BIS Working Papers

No 1105

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by Egemen Eren, Andreas Schrimpf and Fan Dora Xia

Monetary and Economic Department

June 2023 (revised December 2025)

JEL classification: E58, G11, G21, G23, H63

Keywords: government debt, demand, price elasticity, quantitative easing, quantitative tightening

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This publication is available on the BIS website (www.bis.org).

ISSN 1020-0959 (print)
ISSN 1682-7678 (online)

The Demand for Government Debt*

Egemen Eren[†] Andreas Schrimpf[‡] Fan Dora Xia[§]

October 3, 2025

Abstract

We document that the sectoral composition and marginal buyers of government debt differ notably across jurisdictions and over time. We use instrumental variables derived from monetary policy surprises to estimate the demand elasticities of various sectors. In the United States, commercial banks and mutual funds exhibit the most price-elastic demand, while the foreign official sector has a price-inelastic demand. Based on these estimates and under certain assumptions, we find that a 1% increase in the central bank holdings of US Treasuries results in around 8 to 13 basis point drop in long-term yields depending on the market composition. Elasticities of individual sectors do not differ in a statistically significant manner when the central bank share in the Treasury market increases or decreases. However, different market compositions during various quantitative easing (QE) and quantitative tightening (QT) programs have led to an asymmetric effect with the impact of QE on yields being greater than that of QT. Our results suggest that, overall, the demand for US Treasuries is considerably more elastic than for equities, corporate bonds and emerging market sovereign bonds found in the literature. We also repeat the analysis for other jurisdictions and compare estimates for different sectors.

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*Forthcoming at *Management Science*. We are grateful to Lukas Schmid (editor), two anonymous referees, Ryan Banerjee, Claudio Borio, Stijn Claessens, Sebastian Doerr, Mathias Drehmann, Chuck Fang, Zhiyu Fu, Bryan Hardy, Kristy Jansen, Chanik Jo, Andras Lengyel, Benoit Nguyen, Hyun Song Shin, Emil Siriwardane, Frank Smets, Karamfil Todorov, Dimitri Vayanos, Egon Zakrajsek, Haonan Zhou and seminar participants at the Chicago Booth Treasury Markets Conference 2025, BoC-Chicago Fed-SF Fed Conference 2025, EFA 2024, SFS Cavalcade Asia-Pacific 2024, York Asset Pricing Workshop 2024, BOJ/IMES-BOK/ERI Research Workshop 2025, Korea University, Bank of England, University of Bristol, BIS and ESM for helpful comments and discussions. Alexis Maurin provided excellent research assistance. The views in this article are those of the authors and should not be attributed to the BIS. All errors are our own.

[†]Bank for International Settlements (BIS); E-mail: `egemen.eren@bis.org`
[‡]Bank for International Settlements (BIS) and CEPR; E-mail: `andreas.schrimpf@bis.org`
[§]Bank for International Settlements (BIS); E-mail: `dora.xia@bis.org`

1 Introduction

Government debt markets – especially those for US Treasuries – are central to the global financial system. In response to the Great Financial Crisis (GFC) and the Covid-19 pandemic, many major central banks implemented quantitative easing (QE) policies, primarily through government debt purchases, to lower long-term interest rates after short-term interest rates hit their lower bound. As balance sheet normalization — or quantitative tightening (QT) – proceeds, the effects of these policies on government debt markets remain uncertain, largely due to a limited understanding of investor demand for government debt.

In this paper, we present a sector-level analysis of government debt holdings. The demand elasticity of these sectors, shaped by their respective objectives and constraints, plays a key role in how equilibrium prices respond to central bank purchases of government debt, and thus how QE and QT transmit to financial markets and the broader economy. To our knowledge, ours is among the first studies to apply a demand-based asset pricing framework to major government debt markets, including the US Treasury market.

We first document that the sectoral composition and marginal buyers of government debt vary significantly across jurisdictions – the United States, Japan, the United Kingdom, and the Euro area – and over time. Central bank footprint in these markets has increased markedly. Central banks at times absorbed around half of a unit of government bond issuance in the United States, Euro area and the United Kingdom, and around twice the issuance in Japan. In the United States, the share held by foreign official investors declined sharply after the Covid-19 crisis. In the Euro area, the central bank by and large replaced all investors, in particular foreigners. In Japan, increased central bank purchases were offset by reduced bank holdings, suggesting banks were primary sellers, while foreign investors expanded their presence. In the United Kingdom, the central bank balance sheet expanded rapidly following the GFC as major domestic investors lost market share to the central bank as foreign investors kept and slightly increased their share.

Just as the composition of holders of government debt changed during episodes of QE, they also changed during periods of QT. In the United States, the main sectors that were absorbing government debt during QT1 and QT2 were the household sector (which is mainly a proxy for the hedge fund sector), money market funds (MMFs) and foreign private investors. While commercial banks were active as buyers during QT1, their footprint declined during QT2. In the United Kingdom, foreign investors stepped in to replace the central bank during the QT program.

These aggregate trends reflect equilibrium outcomes shaped by the broad macroeconomic environment. In the second part of our analysis, we estimate the elasticity of demand of different investor groups in response to changes in long-term yields in government debt in order to better gauge the impact of QE and QT on their demand. To do so, we take a demand system perspective. We construct instruments for government bond yields based on monetary policy surprises for use in two-stage least square regressions. Due to data availability and quality, we primarily focus on the United States, but we report results for other jurisdictions as well, compare estimates, and discuss the limitations.

For the United States, we construct an instrument for long-term yields by extracting the first principal component from multiple measures of monetary policy surprises, following Nakamura and Steinsson (2013), Jarociński and Karadi (2020), Swanson (2021), Bu et al. (2021), and Kearns et al. (2022). Rather than relying on any single measure, this approach aggregates information across all series, yielding a more powerful instrument. These surprises are derived from asset price movements in short intra-day windows around monetary policy announcements to overcome endogeneity issues. Since the main investor sectors we analyze – commercial banks, asset managers, pension funds, and insurance companies – adjust their portfolios at lower frequencies, they are unlikely to drive these high-frequency market responses, which is important for the exogeneity of the instrument. Moreover, our instrument is only weakly correlated with central bank information shocks identified in the literature and thus is unlikely to reflect macroeconomic signals that could directly affect holdings.[1] Finally, we show that the instrument loads strongly on monetary policy surprises that are related to the longer-end of the yield curve, indicating a close link to central bank balance sheet policies that influence the supply of government bonds.

Our second-stage regressions indicate that most sectors in the United States have downward-sloping demand curves with respect to prices. Since prices and yields move in opposite directions,

[1] In robustness checks, we completely orthogonalize our instrument to series of central bank information shocks.

this means that these investors' demand more when the yield goes up, and *vice versa*. Commercial banks and investment funds (in particular, open-ended mutual funds) are the sectors with the most elastic demand. We also find that foreign private investors, pension funds and insurance companies have downward-sloping demand, albeit to a lesser degree. Importantly, we also find that the demand functions of the foreign official sector (ie. mostly reserve managers at central banks) and exchange-traded funds (ETFs) are price-inelastic suggesting that it is mainly non-price factors that drive their demand.

Based on these estimates and under certain assumptions, we find that a 1% change in the residual supply of Treasuries due to central bank balance sheet policies translates into an impact of around 8-13 basis points on long-term yields in the United States.[2] These estimates suggest that the demand in the market for US Treasuries is more elastic than estimates of the aggregate equity market (Gabaix and Koijen, 2023), the corporate bond market (Chaudhary et al., 2023) and the markets for EME sovereign bonds (Fang et al., 2022).

We repeat our analysis for Japan and the United Kingdom, and rely on estimates in Koijen et al. (2021) for the Euro area to compare the elasticities of different sectors across jurisdictions. This exercise provides useful insights even though it is important to caveat that our instrument for these jurisdictions do not have several desirable features of those for the United States and thus not as reliable. We overall find that most sectors exhibit a more elastic demand for government bonds in Japan with the exception of insurance companies and pension funds. In the United Kingdom, most notably, we find that insurance companies and pension funds, as well as the category "other financial institutions", which includes certain investment funds, have upward sloping demand curves, which could amplify price movements both during QE and QT.

Using our framework, we also study the potential asymmetry between QE and QT in the United States, which has attracted considerable interest in academic and policy circles with mixed evidence (see, e.g. D'Amico and Seida, 2024; Du et al., 2024; Jiang and Sun, 2024). In order to improve on the

[2] Since our instrument is constructed from monetary policy surprises measured in high-frequency windows around policy announcements, the exclusion restriction may not hold for market makers in bond markets (ie. broker-dealers) or certain types of hedge funds. We hence omit these latter groups from our analysis since our identification assumption would not hold for them. To construct estimates for the aggregate response, we assume a fully inelastic demand function for these sectors, which yields an upper bound estimate for the aggregate impact on yields assuming their demand is not upward-sloping.

limitations posed by the small number of observations for which QT has taken place, we compare the elasticity of demand of different sectors during periods in which the central bank share increases or decreases in the US Treasury market. We cannot reject the hypothesis that the elasticities of individual investor groups in US Treasury securities are the same when the central bank share increases or decreases. Since the market composition and the overall market elasticity differed between periods of QE and QT, however, our results indeed suggest an asymmetric response. The Treasury market composition was tilted more towards inelastic sectors as QE was taking place compared to QT. Hence, the impact of QE on yields was more pronounced than that of QT.

We perform several counterfactual analyses. First, we assign the actual market shares observed during QT periods to QE and *vice versa*, to study the impact of these policies with different market compositions. Moreover, we do counterfactual analyses of what the impact of QE and QT programs would have been in the absence of individual sectors. We find that in the absence of the foreign official sector, the sector with the most price inelastic demand, the impact of both QE and QT programs on yields would have been significantly smaller, more so for QE than QT. We also use our findings to provide estimates for several forward-looking counterfactual scenarios of potential divestments of US Treasuries by China and major oil producers.

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 et al. (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 et al. (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 price elasticity of their demand for government.

To our knowledge, we provide the first analysis of the US Treasury market based on a demand-based asset pricing framework. Results from alternative approaches are important to further our understanding of demand in government bond markets where evidence is still relatively scant compared to other markets, such as equities.[3] In complementary work to ours, Jansen et al. (2024) study own- and cross-price elasticities of different sectors across various maturities using an alternative identification strategy. Similar to us, they find that banks and investment funds are some of the most price elastic investors, and the economic magnitudes of the estimated impact of balance sheet policies are close to ours. Chaudhary et al. (2024) find that a 1 percent demand increase for U.S. Treasury notes and bonds results in a 1 percent increase in prices, equivalent to approximately a 10 basis points decline for the ten-year yield.

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 et al. (2021), in particular, show that the main counterparties to the ECB's asset purchase programs 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 QQE policy. They find foreign entities to have responded the most via asset sales to the Bank of Japan, followed by domestic banks. Carpenter et al. (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 rebalancing 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 sheet policies on government bond yields.[4] A large body of literature finds that purchases of

[3] 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; Siani, 2022; Darmouni et al., 2022; Chaudhary et al., 2023).

[4] Another related literature examines the impact of flows by the foreign official sector. See, for example, Bernanke et al. (2004); Warnock and Warnock (2009); Beltran et al. (2013); Ahmed and Rebucci (2022).

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), though our findings suggest a greater impact of QE than many other studies in part due to the measurement of the effect throughout the program rather than those of announcement effects. 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 muted effects on term premia and bond yields – 10 basis points higher in 10-year term premia. D'Amico and Seida (2024) study high-frequency announcement effects and find higher yield sensitivity to QT surprises than QE surprises. On the other hand, Du et al. (2024) focus on the announcements of QT policies and find muted market reaction. Our approach in turn takes a time series approach based on the relationship between prices and quantities rather than an event study approach. Based on this setting, we do not find evidence of an asymmetry in the price elasticity of individual sectors during periods of increasing or declining central bank share in the US Treasury market over longer horizons. However, we find that the market composition has shifted to become more elastic during periods of QT compared to those of QE, suggesting an asymmetric and more muted response of yields in response to QT.

Our work further contributes to the literature that emphasizes 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 et al., 2015; Nagel, 2016; d'Avernas and Vandeweyer, 2023; Doerr et al., 2023; Krishnamurthy and Li, 2023; Acharya and Laarits, 2023). We contribute to this literature by providing new stylized facts on compositional shifts among various investor groups in absorbing US government debt and by estimating their respective price 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. In this section, we document how government debt holdings by different sectors have evolved over time across the different jurisdictions. After briefly describing our holdings data, we lay out 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 holding government debt has changed since the early 2000s and in response to QE and QT programs.

2.1 Holdings data

Our data on investor holdings come from public sources. The data for US Treasuries are from the Flow of Funds accounts compiled by the Board of Governors. The data for Japan are also from the Flow of Funds accounts accessed through the Bank of Japan. The data for the Euro Area are from the ECB data portal. Those for the United Kingdom are from the Office for National Statistics. These data allow us to observe the time series of the holdings of various sectors.

While these data offer a comprehensive coverage of government debt holdings by sector, it is important to note that the 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 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.

2.2 Accounting framework to measure investors' marginal debt absorption

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 et al. (2022). Specifically, for each jurisdiction j, the idea is to 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.[5]

2.3 Shifts in marginal buyers of Treasury bonds in the 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). To illustrate these shifts, we divide the sample into five disjoint periods compare the holdings of different sectors, which correspond to: Pre-GFC (2001Q1-2008Q4), Post-GFC (2009Q1-2017Q3), QT1 (2017Q4-2019Q4),[6] Covid (2020Q1-2022Q2) and QT2 (2022Q3-2024Q2).

Several important observations stand out. First, during the post-GFC period, despite the large-scale asset purchases during various rounds of QE, the average share of the Federal Reserve remained roughly constant. The main reason behind this constant share was a large simultaneous issuance of government debt that was absorbed by other sectors. Only following the asset purchases

[5] 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.

[6] While QT1 officially ended in August 2019, we include the last two quarters of 2019 in this time period.

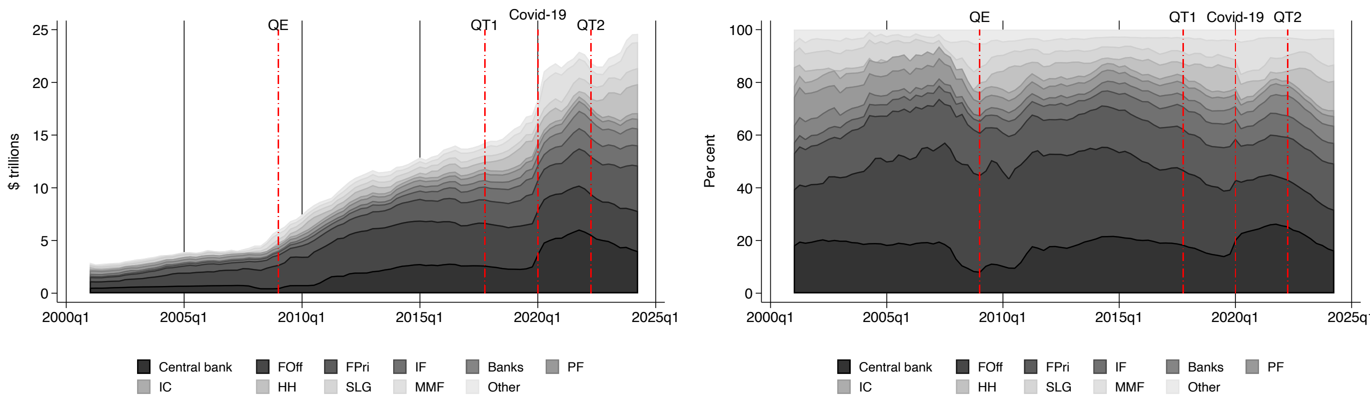
in response to the pandemic did the share of the central bank increase materially. This also provides a perspective on the sheer magnitude of quantitative easing conducted by the Federal Reserve amid the pandemic in relation to the one conducted during the post-GFC period. Second, holdings by the foreign official sector (i.e. reserve managers in foreign central banks or sovereign wealth funds) has been trending down in relative terms since 2017. While the share of these investors once stood at around 33% during the post-GFC period, it has since then fallen to a mere 16.4%. Similarly, the share of pension funds and insurance companies fell since the GFC. This is in part due to higher incentives to take on more risk in the low-for-long interest rate era, tilting portfolios towards riskier assets (in line with the portfolio balance channel of QE). The resulting slack was picked up mostly by the investment fund sector (mutual funds and money market funds) and the household sector. Finally, during periods of sizable balance sheet contraction, the share of households rose the most, while also the share of foreign private investors and commercial banks increased somewhat. The increase in holdings by the household sector needs to be interpreted with special care here, as it is largely composed of hedge funds (Vissing-Jorgensen, 2021). Our findings indicate that in QT phases hedge funds tended to ramp up their holdings of Treasuries, mostly driven by cash-futures basis arbitrage trades (Schrimpf et al., 2020; Barth and Kahn, 2021).

