, 23% for non-monetary financial institutions, and 6% for non-financial corporates. Following the launch of the Eurosystem's public sector bond purchases, these numbers have dropped to 22% for monetary financial institutions, to 4% for foreign investors and to 1% for non-monetary financial institutions, with both of the latter being statistically insignificant.[24]

Even though our estimates cannot directly speak to which sectors were net sellers to the central bank, they are nonetheless consistent with the estimates of Koijen, Koulischer, Nguyen and Yogo (2021) in a study of portfolio rebalancing during initial phases of the ECB's PSPP programme. During that time period, they find based on a confidential security-level dataset that foreign investors sold €0.40, banks sold €0.20 and mutual funds sold €0.06 per unit purchased by the Eurosystem.

Figure 6: Total holdings and shares of different sectors - Japan



(a) Total holdings

(b) Shares

Note: Panels 6(a) and 6(b) show the total market value of the government debt holdings and market shares of each sector in Japan, respectively, between 2005Q1 and 2022Q2 (quarterly data). Central bank refers to the holdings of the Bank of Japan. Banks refer to commercial banks. ICPF refers to insurance companies and pension funds. ROW refers to foreign investors (official and private). HH refers to households. Others combine all other remaining sectors. Qualitative and quantitative easing (QQE) starts at 2013Q2. Covid-19 is at 2020Q1. *Source: Bank of Japan.*

[24] If we exclude 2020 and 2021 from the estimation, we find a large negative coefficient for foreign investors in the QE period compensated mostly by the central bank, which is in line with foreign investors selling bonds to the central bank. This is in line with the findings of Koijen, Koulischer, Nguyen and Yogo (2021).

A.2 Japan

In Japan, the expansion of the central bank's footprint in government debt markets has gone furthest with its share of government bond holdings reaching around 44% since the Covid-19 crisis (Table 6 Panel (a)). Along with the central bank, holdings of Japanese government bonds by foreign investors also increased from 6.7% prior to the qualitative and quantitative easing programme to 13.3% recently.[25] The share of insurance companies and pension funds remained constant at around 20%. As a flip-side of the rise in the Bank of Japan's holdings, the share of banks fell from 38% to 14%, the share of public pensions declined from 9% to 3%, the share of households declined from around 4% to 1%, and the share of all other sectors declined from 11% to 3%.

The estimates of the marginal response by each sector in Japan are consistent with an outsized role by the central bank purchases, especially since the beginning of the QQE programme until the Covid-19 crisis. During this period, the central bank bought 1.85 units of government bonds per unit of increase in the total amount outstanding of government bonds, that is, Bank of Japan purchases even exceeded the amounts of new debt placed by the government in markets. Banks, on the other hand, reduced their holdings by 1.03 units per unit of increase in government debt, suggesting that banks were the major sellers to the central bank in this episode. Since the Covid-19 crisis, the marginal response by each sector has become more balanced even as the marginal role of the central bank still remains the highest.

A.3 United Kingdom

The rate at which the share of the central bank has grown in UK debt markets following QE is only second to Japan. Since then, the share of the central bank has increased from essentially zero to roughly one third of the market (Table 7 Panel (a)). Similar to Japan, the share of foreign investors has also picked up from 19.9% to 28.1%. The share of all other sectors, apart from insurance companies and pension funds, remained roughly stable. In case of the latter, the drop in

[25]The rise in the role of foreign investors might be reflecting the arbitrage trade foreigners do to take advantage of covered interest parity deviations as they swap dollars for Japanese yen and invest those in (mostly short-term) Japanese government debt securities (Rime, Schrimpf and Syrstad, 2022, see, for example,)

Table 6: Average shares and marginal response by sectors over time in Japan

Panel (a): Average shares by sector over different periods

Avg. Share	CB	Banks	ICPF	ROW	PP	HH	Others
Pre-QQE	9.9	38.6	20.6	6.7	9.1	3.7	11.4
QQE	33.7	22.2	22.6	10.4	5	1.4	4.7
Post-Covid	44.2	14.1	20.6	13.3	3.4	1.1	3.3

Panel (b): Marginal holdings by sector over different periods

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	CB	Banks	ICPF	PP	HH	ROW	Others
Pre-QQE * Pct. Ch. Gov. Debt	0.15	0.41***	0.23***	-0.02	0.01	-0.02	0.24*
	(0.14)	(0.06)	(0.09)	(0.05)	(0.04)	(0.06)	(0.13)
QQE * Pct. Ch. Gov. Debt	1.85***	-1.03***	0.09	-0.17**	-0.06	0.07	0.25
	(0.27)	(0.17)	(0.14)	(0.08)	(0.04)	(0.12)	(0.22)
Post-Covid * Pct. Ch. Gov. Debt	0.33***	0.25***	0.13***	0.04**	0.00	0.17***	0.09*
	(0.10)	(0.04)	(0.03)	(0.02)	(0.01)	(0.03)	(0.05)
Observations	69	69	69	69	69	69	69
R-squared	0.60	0.39	0.21	0.13	0.06	0.06	0.04

Note: Panel (a) shows the average share of each investor group over different time periods. Panel (b) reports the coefficients of the OLS regression of Equation 2 for Japan. CB refers to the holdings of the Bank of Japan. Banks refer to commercial banks. ICPF refers to insurance companies and pension funds. ROW refers to foreign investors (official and private). PP refers to public pensions. HH refers to households. Others combine all other remaining sectors. Qualitative and quantitative easing (QQE) starts at 2013Q2. Covid-19 is at 2020Q1. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** p<0.01, ** p<0.05, * p<0.1. *Source: Bank of Japan*

average shares from 63.6% to 28.1% indicates that insurance companies and pension funds were in fact the main sellers to the central bank during the Bank of England's QE.

The marginal response by each sector, reported in Table 7 Panel (b), also shows that the central bank effectively replaced the demand for government debt that came previously from insurance companies and pension funds. When the Bank of England's balance sheet expanded with QE, the marginal response of the central bank to an additional unit of increase in the government debt rose

Table 7: Average shares and marginal response by sectors over time in the UK

Panel (a): Average shares by sector over different periods

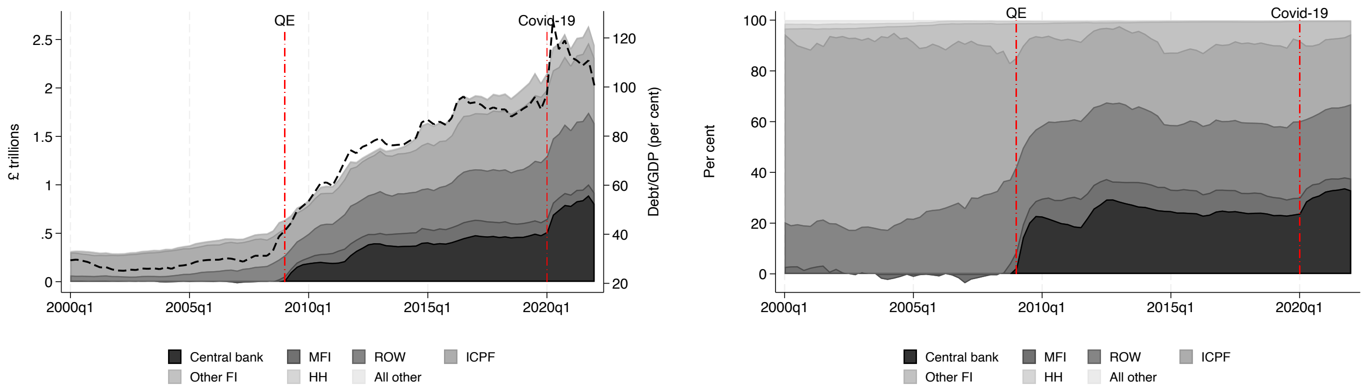
Avg. Share	CB	ROW	ICPF	MFI	HH	OFI	AllOther
Pre-QE	0	19.9	63.6	6.4	1.9	6.9	1.2
QE	23.4	28.7	31.4	8.3	.5	7.5	.2
Post-Covid	31	28.1	28.1	4.9	.2	7.6	.1

