Essays on government bond markets and macroeconomic stabilization

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This thesis aims to identify the effects of macroeconomic stabilization efforts of the recent past, with a particular focus on government bond markets. The first two chapters investigate how Treasury markets react to changes in government debt supply and demand. The next two chapters widen the scope to examine macroeconomic stabilization in a broader context. Chapter 1 proposes an identification scheme for government bond supply shocks and analyses how the additional debt issuance affects the term structure of interest rates. It argues both theoretically and empirically that changes in the debt supply influence yields by affecting risk premia. Chapter 2 shifts its attention to the demand side of Treasury markets. It analyzes the effects of sudden shifts in investor demand for German and Italian government bonds. We aim to understand how different credit risk profiles of the countries impact these effects. Chapter 3 provides fresh insights into the absorption of idiosyncratic output shocks in the US and the Euro Area through private and public risk-sharing channels. It proposes a new empirical framework to trace how these risk-sharing channels evolved over time and attempts to explain the changes with some key macroeconomic and financial factors. Lastly, Chapter 4 examines the effectiveness of government spending during health crises. We provide evidence that during such exceptional times, higher government spending has a more potent, stimulative impact on the economy.

Essays on government bond markets and macroeconomic stabilization   Andras Lengyel

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Universiteit van Amsterdam

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Essays on government bond markets and macroeconomic stabilization

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Contents

Introduction

Developed economies from the mid-1980s experienced a remarkable decrease in macroeconomic volatility. Economists came up with several potential causes for this phenomenon, labeled as the Great Moderation (see Stock and Watson (2002) and Summers et al. (2005) among others). These causes include structural changes in the economy that make it less vulnerable to shocks, a decrease in the magnitude of the shocks hitting the economy, and most importantly, improved policy tools and frameworks to tackle the adverse effects of shocks. Convinced of this latter factor, Robert Lucas, in his 2003 presidential address to the American Economic Association, went as far as stating that the *"central problem of depression-prevention has been solved"* (Lucas Jr (2003)). The consensus among economists was that economic stability is achievable simply through the management of short-term interest rates by central banks.

The end of the 2000s brought an end to both the Great Moderation and the belief that the adjustment of short rates is a sufficient stabilization tool. The immense recession following the Global Financial Crisis of 2007-2008 and the European debt crisis urged central banks to start experimenting with additional tools, such as purchasing government debt in large scales. Subsequently, the historically low interest rate environment in the 2010s and the recession induced by the Covid-19 pandemic led economists and policymakers to re-examine the role of fiscal stimulus as a stabilization tool (see for example Blanchard (2023) and Romer (2021)).

This paradigm shift in both spheres of policymaking spurred unprecedented activity in government bond markets. Central banks became the largest buyers of Treasuries. At their respective peak, the US Federal Reserve (FED) owned 21% of the US Treasury market, the European Central Bank (ECB) owned 29% of the German and 26% of the Italian bond market, and the Bank of England (BoE) owned 38% of the outstanding stock of UK government debt.[1] Additionally, debt issuance picked up with Covid, and the outstanding

[1] See the updated database of Arslanalp and Tsuda (2014) at https://www.imf.org/~/media/Websites/IMF/imported-datasets/external/pubs/ft/wp/2012/Data/_wp12284.ashx.

stock of government debt reached historically high levels. However, supply and demand conditions in government bond markets did not only shift due to fiscal and monetary policy. The last decade has also seen some abrupt changes in the activity of the private sector. Eurozone periphery countries have seen investors' demand suddenly drying up at the peak of the Eurozone debt crisis, while the UK experienced a somewhat similar episode after the government published the mini-budget of September 2022. Finally, financial regulation after the GFC also shifted the demand of private banks, forcing them to hold more government bonds than before.

The first two chapters of this dissertation aim to make a step in the direction of understanding the effects of large shifts in the demand or supply conditions in government bond markets. Chapter 1 looks at how the additional debt issued by the government affects the term structure of interest rates. It identifies Treasury supply shocks using intraday high-frequency data, by exploiting the institutional setup of the UK government bond primary market. It finds that supply shocks have a positive effect on interest rates. The effect increases with maturity, so longer-term yields respond more. We decompose yields into expectations components and risk premia, showing that the reaction is primarily due to the risk premia component of yields. The chapter also documents that risk premia reacts stronger during market stress, and that the effects localize, i.e., yields react more to the shock around the location of the shock in the maturity space. To rationalize these findings, the chapter extends the Greenwood and Vayanos (2014) framework with inflation risk and includes two types of bonds (nominal and inflation-linked). It argues that bond supply shocks transmit via the repricing of duration and inflation risks in the economy, i.e., via increasing risk premia. Furthermore, the effects are stronger when investors' risk aversion is higher, confirming the empirical findings. These results have important implications for central banks trying to decrease the size of their balance sheet through Treasury sales, and for fiscal authorities who finance their additional spending through debt issuance.

Chapter 2, on the other hand, focuses on the demand side of public debt markets. In the chapter, we use intraday government bond futures price changes around German and Italian Treasury auctions to identify unexpected shifts in the demand for public debt. Estimates show that positive demand shocks lead to large negative movements in Treasury yields. We document significant spillover effects into Treasury bond, equity, and corporate bond markets of other Eurozone countries. We find interesting differences in the effects of demand shocks between the two countries. Main euro area stock indices drop following German demand shocks, while Treasury CDS spreads increase. On the other hand, Italian demand shocks lead to positive responses of equity prices and decreases in CDS spreads.

This is consistent with the "safe-haven" status of German bonds versus the "high-debt" status of Italian Treasuries. Increased demand for German debt is associated with "flight-to-safety" capital movements, while higher demand for Italian debt signals a better outlook for the Italian economy and its public finances. Our results also suggest that all of these effects are stronger during periods of high financial stress.

The final two chapters study macroeconomic stabilization in a broader context. Chapter 3 studies how countries use international channels to mitigate the impact of an output shock. It analyzes how these so-called risk-sharing channels have evolved over time in the United States and the Euro Area. In particular, we focus on the capital channel (income from cross-border ownership of productive assets), the credit channel (interstate or cross-country bank lending), and the fiscal channel (federal or international fiscal transfers). We offer three main contributions. First, we propose a time-varying parameter panel VAR model, with stochastic volatility, which allows us to formally quantify time variation in risk-sharing channels. Second, we develop a new test of the complementarity vs. substitutability hypothesis of the three risk-sharing channels, based on the correlation between the impulse responses of these channels to idiosyncratic output shocks. Third, for the United States, we explain time variation in the risk-sharing channels based on some key macroeconomic and financial variables. We find that the overall level of risk-sharing has significantly increased in the US and also, slightly in the EA over the past decades. This improvement is mostly due to private risk-sharing channels. However, while in the US the role of the capital channel is the most sizable, we find that in the EA the credit channel dominates. We also find strong substitution effects between capital market smoothing and credit market smoothing in both regions, suggesting substitutability between these channels. In the case of the US, we also find evidence for complementarity between the private (e.g., credit and capital) and the public (e.g., fiscal) channels. Finally, we find evidence that the fiscal channel is stronger under weak economic conditions and when there is ample fiscal space. The capital channel is also stronger in downturns, while the credit channel is improved by stronger financial integration. These results have important implications for the reform of the European Monetary Union's (EMU) governance framework. They imply that there is substantial room to strengthen the shock absorption capacity of the EMU. The Capital Market Union and the Credit Market Union could reinforce risk-sharing and improve the resilience of the EA to macroeconomic shocks by contributing to smooth output shocks through the capital channel and the credit channel. Furthermore, the positive externalities between the fiscal channel and the private risk-sharing channels imply that progress on the Fiscal Union could also have a beneficial effect on private risk-sharing mechanisms in the EMU.

Finally, Chapter 4 focuses on one specific macroeconomic stabilization policy, fiscal expansions. This research is motivated by the fiscal packages deployed to cushion the economic fallout from the COVID-19 pandemic. We take a historical look at past pandemics and epidemics and the effectiveness of public sector support in response to these health crises. The chapter assesses how fiscal multipliers could vary during and after health crises. In particular, how factors such as social distancing and uncertainty could affect contemporaneous multipliers and how near-term multipliers are affected as economies re-open. The chapter shows that cumulative fiscal multipliers one year after a health crisis are about fifty percent larger than during normal times, particularly in advanced economies. The results also suggest that the reason behind this is the fact that health crises episodes are often characterised by heightened uncertainty and suppressed demand. Fiscal expansions tend to be more effective in such environments. Our findings suggest that large-scale fiscal support deployed at the onset of the COVID-19 pandemic could have larger than usual lingering impacts on economic activity, which need to be accounted for when calibrating policies.

Chapter 1

Treasury Supply Shocks and the Term Structure of Interest Rates in the UK*

1.1 Introduction

How does the additional issuance of government debt affect the term structure of interest rates? The question is both important and topical given the rapid increases in government deficits and debt levels across the globe due to the Covid-19 pandemic. At the same time, the interest rate environment is on the rise, due to central banks' efforts to fight the inflationary pressure of the recovery. The resulting increased public debt service cost requires the active management of the debt and a good understanding of the financial market effects of debt issuance.

The effect of government debt issuance on interest rates is not well established in the empirical literature. Surveys on the effects of fiscal deficits on interest rates by Gale and Orszag (2003) and Engen and Hubbard (2004) found around the same number of papers with positive and significant effects as the number of papers with insignificant effects. The reason is that identification is difficult. While interest rates are available at every point in time, budgetary variables are only available annually or at best quarterly frequency. Reverse causality, common factors, and anticipation effects are all complicating the problem. For example, agents often anticipate and price public policies in advance, making it difficult to time and measure their true causal effects. Moreover, factors that affect both interest rates and the deficit can lead to finding a spurious relationship: while

* This chapter is based on Lengyel (2022).

the central bank cuts the policy rate in a recession, the fiscal deficit and the debt issuance increase. Our goal in this paper is to get around these issues and uncover the causal effect of debt supply on interest rates. This has a direct policy relevance. Governments have ever-increasing financing needs, which tend to pick up during crises. This constantly provides markets new supply of debt to be absorbed. The new supply directly impacts asset prices and the funding costs of the government, firms, and households. This ultimately affects economic activity. Therefore, fiscal authorities need a thorough understanding of how their funding decisions affect markets. In this paper, we intend to provide them with quantitative estimates of the impact of their new debt issuances on asset prices.

Our first contribution is to propose a novel identification of government bond supply shocks. The identification exploits the institutional features of the Debt Management Office (DMO) of the United Kingdom. We focus on the *announcements* of the supply of upcoming bond auctions. We follow intraday bond futures price movements in a narrow event window around the announcements to capture the information content of the announcements. This information content is high, as the UK DMO does not provide information about the volume of the upcoming auctions before these releases.[1] Furthermore, the announcements contain information solely on the supply side of the bond market. This institutional framework provides an ideal setting to apply the high-frequency identification scheme, that was initially proposed to identify monetary policy shocks (Kuttner (2001)). Price movements in a narrow event window around the auction announcements can be related to information about future bond supply. We interpret these price movements as shocks to the supply of government bonds. The second contribution of the paper is to study the effects of debt supply shocks on the term structure of interest rates. This adds to the empirical literature that studies the effect of government debt supply on interest rates. The fiscal policy literature tends to relate these two variables at the quarterly or annual frequency. We, on the other hand, estimate this relation at the daily frequency. Moving higher in frequency has the advantage of allowing for an (arguably) cleaner identification, but it comes at the expense of being more restricted in the range of addressable research questions. We, therefore, focus only on financial variables in this paper.

We find that debt issuance has significant positive effects on nominal interest rates: a positive standard deviation bond supply shock increases nominal rates by 1-1.5 basis points. Longer maturities respond more, so the slope of the yield curve increases. This effect spills over to equity and corporate bond markets. To give more intuition on the size of

[1] This is in contrast with other countries like the US, where the Treasury provides estimated future auction sizes every quarter.

the effect, we provide a back-of-the-envelope exercise. On the 11th of March 2020, the UK government announced a Covid-19 support package of £12bn. According to our estimates, an unexpected debt issuance announcement of this size would raise nominal yields by 13-19 basis points. These estimates are in line with actual changes in benchmark yields on the day of the announcement. Next, we study the mechanism of how supply shocks transmit to the term structure of interest rates. We find that the supply shock increases real rates almost as much as nominal rates. This implies that the main driver of the effect is not the higher inflation outlook of investors. To investigate further, we decompose yields into expectations and risk premia components with the Affine Term Structure Model (ATSM) of Abrahams et al. (2016). We find that over two-thirds of the response of long-term yields are attributed to risk premia components. Additional government debt supply raises both the real term premium and the inflation risk premium. Interestingly, expected inflation is unaffected.

We illustrate these empirical findings in an equilibrium term structure model with supply effects. The model extends the framework of Greenwood and Vayanos (2014) with exogenous inflation. It has two types of bonds: nominal and inflation-linked. The supply of nominal bonds is stochastic. Shocks to this supply are absorbed by risk averse investors, holding more inflation and interest rate risks in their equilibrium portfolio. As investors become more exposed to these risk factors, they require higher compensation to hold these risks in their portfolios. This drives up risk premia, and consequently, yields. Inflation-linked bonds are unaffected by inflation risk, so their yield is less affected by the shock compared to nominal bond yields, consistent with our empirical finding. The mechanism in the model is closely linked to investors' limited risk-bearing capacity. Supply effects are stronger when investors are more risk averse, as they require even higher compensation for a given amount of risk. We test this prediction empirically, by exploring state dependence in the effects of the high-frequency supply shock. Consistent with the model's prediction, we find that yields react stronger in times of market stress. This is driven by the higher reaction of risk premia. Furthermore, during market stress supply shocks have localized effects, i.e., yields react stronger in the maturity segment of the new debt. This is consistent with the effects of bond demand shocks in Vayanos and Vila (2021) and Droste et al. (2021). We also explore the effect of supply shocks in times when the policy rate is at the effective lower bound (ELB). When the policy rate is constrained, short-term rates respond less, while long-term rates react stronger to supply shocks. The main driver here is again rising risk premia and not higher expectations about short-rates or inflation.

The rest of the paper is organised as follows. Section 1.2 connects our paper to the literature. Section 1.3 explains our identification in two steps. Section 1.3.1 describes the institutional framework of the UK government bond primary market, while Section 1.3.2 explains how we exploit this to construct the supply shock. Section 1.4 analyses the effect of the supply shock on yields. Section 1.4.2 demonstrates that the supply shock transmits by affecting risk prices. Section 1.4.3 presents an equilibrium asset pricing model where we illustrate this effect. Section 1.5 investigates the role of non-linearities. Lastly, Section 1.6 concludes.

1.2 Related literature

We identify government bond supply shocks using the high-frequency identification (HFI) method. HFI was developed initially to study the effects of monetary policy shocks (Kuttner (2001), Gürkaynak et al. (2005)). It has recently been applied to identify oil price shocks (Känzig (2021)), carbon policy shocks (Känzig (2023)), and Treasury demand shocks (Droste et al. (2021), Lengyel and Giuliodori (2022)). The latter is the application most similar to ours. Droste et al. (2021) identifies Treasury *demand* shocks of large institutional investors by following Treasury futures prices around Treasury auction result releases. We focus, on the other hand, on the *announcements* of auctions. In this aspect, the paper by Simon (1991) is closely related to ours. Simon (1991) analyses the announcements of US government cash-management bills in an event study. The focus of both Droste et al. (2021) and Simon (1991) is on the segmentation of Treasury markets and the localized effects of supply and demand conditions. While we do find some evidence for segmentation, our main focus is on risk pricing and the transmission of government debt supply shocks. Our paper is also similar to D'Amico and Seida (2020), in the sense that our identification isolates the expected and the unexpected components of the announcements of changes in the supply of bonds to see the reaction of yields.

During the finalization of our results, we became aware of Phillot (2021). Similar to us, he proposes the identification of US Treasury supply shocks by following futures price movements around auction supply announcements. In line with our results, he finds empirically that the supply shock is followed by a positive shift in the yield curve, higher inflation compensation, rising stock prices, and corporate bond yields. The main differences compared to our paper are the following. Firstly, we follow *intraday* futures price movements in a one-hour event window around announcements, while Phillot (2021) records daily price differences between the price on the announcement day and the price on the previous day. Our narrower event window means that our shock series is less affected by potential confounding factors contaminating the results. What allows us to go higher in frequency is focusing on the United Kingdom instead of the US. In the UK the exact publication time of the announcements is known, while in the US only the date is known. This allows us to zoom in on intraday price movements in a short event window around the release time of the announcements. An additional benefit of focusing on the UK is that UK auction announcements have higher information content about bond supply compared to the US, as explained more in detail in Section 1.3.1. Apart from the identification, an important difference between our papers is that our analysis focuses more on the mechanism of how debt supply affects the term structure of interest rates. We first break down yields into their components with an empirical ATSM and study the reaction of each component separately. Then, we illustrate these empirical findings in an equilibrium asset pricing model.

Supply and demand conditions in Treasury markets have gained much attention with central banks' QE operations. The general finding is that official demand for bonds decreased the level and slope of the yield curve (Hamilton and Wu (2012), Li and Wei (2013), McLaren et al. (2014) and others). The underlying mechanism is explained by the preferred-habitat theory of interest rates (Modigliani and Sutch (1966), Vayanos and Vila (2021)). This theory argues that changes in the demand and supply conditions are transmitted through bond risk premia. Risk premia has a positive relationship with the slope of the term structure, as long-term bonds are more exposed to risks. Our study brings evidence in line with this literature, connecting bond supply with the level and slope of the yield curve, and bond risk premia. However, in contrast with the empirical literature on QE, we do not focus on changes in central bank demand for bonds, but on the supply from the Treasury. We look at episodes when the Treasury increases the outstanding stock of bonds through new issuances. Nevertheless, the mechanism we explain our findings is the mirror image of the one used to explain the effects of QE (Vayanos and Vila (2021), Greenwood and Vayanos (2014)).

We find that changes in Treasury supply transmit to yields by affecting bond risk premia. Therefore, the spending and financing decisions of the government have direct effects on risk pricing. In this regard, our paper is connected to the strand of literature that establishes a connection between measures of fiscal policy and risk pricing. Studies have found that bond risk premia is affected by the level of fiscal expenditures and the uncertainty around it (Bretscher et al. (2020), Horvath et al. (2021), Kučera et al. (2022), Bayer et al. (2020)), the government debt ratio (Alesina et al. (1992), Greenwood and

Vayanos (2014), Nguyen (2018)) and the maturity structure of the debt (Chadha et al. (2013), Greenwood and Vayanos (2014), Corhay et al. (2021)). The government debt ratio was also found to influence equity- and credit-risk premia (Gomes et al. (2013), Liu (2023)) as well as the liquidity premium (Krishnamurthy and Vissing-Jorgensen (2012), Bayer et al. (2020)). Our paper contributes to this literature, by linking Treasury debt issuance with the term premium and the inflation risk premium.