Our accounting framework sheds further light on the evolution of holdings. 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 (suppressing the jurisdiction superscript j here for simplicity):

$$\frac{H_t^s - H_{t-1}^s}{D_{t-1}} = \alpha^s + \boldsymbol{\beta^{s\prime}} \boldsymbol{\iota_t} \times \frac{D_t - D_{t-1}}{D_{t-1}} + \varepsilon_t^s, \tag{2}$$

where $\boldsymbol{\iota}_t = (\iota_{\text{Pre-GFC},t} \quad \iota_{\text{Post-GFC},t} \quad \iota_{\text{QT1},t} \quad \iota_{\text{Covid},t} \quad \iota_{\text{QT2},t})'$ denotes a vector that collects dummy variables taking the value 1 if period t belongs to a particular phase, and 0 otherwise. As key phases we distinguish the periods before the GFC, the phase of central bank balance sheet expansion after the GFC (following the start of QE), the first QT period, Covid-19 crisis and its aftermath, and the second phase of QT following the post-pandemic inflation surge. The corresponding coefficients

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 2001Q1 and 2024Q2 (quarterly data). Central bank refers to the holdings of the Federal Reserve. Foff and FPri refer to the holdings of the foreign official and private sectors, respectively. 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. Pre-GFC is between 2001Q1 and 2008Q4, Post-GFC is between 2009Q1 and 2017Q3, Post-QT1 is between 2017Q4 and 2019Q4, Post-Covid is between 2020Q1 and 2022Q2 and Post-QT2 is between 2022Q3 and 2024Q2. *Source: Federal Reserve.*

in turn are collected in the vector $\boldsymbol{\beta^s} = \begin{pmatrix} \beta^s_{\text{Pre-GFC}} & \beta^s_{\text{Post-GFC}} & \beta^s_{\text{QT1}} & \beta^s_{\text{Covid}} & \beta^s_{\text{QT2}} \end{pmatrix}$. We interpret these coefficients as the marginal response of sector s during the respective time period. And, as before, the coefficients sum to zero across sectors, ie. $\sum_s \beta^s_{Pre-GFC} = \sum_s \beta^s_{Post-GFC} = \sum_s \beta^s_{QT1} = \sum_s \beta^s_{Covid} = \sum_s \beta^s_{QT2} = 1$.

We report the regression results of Equation (2) for the United States in Table 1 Panel (b). The rows show the OLS estimates of $\beta^s_{Pre-GFC}$, $\beta^s_{Post-GFC}$, β^s_{QT1}, β^s_{Covid} and β^s_{QT2}, respectively. Each column shows the results for sector s denoted in the column heading. Before the GFC, for every additional change in government debt, 31% was absorbed by the foreign official sector, 26% by the domestic household sector, 21% by MMFs, and 9% was absorbed by foreign private investors.[7]

This pattern of debt supply absorption changed substantially after the GFC, and especially so

[7]Another 15% was absorbed by "other" sectors combined. 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.

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	FOff	FPri	PF	IF	Banks	SLG	MMF	HH	IC	Other
Pre-GFC	18.1	30.3	16.3	8	4.4	4.1	7.6	4	-.6	4.7	3.1
Post-GFC	17.3	33.8	14.7	4.3	5.9	3.4	4.8	4.8	5	2.7	3.4
QT1	16	26.5	15.5	4.8	8.8	4.7	4	6	7.9	2.4	3.5
Covid	24.4	19.4	14.4	4	7.7	6	4.6	9.8	4.4	1.9	3.4
QT2	20.2	16.4	16.8	3.6	7.9	6.7	6.2	7.3	9.7	1.9	3.3

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

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
VARIABLES	CB	FOff	FPri	IF	Banks	PF	IC	HH	SLG	MMF	Other
Pre-GFC * Pct. Ch. Gov. Debt	-0.02	0.31***	0.09***	0.01**	-0.01	0.00	0.02***	0.26***	-0.02*	0.21***	0.15***
	(0.05)	(0.04)	(0.03)	(0.01)	(0.01)	(0.01)	(0.01)	(0.04)	(0.01)	(0.06)	(0.02)
Post-GFC * Pct. Ch. Gov. Debt	0.19***	0.28***	0.04	0.08***	0.06***	0.02***	0.02**	0.27***	-0.00	0.01	0.04**
	(0.07)	(0.06)	(0.05)	(0.01)	(0.01)	(0.01)	(0.01)	(0.07)	(0.01)	(0.06)	(0.02)
QT1 * Pct. Ch. Gov. Debt	-0.02	-0.01	0.17***	0.08***	0.10***	0.04**	-0.00	0.33***	-0.02	0.26***	0.10***
	(0.16)	(0.06)	(0.05)	(0.01)	(0.01)	(0.02)	(0.01)	(0.11)	(0.02)	(0.04)	(0.03)
Covid * Pct. Ch. Gov. Debt	0.51***	-0.02	0.01	-0.01	0.07***	0.00	0.01**	-0.01	0.04***	0.33***	0.08***
	(0.07)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.03)	(0.01)	(0.04)	(0.01)
QT2 * Pct. Ch. Gov. Debt	-0.08	-0.00	0.11***	0.06***	0.04	0.05***	0.01**	0.32***	0.03*	0.43***	0.04**
	(0.16)	(0.05)	(0.02)	(0.01)	(0.02)	(0.01)	(0.01)	(0.08)	(0.01)	(0.06)	(0.02)
Observations	93	93	93	93	93	93	93	93	93	93	93
R-squared	0.52	0.37	0.04	0.26	0.19	0.07	0.11	0.29	0.08	0.46	0.27

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. FOff refers to foreign official sector. FPri refers to foreign private investors. IF refers to investment funds (open-ended mutual funds, exchange-traded funds and closed-end funds). Banks refer to commercial banks. PF refers to pension funds. IC refers to insurance companies. HH includes the direct and indirect holdings (e.g. through hedge funds) of households. SLG refers to state and local governments. MMF refers to money market funds. Other refers to all other sectors combined. Data are quarterly. Pre-GFC is between 2000Q1 and 2008Q4. Post-GFC is between 2009Q1 and 2017Q3. QT1 is between 2017Q4 and 2019Q4. Covid is between 2020Q1 and 2022Q2. QT2 is between 2022Q3 and 2024Q2. 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.

with the advent of central bank balance sheet policies. Between the start of the first QE program until the first QT program, the central bank increased its marginal absorption of debt supply from virtually zero to 19%, investment funds from 1% to 8% and commercial banks from zero to 6%. The marginal role of foreign private investors, MMFs and other investors, by contrast, declined.

This reflects the changing investment and regulatory landscape as broker-dealers scaled down their market making activities, commercial banks increasing their holdings to comply with liquidity regulations and the mutual fund sector receiving large inflows.

During the first quantitative tightening phase, as the central bank halted its active government bond purchases, for every dollar of new government debt, households (in particular hedge funds) bought 33 cents, MMFs 26 cents, foreign private investors 17 cents, commercial banks 10 cents, investment funds 8 cents, and pension funds bought 4 cents.

During the massive balance sheet expansion following the Covid-19 crisis, however, only two sectors – the central bank and MMFs[8] – jointly account for the absorption of 84% of a one unit increase government debt supplied. The marginal role of almost all other sectors declined.

Finally, during the second QT period, similar to the previous one, most of the new issuance was absorbed by households (ie. effectively the hedge fund sector), foreign private investors, MMFs and investment funds with limited role for commercial banks this time around.

2.4 Summary of findings for the 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.[9] There are a few commonalities, particularly regarding the expanding role of the central bank, along with several idiosyncratic developments.

Similar to the United States, central banks in other major currency areas increased their footprints in their respective sovereign bond markets substantially via various rounds of QE. In the Euro area, the central bank has absorbed about half of every unit of new government debt since the launch of its public sector asset purchase program (PSPP), raising its holdings share from less than 3% to 18%. In the United Kingdom, the central bank has taken down about 40-50% of new debt, leading to a share of about one third on average. The expansion of the central bank's role

[8] 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 et al., 2023, for the role of MMFs in the Treasury bill market and the ON RRP facility).

[9] We provide an overview of the results in this section. We refer the interested reader to Section IA.I of the Internet Appendix for the detailed analysis and results.

is most extreme in Japan. Between QQE and the Covid-19 outbreak, the Bank of Japan bought 1.79 units for every unit increase in the outstanding amount of government bonds. This indicates that the central bank absorbed a large share of the new issuance as well as the holdings of other participants. While the pace slowed after Covid-19, the central bank accounted for almost half of the market share in the post-pandemic period.

The compositional shift among non-central bank holders is more heterogeneous across regions. In the Euro area, the marginal absorption by non-central-bank players declined, including for foreign investors (see Koijen et al., 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.[10] In the United Kingdom, the central bank balance sheet expanded rapidly following the GFC, with major domestic investors losing market share to the central bank while foreign investors maintained and even slightly increased their share.

3 The demand elasticities of different types of investors

As central banks embarked on quantitative easing programs to bring down 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). Due to data availability and quality, we focus primarily on the United States, but report a summary of the results and comparisons for other jurisdictions as well in Section 3.7.

This quantification is crucial for assessing the impact of central bank asset purchases and the subsequent unwinding of the balance sheets. The framework can also shed light on the potential

[10] A likely reason for these increased holdings has been foreign investors engaging in Covered Interest Parity Arbitrage (Du et al., 2019; Rime et al., 2022). To do so, foreign investors enjoying favorable funding costs would acquire US dollar funding in US wholesale money markets (say from MMFs), swap them into yen using an FX swap, and park the proceeds in the Bank of Japan's deposit facility or (short-term) government securities.

impact of the changes in the investor composition, such as China reducing their holdings of US Treasuries.

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 $\mathbf{X_t}$:

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

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 $\boldsymbol{\beta_2^{s'}}$ captures 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, and hence 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.[11] Faced with this challenge, empirical researchers often resort to a parametric specification of the outside asset which aims to capture other investment opportunities investors have.

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:

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

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 et al., 2021; Fang et al., 2022; Jansen, 2023). This takes a linear form with $log(H(0)_t^s) = \phi^s + \boldsymbol{\gamma^{s'}} \; \mathbf{W_t}$, where $\mathbf{W_t}$ is a vector that is meant to capture factors that affect the holdings of the outside asset, and $\boldsymbol{\gamma^{s'}}$ 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 Flow of Funds statistics about other financial assets which is available for some sectors. We use debt securities other than Treasuries as the outside asset. That said, information on the total assets of some key investor groups such as foreign official and foreign private investors is unfortunately missing from this dataset. Hence, for foreign investors we cannot use information on holdings of outside assets; for these investors, we thus only use a parametric specification as is common in the literature.

[11] In this case, one approach would be to include time fixed effects, but such an approach is not feasible since elasticities are estimated from time-series variation in yields.

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 (5) for *all sectors* with two-stage least squares:

$$log(H_t^s) = (\alpha^s + \phi^s) + \beta_1^s \; Y_t^8 + \boldsymbol{\beta_2^{s'}} \; \mathbf{X_t} + \boldsymbol{\gamma^{s'}} \; \mathbf{W_t} + t + t^2 + \eta_t^s. \tag{5}$$

In addition, for the subset of sectors included in the Flow of Funds dataset, we make use of our knowledge of the holdings of debt securities other than Treasuries 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 debt securities. We then estimate Equation (6) for these sectors using two-stage least squares:

$$log(H_t^s) - log(H(0)_t^s) = \alpha^s + \beta_1^s \; Y_t^8 + \boldsymbol{\beta_2^{s'}} \; \mathbf{X_t} + t + t^2 + \eta_t^s. \tag{6}$$

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 average duration of assets held by most of the investors of US Treasuries as documented by Tabova and Warnock (2022).[12] We include a (quadratic) trend in order to get identification from the deviations from the trend.[13]

Crucially for our analysis, we use instrumental variables and estimate β_1^s using 2SLS regressions in order to address endogeneity between the latent demand and yields. Here we use cleanly identified high-frequency monetary policy surprises as instruments for government bond yields

[12] In the Internet Appendix, we estimate Equation (5) using different yields (5- and 10-year zero-coupon yields and get similar results.

[13] 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.

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 et al., 2021, for a framework 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 et al. (2021) as they include the yield on US Treasury bonds when they parameterize the outside asset for European investors. Including the VIX is 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.[14]

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.[15] For our baseline results, we assume 8 years as the average modified duration for each sector.[16] We calculate the percentage change in the price of their holdings as:

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

[14]See Section 3.6 for a discussion of the robustness checks and the Internet Appendix.

[15]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. This is likely to lead to a downward bias in our elasticity estimate, making the demand curve appear upward sloping with respect to prices.

[16]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 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.

Using the total holdings by each sector, we translate these percentage changes into those effects in quantities 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 propose high-frequency monetary policy surprises as instruments for government bond yields. These surprises are typically measured as yield changes of a range of interest rates, inferred from either cash bonds or futures, within a short intra-day window around monetary policy announcements. This high-frequency identification approach is well-suited to address endogeneity concerns arising from the fact that yields are jointly endogenous with latent demand since the major investors we are interested in are slow-moving and not the ones behind the high-frequency price reaction (see below). Hence, these surprises can be considered exogenous to the actions of the main investors we are interested in. To match our quarterly holdings data, we aggregate monetary policy surprises, measured within short windows, over each quarter.

The literature has proposed a number of such monetary policy surprises. In principle, these surprises satisfy the relevance conditions since they are known to move 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 relatively small, in the order of a few basis points (Nakamura and Steinsson, 2018). Therefore, while these instruments are relevant, we might face a weak instruments problem.

Identification considerations. Our identification argument is that monetary policy surprises are uncorrelated with the latent demand of major investor groups for government bonds. There could be several potential concerns with this identification strategy, however, which we address in the following discussions.

One threat is simultaneity, which could occur if portfolio adjustments by those investors whose demand elasticity we are interested in measuring were in fact the ones behind the price action around the monetary policy event. We do not see this as a major concern. Even if these changes partly result from actual trading, it is unlikely due to the flow from the main investor groups we are interested in, such as, reserve managers, commercial banks, mutual funds, pension funds, insurance companies etc. Such real-money investors are known to typically rebalance their portfolios in a slow manner. The estimated surprises in our data, by contrast, reflect price changes in a very narrow window (less than one or two hours) around monetary policy events. Undoubtedly, in these short periods market makers in fixed income markets (broker-dealers) will adjust their quotes, while faster investor types – such as certain types of hedge funds – may seek to benefit by trading during these short windows around monetary policy announcements. Our identification strategy may hence not work for these types of players. To keep our identification clean from these concerns, we do not include the household sector, broker-dealers and other related entities in our elasticity estimation. Instead, we focus on the foreign official investors, foreign private investors, commercial banks,[17] investment funds, pension funds, insurance companies, as well as state and local governments.