Panel (b): Marginal holdings by sector over different periods

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	CB	ROW	ICPF	MFI	HH	OFI	All Other
Pre-QE * Pct. Ch. Gov. Debt	-0.00	0.17***	0.47***	0.12***	0.01***	0.23***	-0.00
	(0.03)	(0.02)	(0.08)	(0.03)	(0.00)	(0.05)	(0.00)
QE * Pct. Ch. Gov. Debt	0.38***	0.18***	0.21***	0.10***	0.00***	0.12	-0.00
	(0.05)	(0.04)	(0.03)	(0.03)	(0.00)	(0.12)	(0.00)
Post-Covid * Pct. Ch. Gov. Debt	0.51***	0.14***	0.13***	-0.00	0.00***	0.22***	-0.00
	(0.07)	(0.03)	(0.03)	(0.01)	(0.00)	(0.02)	(0.00)
Observations	140	140	140	140	140	140	140
R-squared	0.49	0.35	0.57	0.24	0.40	0.28	0.00

Note: Panel (a) shows the average share of each investor group over different time periods. Panel (b) reports the coefficients of the OLS regression of Equation 2 for the United Kingdom. CB refers to the holdings of the Bank of England. MFI refers to monetary financial institutions, such as banks and money market funds. ROW refers to foreign investors (official and private). ICPF refers to insurance companies and pension funds. HH refers to households. All other refers to a combination of all other remaining sectors. QE starts at 2009Q1 and Covid-19 is at 2020Q1. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** p<0.01, ** p<0.05, * p<0.1. *Source: Office of National Statistics*

Figure 7: Total holdings and shares of different sectors - United Kingdom



(a) Total holdings

(b) Shares

Note: Panels 7(a) and 7(b) show the total market value of the government debt holdings and market shares of each sector in the United Kingdom, respectively, between 2000Q1 and 2022Q1 (quarterly data). Central bank refers to the holdings of the Bank of England. MFI refers to monetary financial institutions, such as banks and money market funds. Negative values for MFIs are due to market making activities. ROW refers to foreign investors (official and private). ICPF refers to insurance companies and pension funds. HH refers to households. All other refers to a combination of all other remaining sectors. QE starts at 2009Q1 and Covid-19 is at 2020Q1. *Source: Office for National Statistics.*

to 38%. Since March 2020, this response has further increased to 51%. As a mirror image, the marginal purchase of insurance companies and pension funds of a unit increase in government debt first declined from 47% to 21% during the QE period, which further declined to a mere 13% since the Covid-19 crisis.

It is well-established by now that the demand for duration by the UK pension and insurance sector, however, overall increased during the period of low interest rates. As Domanski, Shin and Sushko (2017) point out, as the duration of liabilities was rising faster than that of assets in the period when interest rates fell, this convexity effect meant that these players faced pressure to engage in a hunt for duration. The shrinking holdings of government bonds by these sectors documented above indicates that the demand for duration was largely met synthetically through derivatives and leveraging up long-maturity gilt holdings via repo – a factor which amplified the liquidity crisis in the gilt market of October 2022 (see, e.g. Pinter, 2023, for a forensic analysis).

B Additional tables and figures

B.1 Results with all controls

In this section, we report the full estimation results of the first stage and the second stages of Equations (6) (Table 8 reports the first-stage and Table 10 reports the second stage) and (7) (Table 9 reports the first-stage and Table 11 reports the second stage).

In particular, the elasticities with respect to the other variables included in the regressions give further information on the investment behaviors of various sectors. For example, the coefficient in the second stage for the 5-year German yield is negative, consistent with it being defined as the outside asset. In addition, for many sectors the coefficient for log GDP is negative suggesting lower holdings of Treasuries in boom periods.

Table 8: First stage results with alternative specifications of monetary policy surprises - With control variables