1.3 Constructing the supply shock measure

In this section, we explain our identification in two steps. First, we briefly describe the institutional framework of the UK Debt Management Office and the bond issuing process. For more details see DMO (2021). Then, we outline how we apply HFI and isolate supply shocks in this setting.

1.3.1 Description of the UK primary bond market

The DMO is the institution responsible for the UK government's debt management policy. It carries out this duty by issuing debt securities denominated in pound sterling. The securities with maturity within a year are called bills, while the securities with maturity over a year are called "gilts" or "gilt-edged securities". Gilts make up the largest proportion of government debt, around 86%.[2] The DMO issues two types of gilts: Conventional and index-linked. Conventional gilts are nominal bonds i.e., interest payments and coupon repayments are fixed in nominal terms. They constitute around three-quarters of the gilts issued by the DMO. Index-linked gilts are securities with coupon and final redemption payments linked to inflation, more specifically to the UK Retail Price Index (RPI). They constitute around a quarter of the debt issued by the DMO. The primary means of issuing gilts is through regular auctions, with over 75% of the overall gilt sales. The remaining part is issued through syndicated gilt offerings or mini gilt tenders. The annual financing remit, set by the UK Treasury, outlines the gilt sales required from the DMO for the upcoming financial year. The document specifies the total amount of gilt sales and the breakdown between index-linked gilts and conventional gilts in different maturity buckets. It is published every year in mid-March as the financial year runs from the 1st of April until the 31st of March. Occasionally the remit is revised in April when the central government's final net cash requirement for the previous financial year is pub-

[2] See Figure A1 in Appendix A.

lished. Furthermore, the remit is usually revised in November or December when the UK government publishes its budget together with forecasts of public finances. The remit contains the Gilt Auction Calendar, stating the dates of the auctions in the next financial year. Furthermore, the document states the amount of gilts to be issued and the number of planned auctions in four categories. The four categories are index-linked gilts, and three conventional gilt maturity buckets: short, medium, and long conventional gilts with 0-7, 7-15, and 15+ years to maturity. Therefore, the information in the remit gives investors an idea about the average size of the coming auctions in each category. An example of the DMO Financing Remit is displayed in Figure A2 in Appendix A.

The DMO announces its auction plan for the next quarter on the last business days of March, May, August, and November in an operations calendar. An operations calendar is shown in Figure A3 in the Appendix for an example. This calendar publishes the dates of the coming auctions, mini-tenders, and syndicated issuances in the next quarter. The document also specifies the maturity year and the interest rate of the issuance. Importantly, it does not provide information about the size of the auction. This is in contrast with the US, where the Treasury gives preliminary estimates of future auction volumes every quarter.[3]

The auction announcements are published at 3:30 pm, usually on the Tuesday in the week preceding the auction. This press release contains all the pertinent information about the issuance. Importantly, this is the time investors learn the exact size of the auction. Additional information released in the statement are ISIN, SEDOL codes, coupon payments, and the terms and conditions of the auction. An example announcement of a 10-year gilt auction published on the 21st of April 2015 is displayed in Figure A4 in the Appendix. Progress reports on the financing remit are often included in these announcements. These contain information on the remaining amount of gilts to be issued and the number of auctions to be held in the rest of the fiscal year. See Figure A5 in the Appendix for an example.

1.3.2 High-frequency surprises

To study the effects of debt issuance on interest rates, one can regress daily yield changes on the announced volumes. However, most of the announced new debt either covers the refinancing of maturing bonds or finances public expenditures that are known before the announcement. In other words, a large share of new issuances is anticipated by

[3] See the US Quarterly Refunding Press Conference: https://home.treasury.gov/policy-issues/financing-the-government/quarterly-refunding.

markets. Then, the effect is already priced in by the time of the announcement and the regression coefficients will not reflect the true causal effect.

A second option would be to use the surprise component of the announcements in the regression. Unfortunately, surprises are not observable. What is available is the required average future auction size to meet the DMOs' yearly financing remit. This quantity is published in the auction announcement press releases.[4] It is calculated as: $\frac{\text{Remaining gilt sales}_t}{\text{Number of auctions remaining}_t}$. We can use this as a proxy for investors' expectations of the announcement. We subtract this from the actual announced volume and label it as the *surprise volume*. While this is arguably a better measure, it is still prone to the issue of anticipation, as investors form their expectations based on much more information than the DMOs' progress. Therefore, the surprise volume series cannot be a true shock, which is confirmed by the fact that the series is autocorrelated.[5] To overcome these difficulties and capture unexpected changes in the supply of bonds, we opt for high-frequency identification. Nevertheless, below we will make use of the announced volume and the surprise volume series to support the validity of our identification.

We use high-frequency identification to isolate anticipated and unanticipated policy changes, as in the monetary policy literature (Kuttner (2001), Nakamura and Steinsson (2018)). Most similar to our application is Droste et al. (2021), who identify Treasury *demand* shocks by following futures price movements around the publication of US auction results. In contrast, we identify Treasury *supply* shocks by following futures price movements around announcements of bond issuance volumes. We restrict our attention to announcement days with conventional nominal gilt announcements only (and no tenders or index-linked gilt auctions).

Data on auction announcements are sourced from the DMO. The dataset starts on the 15th of May 2001 (the date of the first announcement on the DMOs' website) and ends on the 31st of December 2019. It contains 400 auctions over 360 announcement days. As explained in Section 1.3.1, these announcement days are usually, but not always the Tuesdays of the week preceding the auction. First, we collect auction dates from the auction results section of the DMOs' website. Then, we match each auction with the corresponding press release of the announcement. This document contains the announced volume, as well as a progress report with the remaining issuance volume and the remaining number of auctions in the fiscal year. In the few cases when the press release is not available, we use the dates and times specified in the DMOs' operations calendar and

[4] See Figure A5 in Appendix A for an example.
[5] See Figure A6 in Appendix A.

obtain the announced volumes from the auction results.

We record high-frequency gilt futures price movements around the announcements, as it is conventional in the high-frequency identification literature. Futures prices have many advantages for this application compared to spot prices or when-issued prices. Futures contracts trade on exchanges, while bonds trade over the counter. Therefore, the quality and the availability of price data is much better. Futures are also much more liquid than their cash counterparts, and futures markets tend to lead price discovery ahead of the spot (Garbade and Silber (1983), Di Gangi et al. (2022)).

We use intra-day gilt futures front contract prices to identify supply shocks, purchased from tickdatamarket.com. The contracts are traded on the London ICE exchange. There are four futures contracts written on UK government bonds: short, medium, long, and ultra-long. These can be satisfied with bonds with remaining maturities of 1.5–3.25, 4–6.25, 8.75–13, and 28–37 years, respectively. Data on the short and the medium contract are available from 2010 onward, while the long contract is available from 2001 onward. The ultra-long contract is much less liquid and we have data only between 2014 and 2016. Results are unchanged if we leave out the ultra-long contract from the analysis.

As described above in detail in Section 1.3.1, the DMO releases precise information about an upcoming action usually on Tuesday of the preceding week at 3.30 pm. These occasions are the first time the DMO discloses information about the size of the auction. Prior to this, investors can only speculate on the volume based on the remaining issuance volume and the number of auctions left for the year. The announcements contain information about (among other things) the volume, the coupon, and the exact maturity of the upcoming issuance. An example announcement is displayed in Figure A4. In other words, the announcements contain information solely about the supply side of the market. Price changes in a narrow window around the announcement should reflect revisions in investors' bond supply expectations. We interpret these as the supply shocks.[6]

The supply shock $S_t^{(m)}$ on announcement day t in maturity segment m is measured as the difference between the (log) futures price after and before the publication of the press release. More explicitly:

[6] Assuming liquidity premia does not change in the narrow event window. While liquidity conditions of Treasury futures are systematically priced, the liquidity premium is considered to move at lower frequencies (see Piazzesi and Swanson (2008) and Nakamura and Steinsson (2018)).

$$S_t^{(m)} = \left(ln(P_{t,post}^{(m)}) - ln(P_{t,pre}^{(m)}) \right) \times 100 \qquad m \in \{\text{short, medium, long, ultra-long}\} \tag{1.1}$$

where $P_{t,post}^{(m)}$ is the futures price 30-minutes after the announcement and $P_{t,pre}^{(m)}$ is the futures price 30-minutes before the announcement.[7] We use the five-minute centered moving average of the price to smooth out noise in the data. In minutes with no trading activity, we use the midquote: the average of the lowest bid price and the highest ask price. We record the price difference in Equation (2.1) for all four futures contracts, regardless of the maturity of the bond announced. Ideally, we would like to have time series that track shifts in the supply at every maturity point of the term structure. However, we can only proxy the shifts by price movements at the four points where futures contracts are available.

An illustrative example is the 10-year conventional gilt auction held on the 29th of April 2015. The exact size of the auction was published at 3:30 pm on the 21st of April (see the press release in Figure A4 in the Appendix). The volume was £3000 million, which was 10% larger than the average future auction size implied by the DMOs' progress report, published a week earlier (see the medium bucket in Figure A5). The release of this information about lower supply was followed by a marked increase in the price of all futures contracts, as displayed in Figure A7.

The time series of the four supply shocks are displayed in Figure A8. The four $S_t^{(m)}$ series are highly correlated, so we found it convenient to compress these series into one variable by extracting the first (probabilistic) principal component. We label this series S_t without a superscript. The interpretation of S_t is an unexpected, non-maturity-specific shift in the supply of government bonds. The mean of S_t is 0.001 with a standard deviation of 0.132. We normalize it to have zero mean and unit variance and use it in our regression analysis as our explanatory variable.[8] The dependent variables in the regressions are daily yield changes. By moving from intraday to daily frequency, we intend to capture responses that might take longer to materialize than the one-hour length of the event window.

Table A1 reports the descriptive statistics of the supply shocks. The means are very close to zero, suggesting that the shocks are not systematic. The ultra-long contract has a positive mean, most likely due to the short sample and the low liquidity of the contract.

[7] Our results are robust to both narrower and wider event window specifications. These results are available upon request.

[8] Results using the maturity-specific surprises $S_t^{(m)}$ are similar and available upon request.

The standard deviations increase with the maturity of the contracts. Table A1 shows that the series are strongly correlated. The ACF in Figure A9 shows no serial autocorrelation. This assures us that the shocks are not just due to shifts in the timing of the DMO's issuance plan.

It is important to make sure that no other relevant information is released around the announcements that could contaminate the identification of the supply shock. The most important drivers of Treasury yields are macroeconomic news releases, monetary policy decisions, and government bond auction results according to Fleming and Remolona (1997). The times of these events are all outside of our event window, but our results are robust to omitting announcement days that coincide with either one of these events.[9] Our event window starts at 3.00 pm. Macroeconomic data releases are published at 7:30 am or 9:00 am by the statistical office. Monetary policy announcements are published at 12:00, with a press conference held at 12:30. The DMO is also very transparent about releasing public announcements. Market-sensitive information is usually announced between 7.30 am and 8.00 am. On auction days, the bidding process closes at 10.00 am or 10.30 am, and the results are published shortly after. Post Auction Option Facility[10] results are published at the end of the take-up window closure at 1.00 pm or 2.00 pm. For more information, see DMO (2021).

We identify the supply shock S_t as price movements within a narrow event window around announcements. The assumption is that these price movements are the equilibrium responses to underlying shifts in the supply. To verify that these market responses are related to actual changes in the supply, we link our high-frequency shock to observable movements in supply. We can use two available observable measures from the auction announcement press release documents. The announced volume, and the "surprise volume" series. In Section 1.3 we discussed that these series are not ideal to analyse the effects of variations in the supply. Nevertheless, we can still use them to validate our high-frequency identification, by relating them with S_t.

First, we regressed the announced volumes on the high-frequency supply shock S_t but did not find a significant relationship between the two variables. Next, we regressed the "surprise volume" series on S_t. The estimated coefficient is significant and negative, implying that higher-than-expected supply is associated with a decrease in the futures price within the event window. Table A2 displays these results in the left column. The right column reports the results when we use a one-day event window as in Phillot (2021),

[9] These results are available upon request.

[10] Since the 1st of June 2009, all successful UK gilt auction bidders have the option to purchase up to 10-15% of the bond they have bought, at the published average auction price.

instead of the one-hour window of S_t. The insignificant coefficients imply that using a narrower event window captures better the price movements that are related to the surprise component of the announced volumes. Furthermore, a regression of the daily surprise on the intraday surprise (reported in Table A3) yields a very low R^2, suggesting that there must be other important drivers of prices on announcement days other than the press release. These results underline our argument to use intraday supply shocks instead of daily supply shocks. This is underlined by Kerssenfischer and Schmeling (2022), illustrating how multiple different news events drive yields within a day.

1.4 Bond supply effects on the term structure of interest rates

1.4.1 Effect on nominal and real yields

To assess how unexpected shifts in the supply of bonds affect interest rates, we regress the supply shock S_t onto interest rates at each maturity:

$$\Delta R_t^{(m)} = a^{(m)} + b^{(m)} S_t + \varepsilon_t^{(m)} \tag{1.2}$$

Where $\Delta R_t^{(m)} = R_t^{(m)} - R_{t-1}^{(m)}$ is the change in the Bank of England zero-coupon-curve at maturity m relative to the previous day. The coefficients of interest are the estimated $b^{(m)}$, which capture the effect of the supply shock on the term structure.

The responses to an unexpected standard deviation increase in the (non-maturity-specific) supply of government bonds are displayed in Figure A10. The blue line shows that an increase in the supply of bonds raises nominal interest rates between 1 and 1.5 basis points. Rates at longer maturities respond stronger, implying an increase in the slope of the yield curve. The effect persists in benchmark rates until the next week when the announced auction takes place (see Figure A13). The magnitude of the effect is similar to the responses to demand shocks, found by Droste et al. (2021) in the US and Lengyel and Giuliodori (2022) in Germany and Italy. This is in line with D'Amico and Seida (2020), who found that Treasury yields reacted similarly to the FED's QE and QT announcements. Figure A11 in the Appendix presents similar IV results, where S_t is instrumented by the announced volume made on day t and the "surprise volume". Our results are also robust to adding control variables, such as the short-term interest rate and inflation (implied by the model in Section 1.4.3) or weekday dummies.[11]

To offer some intuition on the size of this effect, we provide a back-of-the-envelope calculation on a fiscal expansion announcement during the Covid-19 pandemic. On March 11, 2020, the UK government announced a fiscal stimulus package of £12bn.[12] We can translate this quantity into a high-frequency futures price surprise, using the regression results of Table A2. Then, we can obtain an estimate of the reaction of the term structure to an unexpected change in the supply of bonds of the size of the package with the results in Figure A10. These imply that an unexpected new £12bn issuance of nominal bonds is associated with a $12 \times -0.15 = -1.8$ change in the bond futures price, a roughly 13 standard deviation event. This in turn would increase nominal yields by around $13 - 19$ basis points. While this is a huge out-of-sample exercise, actual changes in long-term yields on the announcement day were in the ballpark, between 4-14 basis points. It is important to note, however, that our calculation assumes that the announced package is fully unanticipated and financed entirely by new debt issuance. In reality, the announcement was at least partially anticipated by the press, implying that some of the effects have already been priced in before the announcement.[13]

What could be the reason behind the reaction of the yield curve? Ang et al. (2008) finds that about 80% of the variations in US nominal yields are attributable to changes in expected inflation and the inflation risk premium. The sum of the two is called the inflation compensation, the additional return investors require for being exposed to inflation. To assess if the reactions in nominal yields are due to a change in the inflation compensation, we regress S_t onto the *real* zero-coupon-curve of the Bank of England. The real term structure is constructed using inflation-linked bonds and is available for maturities over 25 months. The spread between a (comparable maturity) nominal and inflation-linked bond is called the breakeven inflation rate. This is a market-based measure of the inflation compensation. Figure A10 shows the response of real rates in red, and the response of breakeven rates in grey. Real rates react with increases of 1-1.2 basis points. This implies moderate, 0-4 basis points increases in breakeven rates and inflation compensation. Inflation swap rates, a different market-based measure of inflation compensation, show similar responses.[14]

These results suggest that the reason behind the reaction of nominal yields to the

[11] These results are available upon request.

[12] See: https://www.gov.uk/government/speeches/budget-speech-2020.

[13] See: https://www.reuters.com/article/uk-britain-sterling-close-idUKKBN20W2IV.

[14] See Figure A12 in the Appendix. An inflation swap contract exchanges a fixed rate against the realized average inflation rate at maturity. It is a market-based measure of the inflation compensation, which is less affected by market liquidity conditions (ECB (2018)).

supply shock is not a change in investors' inflation outlook. Therefore, to get a better understanding of the transmission of the shock, we break down nominal yields into their components in the next section and analyse how each component reacts to the shock.

1.4.2 Supply effects on expected short rates and risk premia

According to the expectations hypothesis, the response of long-term rates could be the result of either higher expected future short rates or higher risk premia. Using quarterly data and recursive identification, Dai and Philippon (2005) found risk premia to account for one third of the reaction of long-term rates to a shock to the fiscal deficit. Laubach (2011), at the same frequency, found that fiscal deficits mostly affect the short rate and inflation, with small movement in risk premia. Similar investigations in the empirical monetary policy literature suggest that high-frequency monetary policy shocks primarily influence expected short rates, with some effect on term premia at longer horizons (Hanson and Stein (2015), Abrahams et al. (2016), Nakamura and Steinsson (2018)). This paper, on the other hand, traces the effects of high-frequency Treasury supply shocks.

To get a better understanding of why government debt issuance affects interest rates, in this section, we first decompose yields into the average expected nominal short rate and the nominal term premium. Then, to shed light on the role of inflation, we further decompose the nominal short rate into the real short rate and expected inflation, and the nominal term premium into real term premium and inflation risk premium. We assess how each term is affected by the supply shock. Lastly, we attempt to clean our results from the relative liquidity effects of nominal and inflation-linked bonds, that might contaminate our inflation expectations and inflation risk premium variables.