A second threat to identification would arise if our monetary policy surprises identified at high frequency were to contain information about the state of the economy, that is if they were to be dominated by *central bank information effects* (e.g. Jarociński and Karadi, 2020).[18] In this case, the exclusion restriction would be violated since the monetary policy surprises do not necessarily alter holdings *only* through the changes in yields, but holdings might respond directly to the information revealed by the central bank. Therefore, even if the portfolio rebalancing occurs outside of the window in which the monetary policy surprises are constructed, the response may still be correlated with the information release. To guard against this possibility, we construct monetary policy surprises that are uncorrelated with the estimates of information shocks.

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 as central bank mandates and doctrine prescribe to set policy to stabilize the economy in case of inflation shortfalls from target and output deviating from potential.

[17] Note that the sectorial information for banks at our disposal excludes broker-dealers which act as market makers in fixed income markets

[18] Note that the existence of "information shock" is still in debate. For instance, it has been recently challenged by Bauer and Swanson (2023).

Finally, to the extent that the monetary policy surprises also affect the control variables, our estimates for the response of demand to Treasury yields could be biased. To alleviate this concern, we lag all controls in the baseline regression by one quarter to make sure that they are recorded prior to the realization of the monetary policy surprises and hence are not affected by the monetary policy surprises in the given quarter. We discuss this further in Section 3.6. Moreover, the instrument we use has a high loading on monetary policy shocks that primarily measure the responses at the longer-end of the yield curve. As such, our measure does capture shocks about the conduct of balance sheet policies, which alter the supply of government bonds available.

IV construction based on monetary policy surprises. There are multiple measures of monetary policy surprises proposed in the literature. Rather than taking a prior stance on which ones to use as instruments, we take a data-driven approach and select an instrument that satisfies the following three criteria. *First*, we would like to avoid the problem of weak instruments. Therefore, our first criterion is to have a strong first-stage result. *Second*, since we are primarily interested in analyzing the most recent period and study episodes of central bank balance sheet expansion and contraction, we favor monetary policy surprises that are available during the later part of the sample (covering both QT periods). *Third*, in an attempt to have an instrument that satisfies the exclusion restriction, we favor an instrument that is uncorrelated with the series of central bank information shocks identified in the literature (Jarociński and Karadi, 2020).

In order to construct an instrument that satisfies these criteria, we use various measures of monetary policy surprises individually and also combine them through a principal component analysis. The goal of the latter approach is to make use of information contained in all of the series while keeping the dimensionality in check since we have a relatively short sample period. We report the first-stage results for all individual surprise measures and select our baseline instrument for the second stage according to the satisfaction of the criteria above. Additional diagnostic results are relegated to the Internet Appendix.

We rely on monetary policy surprises that are commonly used in the literature from five sources. The first source is Swanson (2021), who uses a factor analysis of changes in various asset prices, including fixed income instruments along the term structure, around monetary policy news releases to separately identify surprise changes in the federal funds rate, forward guidance, and large-scale asset purchases (LSAPs). The second source is Bu et al. (2021), who develop a heteroskedasticity-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. The third is Kearns et al. (2022), who share a similar goal as Swanson (2021), but take a simpler approach to construct target rate, path, and long-rate surprises. The fourth is Jarociński and Karadi (2020), who decompose monetary policy surprises at the short-end of the yield curve into monetary policy and information shocks using high-frequency co-movement between interest rates and stock prices. Following the third criterion above, we only use the monetary policy surprise and not the information shock.[19] Finally, Nakamura and Steinsson (2018) construct the monetary policy surprise as the first principal component of high-frequency rate changes based on Fed funds and Eurodollar futures with expiry up to a year.[20]

In addition to using the respective monetary policy surprises individually, we also generate a composite monetary policy surprise time series using a principal components analysis, for the reasons outlined above. One data challenge is that various monetary policy surprise series are available for different sample periods. Recent updates are available for three of the monetary policy surprise series, that is, the ones constructed by Jarociński and Karadi (2020), Bu et al. (2021), and Kearns et al. (2022). Relying on these allows us to estimate the model using a sample period that runs until 2024Q2, which is desirable as it covers the recent QT2 episode (our second criterion above).[21] We take 2004Q3 as the beginning of our sample. Doing so allows us to have a common sample across several measures available that covers the key periods we are interested in. Moreover, we find that the instrument is weaker in the earlier periods (unreported).

[19] 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 deal with this issue by selecting an instrument that is uncorrelated with the identified information shocks as our baseline.

[20] For the updated series calculated in Acosta et al. (2024), they use SOFR futures instead of Eurodollar futures from January 2022 onward, given Libor's discontinuation following the benchmark interest rate reform.

[21] Swanson (2021) surprises end in mid-2019, Nakamura and Steinsson (2018) surprises, which are updated by Acosta et al. (2024) run until 2022Q3.

We report the results of the first-stage regression for the estimation of Equation (5) in Table 2. Results of the first stage with Equation (6) in turn are shown in the Internet Appendix.[22] In column (1), we report the results using the Swanson (2021) series. In column (2), we report the results using the Bu et al. (2021) surprises. In column (3), we report the results using the Kearns et al. (2022) series. In column (4), we report the results using the Jarociński and Karadi (2020) monetary policy surprise series. In column (5), we report the results using the Nakamura and Steinsson (2018) monetary policy shock series.

The results in Table 2 indicate that, in most cases, monetary policy surprises taken individually yield weak instruments and thus strengthen the case of using the PCA approach when constructing our main instrument. While in many instances the coefficients on individual surprise series have the economically meaningful sign and are statistically significant, in most of these specifications (in columns (1)-(4)), 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 our instruments are weak.[23] 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.[24]

We proceed with the monetary policy surprises constructed using principal components analysis. First, we take the first principal component of all series that are available through 2024Q2 (Bu et al., 2021; Jarociński and Karadi, 2020; Kearns et al., 2022). The estimated coefficient is positive and significant and the effective F-statistic becomes 47, and hence we reject the null hypothesis of weak instruments. Moreover, this composite surprise series has the lowest correlation with central bank information shocks (8%).[25] We therefore select this instrument PCA 1 (JK MP, BRW, KSX) to be

[22]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.

[23]We also repeated this exercise to check whether specifying the government bond yield and other control variables in changes instead of levels would be preferable and yield a stronger first-stage. However, doing so also leads to a failure to reject the null hypothesis of weak instruments. Therefore, we stick to our specification in Equations (5) and (6).

[24]The only exception is the Nakamura and Steinsson (2018) surprise series. Using this surprise series yields a very strong first stage with an effective F-statistic of 45.6. However, the downside of this series is it ends earlier and it has a 30% correlation with the central bank information shocks of Jarociński and Karadi (2020) (Table 3).

[25]We also run robustness checks by regressing this series on the information shocks and using the residuals as our instrument. The results are similar.

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	0.0012***						
	(0.0003)						
Swanson FG	0.0006**						
	(0.0003)						
Swanson LSAP	-0.0002						
	(0.0002)						
BRW		0.0005**					
		(0.0002)					
KSX (3M)			0.0009***				
			(0.0003)				
KSX (2Y)			0.0001				
			(0.0002)				
KSX (10Y)			0.0006**				
			(0.0002)				
JK MP				0.0012***			
				(0.0003)			
Nakamura-Steinsson					0.0012***		
					(0.0002)		
PCA 1 (JK MP, BRW, KSX)						0.0011***	
						(0.0001)	
PCA 1 (JK MP, BRW, KSX, NS)							0.0012***
							(0.0001)
Observations	60	80	80	80	73	80	73
R-squared	0.9207	0.9038	0.9083	0.9078	0.9030	0.9085	0.9045
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	2004q3-2024q2	2004q3-2024q2	2004q3-2024q2	2004q3-2022q3	2004q3-2024q2	2004q3-2022q3
Effective F-stat	4.84	4.46	8.33	13.71	45.60	47.01	68.65
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	15.15	23.11	18.68	23.11	23.11	23.11	23.11

Note: This table reports the estimated coefficients of the first-stage regression that provides the basis for the ensuing second-stage estimation of Equation (5). The sample period varies depending on the availability of data across different monetary policy surprises and is reported in each column. All monetary policy surprise series are 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

the baseline as it satisfies our main criteria outlined.[26] In Table IA.IV of the Internet Appendix, we show that PCA 1 (JK MP, BRW, KSX) has a high loading on monetary policy shocks that primarily measure the responses at the longer-end of the yield curve such as the shocks from Bu

[26]Nonetheless, we also run the first stage adding the Nakamura and Steinsson (2018) shock series. This first stage gives an even greater effective F-statistic (68.6), which however comes at the cost of a shorter sample period and a greater correlation with information shocks. We report the second-stage results with this instrument in the Internet Appendix. They are broadly similar to our main results.

Table 3: Correlation of alternative monetary policy surprises with central bank information shocks and each other

	JK CBI	Nakamura-Steinsson	PCA 1 (JK MP, BRW, KSX)	PCA 1 (JK MP, BRW, KSX, NS)
JK CBI	1.00			
Nakamura-Steinsson	0.30	1.00		
PCA 1 (JK MP, BRW, KSX)	0.08	0.78	1.00	
PCA 1 (JK MP, BRW, KSX, NS)	0.15	0.90	0.98	1.00

Note: This table reports the correlation coefficients of alternative monetary policy surprises with central bank information shocks and each other.

et al. (2021) and the 10-year shocks from Kearns et al. (2022). This reinforces the point above that our shock measure is related the conduct of balance sheet policies, which alter the supply of government bonds available.

In the Internet Appendix, we report various further results and diagnostic checks for our main instrument. These include the details of the principal components analysis and the residuals from running a regression of these shocks on the central bank information shocks.

3.4 Second-stage results

We replace the government bond yields in Equation (5) by the fitted values recovered from the first stage to consistently estimate β_1^s in the second stage. A higher estimate 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. $\widehat{\beta_1^s}$ %. 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 (5) in the first panel of Table 4.[27] For robustness, we report estimates of Equation (6) for the subset of sectors for which we have

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

information on other financial assets in the second panel of the same table. We use holdings of debt securities other than Treasuries as the outside asset. The results are broadly similar.

Our estimates indicate that the demand by banks and various asset managers is most elastic. We find that the yield elasticity of demand is positive and statistically significant for commercial banks (32.76), investment funds (22.04), foreign private investors (19.31), insurance companies (12.42) and pension funds (10.85), with the elasticity estimates indicated in parentheses.[28] These estimates indicate that these investor groups have a downward-sloping demand curve (with respect to prices). Commercial banks and investment funds stand out as as being the most price sensitive compared to other entities.[29]

These results are consistent with findings in other related work. Using data from the Netherlands, Jansen (2023) also finds that banks are the most price elastic sector. And Jansen et al. (2024) find that the elasticity of investment funds and commercial banks is larger than those of pension funds and insurance companies. Interestingly, we find the elasticity estimate for the foreign official sector to be statistically indistinguishable from zero. The result means that reserve managers in central banks are not particularly sensitive to changes in the yield of government debt. This is not too surprising as reserve accumulation is likely influenced by autonomous factors, notably the motive to build precautionary buffers of US Treasury bonds due to their safety and liquidity features.

3.5 Price elasticity of demand

We follow Koijen and Yogo (2019) and Koijen et al. (2021) and define the price elasticity of demand of sector s for Treasuries as:

[28] We provide a more detailed analysis of the investment fund sector in Section IA.V of the Internet Appendix. The analysis suggests that while open-ended mutual funds drive this elasticity, while ETF demand is inelastic (see Chaudhary et al., 2024, for a similar finding).

[29] Acknowledging the identification challenges, we repeat the analysis for the household sector and find a coefficient that is statistically indifferent from zero (unreported). This inelasticity could be rationalized by the business model of hedge funds. They are typically involved in arbitraging the changes in the cash-futures basis rather than responding to changes in yields. For the counterfactual scenarios below, we assume both the household and the broker-dealer sectors to be inelastic. If they have elastic downward-sloping demand, our estimates for the responses of yields can be taken as an upper bound.

Table 4: The yield elasticity of demand across different sectors in the United States

Panel (a): Using a parametric specification for the outside asset (Equation (5))

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	1.53	19.31*	22.04***	32.76***	10.85*	12.42***	1.51
	(7.01)	(9.93)	(6.73)	(11.38)	(5.79)	(3.76)	(9.40)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.83	0.03	0.00	0.00	0.05	0.00	0.87
Underidentification LM stat (p-val)	0.06	0.06	0.06	0.06	0.06	0.06	0.06

Panel (b): Using portfolio information to control for the outside asset (Equation (6))

VARIABLES	(1) log(IF)-log(DebtOA$_{IF}$)	(2) log(Banks)-log(DebtOA$_{Banks}$)	(3) log(PF)-log(DebtOA$_{PF}$)	(4) log(IC)-log(DebtOA$_{IC}$)	(5) log(SLG)-log(DebtOA$_{SLG}$)
8Y Yield (ZC)	20.09***	33.34***	12.70***	12.72***	6.10
	(5.45)	(6.13)	(4.16)	(3.92)	(7.87)
Observations	80	80	80	80	80
Controls	**X**	**X**	**X**	**X**	**X**
Anderson-Rubin Wald test (p-val)	0.000	0.000	0.001	0.001	0.441
Underidentification LM stat (p-val)	0.090	0.090	0.090	0.090	0.090

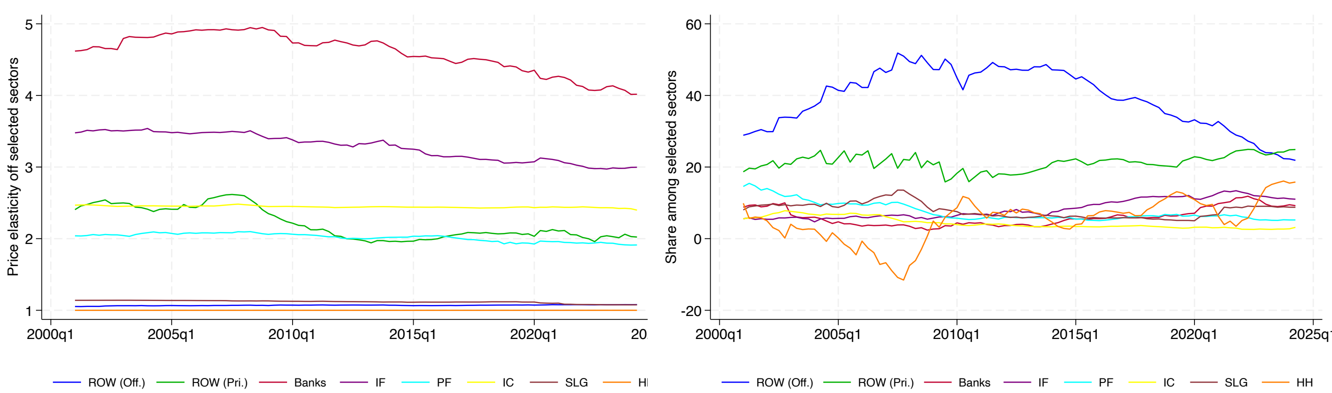
Note: This table reports the coefficients of the second-stage regression specified in Equation (5) in Panel (a) and Equation (6) in Panel (b) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The sample period is between 2004q3 and 2024q2. Anderson-Rubin test has the null hypothesis that the 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.

$$-\frac{\partial log(Q_{s,t})}{\partial log(P_{s,t})} = 1 + \frac{\beta_1^s}{m_t} \times (1 - w_{s,t}),$$

where m_t is the maturity (we use 8-years throughout to approximately match our yield estimate). A higher coefficient β_1^s on the yield implies a greater price elasticity of demand. Finally, this calculation depends on the portfolio weight of Treasuries ($w_{s,t}$). For each domestic sector, we use the debt securities other than Treasuries as the outside asset obtained from the Flow of Funds.