VARIABLES	(1) 8Y Yield (ZC)	(2) 8Y Yield (ZC)	(3) 8Y Yield (ZC)	(4) 8Y Yield (ZC)	(5) 8Y Yield (ZC)	(6) 8Y Yield (ZC)	(7) 8Y Yield (ZC)
Swanson FFR	-0.0002						
	(0.0003)						
Swanson FG	0.0011***						
	(0.0004)						
Swanson LSAP	0.0001						
	(0.0001)						
BRW		0.0005*					
		(0.0003)					
KSX (3M)			0.0012***				
			(0.0004)				
KSX (2Y)			-0.0006				
			(0.0007)				
KSX (10Y)			0.0010**				
			(0.0004)				
JK MP				0.0000			
				(0.0003)			
JK CBI				0.0014***			
				(0.0003)			
PCA 1 (BRW, Swanson, JK)					0.0003		
					(0.0002)		
PCA 1 (All)						0.0005***	
						(0.0001)	
PCA 1 (BRW, KSX)							0.0008***
							(0.0001)
log GDP	0.0370	0.0193	0.0496	0.0490	0.0407	-0.0380	0.0324
	(0.0429)	(0.0307)	(0.0506)	(0.0369)	(0.0468)	(0.0495)	(0.0543)
GDP growth	0.2061***	0.0088	0.0047	0.1190*	0.1911***	0.1924***	0.0224
	(0.0589)	(0.0373)	(0.0444)	(0.0635)	(0.0568)	(0.0355)	(0.0502)
inflation	0.0029	-0.0383	-0.0029	-0.0426	0.0170	-0.0604	-0.0456
	(0.0545)	(0.0585)	(0.0766)	(0.0817)	(0.0654)	(0.0578)	(0.0691)
log broad dollar index	0.0185**	0.0149**	0.0475***	0.0188***	0.0175**	0.0410**	0.0499***
	(0.0075)	(0.0071)	(0.0096)	(0.0061)	(0.0078)	(0.0178)	(0.0115)
log VIX	-0.0017	-0.0067***	-0.0055***	-0.0016	-0.0022	-0.0026*	-0.0066***
	(0.0019)	(0.0016)	(0.0007)	(0.0016)	(0.0018)	(0.0015)	(0.0007)
5Y German yield (zc)	0.6543***	0.5543***	0.7110***	0.6616***	0.6440***	0.8761***	0.7418***
	(0.1103)	(0.1115)	(0.1135)	(0.1007)	(0.1136)	(0.1246)	(0.1411)
trend	-0.0026**	-0.0015*	0.0017**	-0.0029***	-0.0027**	-0.0006	0.0023***
	(0.0012)	(0.0008)	(0.0008)	(0.0010)	(0.0013)	(0.0016)	(0.0009)
trend squared	0.0000***	0.0000**	-0.0000***	0.0000***	0.0000***	0.0000	-0.0000***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Observations	102	112	74	102	102	60	70
R-squared	0.9548	0.9427	0.9031	0.9549	0.9512	0.9207	0.9058
Sample	1994q2-2019q2	1994q2-2021q4	2004q3-2022q4	1994q2-2019q2	1994q2-2019q2	2004q3-2019q2	2004q3-2021q4
Effective F-stat	4.80	3.36	3.22	5.68	1.36	43.62	27.54
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	18.89	23.11	19.46	9.96	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation 6. The sample period varies depending on the availability of data across different monetary policy surprises. All reported right-hand side variables are normalized. Effective F-stat is calculated using the methodology in Olea and Pflueger (2013). The final row reports the critical values of a test of weak instruments with a 5% confidence level and a 10% worst-case bias. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 9: First stage results with alternative specifications of monetary policy surprises - using information on outside asset - with control variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)
Swanson FFR	-0.0000						
	(0.0005)						
Swanson FG	0.0010**						
	(0.0005)						
Swanson LSAP	-0.0004						
	(0.0004)						
BRW		0.0002					
		(0.0004)					
KSX (3M)			0.0019***				
			(0.0006)				
KSX (2Y)			-0.0014*				
			(0.0007)				
KSX (10Y)			0.0020***				
			(0.0007)				
JK MP				0.0002			
				(0.0005)			
JK CBI				0.0011***			
				(0.0004)			
PCA 1 (BRW, Swanson, JK)					0.0002		
					(0.0004)		
PCA 1 (All)						0.0006***	
						(0.0002)	
PCA 1 (BRW, KSX)							0.0010***
							(0.0003)
log GDP	0.1639***	0.1239***	0.2003***	0.1759***	0.1681***	0.1551***	0.1542***
	(0.0369)	(0.0240)	(0.0343)	(0.0269)	(0.0400)	(0.0562)	(0.0405)
GDP growth	0.0776	-0.0215	-0.0661**	0.0103	0.0877	0.0981	-0.0154
	(0.0736)	(0.0337)	(0.0261)	(0.0429)	(0.0722)	(0.0779)	(0.0396)
Inflation	-0.0001	-0.0008	-0.0008	-0.0005	0.0000	-0.0005	-0.0007
	(0.0009)	(0.0007)	(0.0009)	(0.0009)	(0.0007)	(0.0009)	(0.0006)
log broad dollar index	-0.0013	-0.0056	0.0303**	-0.0013	-0.0026	0.0125	0.0343**
	(0.0075)	(0.0059)	(0.0135)	(0.0041)	(0.0072)	(0.0173)	(0.0141)
trend	-0.0059***	-0.0042***	-0.0021**	-0.0061***	-0.0059***	-0.0042**	-0.0008
	(0.0011)	(0.0007)	(0.0009)	(0.0009)	(0.0012)	(0.0021)	(0.0009)
trend squared	0.0000***	0.0000***	-0.0000	0.0000***	0.0000***	0.0000	-0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Observations	102	112	74	102	102	60	70
R-squared	0.9176	0.9078	0.8094	0.9164	0.9138	0.7917	0.8115
Sample	1994q2-2019q2	1994q2-2021q4	2004q3-2022q4	1994q2-2019q2	1994q2-2019q2	2004q3-2019q2	2004q3-2021q4
Effective F-stat	2.26	0.34	8.80	2.01	0.30	8.47	15.73
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	15.80	23.11	18.73	16.66	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation 6. The sample period varies depending on the availability of data across different monetary policy surprises. All reported right-hand side variables are normalized. Effective F-stat is calculated using the methodology in Olea and Pflueger (2013). The final row reports the critical values of a test of weak instruments with a 5% confidence level and a 10% worst-case bias. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** p<0.01, ** p<0.05, * p<0.1.

Table 10: Yield elasticity of demand across different sectors in the United States: Second stage results - Panel (a) of Table 3- with control variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(PF)	log(IF)	log(HH)	log(Banks)	log(SLG)	log(IC)	log(Other)
8Y Yield (ZC)	1.49	24.34***	23.28***	21.36***	2.98	31.35***	6.73*	16.39***	-29.75*
	(2.58)	(7.78)	(8.72)	(7.26)	(43.32)	(11.45)	(3.84)	(4.72)	(16.36)
log GDP	-2.69***	0.29	0.74	-1.70*	-25.22***	-3.43*	1.52***	-1.71***	-7.26***
	(0.52)	(1.18)	(1.16)	(0.95)	(4.17)	(1.87)	(0.45)	(0.61)	(1.90)
GDP growth	1.43	-3.97	-5.42*	1.93	7.52	7.66*	-1.12	0.45	0.01
	(1.64)	(3.00)	(3.02)	(3.03)	(8.40)	(4.07)	(1.26)	(2.12)	(4.34)
inflation	0.25	3.62***	0.72	3.27*	4.25	5.02**	-0.46	2.69*	4.32
	(1.17)	(1.39)	(1.09)	(1.83)	(12.06)	(2.38)	(1.11)	(1.55)	(9.60)
log broad dollar index	-0.83***	-0.63	-1.16***	-0.68	-1.83	-0.33	0.28	-0.62	-0.48
	(0.17)	(0.60)	(0.36)	(0.45)	(1.16)	(0.75)	(0.18)	(0.45)	(1.18)
log VIX	-0.04	-0.01	0.01	0.03	0.31**	0.05	-0.03	-0.05	0.18**
	(0.05)	(0.06)	(0.02)	(0.05)	(0.14)	(0.10)	(0.02)	(0.05)	(0.08)
5Y German yield (zc)	-5.80**	-28.19***	-21.58**	-23.23***	-19.37	-31.56***	3.35	-19.07***	31.07**
	(2.87)	(7.50)	(9.76)	(8.61)	(41.85)	(11.60)	(3.27)	(5.52)	(14.62)
trend	0.21***	0.15***	0.01	0.12***	-0.14	0.13	0.13***	0.03	0.09
	(0.02)	(0.04)	(0.03)	(0.02)	(0.19)	(0.08)	(0.02)	(0.03)	(0.12)
trend squared	-0.00***	-0.00***	-0.00	-0.00***	0.00*	-0.00	-0.00***	-0.00	0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	60	60	60	60	57	60	60	60	60
Anderson-Rubin Wald test (p-val)	0.56	0.00	0.00	0.00	0.95	0.00	0.04	0.00	0.06
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 11: Yield elasticity of demand across different sectors in the United States: Second stage results - Panel (b) of Table 3- with control variables