Decomposing nominal and real yields

We use the affine term structure model (ATSM) of Abrahams et al. (2016) to jointly price nominal and inflation-linked bonds. For details of the model and the estimation see Appendix C. The model assumes that bond yields and the market price of risks are affine functions of the state variables, which are assumed to be observable. Hence, the log prices of a nominal ($P_t^{(\tau)}$) and an inflation-linked ($P_{t,R}^{(\tau)}$) zero-coupon risk-free bonds with remaining time to maturity τ follows:

$$\log P_t^{(\tau)} = A_\tau + B'_\tau X_t \qquad \log P_{t,R}^{(\tau)} = A_{\tau,R} + B'_{\tau,R} X_t$$

under the pricing measure, where X_t is the vector of pricing factors, assumed to follow an autoregression. Bond prices and yields are related through

$$y_t^{(\tau)} = -\frac{\log P_t^{(\tau)}}{n} \qquad\qquad y_{t,R}^{(\tau)} = -\frac{\log P_{t,R}^{(\tau)}}{n}$$

By imposing no arbitrage, expressions for the pricing coefficients $A.$ and $B.$ can be obtained, where the pricing coefficients are non-linear, recursive functions of the parameters driving the factors, the short rate, inflation, and the risk prices.

A τ-period nominal bond yield can be decomposed into the average expected nominal short rate over the next τ periods and the nominal term premium $\mathrm{TP}_t^{(\tau)}$. More explicitly:

$$y_t^{(\tau)} = \frac{1}{\tau}\sum_{i=0}^{\tau} E_t r_{t+i} + \mathrm{TP}_t^{(\tau)} \tag{1.3}$$

This can be further decomposed into the average expected real short rate, the average expected inflation, real term premium $\mathrm{TP}_{t,R}^{(\tau)}$ and inflation risk premium $\mathrm{IRP}_t^{(\tau)}$:

$$y_t^{(\tau)} = \frac{1}{\tau}\sum_{i=0}^{\tau} E_t(r_{t+i,R} + \pi_{t+i}) + \mathrm{TP}_{t,R}^{(\tau)} + \mathrm{IRP}_t^{(\tau)} \tag{1.4}$$

The interpretation of $\mathrm{TP}_{t,R}^{(\tau)}$ is the compensation investors require today to hold (real) interest rate risk for the next τ periods, while the interpretation of $\mathrm{IRP}_t^{(\tau)}$ is the compensation investors require to hold inflation risk for the next τ periods.

The elements of Equations (1.3) and (1.4) can be obtained as the following. Setting the price of risk parameters to zero, one can obtain the risk-adjusted counterparts of the pricing recursion coefficients $\tilde{A}.$ and $\tilde{B}.$. Bond yields calculated with these coefficients are interpreted as the time t expectation of average future short rates over the next τ periods. This would be the prevailing yield if all investors were risk neutral. The difference between the risk-adjusted expected nominal and the risk-adjusted expected real short rate is the average expected future inflation over the next τ periods. The nominal (real) term premium can be obtained by subtracting the nominal (real) expected short rate from the fitted yield. The inflation risk premium is obtained as the difference between the fitted breakeven inflation and the inflation expectation.

Reaction of yield components

Which components account for the strong response to the supply shock? To answer this question, we regress the supply shock on each component obtained above. First, we look at the response of expected nominal short-term rates and the nominal term premium. Then, the expected real short-term rates, expected inflation, the real term premium, and the inflation risk premium.

The top panel of Figure A20 shows the reaction of the nominal term premium and expected nominal short rates to the supply shock. The response of yields is given by the sum of the two bars. The figure shows that a standard deviation increase in bond supply raises 10-year yields by about 1.4 basis points. Around 1-basis point increase comes from the reaction of the term premium, and 0.4 basis point increase comes from higher expected short rates. Next, to shed more light on the role of inflation, we look at the average real short rate, the expected average inflation, the real term premium, and the inflation risk premium. The reaction of each component to the supply shock is displayed in the bottom panel of Figure A20. It shows that most of the reaction of the nominal term premium is due to the response of the real term premium. Interestingly, inflation expectations are unaffected, while the inflation risk premium displays a modest increase. In other words, additional government debt issuance mostly raises the compensation investors require to hold interest rate and inflation risks. This is in contrast with Dai and Philippon (2005) and Laubach (2011), who find that short rate expectations respond more to higher fiscal deficits. They conduct their analysis, however, at a much lower frequency and focus on deficits rather than bond supply shocks. Phillot (2021) finds that bond supply shocks raise breakeven inflation rates. Our results demonstrate that this reaction is not due to higher inflation expectations, but due to a higher inflation risk premium. One interpretation of this result is the following. The government might issue more debt to finance expansionary fiscal policy. Any inflationary pressure from this policy is expected to be fully offset by the central bank, keeping inflation expectations unchanged, but raising real short rate expectation.

Apart from duration and inflation risks, the behaviour of credit risk is also of interest. Excessive debt issuance by the government can lead to debt repayment issues, raising the credit risk of the government. When markets price higher credit risk, it is reflected in higher credit default swap (CDS) rates. CDS rates can be interpreted as the insurance premium paid to insure against the default of the bond issuing entity. We assess if this is the case by regressing the high-frequency supply shock on daily CDS rate changes written on UK Treasuries from Refinitiv. Table A4 shows that increased bond supply does not have a positive effect on CDS rates, implying no increase in the credit risk of the UK government priced in CDS rates.

Next, we look at the behaviour of corporate bond yields and corporate yield spreads at various maturity buckets from Refinitiv. We found spillover effects of the government bond supply shock into corporate bond markets, consistent with the effects of demand shocks, found in Droste et al. (2021) and Lengyel and Giuliodori (2022). Corporate bonds react strongly, with yields increasing between 0.7 and 1.5 basis points in all maturity segments. Figure A21 shows the reaction of AAA, AA, A and BBB rated corporate bonds, categorized into maturity buckets of 3-5, 5-7, 7-10, 10-15 and 15+ years to maturity. Corporate bonds with better rating react more to changes in the supply of government bonds. In terms of remaining maturity, bonds that mature between 7 and 15 years have the strongest reaction to the supply shock, which is the same segment where Treasury bonds respond the most. The reaction of BBB-AAA spreads is reported in Table A5. The regression coefficients are insignificant in all maturity buckets. These results on CDS and corporate bonds suggest that repricing of credit risk is unlikely to be a transmission channel of the supply shock on interest rates. Stock prices react positively. The FTSE 100 index gains 0.166 (0.067) percent after a standard deviation increase in the supply of government bonds.

Overall, our results imply that an important transmission channel of the effects of government debt issuance is the repricing of risks in the economy. The bond supply shock affects markets' perception of duration and inflation risks, and this changes the equilibrium price of these risks. This is similar to the transmission of monetary policy shocks (Hanson and Stein (2015), Abrahams et al. (2016)), which increases the term premium. However, an interesting difference is that monetary policy shocks co-move negatively with the inflation risk premium, while bond supply shocks co-move positively. This is because a positive monetary policy shock is contractionary and disinflationary, while a positive bond supply shock indicates a more expansionary and inflationary stance of fiscal policy.

Adjusting for inflation-linked bond illiquidity

The liquidity of inflation-linked bonds and nominal bonds tend to differ, and the relative liquidity is systematically priced (Pflueger and Viceira (2016)). If the inflation-linked bond relative liquidity effect is priced, it can contaminate the response of inflation-related indicators, as they are derived using the spread between nominal and inflation-linked bonds. According to Joyce et al. (2010), the liquidity premium is unlikely to have had a big influence on UK yield curve dynamics over the period of 1992-2008 and ignores it (together with Evans (1998), Risa (2001), Abrahams et al. (2016) and others). Others quantified it to be low and decreasing over time but jumping higher in crisis periods (see Kaminska et al. (2018), Bekaert and Ermolov (2023)).

We attempt to account for this effect in a robustness exercise, by expanding the state space X_t, by including a liquidity factor L_t, as Abrahams et al. (2016) did for the US. Working at the daily frequency substantially reduces the number of potential liquidity proxies to use. We follow Pflueger and Viceira (2016), Kaminska et al. (2018) and Bekaert and Ermolov (2023) and use the 5-year inflation-swap spread ISS_t^{5Y} as our liquidity proxy.[15] We standardize it and add to each observation the negative of the minimum of the series to ensure the positivity of the index. Then, before calculating inflation-related variables in the model, we subtract the effect associated with L_t from yields. The availability of inflation swap data is from the 29th of June 2007 to the 31st of December 2019, and it is displayed in the bottom right panel of Figure A14 in the Appendix A. It shows a steep increase during the financial crisis and elevated levels during the European debt crisis.

We estimate the effect of the supply shock on the new series and Figure A22 reports the results. Once we account for liquidity effects, almost all of the reaction of yields is attributed to movements in the nominal term premium. The reaction of the real term premium dominates with a minor effect on inflation risk premium. The expectation component only reacts at short horizons. The reason for the muted response of the expectations variables is that they are obtained by setting the prices of risks to zero in the ATSM model when calculating the risk-neutral yields and breakevens. In this exercise, we quantified the price of an additional risk factor: liquidity risk.

1.4.3 A term structure model of nominal and real bonds with supply effects

To summarize our empirical results, the additional supply of nominal bonds raises nominal and real yields, mostly due to increases in risk premia. When investors are faced with a higher supply of government bonds, they require higher compensation for holding interest rate and inflation risks. In the next section, we examine this effect through the lens of a theoretical framework. The aim of the model is to illustrate the mechanism rather than provide a complete structural explanation of the mechanism.

The model builds on the Greenwood and Vayanos (2014) version of the Vayanos and Vila (2021) model. This section provides the intuition, while the complete model is spelled out in the Appendix. We extend the Greenwood and Vayanos (2014) model with inflation risk and include two types of bonds: a continuum of nominal bonds and a continuum of

[15] This is constructed as the difference between the inflation swap rate ISR_t^{5Y} and the breakeven inflation rate $ISS_t^{5Y} = ISR_t^{5Y} - (y_t^{5Y} - y_{t,R}^{5Y})$. Liquidity premium in inflation swap rates is considered to be negligible, therefore, the rates only represent expected inflation and inflation risk premium (see ECB (2018) and Bekaert and Ermolov (2023)). In the absence of liquidity risk premium in breakeven rates, the spread between the swap rate and the breakeven rate should be zero.

inflation-linked bonds. These are supplied by the government in a price inelastic manner. Marginal investors in the model are short-lived risk averse arbitrageurs. They absorb shocks to the supply of bonds and ensure that the term structure of interest rates is smooth and arbitrage free. Arbitrageurs require additional returns for holding the bonds compared to the risk-free short rate, as unexpected shocks can result in the bonds underperforming relative to the short rate.

To the best of our knowledge, there are two papers with similar setups. Saúl (2012) derives the breakeven inflation rate in a model with preferred-habitat investors, as in Vayanos and Vila (2021). Bond prices are determined through the interaction between arbitrageurs and preferred-habitat investors. Preferred-habitat demand for bonds is non-stochastic. This is in contrast to our setup of exogenous bond supply, which is subject to shocks. Our focus is specifically on this additional stochastic risk factor, and we analyse how equilibrium bond prices are affected by this supply risk. Diez de los Rios (2020) constructs a discrete-time version of the Greenwood and Vayanos (2014) model, with both nominal and real bonds in fixed supply. Inflation is endogenous, determined by a Taylor-rule type equation. The focus is on demonstrating how an increase in the bond supply can lead to higher inflation. Our continuous time model has exogenous inflation, to illustrate how additional bond issuance transmits to yields, by altering the price of inflation and duration risks.

Bond yields in the model react positively to the supply shock, with the effect stronger at long horizons, just like in our empirical results. The supply shock raises both the duration risk premium and the inflation risk premium. Furthermore, we also find a positive relationship between the supply shock and the breakeven inflation rate. The intuition is the following. In equilibrium, risk prices are increasing in the sensitivity of investors' portfolios to the risk factors. Expected excess returns of bonds are, therefore, also increasing in this sensitivity. As the outstanding amount of nominal bonds increases, the amount of duration and inflation risks borne by arbitrageurs also increases. The higher sensitivity of their portfolio to the risk factors raises the price of these factors and the risk premiums. This in turn raises the term premium and the inflation risk premium, raising bond yields. As inflation-linked bonds are free from inflation risk, their yield does not rise as much as nominal yields, resulting in higher breakeven inflation rates.

This mechanism is linked to the limited risk bearing capacity of investors. When risk aversion is high in the model, yields and risk prices become more responsive to the supply shock. In the next section, we test this prediction empirically and explore further state dependencies in the effects of the shock.

1.5 Non - linearities

The model in Section 1.4.3 suggests that the response of yields to the supply shock is higher in states when risk aversion is high. This is the same result found in Greenwood and Vayanos (2014) and Vayanos and Vila (2021) found in the context of demand shocks. We test this non-linearity empirically. He and Krishnamurthy (2013) suggest that risk aversion is higher in a crisis and periods of financial market stress. Therefore, we use a country-level composite indicator of systemic stress in the financial system: the CISS index of Hollo et al. (2012). We construct a financial stress indicator variable I_t, that is equal to one when the CISS index is above its 75th percentile and zero otherwise.[16] The indicator is displayed in Figure A23. We estimate the state-dependent version of Equation 2.3:

$$\Delta R_t^{(m)} = I_t\left[a_1^{(m)} + b_1^{(m)} S_t\right] + (1 - I_t)\left[a_0^{(m)} + b_0^{(m)} S_t\right] + \varepsilon_t^{(m)} \tag{1.5}$$

The findings are reported in Figure A24. In normal times a standard deviation supply shock raises nominal yields up to 1.2 basis points. On the other hand, during market stress periods the reaction is as high as 1.9 basis points at long maturities. The reason behind this is that in turbulent times, the term premium becomes much more responsive to the supply shock. Long-term bonds are more sensitive to risks and market stress periods are characterized by a steep increase in risk prices. The interpretation of this result through the lens of the model in Section 1.4.3 is that when investors are more risk averse, they require even higher compensation for a given amount of risk. This is consistent with the findings on Treasury demand shocks, which are documented to have stronger effects in times of market stress (Droste et al. (2021), Lengyel and Giuliodori (2022)).

The preferred-habitat theory of bond yields by Vayanos and Vila (2021) and Droste et al. (2021) predicts that when risk aversion is low, demand shocks affect interest rates similarly across the maturity space. However, when investors' risk aversion is high, a shock at a specific maturity segment has more concentrated effects at nearby maturities. In other words, the shock has a localized effect. We test this prediction in the context of supply shocks, by first restricting the announcements sample to only include announcements of short- and medium-maturity bonds (0-15 years according to the DMOs' classification) and estimating Equation (1.5). Then, we restrict the announcements sample to only include announcements of long-maturity bonds (15+ years) and estimate again Equation (1.5). The results are reported in Figure A25. It shows that when markets are calm, the effect

[16] Our results are robust to a wide range of this threshold and they are available upon request.

of the shock is similar across maturities. However, when markets are under stress, short- and medium-maturity bond announcements have a larger effect on the short end of the yield curve, while long-maturity bond announcements have a larger effect on the long end of the curve. This localization effect of bond supply changes is in line with the findings of McLaren et al. (2014) on QE programs in the UK, and Droste et al. (2021) in the US.

Next, we analyse the effects of the supply shock in periods when the monetary policy rate is at the effective lower bound.[17] We construct a dummy variable that takes the value one when the Bank of England policy rate was below 0.5%, and zero otherwise. Figure A23 shows the time series of the variable. The estimation results are reported in Figure A26. Short-maturity rates show a weaker response to the supply shock when they are constrained by the lower bound. At the same time, long-maturity rates react more stronger. The main reason is that average expected short rates do not react as much as in normal periods. Risk premia on the other hand, is more responsive: around 85% of the reaction is due to these components. It is important to note, that ELB periods were often characterized by elevated market stress levels, which might also drive these results.

Overall, in the sub-sample that is characterized by market stress and the ELB, the term premium and the inflation risk premium are more responsive, and short rates are less responsive to supply shocks. As long-term bonds are more sensitive to risk premia, the slope of the yield curve becomes steeper after an increase in government debt issuance. Furthermore, in states of high risk aversion, the localization effect of supply shocks can be observed.

1.6 Conclusion

In this paper, we identify government bond supply shocks by recording intraday price movements around government bond auction volume announcements. We apply this high-frequency identification to the UK Debt Management Offices announcements and study how additional debt issuance affects the term structure of interest rates. We find that a standard deviation bond supply shock increases nominal yields by 1-1.5 basis points. Real rates rise by 1-1.2 basis points, implying a modest reaction of the inflation compensation.

To study the transmission of the shock, we decompose yields into expected short rates and risk premia. We find that the shock mostly affects risk premia components, with smaller effects on future expected average short-term rates and no effect on expected inflation. Both the real term premium and the inflation risk premium react positively to

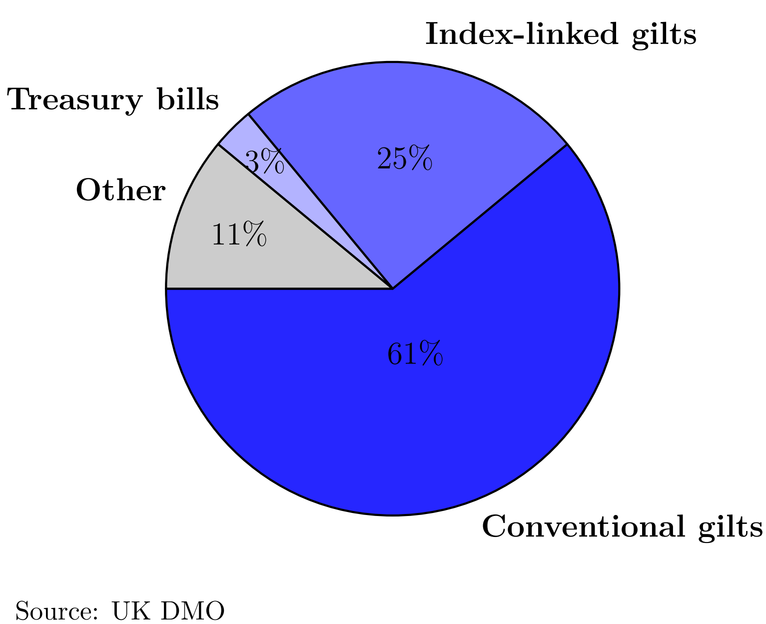
[17] We did not find differences in the effect of the supply shock based on the sign of the shock.

higher bond supply. We reconcile these results in an equilibrium term structure model, where risk averse investors absorb shocks to the supply of nominal bonds. Their equilibrium portfolio becomes more sensitive to duration and inflation risks, driving up the price of these risk factors. This in turn raises risk premia and yields. As inflation-linked bonds are unaffected by inflation risk, the breakeven inflation rate goes up.

The model also predicts that when risk aversion is high, the effects of the supply shock are more pronounced. In line with this, we find empirically that yields react stronger to the supply shock during times of financial market stress and at the effective lower bound. The increase is driven by higher risk premia, consistent with the equilibrium model. Furthermore, we find evidence for the localization of the effect during market stress periods.