For the foreign private sector, we first compute the share of private investors in total US Treasury holdings and apply this share to debt security holdings of foreigners obtained from the Flow of Funds. We combine this information with that on holdings of Treasuries to compute portfolio weights. For the foreign official sector, we simply use the dollar share of reserves obtained from the IMF COFER dataset. We use the shares of each sectors among those long-term holders we are interested in to calculate the weighted elasticities of these sectors combined, shown in Figure 2(b).

Figure 2: Price elasticity of demand and shares of sectors among major long-term investors



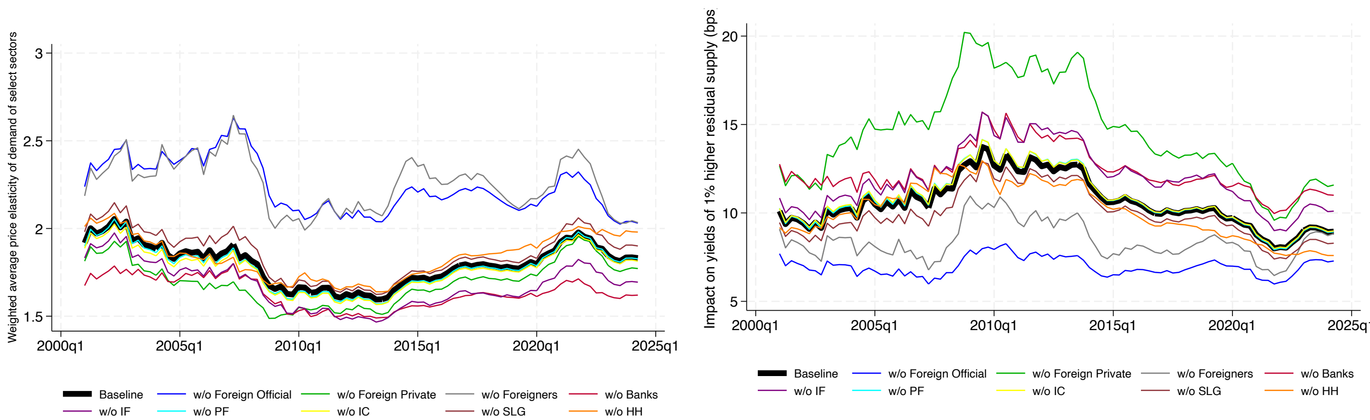
(a) Price elasticity of select sectors

(b) Market shares of select sectors over time

Note: The left-panel shows the price elasticity of demand of select sectors over time computed similarly as in Koijen and Yogo (2019). The right-panel shows the evolution of the shares of select sectors in the US government bond market.

These estimates suggest a weighted price elasticity of demand between 1.6 and 2.1 depending on the market composition. This approximately translates into an impact between around 8 to 13 basis points on 8-year yields of a 1% change in the residual supply, respectively as shown in the thick black lines of Figure 3. A higher price elasticity corresponds to a lower impact on yields. Other lines in these figures recalculate counterfactual elasticities and the impact on yields by taking one sector out at a time. For example, the weighted average price elasticity without the foreign official sector, as shown by the blue line, would be around 2.5, resulting in a much lower impact on yields. On the other hand, the counterfactual elasticity would have been much lower without the foreign private sector (green lines), resulting in a much higher impact on yields (with analogous interpretations in the case of the "take one out" analysis of other sectors).

Figure 3: Weighted price elasticity of demand, the approximate impact of a 1% higher residual supply on long-term yields and counterfactuals



(a) Weighted average price elasticity

(b) Approx. impact of a 1% higher residual supply on 8-year yields

Note: The left-panel shows the weighted average price elasticity of demand of the major long-term holders of US Treasuries. The right-panel computes the approximate impact on yields of 1% higher residual supply due to central bank operations computed as the weighted average response of these sectors. The thick black lines show the baseline estimates. Other lines are the counterfactual estimates taking one sector out at a time.

How do these estimates compare to those in the literature for government bonds and other assets? Chaudhary et al. (2024) find that 1% increase in demand increases 10-year US Treasury yields by 10 basis points. The estimates are relatively close to each other. Our results and those in Jansen et al. (2024) and Chaudhary et al. (2024) suggest that the demand for US Treasuries is more elastic than those for the aggregate equity market (Gabaix and Koijen, 2023) and the estimates for the corporate bond market (Chaudhary et al., 2023). The weighted average price elasticity for the Euro area is estimated to be around 3 (Koijen et al., 2021). For emerging market government bonds, Fang et al. (2022) find that 1% higher debt leads to 58 basis point yield increase, suggesting a significantly lower elasticity for emerging market government bonds.

We can also use these estimates to provide estimates for several counterfactual scenarios. There has been a recent debate on the potential impact of China and the Gulf States reducing their investments in US Treasuries or the decline in investments of oil revenues into US Treasuries, respectively. In the former case, our estimates suggest that China completely withdrawing from the US Treasury market would have an impact of around 26 basis points on long-term yields as China's current holdings account for a share of around 3% of total marketable US Treasury debt. Similarly, Gulf States' holdings also account for around 1.2% of the total outstanding US Treasuries, suggesting an impact of around 10 basis points should these countries complete disinvest. Here, we assume that the central bank does not participate in absorbing these flows and the effective increase in the supply would be absorbed by non-central bank players only. If the central bank also plays a role in absorbing these in an elastic manner, these estimates can be viewed as an upper bound. These estimates suggest that other participants would absorb US Treasuries under these counterfactual scenarios with a limited impact on prices absent other structural shifts.

3.6 Robustness checks and other considerations

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

We report the elasticities of different sectors with respect to variables used in $\mathbf{X_t}$ and $\mathbf{W_t}$ in the Internet Appendix (Section IA.III1). Important take-aways are (i) that the coefficient in the second stage for the 5-year German yield is negative, consistent with it being defined as the outside asset, and (ii) that 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.

In Section IA.III2 of the Internet Appendix, we vary the control variables used in the regression Equation (5). 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. We also consider a specification where we control for a linear trend instead of a quadratic trend. In another specification, we lag all the control variables used in the baseline regression by one quarter in order to make sure that they are recorded prior to the realization of the monetary policy surprises. We do so to alleviate the concern that the impact of the instrument on the endogenous variable could be mediated through the control variables. Our results remain qualitatively and quantitatively similar.

We vary the instruments in these regressions in robustness exercises reported in the Internet Appendix. In Section IA.III3, we report the results using PCA 1 (JK MP, BRW, KSX, NS) as an instrument instead of PCA 1 (JK MP, BRW, KSX) used in the baseline. In Section IA.III4, we take the residuals from regressing PCA 1 (JK MP, BRW, KSX) on JK Central Bank Information shocks and use them as the instrument. In Section IA.III5, we use the first two principal components (PCA 1 (JK MP, BRW, KSX) and PCA 2 (JK MP, BRW, KSX)) as instruments.

In Section IA.III6 of the Internet Appendix, 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 evidence in Tabova and Warnock (2022). We show that our main results also qualitatively hold if we assume the modified duration to be 5 or 10. In Section IA.III7 of the Internet Appendix, we use 5-year and 10-year zero-coupon government bond yields instead of the 8-year zero-coupon yields in the baseline. In Section IA.III8, we vary both yields and valuation adjustments to investigate the sensitivity of our results to these assumptions.

In Section IA.IV of the Internet Appendix, we provide further details about the principal components analysis that we use to generate our composite monetary policy surprise instruments.

3.7 A comparison of elasticities across jurisdictions

We focus our main analysis on the United States primarily due to the availability of multiple measures of monetary policy shocks, which allows us to avoid a weak instrument problem and also construct an instrument that is free from central bank information shocks. For other jurisdictions, we cannot credibly solve these issues to the same extent. Nevertheless, we provide some tentative estimates for Japan and the United Kingdom in Section IA.II in the Internet Appendix, acknowledging and discussing the limitations. In this section, we compare our elasticity estimates for various sectors in the United States to those in Japan, the United Kingdom and the Euro area (for which we do not provide any estimates, but report those in Koijen et al. (2021)).

Our monetary policy surprises for the the UK and Japan come from various sources. For Japan, we make use of monetary policy surprises identified in Kubota and Shintani (2022) and Kearns et al. (2022) to construct an instrument (subject to both the weak instrument and information effect problems). For the United Kingdom, we use the QE surprises identified in Braun et al. (2025) as an instrument. While this instrument does provide a strong first stage, it is not free of central bank information shocks. We compare different estimates for sectors in different jurisdictions here and refer the interested reader to Section IA.II for details. With all these caveats in mind, we acknowledge that the estimates for Japan and the United Kingdom might still contain some bias and recommend caution.

For Japan, we find that, the "Others" sector in the Flow of Funds, which comprises sectors other than the central bank, banks, insurance companies and private and public pension funds, and households, has the highest elasticity - though it is a rather small sector. Households also exhibit a high elasticity. Public pension funds and banks also turn out to be fairly elastic investors. Interestingly, we find a higher point estimate for banks in Japan compared to those in the United States. We also find a somewhat higher point estimate for foreign investors (official and private combined) compared to the United States, even though the estimated coefficient is statistically insignificant. We find that insurance companies and private pension funds are inelastic investors in Japanese government bond markets.

For the United Kingdom, we find a statistically significant downward-sloping demand function for foreign investors. We find that banks (classified as Monetary Financial Institutions - MFI) and households exhibit similar elasticity to foreigners. However, the estimate is statistically insignificant for these sectors. An interesting feature of UK government bond markets is the estimated upward-sloping demand curves for the insurance company and pension fund (ICPF) sectors as well as Other Financial Institutions (OFI) which include investment funds.

Upward-sloping demand curves in the context of financial assets can be due to several factors and act as shock amplifiers in financial markets. ICPFs typically have long-duration liabilities and have to match them with long-duration, low-risk assets, such as government bonds (e.g. Domanski et al., 2017). Similarly, OFIs could receive inflows as other investors sell government debt to the central bank and they might invest these inflows further into government debt (e.g. Fang and Xiao, 2025). This would create a positive feedback loop pushing yields even further during QE periods and create feedback loops amplifying downward price pressures during QT periods. Indeed, such amplification dynamics could have played a role during the UK gilt crisis in 2022.

For the Euro area, Koijen et al. (2021) find that foreign investors have the highest elasticity, followed by mutual funds and banks. Similar to our finding for the United Kingdom, they also find an upward sloping demand curve for the ICPF sector in the Euro area.

All in all, the variation in these estimates suggest that the heterogeneity of the demand elasticities of different sectors plays an important role in how central bank balance sheet policies impact asset prices and these policies can change the market composition consistent with our findings in Section 2.

4 Is there an asymmetry when the central bank share increases or falls?

In this section, we study whether the estimated elasticities meaningfully differ during periods of quantitative tightening—a question which has attracted considerable interest in academic and policy circles lately (see, e.g. D'Amico and Seida, 2024; Du et al., 2024; Jiang and Sun, 2024). While this is a question with great policy relevance, given that we only have a limited number of observations for which we can analyze the impact of QT, any such study faces small sample issues. With these caveats in mind, we nevertheless use our framework in order to shed some light on this question. In particular, we estimate the following equation:

$$
\begin{aligned}
log(H_t^s) =& \alpha^s + \beta_1^s \; Y_t^8 \times \mathbb{1}(\Delta \text{CB Share}_t \geq 0) + \beta_2^s \;\; Y_t^8 \times \mathbb{1}(\Delta CB\ Share_t < 0) \\
&+ \beta_3^s \; \mathbb{1}(\Delta CB\ Share_t < 0) + \Gamma^{s'} \; \mathbf{X_t} + \eta^{s'} \; \mathbf{W_t} + t + t^2 + \eta_t^s,
\end{aligned} \tag{7}
$$

where the $\mathbb{1}(\Delta CB\ Share \geq 0)$ takes the value 1 during quarters in which the share of the central bank in the government bond market increased compared to the previous quarter. $\mathbb{1}(\Delta CB\ Share_t <$

0) corresponds to a decrease of the central bank share instead. All other control variables are the same as before.

This specification allows us to extend somewhat beyond just focusing on around 18 observations for which QT1 and QT2 were in place. Since our main object of interest in this section is the absorption of non-central bank players of government bonds, periods with increasing central bank share corresponds to a lower effective supply of government bonds to be absorbed and vice versa. This comparison also allows us the split the sample more evenly as 58% of the sample corresponds to declining central bank share.[30]

Table 5: Second stage results with $\Delta CB\ Share \geq 0$ and $\Delta CB\ Share < 0$ interactions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(IF)	log(Banks)	log(PF)	log(IC)	log(SLG)
$8Y\ Yield\ (ZC) \times \Delta\ CB\ Share \geq 0$	1.67	13.94	26.86***	40.41***	14.72*	17.79**	-6.11
	(9.28)	(11.15)	(8.80)	(14.64)	(8.22)	(8.85)	(11.32)
$8Y\ Yield\ (ZC) \times \Delta\ CB\ Share < 0$	2.27	2.67	36.69**	54.52**	21.17	27.93*	-21.63
	(14.48)	(23.75)	(15.78)	(25.28)	(14.13)	(16.30)	(16.72)
$\Delta\ CB\ Share < 0$	-0.01	0.29	-0.26	-0.42	-0.22	-0.29	0.41
	(0.16)	(0.41)	(0.24)	(0.42)	(0.18)	(0.23)	(0.32)
Observations	80	80	80	80	80	80	80
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Equal coefficients (p-val)	0.92	0.41	0.34	0.37	0.36	0.29	0.23
Anderson-Rubin Wald test (p-val)	0.98	0.05	0.00	0.00	0.12	0.00	0.15
Underidentification LM stat (p-val)	0.13	0.13	0.12	0.13	0.13	0.12	0.13

Note: This table reports the coefficients of the second-stage regression specified in Equation (7) in using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The sample period is between 2004q3 and 2024q2. The $\Delta CB\ Share \geq 0$ and $\Delta CB\ Share < 0$ are dummy variables which are 1 if the central bank share in the government bond market increased or decreased from the previous quarter, respectively. Equal coefficients row reports the p-values of the hypothesis test whether the coefficients for the interaction terms are equal. 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.

We present the regression results in Table 5 using the same instrument for Y_t^8 as before in our baseline analysis. Overall, we cannot reject the null hypothesis that the response of non-central

[30]The results remain qualitatively similar if we include a dummy variable for the post-GFC period (see Table IA.24 in the Internet Appendix). For robustness, we also account for the possibility that QE1 was potentially different from other QE/QT programs since market participants did not anticipate it, whereas some expectations were built in for other programs.

bank participants during increasing or decreasing central bank share is symmetric. That said, we find that for some sectors, such as commercial banks, investment funds, pension funds and insurance companies, the point estimates are larger in magnitude during periods in which the central bank share declines, suggesting somewhat greater elasticity during these periods, albeit the difference is statistically insignificant.

While we do not find evidence that the response of each sector's demand does not differ when the central bank share increases or decreases, changing market shares of sectors could indeed lead to an asymmetric market response to QE versus QT. We estimate that the demand by non-central bank long-term holders was less elastic until 2010 and has become more elastic over time as shown in Figure 3. A consequence of the change in the composition of the market is that QE took place in a relatively more inelastic market compared to QT. This implies an asymmetric market response to QE versus QT, whereas earlier QE programs had a larger price impact.