VARIABLES	(1) log(PF)-log(OA_{PF})	(2) log(IF)-log(OA_{IF})	(3) log(HH)-log(OA_{HH})	(4) log(Banks)-log(OA_{Banks})	(5) log(SLG)-log(OA_{SLG})	(6) log(IC)-log(OA_{IC})
8Y Yield (ZC)	20.52**	18.06**	-8.34	27.34*	5.20	15.44**
	(9.49)	(7.41)	(49.83)	(14.25)	(4.15)	(6.34)
log GDP	-4.39**	-7.11***	-28.24***	-10.57***	1.69*	-6.54***
	(2.11)	(1.52)	(6.46)	(3.43)	(0.98)	(1.18)
GDP growth	-3.59***	3.62**	1.16	10.35**	-3.05***	3.34***
	(1.36)	(1.56)	(6.36)	(4.37)	(0.86)	(1.27)
Inflation	0.00	0.03	0.02	0.04	-0.01	0.03
	(0.01)	(0.02)	(0.12)	(0.03)	(0.01)	(0.02)
log broad dollar index	-0.41	0.10	-1.58	0.74	0.15	0.12
	(0.30)	(0.25)	(1.34)	(0.78)	(0.17)	(0.30)
trend	0.06	0.17***	-0.14	0.21*	0.07*	0.08*
	(0.07)	(0.06)	(0.40)	(0.11)	(0.04)	(0.04)
trend squared	-0.00	-0.00	0.00	-0.00	-0.00**	-0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	60	60	57	60	60	60
Anderson-Rubin Wald test (p-val)	0.002	0.000	0.867	0.003	0.247	0.000
Underidentification LM stat (p-val)	0.094	0.102	0.106	0.089	0.102	0.102

Note: This table reports the coefficients of the second-stage regression specified in Equation (7) using PCA 1 (All) as an instrument for yields. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** p<0.01, ** p<0.05, * p<0.1.

B.2 Checks for the first stage regressions

In this section, we report alternative specifications for the first stage regressions.

In Table 12, we restrict the samples to be the same across different constructs of monetary policy surprises. The results are similar in terms of weak instruments tests. This suggest that our baseline results do not represent differences in the sample periods, but instead using information contained in multiple monetary policy surprise constructs does indeed improve the relevance of the instrument.

In Table 13, we report the results with first differences in order to check whether the monetary policy surprises perform better when we instrument the changes in yield instead of levels of yield with them. In particular, we run the following first-stage regression:

$$\Delta Y_t^8 = \mu_0 + \mu_1^{'} \ \mathbf{MPS_t^i} + \mu_2^{'} \ \mathbf{X_t'} + \mu_3^{'} \ \mathbf{W_t'} + \eta_t^s.$$

where ΔY_t^8 is the quarterly change in the yield, $\mathbf{X_t'}$ includes GDP growth, change in inflation, the percent change in the broad dollar index, and $\mathbf{W_t'}$ is the change in the 5-year zero-coupon German government bond yield and the percent change in the VIX.

Using this specification, we fail to reject the null hypothesis of weak instruments for all monetary policy surprise constructs vectors MPS_t^i. This suggests that using changes would not lead to better instruments, reinforcing our specification in the baseline regressions in the main text.

Table 12: First stage: same sample period

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)	8Y Yield (ZC)
Swanson FFR	0.0012***						
	(0.0003)						
Swanson FG	0.0006**						
	(0.0003)						
Swanson LSAP	-0.0002						
	(0.0002)						
BRW		0.0004					
		(0.0003)					
KSX (3M)			0.0004				
			(0.0004)				
KSX (2Y)			0.0011**				
			(0.0004)				
KSX (10Y)			0.0001				
			(0.0004)				
JK MP				0.0005			
				(0.0004)			
JK CBI				0.0011***			
				(0.0004)			
PCA 1 (BRW, Swanson, JK)					0.0005***		
					(0.0001)		
PCA 1 (All)						0.0005***	
						(0.0001)	
PCA 1 (BRW, KSX)							0.0007***
							(0.0001)
Observations	60	60	60	60	60	60	60
R-squared	0.9207	0.9156	0.9237	0.9210	0.9170	0.9207	0.9212
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Sample	2004q3-2019q2	2004q3-2019q2	2004q3-2019q2	2004q3-2019q2	2004q3-2019q2	2004q3-2019q2	2004q3-2019q2
Effective F-stat	4.81	2.61	5.53	4.34	14.09	43.62	40.26
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	15.05	23.11	15.99	16.62	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation 6. The sample period is restricted to be the same, i.e. 2004q3-2019q2, across specifications. In columns (1)-(4), each of the monetary policy surprise series is standardized. Effective F-stat is calculated using the methodology in Olea and Pflueger (2013). The final row reports the critical values of a test of weak instruments with a 5% confidence level and a 10% worst-case bias. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** p<0.01, ** p<0.05, * p<0.1.

Table 13: First stage: first differences

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	Ch. 8Y Yield (ZC)	Ch. 8Y Yield (ZC)	Ch. 8Y Yield (ZC)	Ch. 8Y Yield (ZC)	Ch. 8Y Yield (ZC)	Ch. 8Y Yield (ZC)	Ch. 8Y Yield (ZC)
Swanson FFR	-0.0004						
	(0.0003)						
Swanson FG	0.0006**						
	(0.0003)						
Swanson LSAP	-0.0000						
	(0.0003)						
BRW		-0.0000					
		(0.0002)					
KSX (3M)			0.0004				
			(0.0004)				
KSX (2Y)			0.0006				
			(0.0008)				
KSX (10Y)			-0.0001				
			(0.0005)				
JK MP				-0.0002			
				(0.0003)			
JK CBI				0.0005			
				(0.0003)			
PCA 1 (BRW, Swanson, JK)					0.0000		
					(0.0002)		
PCA 1 (All)						0.0006***	
						(0.0001)	
PCA 1 (BRW, KSX)							0.0005***
							(0.0002)
Observations	101	111	74	101	101	60	70
R-squared	0.6473	0.6142	0.6597	0.6367	0.6265	0.6758	0.6380
Controls	**X' and W'**	**X' and W'**	**X' and W'**	**X' and W'**	**X' and W'**	**X' and W'**	**X' and W'**
Sample	1994q2-2019q2	1994q2-2021q4	2004q3-2022q4	1994q2-2019q2	1994q2-2019q2	2004q3-2019q2	2004q3-2021q4
Effective F-stat	1.87	0.01	1.47	1.23	0.00	8.91	5.66
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	15.54	23.11	20.40	17.56	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression in Equation B.2. The sample period varies depending on the availability of data across different monetary policy surprises. In columns (1)-(4), each of the monetary policy surprise series is standardized. The left-hand side variable is the quarterly change in the yield, $\mathbf{X_t'}$ includes GDP growth, change in inflation, the percent change in the broad dollar index, and $\mathbf{W_t'}$ is the change in the 5-year zero-coupon German government bond yield and the percent change in the VIX. Effective F-stat is calculated using the methodology in Olea and Pflueger (2013). The final row reports the critical values of a test of weak instruments with a 5% confidence level and a 10% worst-case bias. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** p<0.01, ** p<0.05, * p<0.1.

B.3 Checks using BRWKSX as the instrument

In this section, we repeat the second stage and the scenario analysis using the PCA 1 (BRW, KSX) as the instrument for yields. The first stage estimates in Table 2 suggest that while PCA 1 (All) yields the largest effective F-statistic, we are also able to reject the null hypothesis of weak instruments when we use PCA 1 (BRW, KSX).

An important difference between PCA 1 (All) and PCA 2 (BRW, KSX) is the data availability. While the former is available only between 2004q3 and 2019q2, the latter is available between 2004q3 and 2021q4.

We report the results using PCA 2 (BRW, KSX) during 2004q3 and 2021q4 in Table 14. The results are qualitatively similar in both specifications (6) and (7). Quantitatively, the estimates are somewhat smaller suggesting a lower yield elasticity of demand. However, the estimates are closer to those when we restrict the sample period to be the same as in Table 3. These are reported in Table 15. This might be reflecting a possible change in the demand curves following the Covid-19 crisis or might reflect a noisier estimate compared to the PCA 1 (All) as the effective F-statistic is lower in the first stage. More data that allows PCA 1 (All) to be used in a longer sample period can be useful in this regard.