Appendix A Figures and tables

Figure A1: Composition of UK central government sterling debt in December 2019



Source: UK DMO

Figure A2: DMO Financing Remit for the 2015-16 financial year, published the 18th of March 2015

DMO FINANCING REMIT 2015-16: 18 MARCH 2015

1. The DMO's financing remit for 2015-16 has been published today as part of the Budget 2015 announcements. The main points are summarised below.

A) Debt issuance by the DMO

2. The DMO plans to raise £140.4[1] billion in 2015-16, split as follows:

• Outright gilt sales:	£133.4 billion.
• Net Treasury bill sales (via tenders):	£7.0 billion.

B) Planned gilt sales

3. It is intended that the gilt sales plans will be met through a combination of:

   - £105.2 billion of issuance in 39 auctions; and
   - additional supplementary gilt sales of £28.2 billion (21.1% of total issuance) via a combination of syndicated offerings and, subject to demand, mini-tenders. This will comprise a minimum £24.2 billion via a syndication programme. Any additional sales via syndication can only be of long conventional or index-linked gilts but mini-tenders can be used for issuance of conventional and index-linked gilts across the curve.

4. The planned split of issuance by maturity and type of gilt to be sold via auctions and syndicated offerings is as follows:

   Conventional:

Short:	£33.9 billion (25.4%) in 8 auctions
Medium:	£26.7 billion (20.0%) in 8 auctions
Long:	£37.4 billion (28.0%) in 12 auctions and via syndicated offerings (aiming to raise £28.1 billion by auctions and a current planning assumption of a minimum of £9.3 billion via syndication).
Index-linked:	£31.4 billion (23.5%) in 11 auctions and via syndicated offerings (aiming to raise £16.5 billion by auctions and a current planning assumption of a minimum of £14.9 billion via syndication).

5. The issuance methods to achieve the syndication and mini-tender plans are based on current assumptions. In particular, total financing achieved through each supplementary issuance method will be dependent on market and demand conditions at the time transactions are conducted.

[1] Sales figures in this announcement are in cash terms unless otherwise indicated.

Figure A3: Gilts Operations Calendar for April-May 2015, published the 31st of March 2015

United Kingdom
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www.dmo.gov.uk

31 March 2015

PRESS NOTICE

GILT OPERATIONS CALENDAR: APRIL- JUNE 2015

PLANNED SYNDICATED OFFERING OF AN INDEX-LINKED GILT WITH A MATURITY IN THE 30 YEAR AREA OR LONGER IN JUNE 2015

The UK Debt Management Office ("the DMO") is announcing today that the first syndicated offering of the 2015-16 programme will be the sale of an Index-linked gilt with a maturity in the 30 year area or longer. The DMO expects that, subject to market conditions, the sale will take place in the second half of June 2015. Further details of the sale, including the composition of the syndicate, will be announced in due course.

The DMO also announces that in the period April-June 2015 it plans to hold ten outright gilt auctions as well as the syndicated offering, as set out below.

Auction date	**Gilt**	**Further details announced**[1]
Wednesday 8 April	2% Treasury Gilt 2020	Tuesday 31 March
Thursday 16 April	0⅝% Index-linked Treasury Gilt 2040	Tuesday 7 April
Tuesday 21 April	3½% Treasury Gilt 2045	Tuesday 14 April
Wednesday 29 April	2% Treasury Gilt 2025	Tuesday 21 April
Thursday 14 May	2% Treasury Gilt 2020	Tuesday 5 May
Thursday 21 May	4¾% Treasury Gilt 2030	Tuesday 12 May
Wednesday 27 May	0⅛% Index-linked Treasury Gilt 2058	Tuesday 19 May
Tuesday 2 June	2% Treasury Gilt 2025	Tuesday 26 May
Tuesday 9 June	0⅛% Index-linked Treasury Gilt 2024	Tuesday 2 June
Thursday 11 June	3½% Treasury Gilt 2045	Tuesday 2 June

[1] Further to the announcement on 29 January 2015, as of 31 March 2015 the DMO will no longer be declaring a "When Issued" (WI) trading period in cases where the stock being auctioned is a re-opening of an existing line of gilts with a pre-existing ISIN code. Accordingly, the DMO will no longer be issuing a separate WI ISIN code for new tranches of existing gilts. The 29 January announcement can be found at: http://www.dmo.gov.uk/documentview.aspx?docName=/gilts/press/sa290115.pdf

Figure A4: Auction announcement of the auction of £3,000 million of 2% Treasury Gilt 2025, published the 21st of April 2015

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21 April 2015

PRESS NOTICE

AUCTION OF BRITISH GOVERNMENT STOCK

Auction Details

Auction Date	Wednesday, 29 April 2015
Issue and Settlement Date	Thursday, 30 April 2015
Bidding Convention	Fully paid Bid Price (see Note 1)
Accrued Interest payable with bid	£0.222826 per £100 nominal
Auction Close	10:30am London Time

Details of Security

Title	2% Treasury Gilt 2025
Amount (nominal) for auction	£3,000 million (fungible with previous issue) (see Note 4)
Nominal outstanding after auction	£6,024.9 million
Maturity Date	7 September 2025 at par
Interest Dates	7 March – 7 September
ISIN Code	GB00BTHH2R79
SEDOL Code	B-THH-2R7
Strippable	From 30 April 2015 (see Note 2)
Interest Payable	Gross (see Note 3)
Next Interest Date	7 September 2015 - £0.929348 per £100 nominal (Short First Coupon)

Note 1: Bids may be made on either a competitive or a non-competitive basis. Details of the bidding procedures are set out in the prospectus and in the Information Memorandum. Gilt-edged Market Makers may bid by means of the Bloomberg Bond Auction System to the DMO not later than 10.30 am on Wednesday, 29 April 2015.

Note 2: Following the issue of this further amount of the Gilt, 2% Treasury Gilt 2025 may be stripped and holdings of the Gilt reconstituted: the provisions relating to strips contained in the Information Memorandum will therefore apply except that the minimum stripping unit will be £1,000,000 nominal until the payment of the non-standard first coupon on 7 September 2015. The SEDOL and ISIN codes for the new principal strip are B-WXB-PL9 and GB00BWXBPL93 respectively.

Note 3: Holders may elect to have United Kingdom income tax deducted from interest payments, should they so wish, on application to the Registrar, Computershare Investor Services PLC.

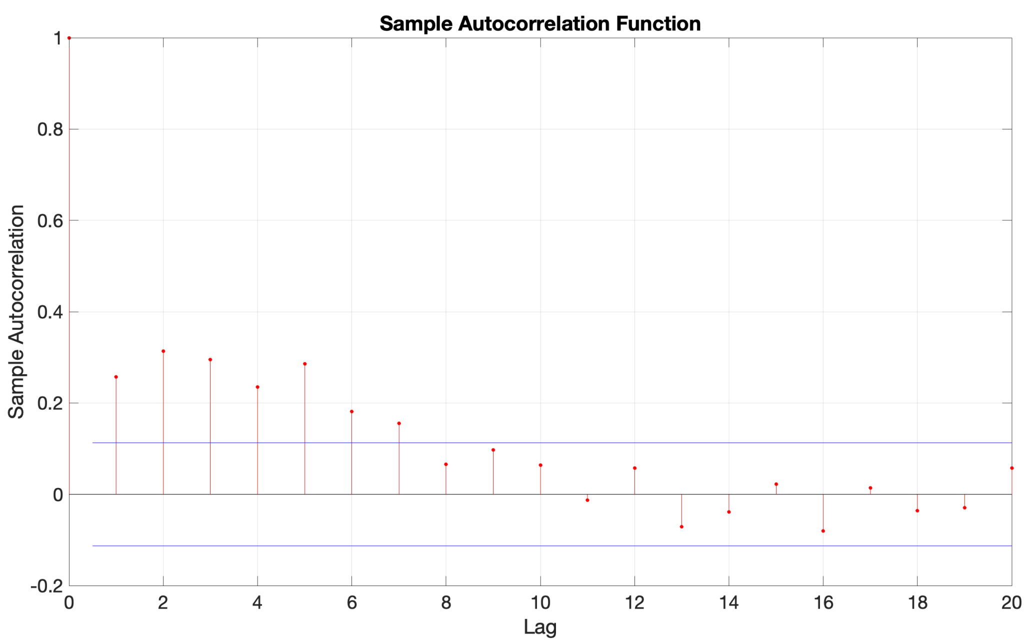
Figure A5: A progress report of the 2015-16 Financing Remit, published on the 14th of April 2015

Remit 2015-16

Gilt sales of £133.4 billion (cash) are planned in 2015-16 and progress against the remit is summarised in the table below (which may not include the amount of gilts issued under the Post Auction Option Facility for the most recent auction, if any).

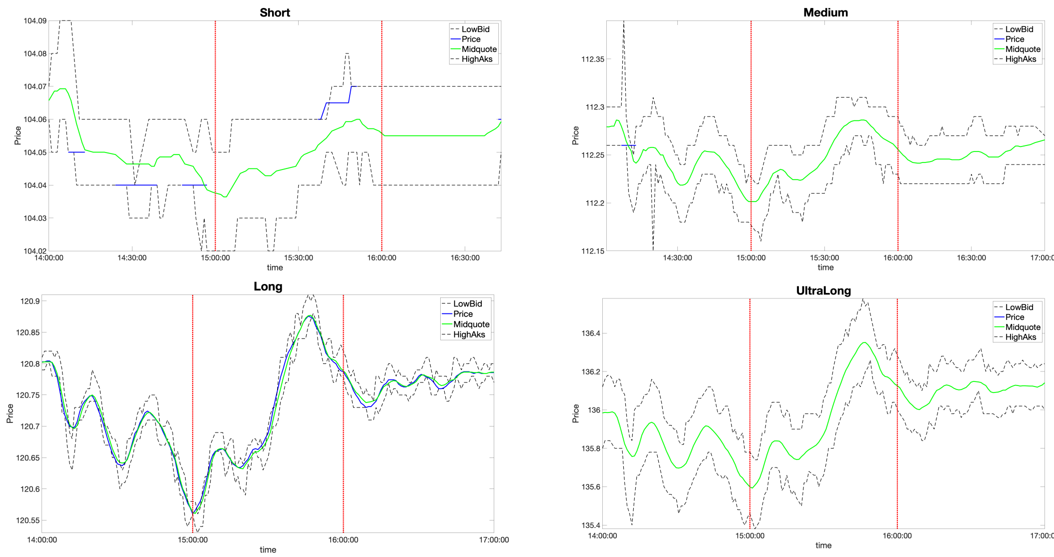
Gilt sales relative to remit plans 14 April 2015 (£ millions)					
	Conventional Gilts			Index-linked gilts	Total
	Short	Medium	Long		
Auction proceeds to-date	4,166	0	0	0	4,166
PAOF proceeds to-date	10	0	0	0	10
Auction and PAOF proceeds to-date	4,177	0	0	0	4,177
Syndication sales to-date	0	0	0	0	0
Mini-tender sales to date	0	0	0	0	0
Total gilt sales to date	**4,177**	**0**	**0**	**0**	**4,177**
Auction sales required to meet plans	29,723	26,700	28,100	16,500	101,023
Number of auctions remaining	7	8	12	11	38
Currently required average auction sizes	4,246	3,338	2,342	1,500	
Planned gilt sales at auctions	**33,900**	**26,700**	**28,100**	**16,500**	**105,200**
Number of auctions scheduled	8	8	12	11	39
Minimum syndication sales plan	0	0	9,300	14,900	24,200
Syndication sales required to meet minimum plan	0	0	9,300	14,900	24,200
Balance of supplementary gilt sales					28,200
Total planned supplementary gilt sales					**28,200**
Total planned gilt sales					**133,400**

Figure A6: Autocorrelation function of the surprise volume series



Note: Sample autocorrelation function of the shock series up to ten lags. Blue lines indicate two standard errors confidence bounds.

Figure A7: Futures price movement around the event window on the 21th of April 2015



Note: Red lines denote the event window, dashed line the lowest bid and highest ask price. The green line is the 5-minutes moving average of the midquote, the blue line is the 3-minutes moving average of the recorded traded price. Announcement was made at 15:30.

Figure A8: Time series of the supply shocks

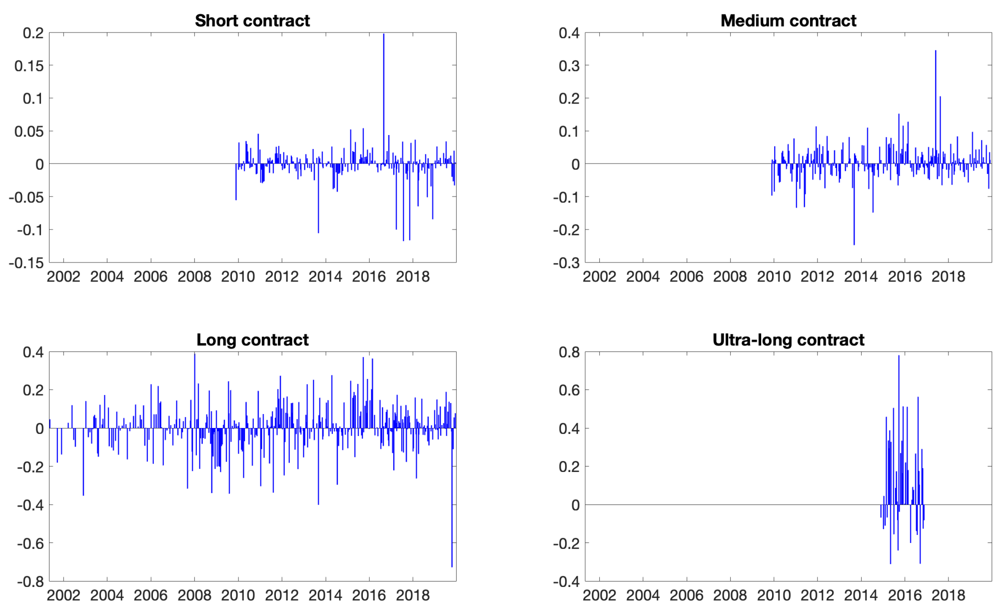
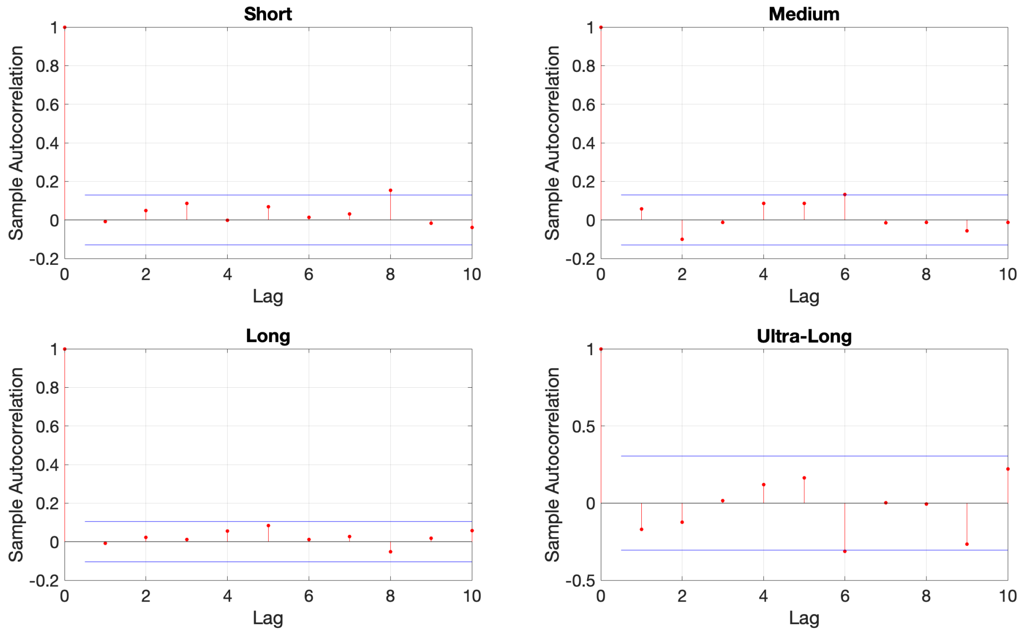
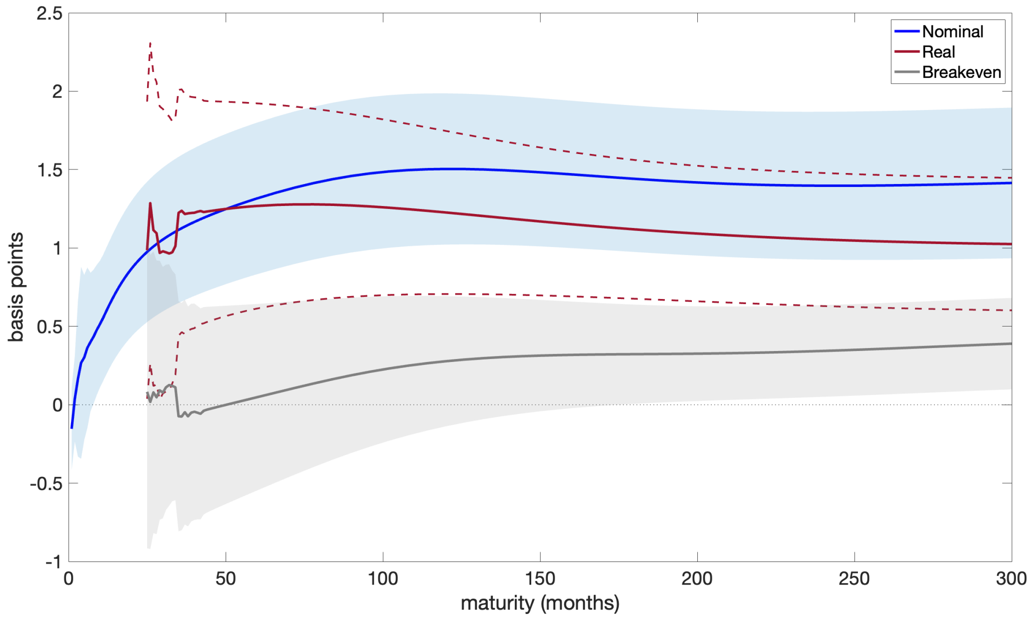


Figure A9: Sample autocorrelation function of the shock series



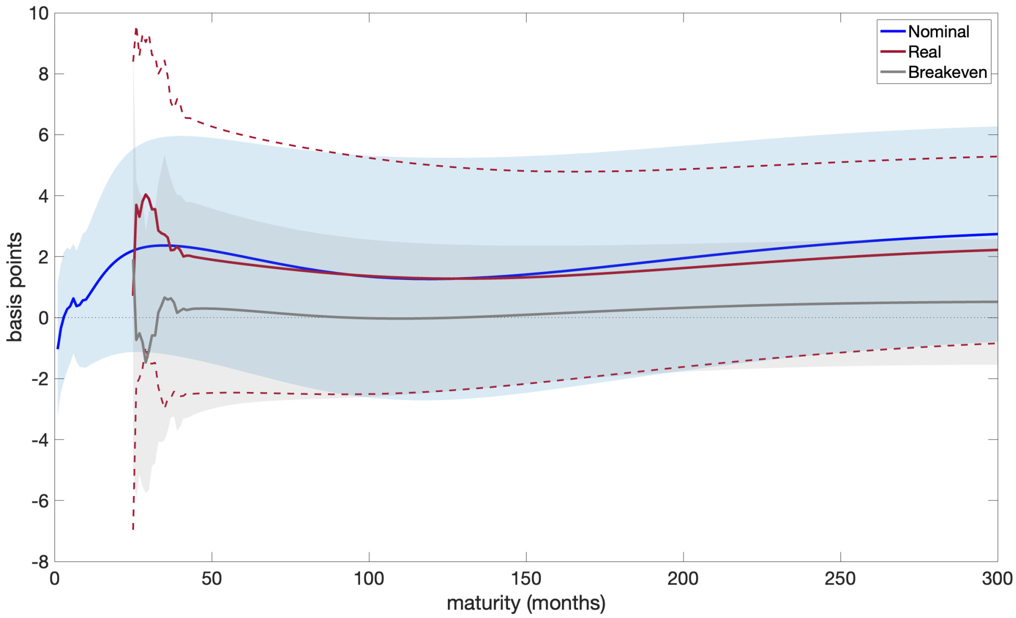
Note: Sample autocorrelation function of the shock series up to ten lags. Blue lines represent two standard errors confidence intervals.