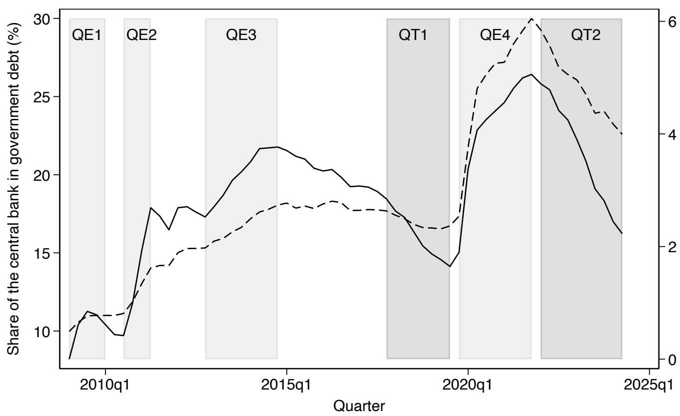
Using our elasticity estimates, we quantify the impact of various quantitative easing and quantitative tightening programs on long-term yields in Figure 4. We use the change in the central bank share as the change in the effective supply for non-central bank players to absorb and use the estimates obtained above to assess the total impact of these programs on long-term yields.[31] Based on these estimates, central bank asset purchases during various QE programs had the following impact on long-term yields: 25 bps during QE1, 113 bps during QE2, 57 bps during QE3 and 121 bps during QE4. On the other hand, in the absence of central bank purchases, the effective increase in the government bond supply led to a 43 bps increase in long-term yields during QT1 and a 80 bps increase in long-term yields until the second quarter of 2024 during QT2.[32]

In a counterfactual analysis, we assigned the market composition observed during the final quarter of the sample, as QT progressed, to each QE program. Similarly, we applied the peak aggregate market inelasticity observed during QE periods to QT programs. Under this scenario, the estimated effects on yields were 17 bps for QE1, 76 bps for QE2, 40 bps for QE3, and 112 bps

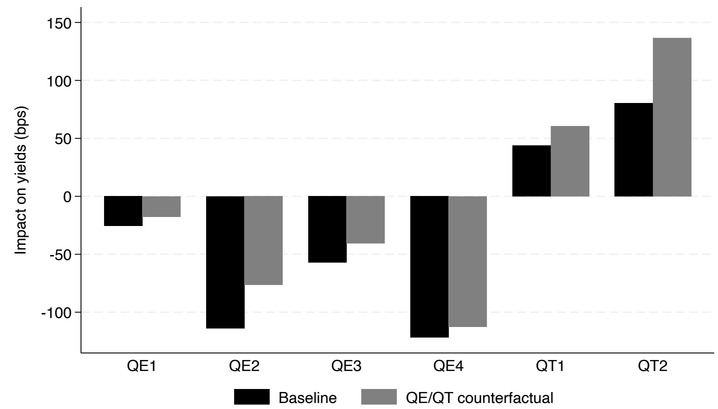
[31] As in the rest of the paper, we assume an elasticity of zero for the household sector and the "other" category, which includes broker-dealers.

[32] The effects we identify relate to purely the impact of debt supply absorption given inelasticity of demand when the policies are implemented. It is possible that these effects do not provide the full picture if announcements of policies also contain signalling effects among others, which is particularly the case for QE1 and QE4 (during Covid).

Figure 4: Impact of QE and QT on yields and counterfactuals



(a) Central bank holdings and share



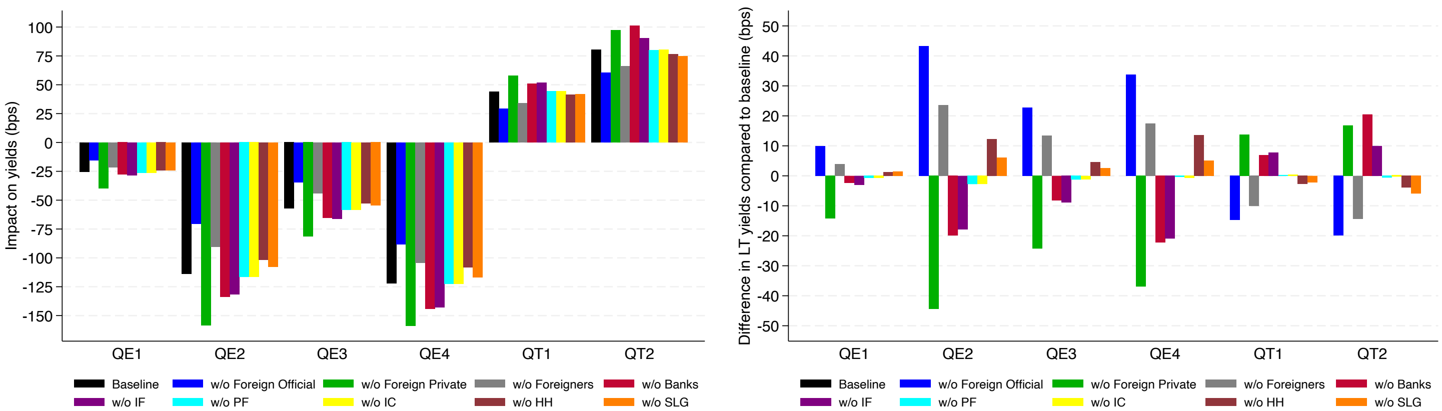
(b) Impact of QE/QT programs on yields and QE/QT counterfactuals

Note: The left-panel shows the evolution of the central bank share (solid line) and total holdings (dashed line, second axis) in the government bond market. The highlighted areas correspond to various QE and QT programs. Using the changes in the central bank share and assuming that they need to be absorbed by non-central bank players, the black bars in the right-panel compute the impact of various QE and QT programs. The gray lines for QE programs assume the market composition during these programs were the same as the last period in our sample. The gray lines for QT assume the most inelastic market composition during the QE programs to compute their counterfactual impact during QT had the market composition remained the same.

for QE4. For QT programs, the estimated impacts were 60 bps for QT1 and 136 bps for QT2. These findings highlight that market composition during central bank balance sheet operations can significantly influence yields and reveal the asymmetric effects of quantitative easing and tightening programs.

In additional counterfactual exercises, in order to gauge the impact of the international role of the dollar in the effectiveness of balance sheet policies, we estimate the price impact of various QE/QT programs with and without the foreign official sector, foreign private sector, without foreigners as a whole, and without each individual sector shown in Figure 5. We find that without the foreign official sector (and assuming their holdings would have been distributed proportionally across sectors depending on their market share) the impact of QE1, QE2, QE3 and QE4 on lowering long-term yields would have been attenuated by 10 bps, 43 bps, 22 bps and 34 bps, respectively. Similarly, the impact of QT1 and QT2 would be attenuated by 15 bps and 20 bps. On the other hand, without the foreign private sector, the impact of these programs would be amplified by 14 bps, 44 bps, 24 bps and 37 bps for QE1, QE2, QE3 and QE4, respectively, and by 14 bps and 17 bps for QT1 and Q2, respectively. Without both of these sectors, the impact would have been closer to the absence of the foreign official sector whereby the impact of the balance sheet policies would be less, but slightly less attenuated.[33]

Figure 5: Impact of QE and QT and counterfactuals with "take one sector out"



(a) Impact on yields with "take one sector out" **(b)** Difference to baseline

Note: The black bars in the left-panel shows the baseline impact of various QE and QT programs on long-term yields. Other bars show the counterfactual estimates of the impact without each individual sector. The right-panel show the difference in yields without each sector compared to the baseline. A positive number suggests that yields during a given QE/QT program would have been higher than the baseline estimate.

5 Conclusion

This paper provides a comprehensive analysis of the demand for government debt across different investor groups. By estimating the demand elasticities using instrumental variables derived from monetary policy surprises, we uncover significant heterogeneity in the elasticity of various investor groups' demand for government debt. Our findings indicate that commercial banks and investment

[33]The absence of banks or investment funds would work similarly to the absence of the foreign private sector even though with a much smaller impact. The absence of the household sector and state and local governments would be similar to the absence of the foreign official sector, similarly with a small impact. The absence of pension funds and insurance companies would have a minimal effect.

funds exhibit the highest elasticity, while the foreign official sector has price-inelastic demand. We also cannot reject the null hypothesis that the response of different sectors to quantitative easing and quantitative tightening is symmetric. However, market composition of sectors with varying elasticities can lead to an asymmetric overall impact of QE and QT. Since QE took place in a relatively more inelastic market compared to QT, our estimates suggest that the overall impact of QE was more pronounced than the reverse effect of QT.

These results have implications for monetary policy strategy and conduct. With balance sheet policies now an integral part of central bank toolkits, 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. Our results across jurisdictions highlight the crucial role of market composition when evaluating the likely impact of these policies.

Our results also have important implications for the interactions of monetary policy with the international role of the dollar. Our findings suggest that QE policies in the past would have been less effective without the foreign official sector. Therefore, with the declining share of the foreign official sector and a more elastic market, future QE policies might be less effective than those in the past.

Our results also have implications for financial stability. Heterogeneous elasticity estimates across sectors imply a change in the composition of government debt holders as central banks normalize balance sheets. Some of these players stepping in as marginal buyers notably non-bank financial intermediaries exposed to liquidity mismatches, feature vulnerable business models. Tracking their absorption role and market footprint is hence important from a financial stability perspective. All these issues are set to become even more pertinent as the amount of debt supply that markets need to absorb is set to increase further given the fiscal outlook in many important jurisdictions, while dealer intermediation remains constrained.

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Supplementary Internet Appendix

to accompany the paper

The Demand for Government Debt

(Egemen Eren, Andreas Schrimpf, Fan Dora Xia)

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

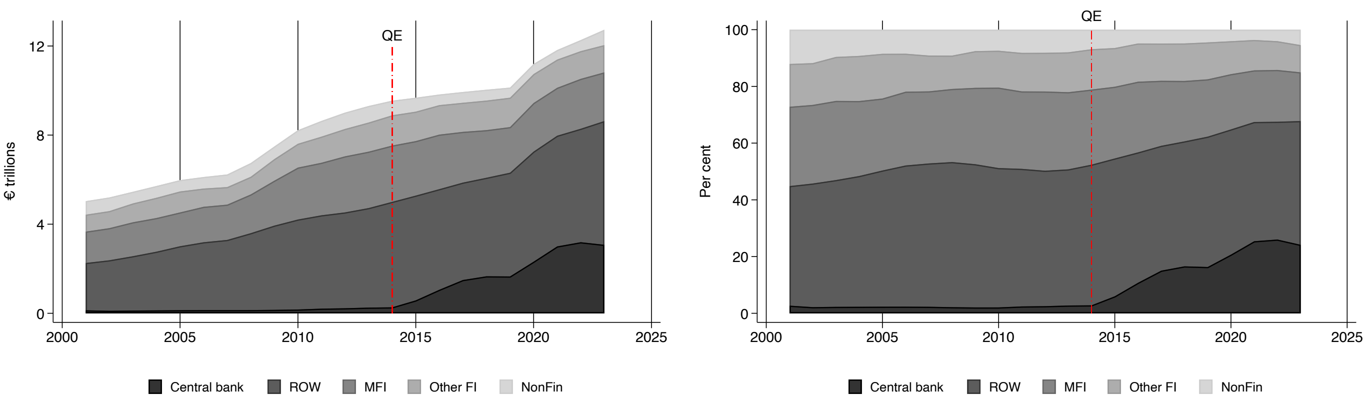
In this section, we repeat the same analysis as in Section 2 for the Euro area, Japan and the United Kingdom.

IA.I1 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 IA.1 Panel (a). The average share of the central bank in the holdings of government bonds has increased from 2.4% to 17.9%. The share of monetary financial institutions, which is comprised to a large extent of commercial banks, has fallen from 26.9% to 20.9%. The share of non-financial corporations has also dropped from 8.9% to 4.7%. The share of foreign investors declined slightly from 47.8% to 44.4%. The share of other financial institutions (which include mutual funds, pension funds, insurance companies) has declined from 14% to 12.1%.

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 IA.1 Panel (b). The central bank has absorbed 53% of every unit of new government debt since the beginning of the ECB/Eurosystem's public sector asset purchases program in March 2015. This is in contrast with the estimate of -13% prior to QE (insignificant and at a very low base as the central bank holdings of government

Figure IA.1: Total holdings and shares of different sectors - Euro Area



(a) Total holdings

(b) Shares

Note: Panels IA.1(a) and IA.1(b) show the total market value of the government debt holdings and market shares of each sector in the Euro area, respectively, between 2001 and 2023 (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.*

debt averaged only 2.7% prior to the quantitative easing program), likely owing to the fact that the Eurosystem had been shedding bond holdings at a time when total government debt supply increased pre-GFC.

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 44% for foreign investors, 47% for monetary financial institutions, 20% for non-monetary financial institutions, and 2% for non-financial corporates (insignificant). Following the launch of the Eurosystem's public sector bond purchases, these numbers have dropped to 24% for monetary financial institutions, to 17% for foreign investors and to 2% for non-monetary financial institutions (insignificant) and 4% for the non-financial sector.

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 et al. (2021) in a study of portfolio rebalancing during initial phases of the ECB's PSPP program. During that time period,

Table IA.1: 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	2.4	47.8	14	26.9	8.9
Post-QE	17.9	44.4	12.1	20.9	4.7

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	-0.13	0.44***	0.20***	0.47***	0.02
	(0.13)	(0.06)	(0.03)	(0.02)	(0.03)
Post-QE * Pct. Ch. Gov. Debt	0.53***	0.17**	0.02	0.24***	0.04*
	(0.16)	(0.08)	(0.04)	(0.01)	(0.02)
Observations	22	22	22	22	22
R-squared	0.53	0.49	0.48	0.81	0.01

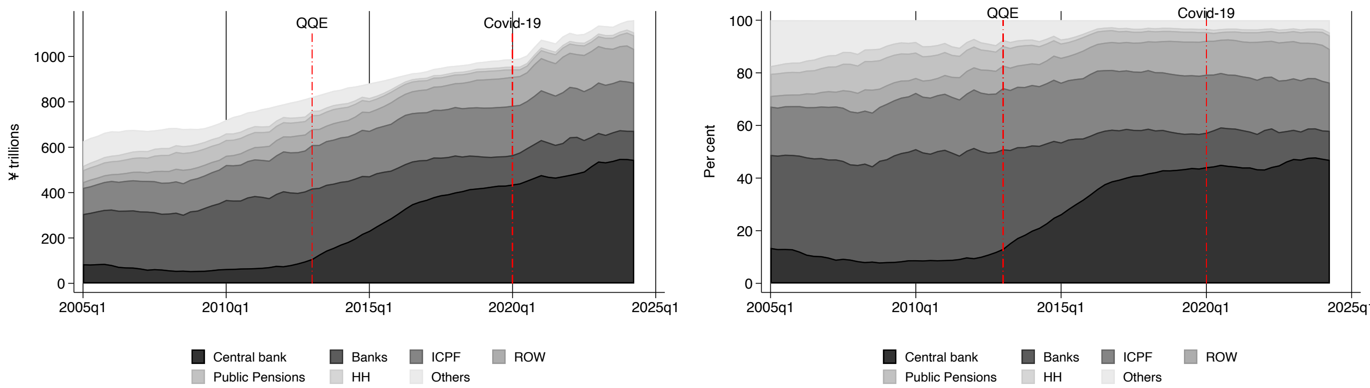
Note: Panel (a) shows the average share of each investor group over different time periods. Panel (b) reports the coefficients of the OLS regression of Equation (2) for the Euro area. CB refers to the holdings of the ECB/Eurosystem. ROW refers to foreign investors (official and private). MFI refers to 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 2001 and 2014. Post-QE is between 2005 and 2023. 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*

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.