We report the results of the scenario analysis conducted in Section 4 using PCA 1 (BRW, KSX) in a longer sample period. These estimates suggest that a size of quarterly QT of \$145 billion corresponds to 10 basis points quarterly change in long-term yields (instead of the \$266 estimates in the main text).

Table 14: The yield elasticity of demand across different sectors in the United States - using PCA 1 (BRW,KSX) as an instrument

Panel (a): Using a parametric specification for the outside asset (Equation (6))

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(PF)	log(IF)	log(HH)	log(Banks)	log(SLG)	log(IC)	log(Other)
8Y Yield (ZC)	-2.98	11.85*	13.40*	12.61***	24.80	16.33**	-6.38	7.60**	-22.55
	(4.27)	(7.07)	(8.08)	(4.00)	(33.84)	(6.63)	(6.37)	(3.29)	(15.30)
Observations	70	70	70	70	67	70	70	70	70
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.48	0.14	0.04	0.00	0.49	0.02	0.35	0.04	0.18
Underidentification LM stat (p-val)	0.08	0.08	0.08	0.08	0.07	0.08	0.08	0.08	0.08

Panel (b): Using portfolio information to control for the outside asset (Equation (7))

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	log(PF)-log(OA$_{PF}$)	log(IF)-log(OA$_{IF}$)	log(HH)-log(OA$_{HH}$)	log(Banks)-log(OA$_{Banks}$)	log(SLG)-log(OA$_{SLG}$)	log(IC)-log(OA$_{IC}$)
8Y Yield (ZC)	9.44	15.90***	17.21	20.21***	-13.12*	12.05***
	(7.53)	(4.02)	(30.83)	(4.88)	(6.80)	(2.83)
Observations	70	70	67	70	70	70
Controls	**X**	**X**	**X**	**X**	**X**	**X**
Anderson-Rubin Wald test (p-val)	0.111	0.001	0.589	0.007	0.068	0.001
Underidentification LM stat (p-val)	0.091	0.091	0.089	0.091	0.091	0.091

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) in Panel (a) and Equation (7) in Panel (b) using PCA 1 (All) as an instrument for yields. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** p<0.01, ** p<0.05, * p<0.1.

Table 15: The yield elasticity of demand across different sectors in the United States - using PCA 1 (BRW,KSX) as an instrument - restricting the sample period to 2004q3-2019q2

Panel (a): Using a parametric specification for the outside asset (Equation (6))

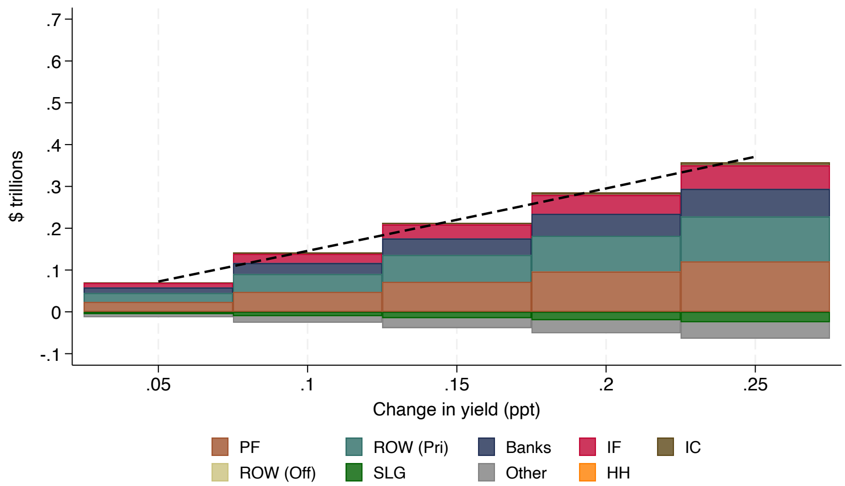
VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
8Y Yield (ZC)	-0.69	22.99***	20.58**	14.60***	-3.53	28.97***	4.07	14.19***	-33.46*
	(3.87)	(7.09)	(8.90)	(5.36)	(49.88)	(10.94)	(4.47)	(3.78)	(18.58)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.86	0.00	0.01	0.00	0.94	0.00	0.35	0.00	0.07
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07

Panel (b): Using portfolio information to control for the outside asset (Equation (7))

VARIABLES	(1) log(PF)-log(OA_{PF})	(2) log(IF)-log(OA_{IF})	(3) log(HH)-log(OA_{HH})	(4) log(Banks)-log(OA_{Banks})	(5) log(SLG)-log(OA_{SLG})	(6) log(IC)-log(OA_{IC})
8Y Yield (ZC)	19.60*	13.10**	-17.34	27.87**	0.87	15.25***
	(10.40)	(5.33)	(65.45)	(13.27)	(6.24)	(5.33)
Observations	60	60	57	60	60	60
Controls	**X**	**X**	**X**	**X**	**X**	**X**
Anderson-Rubin Wald test (p-val)	0.009	0.004	0.785	0.006	0.891	0.000
Underidentification LM stat (p-val)	0.097	0.109	0.112	0.084	0.109	0.109

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) in Panel (a) and Equation (7) in Panel (b) using PCA 1 (All) as an instrument for yields. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Figure 8: Implied absorption capacity by all sectors for different changes in yield in the United States - Instrument: PCA 1 (BRW,KSX) & Sample period: 2004q3-2021q4



Note: This figure plots the relationship between $\widehat{\Delta H}^{s}_{\Delta Y^{8}}$ and ΔY^{8} using Equation (8) based on the second stage estimates in Table **??** where we make valuation adjustments to holdings of each sector assuming a modified duration of 8. The stacked bars correspond to how much demand by each sector would change in response to changes in yield using the elasticity estimates in Table **??** and the latest holdings in Figure 2. The dashed line is the total change in demand obtained by adding up all stacked bars.

B.4 Checks excluding the Jarociński and Karadi (2020) information shocks from the principal components analysis

Jarociński and Karadi (2020) differentiate between monetary policy surprises into monetary policy shocks and "central bank information" shocks. The monetary policy shock moves interest rates and stock prices in opposite directions, while the central bank information shock lead to same directional change in interest rates and stock prices. There is an ongoing debate in the literature whether these represent private information that the Federal Reserve transmits to the market or they simply measure the response of the Fed to news (Bauer and Swanson, 2023). In the latter case, for which we believe the evidence is compelling, there is no threat to our exclusion restriction. However, if the former interpretation is right, it might threaten the exclusion restriction if the Fed's private information affects holdings also through the release of information about the economy. We report the results of the second stage when we exclude the estimated information shocks by Jarociński and Karadi (2020) from the construction of the principal components used as instruments in Table 16. Our results remain similar.