Figure A10: Reactions of the term structure of nominal, real, and breakeven inflation to the bond supply shock



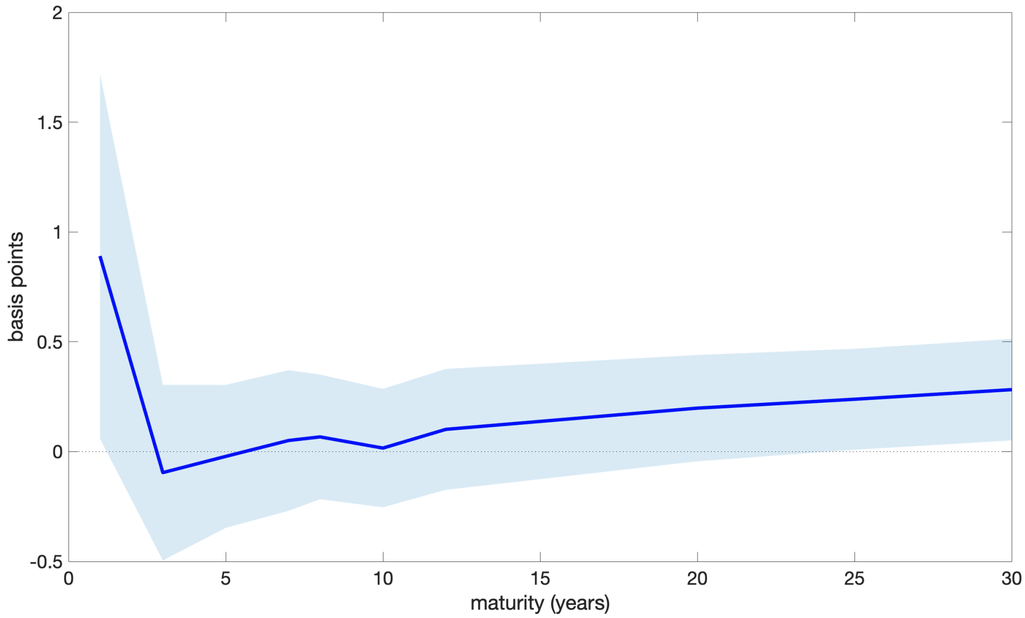
Note: Estimated $b^{(m)}$ coefficients from equation (2.3). Dashed lines and shaded areas are 95% (Newey-West, 10 lags) confidence intervals. Sample: 31.03.2001-31.12.2019.

Figure A11: Reaction of the nominal and real and the breakeven inflation term structures to the bond supply shock - IV regression



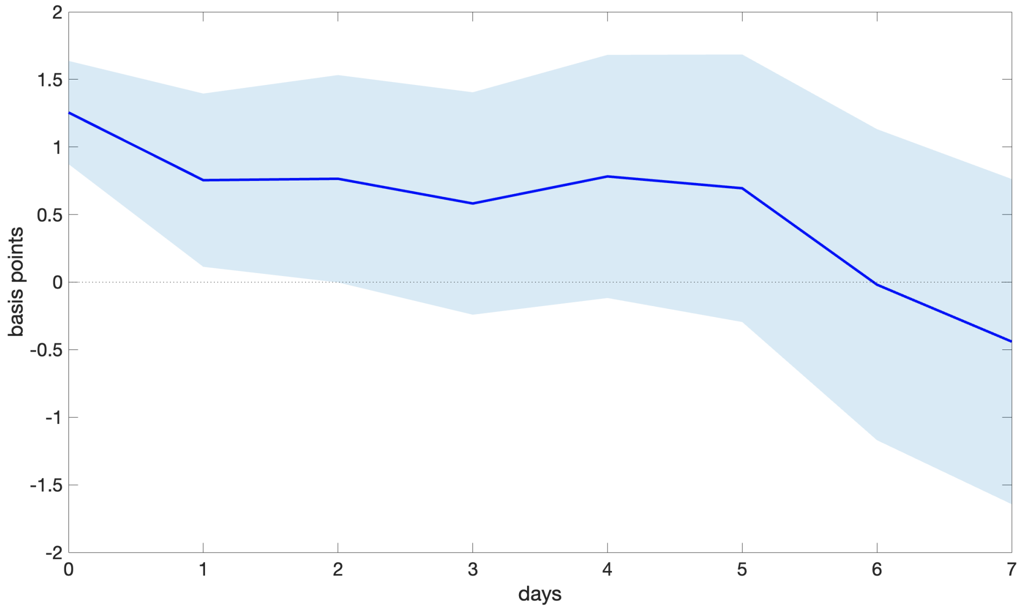
Note: Instrumental variable estimation of (2.3), where S_t is instrumented by the announced volumes and the surprise component of the announced volume. Shaded areas and dotted lines are two standard deviation confidence bands. Sample: 02.05.2006-31.12.2019.

Figure A12: Reaction of the inflation swap curve to the supply shock



Note: Nodes are the estimated $b^{(m)}$ coefficients from equation (2.3). Dependent variables are inflation swap rates from Refinitiv, with maturities of $1, 3, 5, 7, 8, 10, 12, 20, 25,$ and 30 years. Dashed lines are 90% confidence intervals. Sample: 01.05.2009-31.12.2019.

Figure A13: Impulse response of the 10-year benchmark nominal rate



Note: Impulse response of 10-year benchmark rates from long difference regressions, where the dependent variable is $Y_{t+h} - Y_{t-1}$ and h are days. Shaded area is 90% Newey-West (10 lags) confidence interval. Sample: 31.03.2001-31.12.2019.

Figure A14: Time Series of the ATSM Pricing Factors

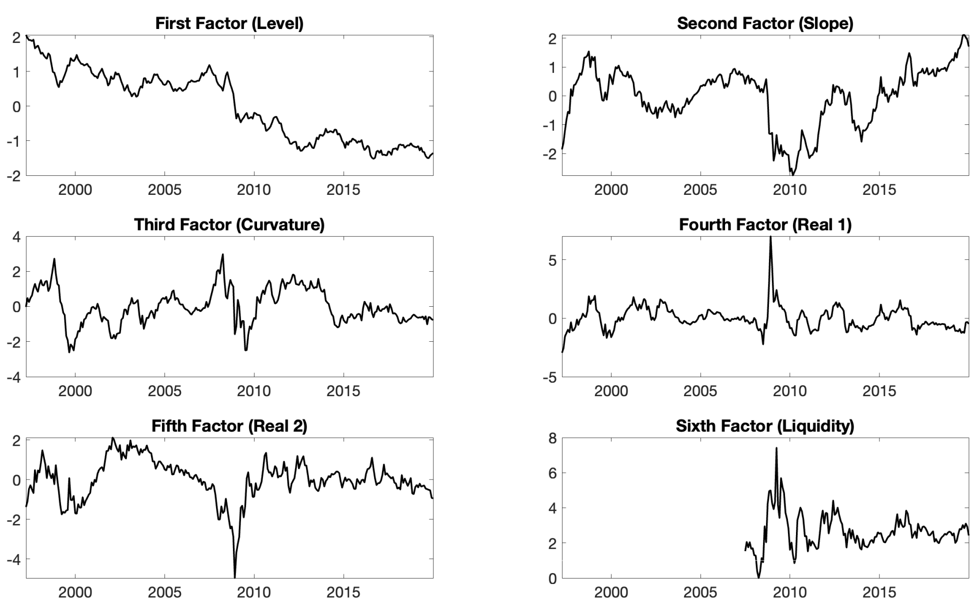


Figure A15: ATSM model fit at 10-years, monthly frequency - nominal yield (left), real yield (right)

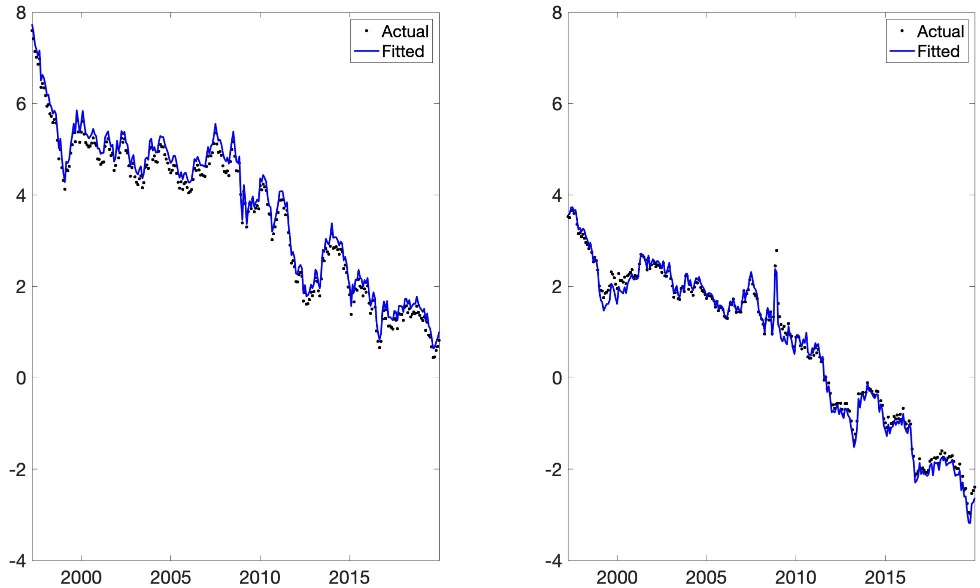


Figure A16: ATSM model fit at 10-years, daily frequency - nominal yield (left), real yield (right)

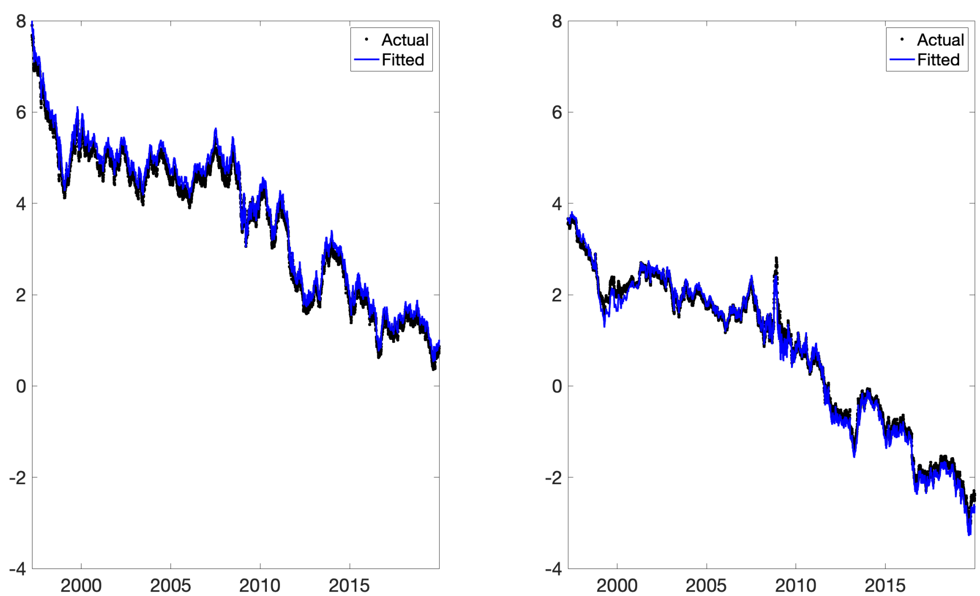
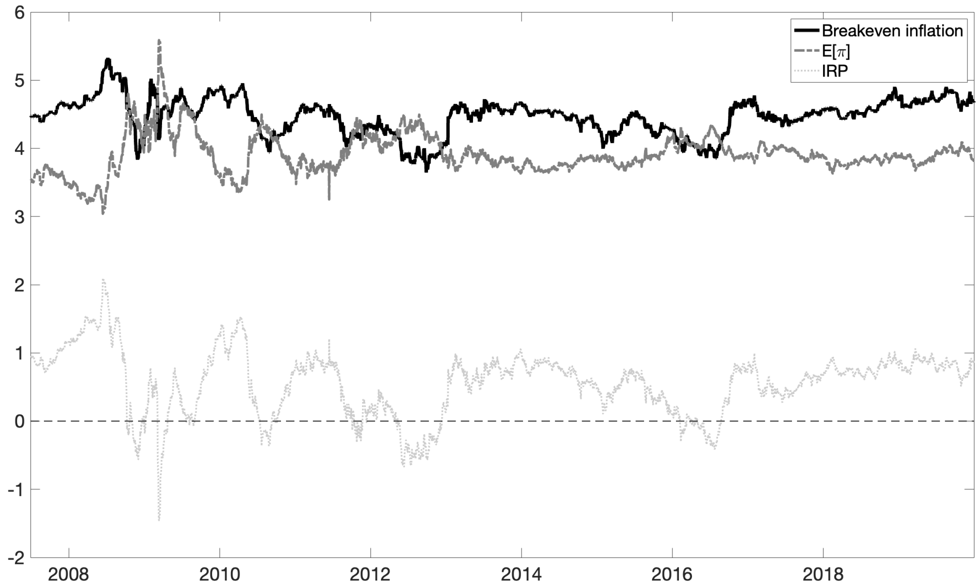


Figure A17: 10-year breakeven inflation rate decomposition at daily frequency, adjusted for liquidity



Note: Decomposition of the ATSM model implied 10-year breakeven inflation rates into expected average inflation and inflation risk premium, where the state space of pricing factors is extended with a liquidity proxy: the inflation swap, breakeven inflation rate spread.

Figure A18: 10-year nominal yield decomposition (left) and real yield decomposition (right) at the daily frequency