IA.I2 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 45% since the Covid-19 crisis (Table IA.2, panel (a)). Along with the central bank, holdings of Japanese government bonds by foreign investors also increased from 7% prior to the qualitative and quantitative easing program to

Figure IA.2: Total holdings and shares of different sectors - Japan



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

Note: Panels IA.2(a) and IA.2(b) show the total market value of the government debt holdings and market shares of each sector in Japan, respectively, between 2005Q1 and 2024Q2 (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.*

14% recently.[34] 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 13%, the share of public pensions declined from 9% to 4%, 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 program until the Covid-19 crisis. During this period, the central bank bought 1.79 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 0.92 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

[34]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 (see, for example, Rime et al., 2022)

Table IA.2: 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	45.4	12.8	20	13.5	3.8	1.1	3.4

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	0.11	0.48***	0.26***	-0.05	0.01	0.03	0.16
	(0.12)	(0.07)	(0.09)	(0.05)	(0.04)	(0.07)	(0.14)
QQE * Pct. Ch. Gov. Debt	1.79***	-0.92***	0.14	-0.21***	-0.06**	0.15	0.11
	(0.22)	(0.19)	(0.14)	(0.07)	(0.03)	(0.13)	(0.25)
Post-Covid * Pct. Ch. Gov. Debt	0.41***	0.18***	0.14***	0.04**	0.00	0.16***	0.07
	(0.10)	(0.06)	(0.04)	(0.02)	(0.01)	(0.05)	(0.06)
Observations	77	77	77	77	77	77	77
R-squared	0.56	0.34	0.24	0.18	0.07	0.05	0.02

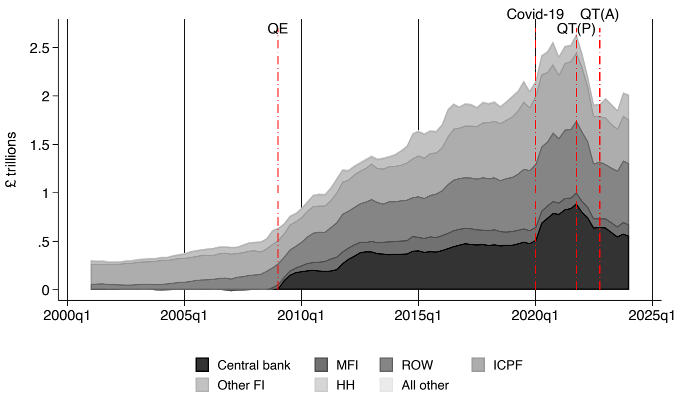
Note: Panel (a) shows the average share of each investor group over different time periods. Panel (b) reports the coefficients of the OLS regression of Equation 2 for Japan. Tha sample runs between 2005Q1-2024Q2. CB refers to the holdings of the Bank of Japan. Banks refer to commercial banks. ICPF refers 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*

crisis, the marginal response by each sector has become more balanced even as the marginal role of the central bank still remains the highest.

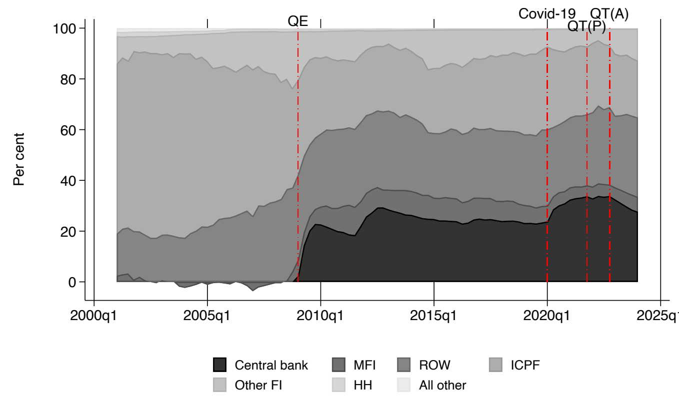
IA.I3 United Kingdom

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 IA.3 Panel (a)). Similar to Japan, the share of foreign investors has also picked up from 24% to 30%. The share of insurance companies and pension funds dropped from around 62% to 24%. The share of banks increased to 5% and other financial institutions declined from 12% to 9%.

Figure IA.3: Total holdings and shares of different sectors - United Kingdom



(a) Total holdings



(b) Shares

Note: Panels IA.3(a) and IA.3(b) show the total market value of the government debt holdings and market shares of each sector in the United Kingdom, respectively, between 2001Q1 and 2024Q2 (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 in 2009Q1, Covid-19 is at 2020Q1, and QT starts in 2022Q1. *Source: Office for National Statistics.*

The marginal response by each sector, reported in Table IA.3 Panel (b), also shows that 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 to 39%. After March 2020, this response has further increased to 53%. It is, however, interesting to note that while the central bank gained market share mostly from insurance companies and pension funds initially, the pace of absorption of this sector of newly issued government bonds has increased compared to the pre-GFC period.

Table IA.3: 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-GFC	0	24.1	62.4	-.2	1.3	11.6	.8
Post-GFC	23.4	28.7	28.2	8.3	.5	10.7	.2
Post-Covid	30.3	27.9	28.9	5	.2	7.6	.1
Post-QT	31.5	30.1	24.4	5.2	.2	8.5	.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-GFC * Pct. Ch. Gov. Debt	-0.00	0.20***	0.13***	0.20***	0.00	0.45***	0.00
	(0.08)	(0.03)	(0.04)	(0.03)	(0.00)	(0.05)	(0.00)
Post-GFC * Pct. Ch. Gov. Debt	0.39***	0.17***	0.22***	0.10***	0.00***	0.11	0.00
	(0.07)	(0.04)	(0.04)	(0.03)	(0.00)	(0.11)	(0.00)
Post-Covid * Pct. Ch. Gov. Debt	0.53***	0.12***	0.20***	-0.01	0.00***	0.14***	0.00
	(0.11)	(0.03)	(0.03)	(0.01)	(0.00)	(0.02)	(0.00)
Post-QT * Pct. Ch. Gov. Debt	0.31***	0.28***	0.27***	0.03***	0.00***	0.12***	-0.00
	(0.08)	(0.02)	(0.03)	(0.01)	(0.00)	(0.04)	(0.00)
Observations	92	92	92	92	92	92	92
R-squared	0.51	0.42	0.44	0.30	0.12	0.35	0.01

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. The sample runs between 2001Q1-2024Q2. 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. Post-GFC starts in 2009Q1 and Covid-19 is in 2020Q1. QT starts in 2022Q1. 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*

IA.II Estimation of elasticities of various sectors in other countries

In this section, we repeat our analysis for US Treasuries for Japanese and UK government bonds. Here, we do not repeat our analysis for the Euro area since we do not have quarterly data for the Euro area. In the main text, we thus resort to the results in Koijen et al. (2021) in order to provide a comparison.

We relegate these results to the Internet Appendix, because of several caveats making the elasticity estimation for Japan and the United Kingdom using our methodology not as reliable as the United States. Most importantly, there are not as many monetary policy shock measures for these jurisdictions. This makes the isolation of central bank information shocks difficult. Moreover, in most cases, we are not able to construct an instrument that yields a strong first stage. We report the results in this section, but these caveats should be taken into consideration while interpreting them.

IA.II1 Japan

For Japan, we use monetary policy shocks from Kubota and Shintani (2022) (henceforth, BOJ KS) and from Kearns et al. (2022) (henceforth BOJ KSX). Both studies use a high-frequency event study approach to isolate monetary policy surprises at different parts of the yield curve. Kubota and Shintani (2022) report surprises for the target monetary policy (short-end) and the path (the longer-end). Kearns et al. (2022) trace the high-frequency responses at 3-month, 2-year and 10-year segments of the yield curve following monetary policy announcements by the Bank of Japan. We take the 8-year zero coupon rates from Bloomberg.

We follow a similar approach as we did for the United States. We present the results of the first stage regressions with alternative monetary policy surprise instruments in Table IA.4. In the first column, we report the results using different shocks of BOJ KS individually. In the second column, we report the results using the shocks of BOJ KSX individually. In the third and fourth columns, we report the results with the first principal component of BOJ KS and BOJ KSX, respectively.

In the last column, we use the first principal component of all the shocks. All cases lead to a weak instruments problem.

While the caveats discussed above are important in interpreting results, we report the second stage results in Table IA.5 to facilitate some tentative comparison to our results for the United States. In order to make the sample period comparable to the one we reported for the United States, we report the second stage results using the PCA 1 BOJ (KSX). For Japan, we find that, the "Others" sector in the Flow of Funds, which comprises sectors other than the central bank, banks, insurance companies and private and public pension funds, and households, has the highest elasticity - though it is a rather small sector. Households also exhibit a high elasticity. Public pension funds and banks are also elastic investors. We find a higher point estimate for banks in Japan compared to those in the United States. We also find a somewhat higher point estimate for foreign investors (official and private combined) compared to the United States, it is, however, statistically insignificant. We find that insurance companies and private pension funds are inelastic investors in Japanese government bond markets.

Table IA.4: First-stage results with alternative specifications of monetary policy surprises for Japan

VARIABLES	(1) JP 8Y Yield (ZC)	(2) JP 8Y Yield (ZC)	(3) JP 8Y Yield (ZC)	(4) JP 8Y Yield (ZC)	(5) JP 8Y Yield (ZC)
BOJ KS Target	-0.0003***				
	(0.0001)				
BOJ KS Path	-0.0002				
	(0.0002)				
BOJ KSX 3m		-0.9889			
		(1.2214)			
BOJ KSX 2y		1.0840			
		(2.0802)			
BOJ KSX 10y		-1.5583*			
		(0.9186)			
PCA 1 BOJ (KS)			0.0008*		
			(0.0005)		
PCA 1 BOJ (KSX)				-0.0003*	
				(0.0002)	
PCA 1 BOJ (KS, KSX)					-0.0002***
					(0.0001)
Observations	60	77	60	77	60
R-squared	0.9471	0.9241	0.9454	0.9229	0.9451
Trend	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes
Sample	2005q2-2020q1	2005q2-2024q2	2005q2-2020q1	2005q2-2020q1	2005q2-2024q2
Effective F-stat	2.86	1.48	2.47	4.21	7.49
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	13.06	16.83	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation 5 estimated for Japan. The sample period varies depending on the availability of data across different monetary policy surprises and based on the availability of Flow of Funds data. All reported right-hand side variables are normalized. Controls include log GDP, GDP growth, inflation, log broad JPY index, log VIX, the quarterly return on the Nikkei stock index and the 8-year zero coupon US Treasury yield. 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 IA.5: Yield elasticity of demand across different sectors in Japan

VARIABLES	(1) log(ROW)	(2) log(Banks)	(3) log(HH)	(4) log(ICPF)	(5) log(PP)	(6) log(Others)
JP 8Y Yield (ZC)	11.92	47.43**	74.35***	-1.70	61.87*	130.35***
	(28.79)	(21.95)	(15.36)	(11.53)	(35.18)	(43.21)
Observations	77	77	77	77	77	77
Trend	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Anderson-Rubin Wald test (p-val)	0.66	0.15	0.03	0.88	0.25	0.00
Underidentification LM stat (p-val)	0.12	0.12	0.12	0.12	0.12	0.11

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 BOJ (KSX) as an instrument for yields. The sample period is between 2005q2 and 2024q2. Controls include log GDP, GDP growth, inflation, log broad JPY index, log VIX, the quarterly return on the Nikkei stock index and the 8-year zero coupon US Treasury yield. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.II2 United Kingdom

For the United Kingdom, we use monetary policy shocks from Braun et al. (2025) (henceforth, BOE BMAS) and from Kearns et al. (2022) (henceforth BOE KSX). Both studies use a high-frequency event study approach to isolate monetary policy surprises at different parts of the yield curve. Braun et al. (2025) report surprises for the target monetary policy (short-end), the path (the middle) and quantitative easing (the longer-end). Kearns et al. (2022) trace the high-frequency responses at 3-month, 2-year and 10-year segments of the yield curve following monetary policy announcements by the Bank of Japan. We take the 8-year zero coupon rates from Bloomberg.

We follow a similar approach as in the rest of the paper. We present the results of the first stage regressions with alternative monetary policy surprise instruments in Table IA.6. In the first column, we report the results using different shocks of BOE BMAS individually. In the second column, we report the results using the shocks of BOE KSX individually. In the third and fourth columns, we report the results with the first principal component of BOE BMAS and BOE KSX, respectively. In the fifth column, we use the first principal component of all the shocks. All cases lead to a weak instruments problem. In this case, however, we are able to resolve the weak instruments problem by using the QE surprises identified by Braun et al. (2025) (i.e. BOE BMAS QE). However, the problems due to the inclusion of information effects remain, and hence again the results should be interpreted with caution.

Keeping these caveats in mind, we report the second stage results in Table IA.7 to facilitate some tentative comparison to our results for the United States. We use the BOE BMAS QE as the instrument in the second stage. For the United Kingdom, we find a statistically significant downward-sloping demand function for foreign investors. We find that banks (classified as Monetary Financial Institutions - MFI) and households exhibit similar elasticity to foreigners. However, the estimate is statistically insignificant for these sectors. An interesting feature of UK government bond markets is we estimate upward sloping demand curves for the insurance company and pension fund (ICPF) sectors as well as Other Financial Institutions (OFI) which include investment funds. This result is similar to the finding of upward sloping demand curves for the ICPF sector in the Euro area (Koijen et al., 2021).

Upward-sloping demand curves in the context of financial assets can be due to several factors and can provide amplification of shocks in financial markets. ICPFs typically have long-duration liabilities and have to match them with long-duration, low-risk assets, such as government bonds (e.g. Domanski et al., 2017). Similarly, OFIs could receive inflows as other investors sell their government bonds to the central bank and they might invest these inflows further into government bond (e.g. Fang and Xiao, 2025). This would create a positive feedback loop pushing yields even further during QE periods and create feedback loops amplifying downward price pressures during QT periods. Such amplification dynamics could have played a role during the UK gilt crisis in 2022.

Table IA.6: First-stage results with alternative specifications of monetary policy surprises for the United Kingdom

VARIABLES	(1) UK 8Y Yield (ZC)	(2) UK 8Y Yield (ZC)	(3) UK 8Y Yield (ZC)	(4) UK 8Y Yield (ZC)	(5) UK 8Y Yield (ZC)	(6) UK 8Y Yield (ZC)
BOE BMAS Path	-0.0001					
	(0.0001)					
BOE BMAS Target	0.0001					
	(0.0001)					
BOE BMAS QE	0.0002***					0.0002***
	(0.0001)					(0.0000)
BOE KSX 3m		1.5144				
		(1.0812)				
BOE KSX 2y		-0.9297***				
		(0.3099)				
BOE KSX 10y		1.8578***				
		(0.7211)				
PCA 1 BOE (BMAS)			-0.0008**			
			(0.0004)			
PCA 1 BOJ (KSX)				0.0001		
				(0.0005)		
PCA 1 BOE (BMAS, KSX)					-0.0004	
					(0.0006)	
Observations	79	79	79	79	79	79
R-squared	0.9312	0.9287	0.9265	0.9240	0.9245	0.9290
Trend	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Sample	2004q3-2024q1	2004q3-2024q1	2004q3-2024q1	2004q3-2024q1	2004q3-2024q1	2004q3-2024q1
Effective F-stat	3.11	2.07	4.89	0.06	0.37	31.94
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	18.60	19.99	23.11	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation 5 estimated for the United Kingdom. The sample period varies depending on the availability of data across different monetary policy surprises and based on the availability of Flow of Funds data. 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 IA.7: Yield elasticity of demand across different sectors in the United Kingdom

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	log(ROW)	log(MFI)	log(OFI)	log(HH)	log(ICPF)	log(Others)
UK 8Y Yield (ZC)	18.51**	18.28	-46.52*	15.23	-15.64***	0.71
	(8.55)	(30.27)	(25.87)	(9.80)	(5.62)	(24.82)
Observations	79	63	79	79	79	79
Trend	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Anderson-Rubin Wald test (p-val)	0.03	0.53	0.09	0.14	0.01	0.98
Underidentification LM stat (p-val)	0.08	0.24	0.08	0.02	0.08	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using BOE BMAS QE as an instrument for yields. The sample period is between 2004q3 and 2024q1. Controls include log GDP, GDP growth, inflation, log broad GBP index, log VIX, the quarterly return on the FTSE 100 stock index and the 8-year zero coupon US Treasury yield. 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.