Table 16: Second stage: Excluding the Jarociński and Karadi (2020) information shocks from the principal components analysis

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(PF)	log(IF)	log(HH)	log(Banks)	log(SLG)	log(IC)	log(Other)
8Y Yield (ZC)	0.03	24.81***	23.85**	21.47***	-1.31	33.18***	5.81	15.32**	-30.77*
	(3.23)	(8.03)	(10.07)	(8.22)	(48.54)	(12.58)	(4.10)	(6.05)	(18.21)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.99	0.00	0.01	0.00	0.98	0.00	0.12	0.01	0.08
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All exc. JK CBI) as an instrument for yields, where we use all monetary policy surprises in the construction of the first principal component except the Jarociński and Karadi (2020) central bank information shocks. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** p<0.01, ** p<0.05, * p<0.1.

B.5 Robustness checks - second stage: two principal components as instruments

In Table 17, we report the results of the second stage when we instrument for yields with the first two principal componets of all monetary policy surprises instead of only the first one reported in the baseline regressions. The results are similar.

Table 17: Second stage: The first two principal components as instruments

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(PF)	log(IF)	log(HH)	log(Banks)	log(SLG)	log(IC)	log(Other)
8Y Yield (ZC)	5.40**	21.66***	23.39***	21.45***	27.93	27.00***	7.72**	19.84***	-28.92**
	(2.57)	(7.71)	(6.66)	(5.87)	(32.88)	(9.03)	(3.88)	(3.72)	(14.66)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.00	0.00	0.00	0.00	0.01	0.00	0.04	0.00	0.06
Underidentification LM stat (p-val)	0.17	0.17	0.17	0.17	0.17	0.17	0.17	0.17	0.17
Hansen J stat (p-val)	0.33	0.17	0.98	0.98	0.22	0.40	0.55	0.45	0.88

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) and PCA 2 (All) as instruments for yields. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Hansen J-statistic test has the null hypothesis that the model is overidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

B.6 Sensitivity to variables used in X_t and W_t

We repeat the baseline regression where we use a broader set of covariates in the characteristics vector ($\mathbf{X_t}$) and the outside asset proxies ($\mathbf{W_t}$). We use log GDP, GDP growth, inflation, log broad dollar index and the total face value of outstanding Treasuries in the asset characteristics vector $\mathbf{X_t}$, and we use log VIX, AAA bond spread, 5-year (zero-coupon) German government bond yield, S&P 500 dividend yield for the outside asset specification $\mathbf{W_t}$. The results, reported in Table 18 are similar to those in the baseline regression.

In Table 19, we keep the $\mathbf{X_t}$ and $\mathbf{W_t}$ as in the baseline regression, but use a linear trend instead of a quadratic trend. The results are again similar.

Table 18: Second stage: Broader set of controls

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(PF)	log(IF)	log(HH)	log(Banks)	log(SLG)	log(IC)	log(Other)
8Y Yield (ZC)	2.76	30.81***	27.18**	18.86***	6.21	19.06***	9.59**	18.95***	-23.27
	(3.03)	(7.82)	(11.49)	(6.02)	(37.34)	(7.23)	(4.37)	(5.83)	(18.09)
log GDP	0.08	1.01	-0.18	-1.68	-11.76**	-5.10***	-0.10	0.70	-5.74*
	(0.72)	(2.63)	(2.41)	(2.05)	(5.04)	(1.90)	(0.78)	(1.98)	(3.07)
GDP growth	2.59	0.04	-3.26	0.44	11.51**	0.49	0.15	2.30	4.09
	(1.67)	(3.26)	(3.40)	(2.85)	(5.66)	(3.14)	(1.31)	(2.49)	(4.39)
Inflation	-0.02***	0.03	0.01	0.03	-0.05	0.07***	0.01	0.01	0.03
	(0.01)	(0.02)	(0.02)	(0.02)	(0.08)	(0.02)	(0.01)	(0.02)	(0.09)
log broad dollar index	-0.43**	-0.84	-1.56***	-0.54	1.06	0.36	-0.30	-0.38	-0.60
	(0.21)	(0.60)	(0.59)	(0.36)	(1.51)	(0.56)	(0.27)	(0.43)	(1.01)
log UST (fv)	1.03***	0.29	-0.37	0.01	6.12***	-0.29	-0.77***	0.87	0.54
	(0.17)	(0.54)	(0.61)	(0.62)	(0.67)	(0.80)	(0.25)	(0.61)	(0.95)
log VIX	-0.05	-0.09*	-0.03	0.06	0.23*	0.19**	-0.06***	-0.08	0.10
	(0.03)	(0.05)	(0.03)	(0.05)	(0.13)	(0.08)	(0.02)	(0.07)	(0.09)
ICE BofA OAS AAA	0.25	4.92	2.21	-1.64	18.04*	-0.03	-1.62	0.57	3.56
	(0.80)	(4.72)	(4.21)	(1.67)	(9.44)	(2.11)	(1.28)	(2.01)	(6.88)
5Y German yield (zc)	0.31	-30.42***	-26.54*	-21.56**	17.32	-25.98**	-3.46	-14.84*	30.57*
	(3.61)	(7.96)	(15.10)	(9.05)	(36.04)	(12.70)	(4.94)	(9.00)	(16.92)
SP500 dividend yield	11.50***	16.49	7.15	-5.87	27.15	-44.19***	6.78***	14.21**	20.53*
	(2.03)	(11.17)	(8.61)	(5.07)	(26.25)	(5.65)	(1.93)	(6.82)	(10.77)
trend	0.11***	0.13	0.05	0.11*	-0.70***	0.16	0.20***	-0.05	0.05
	(0.02)	(0.09)	(0.09)	(0.07)	(0.19)	(0.11)	(0.03)	(0.08)	(0.19)
trend squared	-0.00***	-0.00*	-0.00	-0.00*	0.00***	-0.00	-0.00***	0.00	0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	60	60	60	60	57	60	60	60	60
Anderson-Rubin Wald test (p-val)	0.33	0.00	0.00	0.00	0.87	0.01	0.01	0.00	0.22
Underidentification LM stat (p-val)	0.08	0.08	0.08	0.08	0.09	0.08	0.08	0.08	0.08

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields, where we use log GDP, GDP growth, inflation, log broad dollar index and the total face value of outstanding Treasuries in the asset characteristics vector $\mathbf{X_t}$, and we use log VIX, AAA bond spread, 5-year (zero-coupon) German government bond yield, S&P 500 dividend yield for the outside asset specification $\mathbf{W_t}$. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** p<0.01, ** p<0.05, * p<0.1.