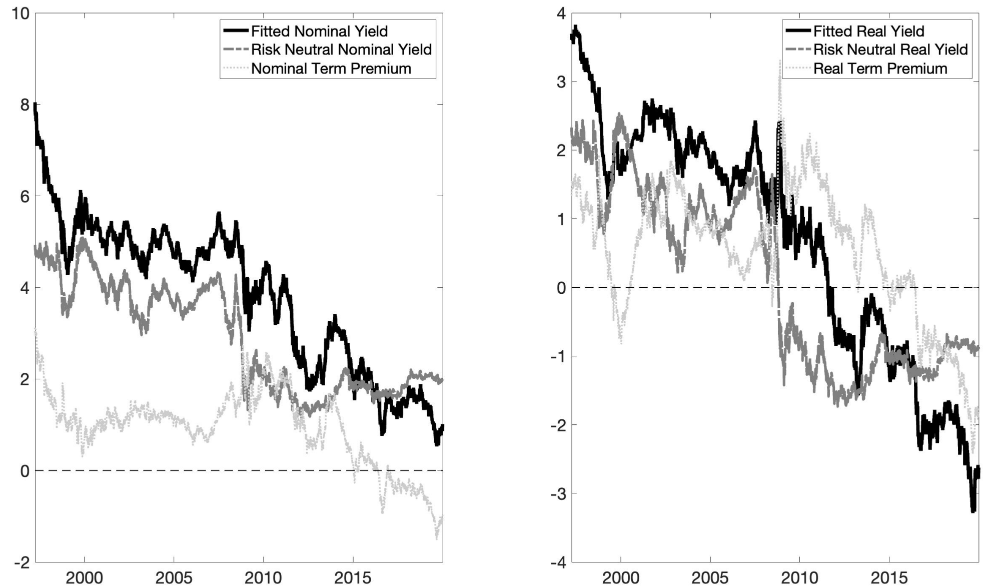


Figure A19: 10-year breakeven inflation rate decomposition at daily frequency

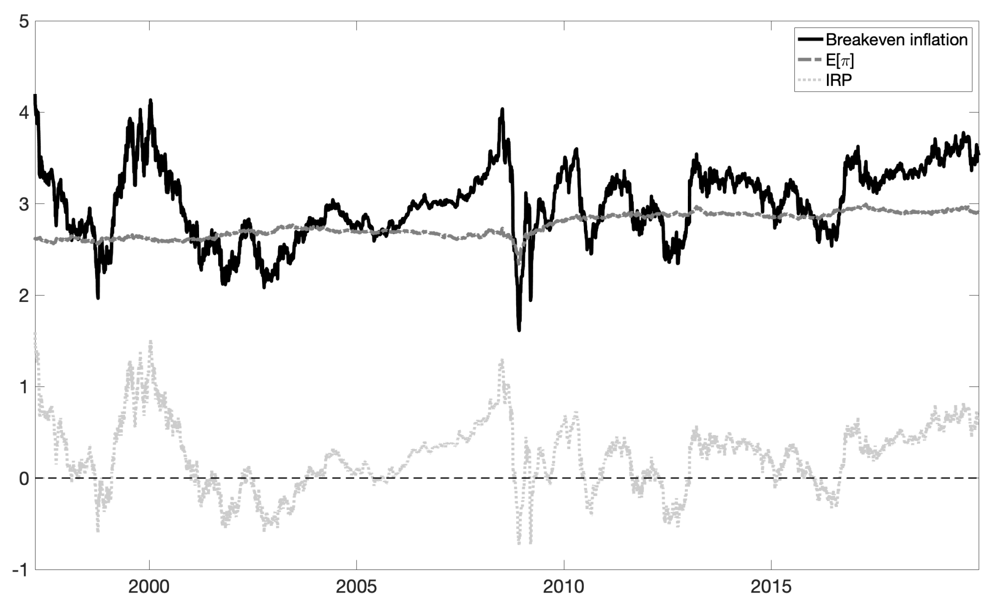
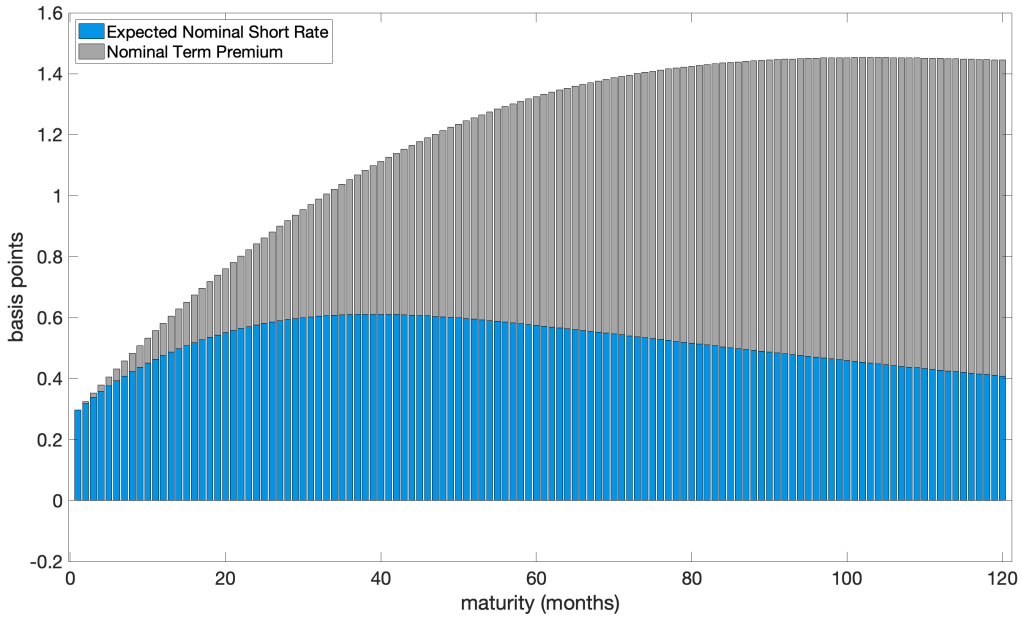
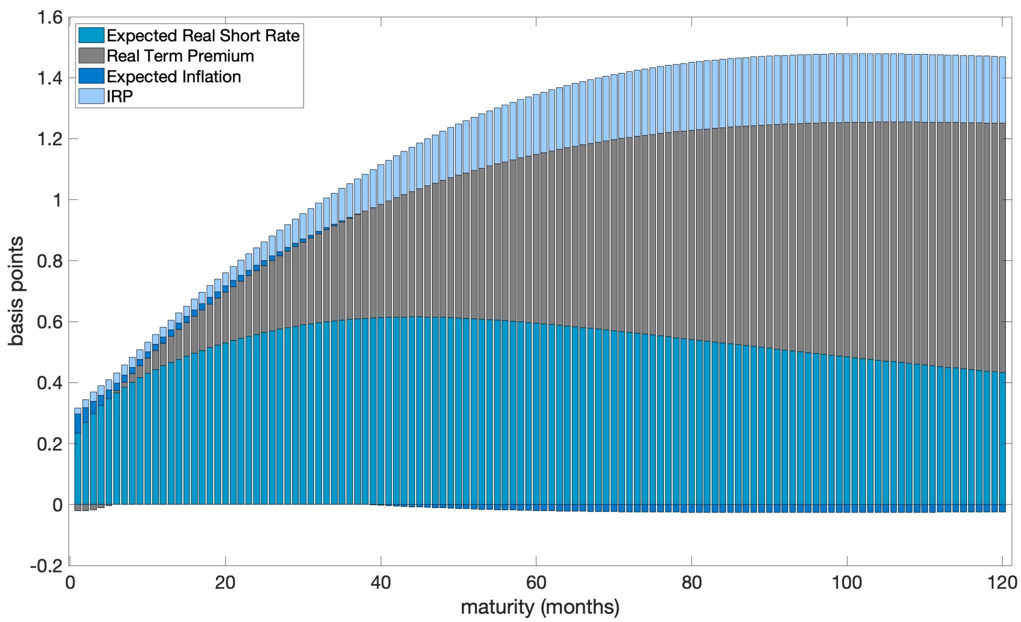


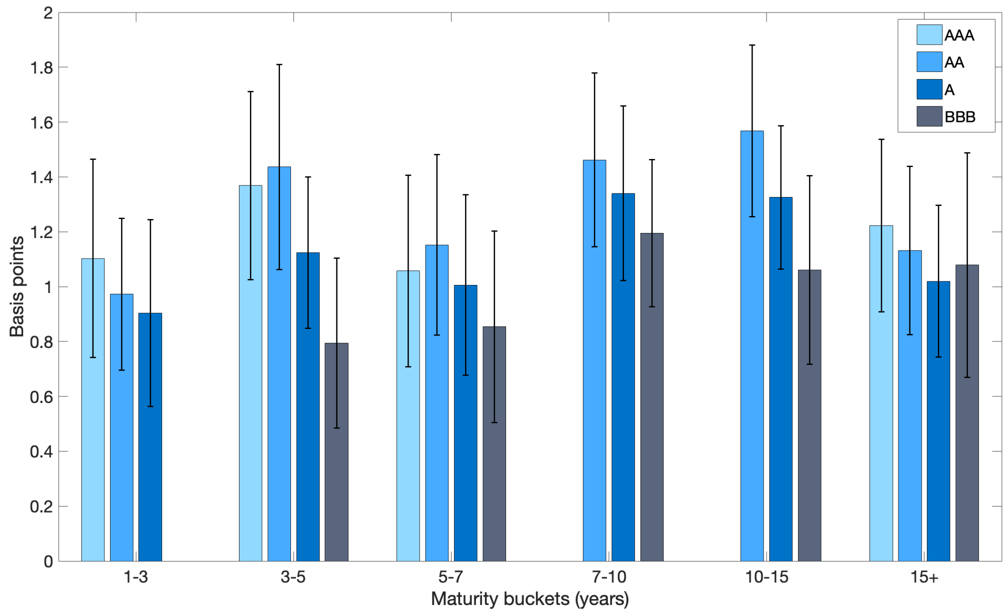
Figure A20: Reactions of the expected nominal short rates and the nominal term premium to the supply shock





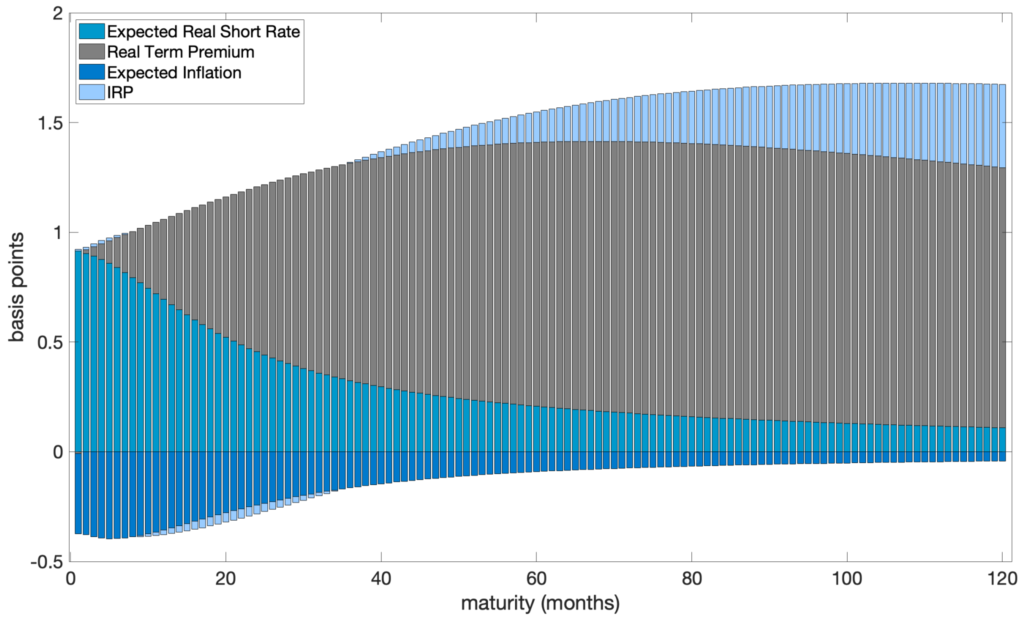
Note: Bars are the estimated $b^{(m)}$ coefficients from equation (2.3). Top panel: the dependent variables are the average expected nominal short rates and the nominal term premium. Bottom panel: the dependent variable is the average expected real short rate, expected inflation, the real term premium and inflation risk premium.

Figure A21: Reactions of corporate bond indices to the supply shock



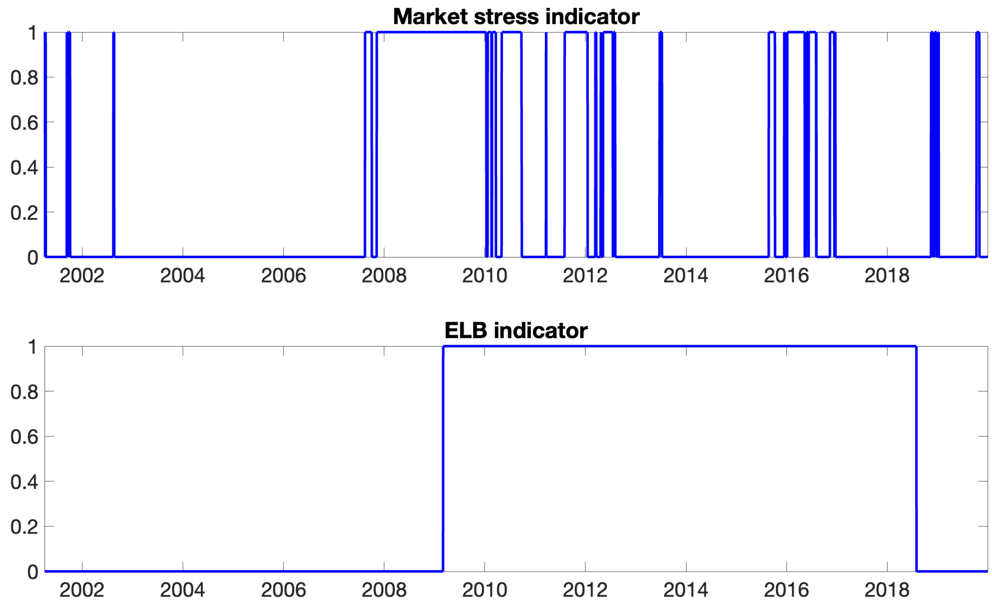
Note: Bars are the estimated $b^{(m)}$ coefficients from equation (2.3), when the dependent variables are AAA, AA, A and BBB rated corporate indices, with remaining maturities between 1-3, 3-5, 5-7,7-10, 10-15 and 15+ years, compiled by Refinitiv. Error bands are 95% (Newey-West, 10 lags) confidence intervals.

Figure A22: 10-year breakeven inflation rate decomposition at daily frequency to the supply shock, adjusted for liquidity effects



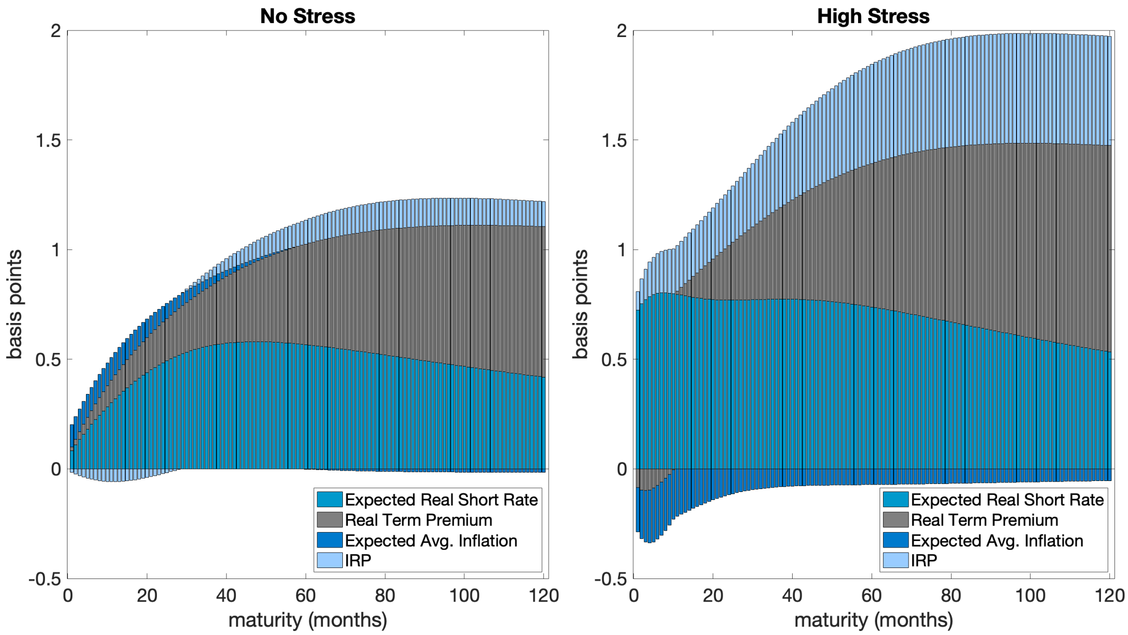
Note: Bars are the estimated $b^{(m)}$ coefficients from equation (2.3), where the dependent variables are the average expected real short rates , the expected inflation, the real term premium, and the inflation risk premium obtained via the ATSM. The state space of pricing factors in the ATSM is extended with a liquidity proxy: the inflation swap, breakeven inflation rate spread.

Figure A23: Times series of state indicators for the non-linear estimation



Note: Time series of the state indicator variables. The financial stress indicator takes the value one when the CISS index is above its 75th percentile. The ELB periods indicator takes the value one when the Bank of England bank rate is below 0.5%.

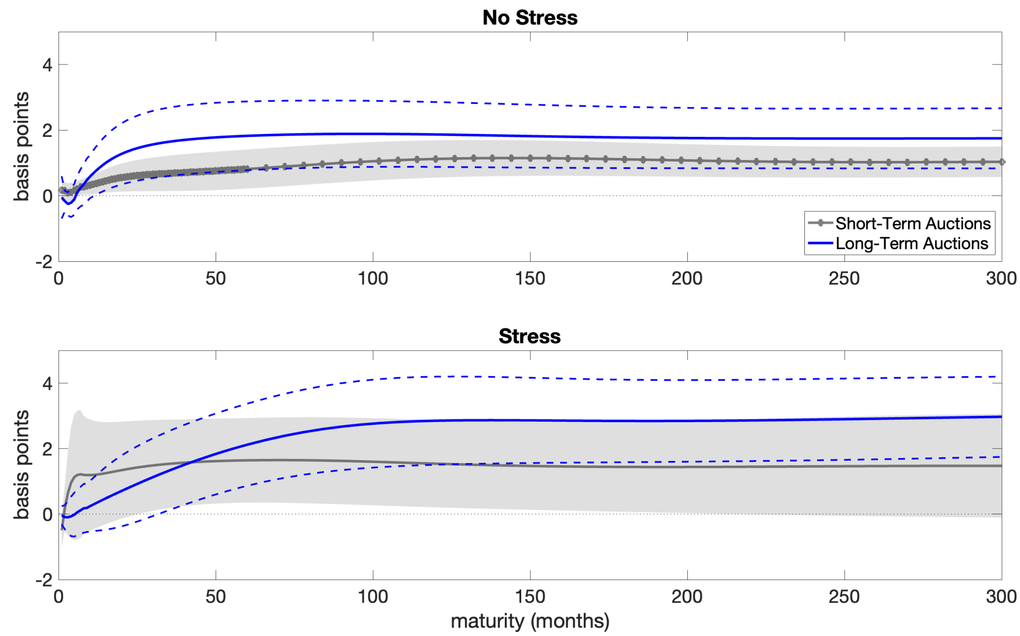
Figure A24: Reactions of yield components in normal times and in high-stress periods



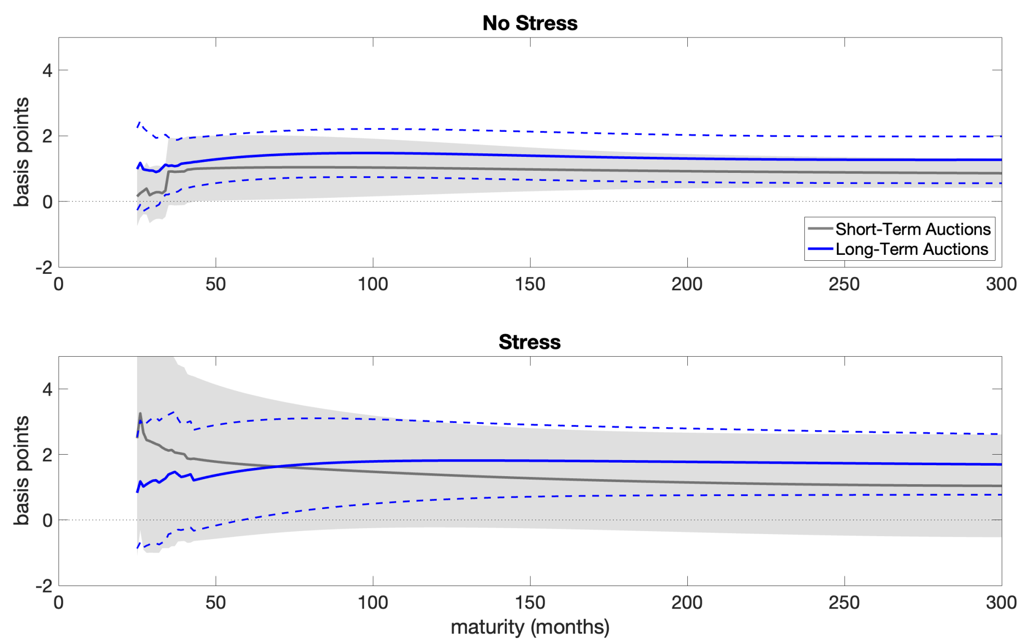
Note: Bars are estimated $b_1^{(m)}$ and $b_0^{(m)}$ coefficients from Equation 1.5 on yield components obtained via the ATSM. I_t indicates periods with the CISS index above its 75th percentile. Dependent variables are the average expected real short rate, the expected inflation, the real term premium, and the inflation risk premium.