IA.III Additional tables and figures

IA.III1 Results with all controls

In this section, we report the full estimation results of the first stage and the second stages of Equation (5), where Table IA.8 reports the first-stage and Table IA.10 reports the second stage, and Equation (6), where Table IA.9 reports the first-stage and Table IA.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 IA.8: First-stage results with alternative specifications of monetary policy surprises - Full results 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	0.0012***						
	(0.0003)						
Swanson FG	0.0006**						
	(0.0003)						
Swanson LSAP	-0.0002						
	(0.0002)						
BRW		0.0005**					
		(0.0002)					
KSX (3M)			0.0009***				
			(0.0003)				
KSX (2Y)			0.0001				
			(0.0002)				
KSX (10Y)			0.0006**				
			(0.0002)				
JK MP				0.0012***			
				(0.0003)			
Nakamura-Steinsson					0.0012***		
					(0.0002)		
PCA 1 (JK MP, BRW, KSX)						0.0011***	
						(0.0001)	
PCA 1 (JK MP, BRW, KSX, NS)							0.0012***
							(0.0001)
log GDP	-0.0340	0.0353	0.0498	0.0431	0.0407	0.0435	0.0412
	(0.0485)	(0.0536)	(0.0488)	(0.0516)	(0.0481)	(0.0509)	(0.0489)
GDP growth	0.1676***	0.0265	0.0137	0.0175	0.0194	0.0182	0.0181
	(0.0430)	(0.0484)	(0.0442)	(0.0437)	(0.0437)	(0.0450)	(0.0448)
inflation	-0.0885	-0.0404	-0.0189	-0.0697	-0.0717	-0.0384	-0.0506
	(0.0587)	(0.0691)	(0.0576)	(0.0618)	(0.0813)	(0.0575)	(0.0792)
log broad dollar index	0.0404**	0.0462***	0.0491***	0.0478***	0.0507***	0.0487***	0.0513***
	(0.0177)	(0.0091)	(0.0090)	(0.0102)	(0.0101)	(0.0096)	(0.0109)
log VIX	-0.0027*	-0.0069***	-0.0056***	-0.0066***	-0.0069***	-0.0062***	-0.0065***
	(0.0016)	(0.0008)	(0.0006)	(0.0006)	(0.0008)	(0.0006)	(0.0007)
5Y German yield (zc)	0.8607***	0.7374***	0.7071***	0.7021***	0.7932***	0.7215***	0.7839***
	(0.1241)	(0.1259)	(0.1210)	(0.1289)	(0.1190)	(0.1237)	(0.1205)
trend	-0.0008	0.0019**	0.0019**	0.0019**	0.0023***	0.0020***	0.0023**
	(0.0015)	(0.0008)	(0.0008)	(0.0008)	(0.0009)	(0.0008)	(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	60	80	80	80	73	80	73
R-squared	0.9207	0.9038	0.9083	0.9078	0.9030	0.9085	0.9045
Sample	1994q2-2019q2	1994q2-2024q2	2004q3-2024q2	1994q2-2024q2	2004q3-2022q3	2004q3-2024q2	2004q3-2022q3
Effective F-stat	4.84	4.46	8.33	13.71	45.60	47.01	68.65
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	15.15	23.11	18.68	23.11	23.11	23.11	23.11

Note: This table reports the coefficients of the first-stage regression for the second-stage Equation 5. 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 IA.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	0.0022***						
	(0.0006)						
Swanson FG	0.0005						
	(0.0005)						
Swanson LSAP	-0.0008						
	(0.0005)						
BRW		0.0001					
		(0.0003)					
KSX (3M)			0.0016***				
			(0.0004)				
KSX (2Y)			-0.0003				
			(0.0004)				
KSX (10Y)			0.0010**				
			(0.0004)				
JK MP				0.0018***			
				(0.0005)			
Nakamura-Steinsson					0.0012**		
					(0.0005)		
PCA 1 (JK MP, BRW, KSX)						0.0013***	
						(0.0002)	
PCA 1 (JK MP, BRW, KSX, NS)							0.0014***
							(0.0004)
log GDP	0.1533**	0.2092***	0.2201***	0.2111***	0.1852***	0.2158***	0.1851***
	(0.0601)	(0.0397)	(0.0313)	(0.0352)	(0.0328)	(0.0361)	(0.0329)
GDP growth	0.0437	-0.0374	-0.0718***	-0.0492	-0.0286	-0.0538	-0.0348
	(0.0847)	(0.0454)	(0.0278)	(0.0332)	(0.0394)	(0.0360)	(0.0382)
Inflation	-0.0008	-0.0021**	-0.0014	-0.0024**	-0.0011	-0.0020**	-0.0008
	(0.0009)	(0.0009)	(0.0009)	(0.0010)	(0.0008)	(0.0009)	(0.0007)
log broad dollar index	0.0165	0.0226**	0.0266***	0.0267***	0.0326**	0.0252***	0.0333***
	(0.0174)	(0.0106)	(0.0090)	(0.0095)	(0.0135)	(0.0095)	(0.0123)
trend	-0.0042*	-0.0026**	-0.0024***	-0.0023**	-0.0013	-0.0024**	-0.0013
	(0.0022)	(0.0010)	(0.0008)	(0.0010)	(0.0009)	(0.0009)	(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	60	80	80	80	73	80	73
R-squared	0.8010	0.7909	0.8132	0.8045	0.7913	0.8026	0.7975
Sample	1994q2-2019q2	1994q2-2024q2	2004q3-2024q2	1994q2-2024q2	2004q3-2022q3	2004q3-2024q2	2004q3-2022q3
Effective F-stat	2.65	0.17	7.89	12.07	8.05	25.62	18.58
Crt. Val. $\alpha = 5\%$ and $\tau = 10\%$	16.85	23.11	16.04	23.11	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 IA.10: Yield elasticity of demand across different sectors in the United States: Second stage results - Panel (a) of Table 4 - with control variables

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	1.53	19.31*	22.04***	32.76***	10.85*	12.42***	1.51
	(7.01)	(9.93)	(6.73)	(11.38)	(5.79)	(3.76)	(9.40)
log GDP	-1.31*	-0.71	-1.80***	-4.41***	-0.51	-2.09***	-0.07
	(0.76)	(0.90)	(0.49)	(1.06)	(0.78)	(0.43)	(0.79)
GDP growth	0.38	-0.79	0.73*	2.04**	-0.22	0.38	0.49
	(0.44)	(0.83)	(0.42)	(0.83)	(0.52)	(0.44)	(0.66)
inflation	0.18	2.47**	3.45***	8.06***	1.94	0.44	4.22
	(0.94)	(1.22)	(1.27)	(2.48)	(1.57)	(1.58)	(3.35)
log broad dollar index	-1.27***	-1.21***	-1.21***	-1.84**	-1.19***	-0.94***	-1.18*
	(0.31)	(0.46)	(0.36)	(0.72)	(0.30)	(0.32)	(0.71)
log VIX	-0.04	0.07	0.06	0.14	0.01	-0.04	-0.01
	(0.06)	(0.08)	(0.06)	(0.13)	(0.07)	(0.03)	(0.07)
5Y German yield (zc)	-2.17	-16.75**	-20.92***	-26.54***	-14.63***	-11.24***	5.65
	(4.50)	(7.40)	(4.25)	(7.69)	(4.31)	(3.16)	(7.69)
trend	0.15***	-0.01	0.03**	-0.12***	-0.08***	-0.05***	-0.04
	(0.02)	(0.03)	(0.02)	(0.04)	(0.02)	(0.01)	(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)	(0.00)	(0.00)
Observations	80	80	80	80	80	80	80
Anderson-Rubin Wald test (p-val)	0.83	0.03	0.00	0.00	0.05	0.00	0.87
Underidentification LM stat (p-val)	0.06	0.06	0.06	0.06	0.06	0.06	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JP MP, BRW, KSX) as an instrument for yields. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table IA.11: Yield elasticity of demand across different sectors in the United States: Second stage results - Panel (b) of Table 4 - with control variables

	(1)	(2)	(3)	(4)	(5)
VARIABLES	log(IF)-log(DebtOA$_{IF}$)	log(Banks)-log(DebtOA$_{Banks}$)	log(PF)-log(DebtOA$_{PF}$)	log(IC)-log(DebtOA$_{IC}$)	log(SLG)-log(DebtOA$_{SLG}$)
8Y Yield (ZC)	20.09***	33.34***	12.70***	12.72***	6.10
	(5.45)	(6.13)	(4.16)	(3.92)	(7.87)
log GDP	-4.67***	-10.24***	-3.54***	-3.32***	-0.54
	(1.15)	(1.78)	(1.02)	(1.20)	(1.62)
GDP growth	1.88***	4.94***	1.28***	1.30***	0.37
	(0.55)	(1.50)	(0.43)	(0.47)	(0.38)
Inflation	0.07***	0.08**	0.02	0.01	0.06
	(0.03)	(0.04)	(0.02)	(0.01)	(0.05)
log broad dollar index	0.55**	-0.47	-0.64	-0.19	-1.40**
	(0.24)	(0.54)	(0.43)	(0.27)	(0.58)
trend	0.13***	0.07	0.00	0.04***	-0.07*
	(0.02)	(0.05)	(0.02)	(0.01)	(0.04)
trend squared	-0.00***	0.00	0.00*	0.00	0.00**
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	80	80	80	80	80
Controls	X	X	X	X	X
Anderson-Rubin Wald test (p-val)	0.000	0.000	0.001	0.001	0.441
Underidentification LM stat (p-val)	0.090	0.090	0.090	0.090	0.090

Note: This table reports the coefficients of the second-stage regression specified in Equation (6) using PCA 1 (JP MP, BRW, KSX) as an instrument for yields. The sample period is between 2004q3 and 2024q2. 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.

IA.III2 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 IA.12 are similar to those in the baseline regression.

In Table IA.13, 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.

In Table IA.14, we lag all control variables used in the regression. The ordering of the sectors remains similar, though the point estimates somewhat decrease.

Table IA.12: Second stage: Broader set of controls

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	3.48	25.47***	20.87***	25.11**	9.43	14.77***	-1.47
	(5.59)	(9.06)	(8.03)	(9.88)	(6.99)	(4.37)	(9.49)
log GDP	-0.35	-0.55	-0.90	-0.37	-1.09*	-1.85***	2.49***
	(0.46)	(1.07)	(0.79)	(1.11)	(0.58)	(0.51)	(0.66)
GDP growth	0.31	-0.45	0.44	0.05	-0.02	0.61**	-0.90***
	(0.29)	(0.56)	(0.42)	(0.60)	(0.27)	(0.27)	(0.34)
Inflation	-0.01	0.02	0.03	0.05**	0.03*	0.00	0.02
	(0.01)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.02)
log broad dollar index	-0.57**	-0.90	-0.66	0.64	-1.63***	-0.72	0.48
	(0.29)	(1.01)	(0.68)	(0.77)	(0.32)	(0.47)	(0.36)
log UST (fv)	0.84***	0.81*	0.42**	1.29***	-0.43***	0.51***	0.88***
	(0.07)	(0.49)	(0.21)	(0.39)	(0.08)	(0.12)	(0.26)
log VIX	-0.02	0.05	0.09*	0.31***	-0.00	-0.05	0.09**
	(0.04)	(0.06)	(0.06)	(0.07)	(0.08)	(0.03)	(0.05)
ICE BofA OAS AAA	0.64	6.56	-1.53	2.70	-4.74*	-1.09	9.63
	(1.87)	(7.87)	(3.75)	(5.69)	(2.49)	(2.71)	(8.28)
5Y German yield (zc)	-2.78	-19.89***	-20.01***	-22.03***	-13.86***	-12.14***	7.14
	(4.53)	(6.55)	(6.77)	(7.74)	(4.88)	(3.50)	(6.95)
SP500 dividend yield	3.37	7.70	-3.73	-53.92***	9.04	12.38	-48.21***
	(2.76)	(20.64)	(11.16)	(12.56)	(7.06)	(8.43)	(15.28)
trend	0.12***	-0.06*	0.03	-0.02	-0.08***	-0.09***	0.05*
	(0.01)	(0.04)	(0.04)	(0.05)	(0.03)	(0.03)	(0.03)
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)
Observations	80	80	80	80	80	80	80
Anderson-Rubin Wald test (p-val)	0.51	0.01	0.00	0.00	0.15	0.00	0.88
Underidentification LM stat (p-val)	0.09	0.09	0.01	0.02	0.09	0.09	0.09

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) 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 2024q2. 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.

Table IA.13: Second stage: Linear trend

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	-10.10	23.19*	24.33***	53.40***	21.27***	20.68***	7.64
	(10.08)	(13.25)	(7.19)	(13.21)	(6.57)	(6.04)	(12.43)
log GDP	-1.61	-0.61	-1.75***	-3.89*	-0.25	-1.88**	0.09
	(1.44)	(1.03)	(0.59)	(2.07)	(1.44)	(0.83)	(1.00)
GDP growth	-0.15	-0.62	0.84**	2.98**	0.26	0.76	0.77
	(0.78)	(0.97)	(0.43)	(1.20)	(0.95)	(0.65)	(0.47)
inflation	-2.51	3.37***	3.98***	12.83***	4.35***	2.35*	5.63
	(2.20)	(1.20)	(1.31)	(3.85)	(1.55)	(1.41)	(4.73)
log broad dollar index	-1.85***	-1.01***	-1.10***	-0.80	-0.66**	-0.52	-0.88**
	(0.43)	(0.30)	(0.30)	(0.60)	(0.33)	(0.37)	(0.44)
log VIX	-0.13	0.10	0.08	0.29**	0.09	0.02	0.04
	(0.08)	(0.10)	(0.07)	(0.14)	(0.08)	(0.05)	(0.10)
5Y German yield (zc)	0.72	-17.72**	-21.49***	-31.67***	-17.22***	-13.30**	4.12
	(7.67)	(8.80)	(4.22)	(9.18)	(3.77)	(5.75)	(7.46)
trend	0.04**	0.03***	0.05***	0.08***	0.02	0.03***	0.02**
	(0.02)	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)
Observations	80	80	80	80	80	80	80
Anderson-Rubin Wald test (p-val)	0.31	0.06	0.00	0.00	0.00	0.00	0.54
Underidentification LM stat (p-val)	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 (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 is that we use a linear trend instead of a quadratic trend. The sample period is between 2004q3 and 2024q2. 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.

Table IA.14: Second stage: Lagged controls

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	4.73*	11.26***	12.93***	16.98***	3.72*	9.37***	8.15**
	(2.46)	(4.29)	(2.04)	(4.04)	(1.91)	(2.64)	(3.52)
Observations	79	79	79	79	79	79	79
Controls	**Lagged X and Lagged W**	**Lagged X and Lagged W**	**Lagged X and Lagged W**	**Lagged X and Lagged W**	**Lagged X and Lagged W**	**Lagged X and Lagged W**	**Lagged X and Lagged W**
Anderson-Rubin Wald test (p-val)	0.11	0.00	0.00	0.00	0.08	0.00	0.08
Underidentification LM stat (p-val)	0.05	0.06	0.06	0.06	0.06	0.06	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 is that we use lag all control variables by one quarter. The sample period is between 2004q3 and 2024q2. 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.

IA.III3 Robustness checks - second stage: using the PCA of JK MP, BRW, KSX and NS surprises as instruments

In this section, we use the PCA 1 (JK MP, BRW, KSX, NS) as an instrument for yields. The first-stage estimation using this instrument corresponds to the last column of the first-stage regressions in Table 2.

Table IA.15: Second stage: using the PCA of JK MP, BRW, KSX and NS surprises as instruments

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(IF)	log(Banks)	log(PF)	log(IC)	log(SLG)
8Y Yield (ZC)	6.32	23.43**	22.94***	31.42***	4.93**	16.60***	0.48
	(5.27)	(9.22)	(3.69)	(6.80)	(2.29)	(2.78)	(6.22)
Observations	73	73	73	73	73	73	73
Anderson-Rubin Wald test (p-val)	0.22	0.00	0.00	0.00	0.04	0.00	0.94
Underidentification LM stat (p-val)	0.09	0.09	0.09	0.09	0.09	0.09	0.09

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX, NS) as an instrument for yields. The sample period is between 2004q3 and 2022q3 (due to unavailability of the NS shocks after 2022q3). 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.III4 Robustness checks - second stage: using the residuals from the regression of the PCA 1 (JK MP, BRW, KSX) on JK Central Bank Information Shocks

In this section, we use the residuals from the regression of PCA 1 (JK MP, BRW, KSX) on JK Central Bank Information shocks as an instrument for yields.