Table 19: Second stage: Linear trend

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
8Y Yield (ZC)	-8.74*	15.90*	22.89***	16.89***	25.53	26.63***	-1.09	15.61***	-28.86
	(5.09)	(9.30)	(7.49)	(5.65)	(34.73)	(7.12)	(5.39)	(4.76)	(17.91)
log GDP	-3.98***	-0.77	0.69	-2.26***	-22.18***	-4.02***	0.54	-1.80***	-7.15***
	(0.83)	(0.72)	(1.09)	(0.66)	(2.60)	(1.35)	(0.89)	(0.64)	(1.75)
GDP growth	0.16	-5.01*	-5.47*	1.37	10.90	7.07*	-2.08	0.35	0.12
	(1.59)	(2.56)	(3.10)	(2.68)	(9.24)	(4.16)	(1.63)	(1.96)	(4.78)
inflation	-0.94	2.63	0.67	2.75*	6.31	4.47**	-1.37	2.60	4.42
	(2.31)	(1.68)	(1.04)	(1.49)	(9.62)	(2.01)	(2.08)	(1.62)	(8.94)
log broad dollar index	-1.48***	-1.16**	-1.19***	-0.97**	-0.29	-0.62	-0.21	-0.67*	-0.43
	(0.30)	(0.50)	(0.39)	(0.38)	(1.50)	(0.47)	(0.24)	(0.39)	(1.02)
log VIX	-0.07	-0.04	0.01	0.02	0.38**	0.04	-0.05*	-0.05	0.18**
	(0.05)	(0.05)	(0.02)	(0.05)	(0.15)	(0.09)	(0.03)	(0.05)	(0.09)
5Y German yield (zc)	2.08	-21.68***	-21.28**	-19.79***	-35.57	-27.93***	9.37**	-18.47***	30.37**
	(5.22)	(7.43)	(8.94)	(7.41)	(36.06)	(7.89)	(4.20)	(5.76)	(14.78)
trend	0.06***	0.02***	0.00	0.05***	0.21***	0.06***	0.01	0.02***	0.11***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.03)	(0.01)	(0.01)	(0.01)	(0.02)
Observations	60	60	60	60	57	60	60	60	60
Anderson-Rubin Wald test (p-val)	0.06	0.07	0.01	0.00	0.48	0.00	0.84	0.00	0.08
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.04	0.07	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields. The difference between this table and the Panel (a) of Table 3 is that we use a linear trend instead of a quadratic trend. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

B.7 Robustness checks - second stage: Alternative valuation adjustments

In this section, we report the results of the baseline regressions when we assume a modified duration of 5 (Table 20) and 10 (Table 21) in the holdings of sectors when we make valuation adjustments to the left-hand side variables of the second stage. The results remain similar.

Table 20: Second stage: Valuation adjustments with 5 year

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
8Y Yield (ZC)	-2.32	20.65***	19.32**	17.45**	0.05	27.22**	2.79	12.47**	-33.27**
	(2.64)	(7.57)	(8.90)	(7.39)	(43.45)	(11.69)	(3.30)	(5.01)	(16.04)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.38	0.01	0.02	0.00	1.00	0.01	0.36	0.01	0.03
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields. The difference between this table and the Panel (a) of Table 3 is that we assume a modified duration of 5 in the portfolios while making the valuation adjustments. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 21: Second stage: Valuation adjustments with 10 year

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
8Y Yield (ZC)	3.75	26.53***	25.65***	23.68***	4.66	33.83***	9.07**	18.72***	-27.65*
	(2.87)	(8.05)	(8.60)	(7.20)	(43.32)	(11.26)	(4.39)	(4.65)	(16.66)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.19	0.00	0.00	0.00	0.91	0.00	0.02	0.00	0.08
Underidentification LM stat (p-val)	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07	0.07

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields. The difference between this table and the Panel (a) of Table 3 is that we assume a modified duration of 10 in the portfolios while making the valuation adjustments. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

B.8 Robustness checks - second stage: Alternative yields

In this section, we report the results of the baseline regressions when we use 5-year zero-coupon yields (Table 22) and 10-year zero-coupon yields (Table 23) to estimate elasticities. The results remain similar.

Table 22: Second stage: using 5 year yield

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
5Y Yield (ZC)	1.20	19.68**	18.82***	17.26***	2.48	25.34***	5.44*	13.25***	-24.05*
	(2.07)	(7.76)	(5.53)	(4.83)	(36.11)	(7.69)	(3.20)	(2.74)	(12.69)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.56	0.01	0.00	0.00	0.95	0.00	0.04	0.00	0.06
Underidentification LM stat (p-val)	0.05	0.02	0.05	0.05	0.05	0.05	0.05	0.05	0.05

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields. The difference between this table and the Panel (a) of Table 3 is that we use the 5-year zero-coupon yield on the right-hand side instead of the 8-year zero-coupon yield in Table 3. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 23: Second stage: using 10 year yield

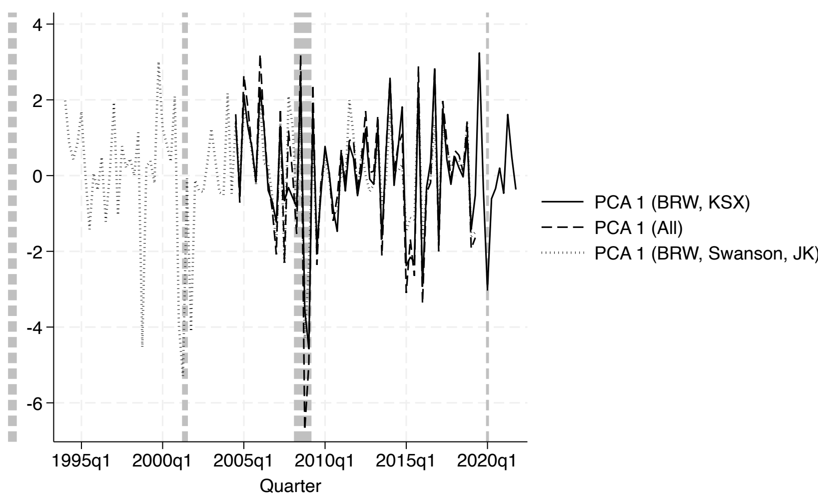
VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(PF)	(4) log(IF)	(5) log(HH)	(6) log(Banks)	(7) log(SLG)	(8) log(IC)	(9) log(Other)
10Y Yield (ZC)	1.68	27.50***	26.30**	24.12***	3.33	35.42**	7.60*	18.52***	-33.61*
	(2.93)	(8.69)	(11.26)	(9.32)	(48.15)	(14.56)	(4.36)	(6.38)	(19.30)
Observations	60	60	60	60	57	60	60	60	60
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Anderson-Rubin Wald test (p-val)	0.56	0.00	0.00	0.00	0.95	0.00	0.04	0.00	0.06
Underidentification LM stat (p-val)	0.08	0.08	0.08	0.08	0.08	0.08	0.08	0.08	0.08

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields. The difference between this table and the Panel (a) of Table 3 is that we use the 10-year zero-coupon yield on the right-hand side instead of the 8-year zero-coupon yield in Table 3. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. Missing observations for HH are due to negative values of holdings (due to short positions) between 2007Q2 and 2007Q4. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

C Principal components analysis

In this section, we report further details of the principal components analysis. In Figure 9, we report the time series of the first principal components of different constructs used in the analysis, i.e. PCA 1 (All), PCA 1 (BRW, KSX), PCA 1 (BRW, Swanson, JK). As we show in Table 24, these different measures are indeed highly correlated with each other. In Table 25, we report the standard deviations of different measures. In Table 26, we report the eigenvalues and the explained percentages of all of the series. The first principal component explains 42% of the variation and the second principal component explains another 26% of the variation. Finally, in Table 27, we report the eigenvectors.

Figure 9: US monetary policy surprises - Alternative first principal components for different sample period availability



Note: The figure shows the first principal components of alternative quarterly monetary policy surprises (all standardized). The solid line is the first principal component of the monetary policy surprise series taken from Bu, Rogers and Wu (2021) and Kearns, Schrimpf and Xia (2022) with the sample period between 2004Q3 and 2021Q4. The dashed line is the first principal component of all series used with the sample period of 2004Q3 and 2019Q2. The dotted line is the first principal component of the Bu, Rogers and Wu (2021), Swanson (2021) and Jarociński and Karadi (2020) series with the sample period between 1994Q2 and 2019Q2.