Figure A25: Localization of the effect of the supply shock during market stress

Panel A: Nominal yields

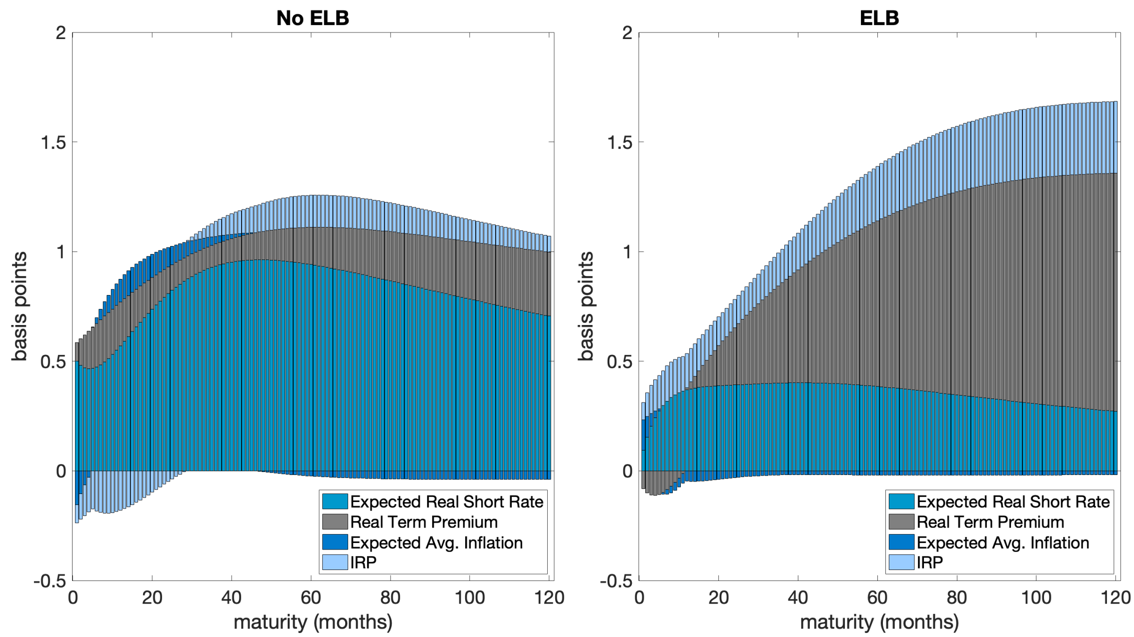


Panel B: Real yields



Note: Estimated $b_1^{(m)}$ and $b_0^{(m)}$ coefficients from Equation 1.5. Panel A. shows the result of nominal yields, Panel B. shows the results of real yields. The top chart of each panel shows the results during normal times, the bottom chart shows the results in market stress, characterized by the CISS index over its 75th percentile. Grey lines show the results when the announcements sample is restricted to the DMOs' short- and medium-maturity bucket (0-15 years), blue lines show the results when the announcements sample is restricted to the DMOs' long-maturity bucket (15+ years). Dashed lines and shaded areas are 95% (Newey-West, 10 lags) confidence intervals.

Figure A26: Reactions of yield components during normal times and in ELB periods



Note: Bars are estimated $b_1^{(m)}$ and $b_0^{(m)}$ coefficients from Equation 1.5 on yield components obtained via the ATSM. I_t indicates periods with the BoE Bank Rate below 0.5%. Dependent variables are the average expected real short rate, the expected inflation, the real term premium, and the inflation risk premium.

Table A1: Descriptive statistics of the high-frequency shocks

	Sample	N	Mean	Std.	Correlations				
					$S_t^{(Short)}$	$S_t^{(Med.)}$	$S_t^{(Long)}$	$S_t^{(U.long)}$	S_t
$S_t^{(Short)}$	24.11.09-31.12.19	238	-0.001	0.026					
$S_t^{(Med.)}$	24.11.09-31.12.19	238	0.004	0.053	0.271				
$S_t^{(Long)}$	15.05.01-31.12.19	360	0.001	0.125	0.293	0.949			
$S_t^{(U.long)}$	25.11.14-15.11.16	45	0.120	0.256	0.280	0.819	0.852		
S_t	15.05.01-18.02.20	360	0.001	0.132	0.311	0.959	0.999	0.853	

Table A3: Regression of the daily surprise on the intraday surprise S_t

	Daily surprises
Intraday surprise S_t	1.110
S.E.	(0.180)
P-value	0.000
R^2	0.111
N	345

Note: Dependent variable is the price surprise on announcement days with one-day event window. Independent variable is the intraday price surprise on announcement days with one-hour event window (S_t). Sample period: 15.05.2001-31.12.2019.

Table A2: Regression of the volume surprise at auctions on the high-frequency shock

	Intraday window	**Daily window**
	Panel (A): Announced volume	
Volume (bn £)	0.002	0.003
S.E.	(0.006)	(0.019)
P-value	0.709	0.883
R^2	0.000	0.000
N	314	314
	Panel (B): Surprise volume	
Surprise volume (bn £)	-0.150**	-0.082
S.E.	(0.059)	(0.202)
P-value	0.014	0.686
R^2	0.019	0.000
N	314	314

Note: Dependent variables are the high-frequency price movements in the event window on auction days. Left column: one-hour event window (S_t), right column: one-day event window. Panel (A): independent variable is the announcement volume. Panel (B) independent variable is the difference between the actual announced volume minus the implied average remaining auction size that is required to meet the DMOs' remit. Sample: 04.04.2006-31.12.2019.

Table A4: Reaction of credit default swaps to the Treasury supply shock

	2 years	**5 years**	**10 years**	**30 years**
Coeff.	-0.003	-0.004	-0.003	-0.003
S.E.	(0.004)	(0.003)	(0.002)	(0.002)
P-value	0.417	0.206	0.217	0.250

Note: Each column is a separate regression of credit default swaps written on 2-years, 5-years, 10-years and 30-years UK government bonds on the high-frequency supply shock, respectively. Sample period from 21.07.2008 to 31.12.2019.

Table A5: Reaction of BBB-AAA corporate bond spreads to the Treasury supply shock

	1-3 years	**3-5 years**	**5-10 years**	**15+ years**
Coeff.	1.229	-0.498	-0.239	0.011
S.E.	(1.399)	(0.412)	(0.189)	(0.149)
P-value	0.380	0.227	0.207	0.940

Note: Each column is a separate regression of UK corporate bond spreads with remaining maturities between 1-3 years, 3-5 years, 5-10 years, and 15+ years compiled by Refinitiv, on the high-frequency supply shock, respectively. Sample period from 31.03.2001 to 31.12.2019.

Table A6: Fit Diagnostics of the ATSM model on monthly data

	n = 36	n = 60	n = 84	n = 120
		Nominal Yield Pricing Errors		
mean	0.297	0.221	-0.257	0.189
std	0.049	0.034	0.050	0.037
ρ^y	0.952	0.796	0.936	0.791
ρ^{xr}	0.220	0.129	0.048	-0.021
		Real Yield Pricing Errors		
mean		-0.110	-0.200	0.046
std		0.160	0.127	0.118
ρ^y		0.796	0.936	0.791
ρ^{xr}		0.129	0.048	-0.021

Note: Time series properties of the ATSM pricing errors implied by the monthly estimation. "Mean" and "std" refers to the sample mean and standard deviation of yield pricing errors; ρ^y denotes first order sample autocorrelation coefficient of the yield pricing errors, ρ^{xr} denotes first order sample autocorrelation coefficient of the excess return pricing errors. Sample period: 1997:03 - 2019:12.

Table A7: Fit Diagnostics of the ATSM model on daily data

	n = 36	n = 60	n = 84	n = 120
Nominal Yield Pricing Errors				
mean	0.304	0.229	-0.248	-0.179
std	0.055	0.039	0.039	0.044
ρ^y	1.000	0.997	1.000	0.997
ρ^{xr}	0.369	0.264	0.205	0.154
Real Yield Pricing Errors				
mean		-0.108	-0.197	0.050
std		0.161	0.131	0.124
ρ^y		0.997	1.000	0.997
ρ^{xr}		0.264	0.205	0.154

Note: Time series properties of the ATSM pricing errors implied by the daily decomposition. "Mean" and "std" refers to the sample mean and standard deviation of yield pricing errors; ρ^y denotes first order sample autocorrelation coefficient of the yield pricing errors, ρ^{xr} denotes first order sample autocorrelation coefficient of the excess return pricing errors. Sample period: 1997:03:31 - 2019:12:31.

Appendix B An equilibrium term structure model of nominal and real yields with supply effects

B.1 The setup

The model is set in continuous time with two types of assets: nominal and inflation-linked (or real) zero-coupon bonds. These bonds have maturities τ in the interval $(0;T]$. An inflation-linked bond with maturity τ pays one unit of wealth at time $t+\tau$. Its time t price is denoted by $P_t^{R,(\tau)}$. A nominal bond with maturity τ pays one unit of currency at time $t+\tau$. Its time t price is denoted by $P_t^{N,(\tau)}$. The bond's spot yields are denoted by $y_t^{R,(\tau)}$ and $y_t^{N,(\tau)}$. They are related to the prices by:

$$y_t^{N,(\tau)} = -\frac{\log P_t^{N,(\tau)}}{\tau} \tag{B1a}$$

$$y_t^{R,(\tau)} = -\frac{\log P_t^{R,(\tau)}}{\tau} \tag{B1b}$$

The instantaneous risk-free real rate is denoted by r_t. It is defined as $\lim_{\tau\to 0} y_t^{R,(\tau)} = r_t$ and it follows the Ornstein-Uhlenbeck process

$$dr_t = \kappa_r(\bar{r} - r_t)dt + \sigma_r dB_{r,t} \tag{B2}$$

This rate can be interpreted as the return of a linear and instantaneously riskless production technology. Instantaneous inflation is also assumed to follow the Ornstein-Uhlenbeck process

$$d\pi_t = \kappa_\pi(\bar{\pi} - \pi_t)dt + \sigma_\pi dB_{\pi,t} \tag{B3}$$

where $\bar{r}, \kappa_r, \sigma_r, \bar{\pi}, \kappa_\pi, \sigma_\pi > 0$ are constants and $B_{r,t}$ and $B_{\pi,t}$ are independent Brownian motions. Parameters $\bar{r}$ and $\bar{\pi}$ are the long-run means of the processes, κ_r and κ_π are the mean-reverting parameters. The volatility parameters are σ_r and σ_π.

Bonds are issued by a government and traded by arbitrageurs and other investors that are not modelled explicitly. Following Greenwood and Vayanos (2014), we treat the supply and demand of the government and other investors as price inelastic.

The amount of bonds supplied by the government, net of other investors' demand, is exogenous. The supply of nominal bonds $s_t^{N,(\tau)}$ is given by a one factor model, as in Greenwood and Vayanos (2014). For simplicity, inflation-linked bond supply $s_t^{R,(\tau)}$ is fixed.

$$s_t^{N,(\tau)} = \zeta^N(\tau) + \theta^N(\tau)\beta_t \tag{B4a}$$

$$s_t^{R,(\tau)} = \zeta^R(\tau) \tag{B4b}$$

The functions $\zeta^N(\tau)$, $\zeta^R(\tau)$, and $\theta^N(\tau)$ are deterministic functions of the maturity of the bonds. The variable β_t is a stochastic nominal bond supply factor that follows the Ornstein-Uhlenbeck process. Its long-run mean is zero, similar to our high-frequency futures price shock series.

$$d\beta_t = -\kappa_\beta \beta_t dt + \sigma_\beta dB_{\beta,t} \tag{B5}$$

The function $\zeta^N(\tau)$ gives the average supply of bonds at maturity τ, while $\theta^N(\tau)$ measures the sensitivity of the nominal bond supply to the supply factor β_t. We assume that $\theta^N(\tau)$ has the following properties:

Assumption 1. *The functions $\theta^N(\tau)$ satisfies*
(i) $\int_0^T \theta^N(\tau) \geq 0$;
(ii) There exists $\tau^* \in [0;T)$ such that $\theta^N(\tau) < 0$ for $\tau < \tau^*$ and $\theta^N(\tau) > 0$ for $\tau > \tau^*$

The first point of the assumption ensures that an increase in β_t does not decrease the total value of bonds supplied to arbitrageurs. The second point allows the possibility that after an increase in β_t the supply of some shorter maturity bonds can decrease, while the total supply of bonds does not decrease. These assumptions ensure that an increase in β_t makes arbitrageurs' equilibrium portfolios more sensitive to inflation and duration risks. We assume $B_{\beta,t}$ is independent of $B_{r,t}$ and $B_{\pi,t}$. Greenwood and Vayanos (2014) considers a case where $B_{\beta,t}$ is correlated with $B_{r,t}$ which is reasonable given their empirical measure of supply. In our case, assuming independence corresponds more our to high-frequency supply shock in Section 1.3.

B.2 Arbitrageurs

Arbitrageurs are assumed to be mean-variance maximizers of their real wealth. They select their portfolio by solving:

$$\max_{\left\{x_t^{N,(\tau)},x_t^{R,(\tau)}\right\}_{\tau\in(0,T]}} E_t[dW_t] - \frac{a}{2}V_t[dW_t] \tag{B6}$$

W_t denotes arbitrageurs' real wealth, a is the coefficient of risk aversion. $x_t^{N,(\tau)}$ and $x_t^{R,(\tau)}$ are the units of wealth invested in the nominal bond and the inflation-linked bond with maturity of τ. Vayanos and Vila (2021) gives the interpretation for this setting that there are overlapping generations of arbitrageurs living over infinitesimal periods. A generation, born in t with wealth W_t, invests from t to $t+dt$ and then consumes and dies at $t+dt$. The corresponding budget constraint to the problem is given by:

$$\begin{aligned} dW_t = \int_0^T \left(x_t^{N,(\tau)} \frac{dP_t^{N,(\tau)}}{P_t^{N,(\tau)}} + x_t^{R,(\tau)} \frac{dP_t^{R,(\tau)}}{P_t^{R,(\tau)}} \right) d\tau - \Big(\int_0^T x_t^{N,(\tau)} d\tau \Big) \pi_t dt \\ + \Big(W_t - \int_0^T \left(x_t^{N,(\tau)} + x_t^{R,(\tau)} \right) d\tau \Big) r_t dt \end{aligned} \tag{B7}$$

The first expression is the return from investing in bonds, as $\int_0^T x_t^{N,(\tau)} d\tau$ and $\int_0^T x_t^{R,(\tau)} d\tau$ are the amount of wealth invested in nominal bonds and real bonds respectively. The second term $\left(\int_0^T x_t^{N,(\tau)} d\tau \right) \pi_t dt$ deflates the return from nominal bonds. Finally, the last expression is the return gained by investing the remaining wealth in the risk-free rate.

B.3 Solving the model

The model is solved by first conjecturing and later verifying that equilibrium spot rates are affine functions of the risk factors. The price of the nominal bond $P_t^{N,(\tau)}$, and the price of the inflation-linked bond $P_t^{R,(\tau)}$ are:

$$P_t^{N,(\tau)} = e^{-[A_r^N(\tau) r_t + A_\beta^N(\tau)\beta_t + A_\pi^N(\tau)\pi_t + C^N(\tau)]} \tag{B8a}$$

$$P_t^{R,(\tau)} = e^{-[A_r^R(\tau) r_t + A_\beta^R(\tau)\beta_t + C^R(\tau)]} \tag{B8b}$$

Lemma 1. *The dynamics of the nominal bond prices and the inflation-linked bond prices are given by*

$$\frac{dP_t^{N,(\tau)}}{P_t^{N,(\tau)}} = \mu_t^{N,(\tau)} dt - A_r^N(\tau)\sigma_r dB_{r,t} - A_\beta^N(\tau)\sigma_\beta dB_{\beta,t} - A_\pi^N(\tau)\sigma_\pi dB_{\pi,t} \tag{B9a}$$

$$\frac{dP_t^{R,(\tau)}}{P_t^{R,(\tau)}} = \mu_t^{R,(\tau)} dt - A_r^R(\tau)\sigma_r dB_{r,t} - A_\beta^R(\tau)\sigma_\beta dB_{\beta,t} \tag{B9b}$$

where instantaneous expected returns $\mu_t^{N,(\tau)}$ and $\mu_t^{R,(\tau)}$ are given by equations (B16a) and (B16b) in Section B.6.

Having derived the price dynamics of the two assets, we can substitute into the budget constraint (B7) and solve the arbitrageurs' optimization problem (B6). This is derived in Section B.6.

Lemma 2. *The first order conditions are given by:*

$$\mu_t^{N,(\tau)} - \pi_t - r_t = A_r^N(\tau)\lambda_{r,t} + A_\beta^N(\tau)\lambda_{\beta,t} + A_\pi^N(\tau)\lambda_{\pi,t} \tag{B10a}$$

$$\mu_t^{R,(\tau)} - r_t = A_r^R(\tau)\lambda_{r,t} + A_\beta^R(\tau)\lambda_{\beta,t} \tag{B10b}$$

where coefficients $\lambda_{i,t}$ are given by:

$$\lambda_{i,t} = a\sigma_i^2 \int_0^T x_t^{N,(\tau)} A_i^N(\tau) + x_t^{R,(\tau)} A_i^R(\tau) d\tau \qquad \textit{for } i = r, \beta \tag{B11}$$

$$\lambda_{\pi,t} = a\sigma_\pi^2 \int_0^T x_t^{N,(\tau)} A_\pi^N(\tau) d\tau \tag{B12}$$

Equations (B10a) and (B10b) are also the no-arbitrage conditions in the model. No arbitrage requires the existence of prices of each risk factor. Then, the expected excess return of each asset over the short rate is equal to the asset's sensitivity to the risk factors times the risk factor's price, summed across all risk factors. The coefficients $\lambda_{i,t}$ are the prices of the risk factors, measuring the expected excess return per unit of sensitivity to each factor. They are determined through equilibrium conditions. Note, that $\lambda_{i,t}$ are proportional to how sensitive the arbitrageurs' portfolio is to factor i. For example, the sensitivity of the portfolio to the short rate is $\int_0^T x_t^{N,(\tau)} A_r^N(\tau) + x_t^{R,(\tau)} A_r^R(\tau) d\tau$. As inflation-linked bonds shield investors from inflation, the inflation risk factor only loads on the nominal bond. Note, that even the real return of the nominal bond is sensitive to inflation risk.