Table IA.16: Second stage:second stage: using the residuals from the regression of the PCA 1 (JK MP, BRW, KSX) on JK Central Bank Information Shocks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(IF)	log(Banks)	log(PF)	log(IC)	log(SLG)
8Y Yield (ZC)	0.99	17.76*	21.69***	32.39***	10.91*	11.91***	0.67
	(7.24)	(9.75)	(6.94)	(11.55)	(5.88)	(3.73)	(9.51)
Observations	80	80	80	80	80	80	80
Anderson-Rubin Wald test (p-val)	0.89	0.05	0.00	0.00	0.05	0.00	0.94
Underidentification LM stat (p-val)	0.06	0.06	0.06	0.06	0.06	0.06	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using the residuals from the regression of the PCA 1 (JK MP, BRW, KSX) on JK Central Bank Information Shocks as an instrument for yields. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.III5 Robustness checks - second stage: two principal components as instruments

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

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(IF)	log(Banks)	log(PF)	log(IC)	log(SLG)
8Y Yield (ZC)	8.49*	25.36**	25.59***	34.31***	8.78*	18.51***	4.80
	(4.81)	(10.87)	(4.96)	(9.28)	(4.62)	(7.10)	(9.19)
Observations	80	80	80	80	80	80	80
Controls	**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	0.01	0.13	0.00	0.23
Underidentification LM stat (p-val)	0.16	0.16	0.14	0.16	0.16	0.03	0.16
Hansen J stat (p-val)	0.19	0.19	0.33	0.68	0.29	0.11	0.18

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) and PCA 2 (JK MP, BRW, KSX) as instruments for yields. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.III6 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 IA.18) and 10 (Table IA.19) in the holdings of sectors when we make valuation adjustments to the left-hand side variables of the second stage.

Table IA.18: Second stage: Valuation adjustments with 5 year

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	-0.21	17.69*	20.18***	30.81***	9.13*	10.65***	-0.28
	(6.40)	(9.21)	(5.96)	(10.73)	(4.92)	(3.46)	(8.89)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.97	0.03	0.00	0.00	0.05	0.00	0.97
Underidentification LM stat (p-val)	0.06	0.06	0.06	0.06	0.06	0.06	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 is that we assume a modified duration of 5 in the portfolios while making the valuation adjustments. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table IA.19: Second stage: Valuation adjustments with 10 year

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
8Y Yield (ZC)	2.28	19.96*	22.85***	33.64***	11.57*	13.16***	2.29
	(7.42)	(10.48)	(7.25)	(11.79)	(6.38)	(4.17)	(9.73)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.76	0.03	0.00	0.00	0.06	0.00	0.81
Underidentification LM stat (p-val)	0.06	0.06	0.06	0.06	0.06	0.06	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 is that we assume a modified duration of 10 in the portfolios while making the valuation adjustments. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.III7 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 IA.20) and 10-year zero-coupon yields (Table IA.21) to estimate elasticities. The results remain similar.

Table IA.20: Second stage: using 5 year yield

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
5Y Yield (ZC)	1.09	13.78*	15.73***	23.38***	7.74**	8.86***	1.08
	(4.99)	(7.22)	(4.65)	(7.50)	(3.95)	(2.57)	(6.70)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.83	0.03	0.00	0.00	0.05	0.00	0.87
Underidentification LM stat (p-val)	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 (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 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 4. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table IA.21: Second stage: using 10 year yield

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
10Y Yield (ZC)	1.87	23.61*	26.95***	40.06**	13.27*	15.18***	1.84
	(8.60)	(12.36)	(8.99)	(15.69)	(7.55)	(5.10)	(11.52)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.83	0.03	0.00	0.00	0.05	0.00	0.87
Underidentification LM stat (p-val)	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 (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 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 4. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.III8 Robustness checks - second stage: Alternative yields and valuation adjustments

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

Table IA.22: Second stage: using 5 year yield and valuation adjustments

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	log(ROW Off)	log(ROW Pri)	log(IF)	log(Banks)	log(PF)	log(IC)	log(SLG)
5Y Yield (ZC)	-0.15	12.62*	14.40***	21.99***	6.52**	7.60***	-0.20
	(4.57)	(6.66)	(3.99)	(6.95)	(3.29)	(2.29)	(6.35)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.97	0.03	0.00	0.00	0.05	0.00	0.97
Underidentification LM stat (p-val)	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 (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 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 4 and do the valuation adjustments based on a 5-year duration rather than 8-year duration. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table IA.23: Second stage: using 10 year yield and valuation adjustments

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
10Y Yield (ZC)	2.79	24.41*	27.94***	41.13**	14.14*	16.09***	2.80
	(9.12)	(12.99)	(9.52)	(16.12)	(8.23)	(5.53)	(11.93)
Observations	80	80	80	80	80	80	80
Controls	**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)	0.76	0.03	0.00	0.00	0.06	0.00	0.81
Underidentification LM stat (p-val)	0.07	0.07	0.05	0.07	0.07	0.07	0.07

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The difference between this table and the Panel (a) of Table 4 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 4 and do the valuation adjustments based on a 10-year duration rather than 8-year duration. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

IA.III9 Robustness checks: QE/QT asymmetry

In this section, we report the robustness checks for the comparison of coefficients during increasing and decreasing share of the central bank in the government bond market.

Table IA.24: Second stage results with $\Delta CB\ Share$ interactions with a post-GFC dummy

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
$8Y\ Yield\ (ZC) \times \Delta\ CB\ Share \geq 0$	4.96	15.20	28.11***	42.15***	13.08	19.65**	-4.81
	(7.36)	(10.57)	(10.69)	(13.52)	(8.77)	(8.37)	(10.33)
$8Y\ Yield\ (ZC) \times \Delta\ CB\ Share < 0$	8.28	5.00	38.97**	57.71**	18.17	31.35**	-19.25
	(10.09)	(22.30)	(19.11)	(24.79)	(15.95)	(15.02)	(17.66)
$\Delta\ CB\ Share < 0$	-0.10	0.25	-0.29	-0.47	-0.17	-0.34*	0.37
	(0.10)	(0.39)	(0.29)	(0.44)	(0.20)	(0.21)	(0.36)
Observations	80	80	80	80	80	80	80
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Equal coefficients (p-val)	0.35	0.43	0.34	0.34	0.53	0.20	0.31
Anderson-Rubin Wald test (p-val)	0.64	0.03	0.00	0.00	0.13	0.00	0.35
Underidentification LM stat (p-val)	0.14	0.14	0.15	0.14	0.14	0.13	0.14

Note: This table reports the coefficients of the second-stage regression specified in Equation (7) in using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The sample period is between 2004q3 and 2024q2. The $\Delta CB\ Share \geq 0$ and $\Delta CB\ Share < 0$ are dummy variables which are 1 if the central bank share in the government bond market increased or decreased from the previous quarter, respectively. Controls also include a post-GFC dummy, which is one after 2008q4. Equal coefficients row reports the p-values of the hypothesis test whether the coefficients for the interaction terms are equal. 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.

Table IA.25: Second stage results with $\Delta CB\ Share$ interactions excluding QE1

VARIABLES	(1) log(ROW Off)	(2) log(ROW Pri)	(3) log(IF)	(4) log(Banks)	(5) log(PF)	(6) log(IC)	(7) log(SLG)
$8Y\ Yield\ (ZC) \times \Delta\ CB\ Share \geq 0$	7.92	19.15	28.33***	32.55**	17.98**	25.25	-7.56
	(9.83)	(12.12)	(9.98)	(13.55)	(7.62)	(17.17)	(12.47)
$8Y\ Yield\ (ZC) \times \Delta\ CB\ Share < 0$	12.74	18.54	39.33**	34.48*	28.35**	41.80	-21.95
	(15.15)	(18.35)	(15.90)	(19.67)	(13.18)	(25.66)	(18.44)
$\Delta\ CB\ Share < 0$	-0.12	-0.01	-0.27	-0.08	-0.31*	-0.45	0.36
	(0.17)	(0.26)	(0.23)	(0.26)	(0.17)	(0.30)	(0.36)
Observations	75	75	75	75	75	75	75
Controls	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**	**X and W**
Equal coefficients (p-val)	0.45	0.94	0.30	0.83	0.19	0.19	0.32
Anderson-Rubin Wald test (p-val)	0.62	0.10	0.00	0.01	0.00	0.00	0.15
Underidentification LM stat (p-val)	0.12	0.12	0.12	0.12	0.12	0.14	0.12

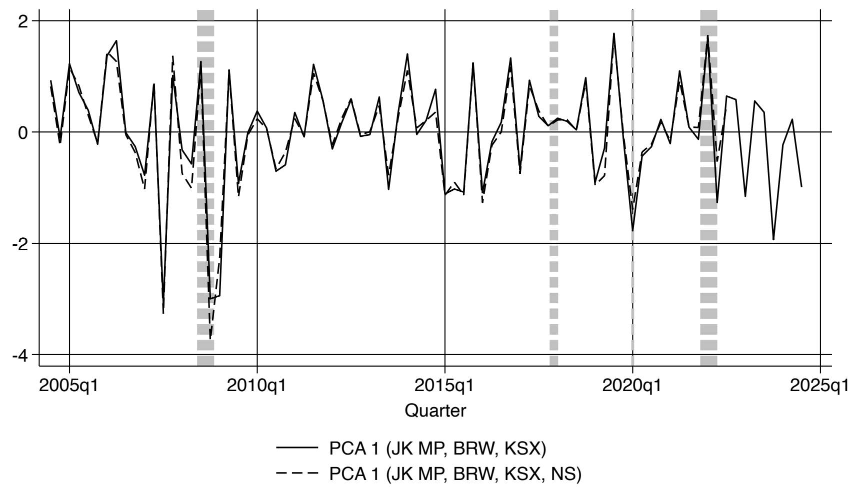
Note: This table reports the coefficients of the second-stage regression specified in Equation (7) in using PCA 1 (JK MP, BRW, KSX) as an instrument for yields. The sample period is between 2004q3 and 2024q2, but it excludes the QE1 period (i.e. between March 2009 and March 2010) since the Federal Reserve's first Treasury purchases were in 2009q1. The $\Delta CB\ Share \geq 0$ and $\Delta CB\ Share < 0$ are dummy variables which are 1 if the central bank share in the government bond market increased or decreased from the previous quarter, respectively. Equal coefficients row reports the p-values of the hypothesis test whether the coefficients for the interaction terms are equal. 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.

IA.IV Principal components analysis

In this section, we report further details of the principal components analysis. In Figure IA.4, we report the time series of the first principal components of different constructs used in the analysis, i.e. PCA 1 (JK MP, BRW, KSX) for the baseline and also PCA 1 (JK MP, BRW, KSX, NS). We standardize both of them in the analysis to have a mean zero and a standard deviation of one. As we show in Table IA.26, these two measures are indeed highly correlated with each other. In Table IA.27, we report the eigenvalues and the explained percentages of all of the series used in our baseline measure PCA 1 (JK MP, BRW, KSX). In our baseline case, the first principal component explains 46% of the variation and the second principal component explains another 23% of the variation. Finally, in Table IA.28, we report the eigenvectors. The PCA 1 (JK MP, BRW, KSX), which we use in our baseline regressions, has a high loading on monetary policy shocks that primarily measure the responses at the longer-end of the yield curve such as the shocks from

Bu et al. (2021) and the 10-year shocks from Kearns et al. (2022). This is reassuring that our measure does capture shocks about the conduct of balance sheet policies, which alter the supply of government bonds available.

Figure IA.4: US monetary policy surprises - PCA 1 (JK MP, BRW, KSX) and PCA 1 (JK MP, BRW, KSX, NS)



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 the monetary policy shock of Jarociński and Karadi (2020), Bu, Rogers and Wu (2021) and Kearns, Schrimpf and Xia (2022) with the sample period between 2004Q3 and 2024Q2. The dashed line is the first principal component when we also include the Nakamura and Steinsson (2018) shocks with the sample period of 2004Q3 and 2022Q3.

Table IA.26: Cross-correlation table

Variables	PCA 1 (JK MP, BRW, KSX)	PCA 1 (JK MP, BRW, KSX, NS)
PCA 1 (JK MP, BRW, KSX)	1.00	
PCA 1 (JK MP, BRW, KSX, NS)	0.98	1.00

Table IA.27: Eigenvalues and explained percentages of PCA (JK MP, BRW, KSX)

	Eigenvalue	Difference	Proportion	Cumulative
Comp1	2.301727	1.127223	0.4603	0.4603
Comp2	1.174504	.4153932	0.2349	0.6952
Comp3	.7591112	.334702	0.1518	0.8471
Comp4	.4244091	.0841613	0.0849	0.9320
Comp5	.3402479	.	0.0680	1.0000

Table IA.28: Eigenvectors of PCA (JK MP, BRW, KSX)

	Comp1	Comp2	Comp3	Unexplained
KSX (3M)	.3490953	.6583496	-.274312	.1533147
KSX (2Y)	.3627327	.4031607	.7860286	.0372381
KSX (10Y)	.5045395	-.4091914	.1085793	.2084663
BRW	.49149	-.4705517	.053696	.1817429
JK MP	.5004348	.1232097	-.5405916	.1838951

IA.V Additional results for investment funds

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. 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.

In Figure IA.5, we report the time series of the total market value of the holdings of different types of investment funds. In Table IA.29, 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 IA.5: Total market value of the holdings of different types of investment funds

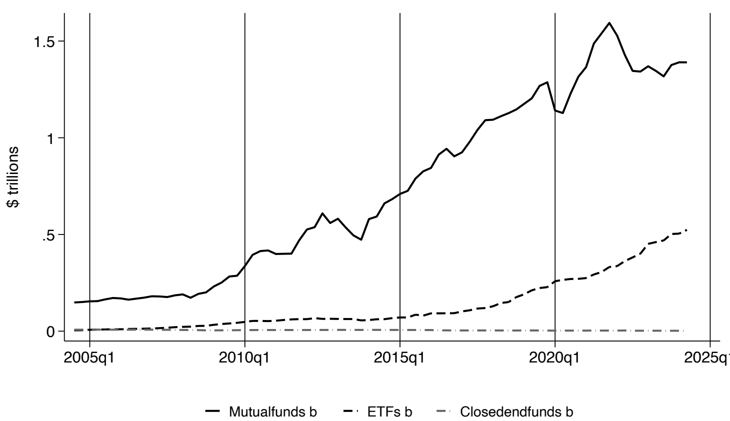


Table IA.29: Yield elasticity estimates for different types of investment funds

VARIABLES	(1) ln(Mutual funds)	(2) ln(ETFs)	(3) ln(Closed-ended funds)
8Y Yield (ZC)	26.8278***	-11.7409	17.8835*
	(8.0165)	(9.1622)	(10.3490)
Observations	80	80	80
Controls	**X and W**	**X and W**	**X and W**
Sample	2004q3-2024q2	2004q3-2024q2	2004q3-2024q2
Anderson-Rubin Wald test (p-val)	0.00	0.19	0.06
Underidentification LM stat (p-val)	0.06	0.06	0.06

Note: This table reports the coefficients of the second-stage regression specified in Equation (5) using PCA 1 (JK MP, BRW, KSX) as an instrument for yields for mutual funds, ETFs and closed-ended funds. The sample period is between 2004q3 and 2024q2. 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. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

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 et al., 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 et al., 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.

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