Table 24: Cross-correlation table

Variables	PCA 1 (All)	PCA 1 (BRW, KSX)	PCA 1 (BRW, Swanson, JK)
PCA 1 (All)	1.00		
PCA 1 (BRW, KSX)	0.95	1.00	
PCA 1 (BRW, Swanson, JK)	0.93	0.85	1.00

Table 25: Summary statistics of the different principal components

Variable	Obs	Mean	Std. Dev.
PCA 1 (All)	60	0	1.95
PCA 1 (BRW, KSX)	70	0	1.52
PCA 1 (BRW, Swanson, JK)	102	0	1.49
PCA 2 (All)	60	0	1.54
PCA 2 (BRW, KSX)	70	0	1.01
PCA 2 (BRW, Swanson, JK)	102	0	1.11

Table 26: Eigenvalues and explained percentages of PCA (All)

	Eigenvalue	Difference	Proportion	Cumulative
Comp1	3.795732	1.430552	0.4217	0.4217
Comp2	2.365179	1.357703	0.2628	0.6845
Comp3	1.007476	.417604	0.1119	0.7965
Comp4	.5898723	.0458679	0.0655	0.8620
Comp5	.5440044	.122319	0.0604	0.9225
Comp6	.4216854	.2455765	0.0469	0.9693
Comp7	.1761089	.0865819	0.0196	0.9889
Comp8	.0895269	.0791123	0.0099	0.9988
Comp9	.0104146	.	0.0012	1.0000

Table 27: Eigenvectors of PCA (All)

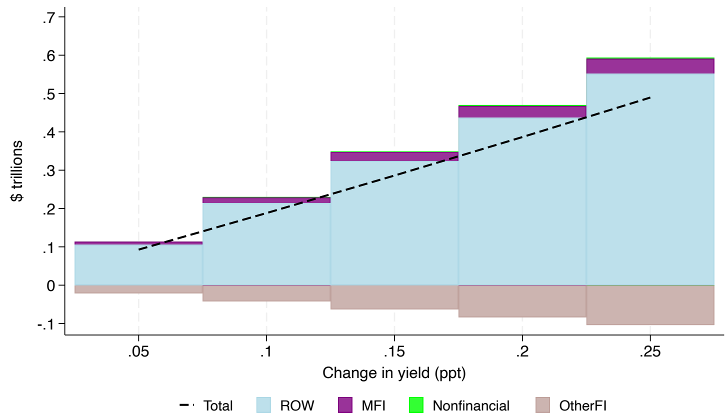
	Comp1	Comp2	Comp3	Unexplained
KSX (3M)	.1917595	.4921141	-.070521	.2826238
KSX (2Y)	.4533107	.0404143	.0826482	.2092679
KSX (10Y)	.4021113	-.2943427	-.3444692	.0617951
BRW	.3649378	-.1532808	.1688911	.4101785
Swanson FFR	.2886219	.4288576	-.2388797	.1913145
Swanson LSAP	-.1121347	.441853	.6477324	.0678142
Swanson FG	.4171747	-.1893306	.1877503	.2191147
JK MP	.4074786	-.0302851	.4355097	.1765052
JK CBI	.1640555	.4797308	-.3734611	.2129985

D QT scenario analysis for the Euro area

In this section, we repeat the scenario analysis of different QT configurations for the Euro area. We use latest holdings information using our data, but use the elasticity estimates from the Table 12 of Koijen, Koulischer, Nguyen and Yogo (2021). Their estimates have a more detailed breakdown of sectors compared to us (e.g. while we see Other FI, they can distinguish between mutual funds, and insurance companies and pension funds). Whenever that is the case, we use information about the shares reported in their Table 4 to weigh the elasticity estimates. For example, we take the total holdings of Other FI in our dataset and use the shares of mutual funds and ICPFs to get at a coarser elasticity estimate. We then proceed similarly as we do for the United States.

The results are reported in Figure 10. Given the large share of foreign investors and a remarkably large elasticity estimate, these estimates suggest that QT would mostly be absorbed by foreign investors in the Euro area. The aggregate estimate adding up all bars is shown with the dashed line. According to these estimates, a QT size of 1€88 billion would lead to a market-clearing rise of long-term yields by 10 basis points.

Figure 10: Implied absorption capacity by all sectors for different changes in yield in the Euro area



Note: This figure plots the scenario analysis of different QT configurations for the Euro area based on the information on the latest holdings of different sectors in Figure 5(b). We use the elasticity estimates in the Table 12 of Koijen, Koulischer, Nguyen and Yogo (2021). Their estimates have a more detailed breakdown of sectors compared to us (e.g. while we see Other FI, they can distinguish between mutual funds, and insurance companies and pension funds). Whenever that is the case, we use information about the shares reported in their Table 4 to weigh the elasticity estimates. The stacked bars correspond to how much demand by each sector would change in response to changes in yield using the elasticity estimates. The dashed line is the total change in demand obtained by adding up all stacked bars.

E Results for investment funds

In Figure 11, we report the time series of the total market value of the holdings of different types of investment funds. In Table 28, we report the elasticity estimates. These results suggest that our results in the baseline regressions for investment funds are mostly driven by mutual funds as the most yield-elastic type.

Figure 11: Total market value of the holdings of different types of investment funds

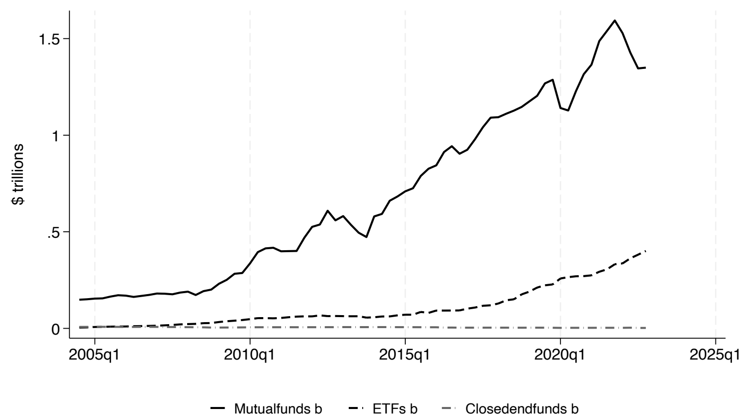


Table 28: Yield elasticity estimates for different types of investment funds

	(1)	(2)	(3)
VARIABLES	ln(Mutual funds)	ln(ETFs)	ln(Closed-ended funds)
8Y Yield (ZC)	25.4024***	-13.7614	15.3606
	(8.5012)	(17.7212)	(11.4673)
Observations	60	60	60
Controls	**X and W**	**X and W**	**X and W**
Sample	2004q3-2019q2	2004q3-2019q2	2004q3-2019q2
Anderson-Rubin Wald test (p-val)	0.00	0.44	0.14
Underidentification LM stat (p-val)	0.07	0.07	0.07

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (All) as an instrument for yields for mutual funds, ETFs and closed-ended funds. The sample period is between 2004q3 and 2019q2. Anderson-Rubin test has the null hypothesis that the estimated coefficient is equal to zero. The underidentification test has the null hypothesis that the model is underidentified. Standard errors are robust to arbitrary heteroskedasticity and autocorrelation with the lag structure automatically selected using the Newey and West (1994) procedure. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.