B.4 Equilibrium term structures

In equilibrium, the supplied amount of bonds will be equal to the investment of the arbitrageurs.

$$x_t^{N,(\tau)} = s_t^{N,(\tau)} \tag{B13a}$$

$$x_t^{R,(\tau)} = s_t^{R,(\tau)} \tag{B13b}$$

We can use the market clearing (B13a), (B13b), the supply (B4a), (B4b) and (B16a), (B16b) to substitute into the first order conditions (B10a), (B10b). This yields two functions that are affine in the risk factors r_t, β_t and π_t, verifying our initial conjecture.

Setting linear terms in r_t, β_t and π_t equal to zero gives five ordinary differential equations (ODEs) in $A_r^N(\tau)$, $A_\beta^N(\tau)$, $A_\pi^N(\tau)$, $A_r^R(\tau)$ and $A_\beta^R(\tau)$. The solutions to these ODEs are stated in Theorem 1 and derived in Section B.6.

Theorem 1. *The nominal bond sensitivities $A^N_r(\tau)$, $A^N_\beta(\tau)$ and $A^N_\pi(\tau)$ are given by*

$$A^N_r(\tau) = \frac{1 - e^{-\kappa_r \tau}}{\kappa_r} \tag{B14a}$$

$$A^N_\beta(\tau) = \frac{Z_r}{\kappa_r}\left[\frac{1 - e^{-\hat{\kappa}_\beta \tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta \tau} - e^{-\kappa_r \tau}}{\kappa_r - \hat{\kappa}_\beta}\right] + \frac{Z_\pi}{\kappa_\pi}\left[\frac{1 - e^{-\hat{\kappa}_\beta \tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta \tau} - e^{-\kappa_\pi \tau}}{\kappa_\pi - \hat{\kappa}_\beta}\right] \tag{B14b}$$

$$A^N_\pi(\tau) = \frac{1 - e^{-\kappa_\pi \tau}}{\kappa_\pi} \tag{B14c}$$

The real bond sensitivities $A^R_r(\tau)$ and $A^R_\beta(\tau)$ are given by

$$A^R_r(\tau) = \frac{1 - e^{-\kappa_r \tau}}{\kappa_r} \tag{B15a}$$

$$A^R_\beta(\tau) = \frac{Z_r}{\kappa_r}\left[\frac{1 - e^{-\hat{\kappa}_\beta \tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta \tau} - e^{-\kappa_r \tau}}{\kappa_r - \hat{\kappa}_\beta}\right] \tag{B15b}$$

Where Z_r, Z_π and $\hat{\kappa}_\beta$ are given by equations (B19a), (B19b) and (B20) respectively in Section B.6. The functions $C^N(\tau)$ and $C^R(\tau)$ are given in Section B.6.

B.5 Analysis of supply effects

In our empirical analysis, we found that the high-frequency supply shock raises both nominal and real yields. In the model bond yields are given by:

$$\begin{aligned} y^{N,(\tau)}_t &= \frac{1}{\tau}\left[A^N_r(\tau) r_t + A^N_\beta(\tau)\beta_t + A^N_\pi(\tau)\pi_t + C^N(\tau)\right] \\ y^{R,(\tau)}_t &= \frac{1}{\tau}\left[A^R_r(\tau) r_t + A^R_\beta(\tau)\beta_t + C^R(\tau)\right] \end{aligned}$$

Therefore, we need to show that $\partial y^{N,(\tau)}_t / \partial \beta_t = A^N_\beta(\tau)/\tau$ and $\partial y^{R,(\tau)}_t / \partial \beta_t = A^R_\beta(\tau)/\tau$ are positive.

Proposition 1. *The effect of a shock to the supply factor β_t on nominal and real yields is positive.*

Proposition 2. *The effect of a shock to the supply factor on duration and inflation risk prices is positive.*

Proposition 3. *The effect of a shock to the supply factor on yields and risk prices increases with risk aversion.*

The proofs follow from Lemma A.1. and Lemma A.2. of Greenwood and Vayanos (2014), and we present it in Section B.6. The intuition is the following. An increase in the supply factor increases the amount of nominal bonds held by investors in equilibrium. This increases the sensitivity of their portfolio to duration and inflation risks, raising the prices of these factors. The increase in duration risk premium and inflation risk premium raises both nominal and inflation-linked bond yields. As inflation risk loads positively only on nominal bonds, the spread between the two type of bonds widen, consistent with our empirical findings.

B.6 Proof of theoretical results

Proof of Lemma 1. *Applying Ito's lemma to (B8a) and (B8b) using (B2), (B5) and (B3), we get:*

$$\begin{aligned}
\frac{dP_t^{N,(\tau)}}{P_t^{N,(\tau)}} =& \Big[\dot{A}_r^N(\tau)r_t + \dot{A}_\beta^N(\tau)\beta_t + \dot{A}_\pi^N(\tau)\pi_t + \dot{C}^N(\tau)\Big]dt - A_r^N(\tau)dr_t - A_\beta^N(\tau)d\beta_t - A_\pi^N(\tau)d\pi_t \\
&+ \frac{1}{2}\Big[(A_r^N(\tau)dr_t)^2 + (A_\beta^N(\tau)d\beta_t)^2 + (A_\pi^N(\tau)d\pi_t)^2\Big] \\
=& \Big[\dot{A}_r^N(\tau)r_t + \dot{A}_\beta^N(\tau)\beta_t + \dot{A}_\pi^N(\tau)\pi_t + \dot{C}^N(\tau)\Big]dt + A_r^N(\tau)[\kappa_r(r_t - \bar{r})dt - \sigma_r dB_{r,t}] \\
&+ A_\beta^N(\tau)[\kappa_\beta\beta_t dt - \sigma_\beta dB_{\beta,t}] + A_\pi^N(\tau)[(\pi_t - \bar{\pi})dt - \sigma_\pi dB_{\pi,t}] \\
&+ \frac{1}{2}\Big[(A_r^N(\tau))^2\sigma_r^2 + (A_\beta^N(\tau))^2\sigma_\beta^2 + (A_\pi^N(\tau))^2\sigma_\pi^2\Big] \\
\frac{dP_t^{R,(\tau)}}{P_t^{R,(\tau)}} =& \Big[\dot{A}_r^R(\tau)r_t + \dot{A}^R\beta(\tau)\beta_t + \dot{C}^R(\tau)\Big]dt - A_r^R(\tau)dr_t - A_\beta^R(\tau)d\beta_t \\
&+ \frac{1}{2}\Big[(A_r^R(\tau)dr_t)^2 + (A_\beta^R(\tau)d\beta_t)^2\Big] \\
=& \Big[\dot{A}_r^R(\tau)r_t + \dot{A}_\beta^R(\tau)\beta_t + \dot{A}_\pi^R(\tau)\pi_t + \dot{C}^R(\tau)\Big]dt A_r^R(\tau)[\kappa_r(r_t - \bar{r})dt - \sigma_r dB_{r,t}] \\
&+ A_\beta^R(\tau)[\kappa_\beta\beta_t dt - \sigma_\beta dB_{\beta,t}] + \frac{1}{2}\Big[(A_r^R(\tau))^2\sigma_r^2 + (A_\beta^R(\tau))^2\sigma_\beta^2\Big]
\end{aligned}$$

where we arrive to (B9a) and (B9b) if we define $\mu_t^{N,(\tau)}$ and $\mu_t^{R,(\tau)}$ as:

$$\begin{aligned}
\mu_t^{N,(\tau)} =& \dot{A}_r^N(\tau)r_t + \dot{A}_\beta^N(\tau)\beta_t + \dot{A}_\pi^N(\tau)\pi_t + \dot{C}^N(\tau) + A_r^N(\tau)\kappa_r(r_t - \bar{r}) + A_\beta^N(\tau)\kappa_\beta\beta_t \\
&+ A_\pi^N(\tau)\kappa_\pi(\pi_t - \bar{\pi}) + \frac{\sigma_r^2}{2}(A_r^N(\tau))^2 + \frac{\sigma_\beta^2}{2}(A_r^N\beta(\tau))^2 + \frac{\sigma_\pi^2}{2}(A_r^N\pi(\tau))^2
\end{aligned} \tag{B16a}$$

$$\begin{aligned}
\mu_t^{R,(\tau)} =& \dot{A}_r^R(\tau)r_t + \dot{A}_\beta^R(\tau)\beta_t + \dot{C}^R(\tau) + A_r^R(\tau)\kappa_r(r_t - \bar{r}) + A_\beta^R(\tau)\kappa_\beta\beta_t \\
&+ \frac{\sigma_r^2}{2}(A_r^N(\tau))^2 + \frac{\sigma_\beta^2}{2}(A_r^N\beta(\tau))^2
\end{aligned} \tag{B16b}$$

□

Proof of Lemma 2. *Simplifying terms yields (B9a) and (B9b). Substituting these into* *the budget constraint (B7):*

$$\begin{aligned} dW_t = & \left[W_t r_t + \int_0^T x_t^{N,(\tau)}[\mu_t^{N,(\tau)} - r_t - \pi_t] + x_t^{R,(\tau)}[\mu_t^{R,(\tau)} - r_t] d\tau \right] dt \\ & - \sigma_r \int_0^T x_t^{N,(\tau)} A_r^N(\tau) + x_t^{R,(\tau)} A_r^R(\tau) d\tau dB_{r,t} - \sigma_\beta \int_0^T x_t^{N,(\tau)} A_\beta^N(\tau) + x_t^{R,(\tau)} A_\beta^R(\tau) d\tau dB_{\beta,t} \\ & - \sigma_\pi \int_0^T x_t^{N,(\tau)} A_\pi^N(\tau) d\tau dB_{\pi,t} \end{aligned}$$

Then, the optimization problem can be written as:

$$\begin{aligned} \max_{\left\{x_t^{N,(\tau)}, x_t^{R,(\tau)}\right\}_{\tau\in(0,T]}} & \int_0^T x_t^{N,(\tau)}[\mu_t^{N,(\tau)} - r_t] + x_t^{R,(\tau)}[\mu_t^{R,(\tau)} - r_t] d\tau \\ & - \frac{a\sigma_r^2}{2}\left(\int_0^T x_t^{N,(\tau)} A_r^N(\tau) + x_t^{R,(\tau)} A_r^R(\tau) d\tau \right)^2 \\ & - \frac{a\sigma_\beta^2}{2}\left(\int_0^T x_t^{N,(\tau)} A_\beta^N(\tau) + x_t^{R,(\tau)} A_\beta^R(\tau) d\tau \right)^2 \\ & - \frac{a\sigma_\pi^2}{2}\left(\int_0^T x_t^{N,(\tau)} (A_\pi^N(\tau)\pi_t(2 + \sigma_\pi + 1) \right)^2 \end{aligned}$$

Point-wise maximization gives the first order conditions (B10a) and (B10b). $\square$

Proof of Theorem 1. *Setting linear terms to zero in (B10a) and (B10b) yields ordinary differential equations that we solve with the initial conditions* $A_r^N(0) = A_r^R(0) = A_\beta^N(0) = A_\beta^R(0) = A_\pi^N(0) = C^N(0) = C^R(0) = 0$*. Identifying terms in* r_t *gives:*

$$\dot{A}_r^N(\tau) + \kappa_r A_r^N(\tau) - 1 = 0 \tag{B17a}$$
$$\dot{A}_r^R(\tau) + \kappa_r A_r^R(\tau) - 1 = 0 \tag{B17b}$$

Identifying terms in β_t *gives:*

$$\dot{A}_\beta^N(\tau) + \hat{\kappa}_\beta A_\beta^N(\tau) = Z_r A_r^N(\tau) + Z_\pi A_\pi^N(\tau) \tag{B18a}$$
$$\dot{A}_\beta^R(\tau) + \hat{\kappa}_\beta A_\beta^R(\tau) = Z_r A_r^R(\tau) \tag{B18b}$$

where

$$Z_r = a\sigma_r^2 \int_0^T \theta^N(\tau) A_r^N(\tau) d\tau \tag{B19a}$$

$$Z_\pi = a\sigma_\pi^2 \int_0^T \theta^N(\tau) A_\pi^N(\tau) d\tau \tag{B19b}$$

And $\hat{\kappa}_\beta$ solves

$$\hat{\kappa}_\beta = \kappa_\beta - a\sigma_\beta^2 \int_0^T \theta^N(\tau) A_\beta^N(\tau) d\tau \tag{B20}$$

Equilibria in the model exist if the arbitrageurs' risk-aversion coefficient a is below a threshold $\bar{a} > 0$. As in Greenwood and Vayanos (2014), we focus on that case, and select the equilibrium corresponding to the largest solution for $\hat{\kappa}$. For more details see Greenwood and Vayanos (2014).

Identifying terms in π leads to:

$$\dot{A}_\pi^N(\tau) + \kappa_\pi A_\pi^N(\tau) - 1 = 0 \tag{B21}$$

The solutions to is (B17a) and (B17b) are (B14a) and (B15a). The solutions to (B18a) and (B18b) are given by (C24) and (B15b), with Z_r, Z_π and $\hat{\kappa}_\beta$ are given by equations (B19a), (B19b) and (B20). The solution to (B21) is (B14c). Identifying constant terms in (B10a) yields

$$\begin{aligned}
&\dot{C}^N(\tau) - A_r^N(\tau)\kappa_r\bar{r} - A_\pi^N(\tau)\kappa_\pi\bar{r} + \frac{\sigma_r^2}{2}(A_r^N(\tau))^2 + \frac{\sigma_\beta^2}{2}(A_\beta^N(\tau))^2 + \frac{\sigma_\pi^2}{2}(A_\pi^N(\tau))^2 \\
&\quad = a\sigma_r^2 A_r^N(\tau) \int_0^T \zeta^N(\tau) A_r^N(\tau) + \zeta^R(\tau) A_r^R(\tau) d\tau \\
&\quad + a\sigma_\beta^2 A_\beta^N(\tau) \int_0^T \zeta^N(\tau) A_\beta^N(\tau) + \zeta^R(\tau) A_\beta^R(\tau) d\tau \\
&\quad + a\sigma_\pi^2 A_\pi^N(\tau) \int_0^T \zeta^N(\tau) A_\pi^N(\tau)
\end{aligned}$$

The solution to $\dot{C}^N(\tau)$ is

$$\begin{aligned}
C^N(\tau) =& \hat{Z}_r \int_0^\tau A_r^N(\tau')d\tau' + \hat{Z}_\beta \int_0^\tau A_\beta^N(\tau')d\tau' + \hat{Z}_\pi \int_0^\tau A_\pi^N(\tau')d\tau' \\
& - \frac{\sigma_r^2}{2}\int_0^\tau (A_r^N(\tau'))^2 d\tau' - \frac{\sigma_\beta^2}{2}\int_0^\tau (A_\beta^N(\tau'))^2 d\tau' - \frac{\sigma_\pi^2}{2}\int_0^\tau (A_\pi^N(\tau'))^2 d\tau'
\end{aligned} \tag{B22}$$

with $\hat{Z}_r$, $\hat{Z}_\beta$ and $\hat{Z}_\pi$ given by

$$\begin{aligned}
\hat{Z}_r &= \kappa_r\bar{r} + a\sigma_r^2 \int_0^T \zeta^N(\tau) A_r^N(\tau) + \zeta^R(\tau) A_r^R(\tau) d\tau \\
\hat{Z}_\beta &= a\sigma_\beta^2 \int_0^T \zeta^N(\tau) A_\beta^N(\tau) + \zeta^R(\tau) A_\beta^R(\tau) d\tau \\
\hat{Z}_\pi &= \kappa_\pi\bar{\pi} + a\sigma_\pi^2 \int_0^T \zeta^N(\tau) A_\pi^N(\tau)(\tau) d\tau
\end{aligned}$$

Identifying constant terms in (B10b) *yields*

$$\dot{C}^R(\tau) = A_r^R(\tau)\hat{Z}_r + A_\beta^R(\tau)\hat{Z}_\beta - \frac{\sigma_r^2}{2}(A_r^R(\tau))^2 - \frac{\sigma_\beta^2}{2}(A_\beta^R(\tau))^2$$

with the solution

$$\begin{aligned} C^R(\tau) &= \hat{Z}_r \int_0^\tau A_r^R(\tau')d\tau' + \hat{Z}_\beta \int_0^\tau A_\beta^R(\tau')d\tau' \\ &- \frac{\sigma_r^2}{2}\int_0^\tau (A_r^R(\tau'))^2 d\tau' - \frac{\sigma_\beta^2}{2}\int_0^\tau (A_\beta^R(\tau'))^2 d\tau' \end{aligned} \tag{B23}$$

□

Proof of Proposition 1. *The effect of a shock to the supply factor to nominal yields is given by:*

$$\frac{\partial y_t^{N,(\tau)}}{\partial \beta_t} = \frac{A_\beta^N(\tau)}{\tau} = \frac{Z_r}{\tau\kappa_r}\Big[\frac{1-e^{-\hat{\kappa}_\beta\tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta\tau} - e^{-\kappa_r\tau}}{\kappa_r - \hat{\kappa}_\beta}\Big] + \frac{Z_\pi}{\tau\kappa_\pi}\Big[\frac{1-e^{-\hat{\kappa}_\beta\tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta\tau} - e^{-\kappa_\pi\tau}}{\kappa_\pi - \hat{\kappa}_\beta}\Big]$$

First, we show that Z_r and Z_π are positive. Then we show that the expression in the brackets are positive.

From Equation (B17a) (B21) $A_r^N(\tau)$ *and* $A_\pi^N(\tau)$ *are positive and they are increasing as:*

$$\frac{\partial A_r^N(\tau)}{\partial \tau} = e^{-k_r\tau} > 0$$

Then, we show that $\int_0^T A_r^N(\tau)\theta^N(\tau)d\tau > 0$ as it can be written an:

$$\begin{aligned} \int_0^T A_r^N(\tau)\theta^N(\tau)d\tau &= \int_0^{\tau^*} A_r^N(\tau)\theta^N(\tau)d\tau + \int_{\tau^*}^T A_r^N(\tau)\theta^N(\tau)d\tau \\ &> A_r^N(\tau^*)\int_0^{\tau^*} \theta^N(\tau)d\tau + A_r^N(\tau^*)\int_{\tau^*}^T \theta^N(\tau)d\tau \\ &= A_r^N(\tau^*)\int_0^T \theta^N(\tau)d\tau \geq 0, \end{aligned}$$

where we used Part (ii) of Assumption 1 in the second step and Part (i) if Assumption 1 in the third step. Therefore $Z_r = a\sigma_r^2\int_0^T A_r^N(\tau)\theta^N(\tau)d\tau > 0$ and analogously for Z_π. Then, we can write $A_\beta^N(\tau)$as:

$$A_\beta^N(\tau) = Z_r\int_0^\tau \frac{1-e^{-\kappa_r\hat{\tau}}}{\kappa_r}e^{-\hat{\kappa}_\beta(\tau-\hat{\tau})}d\hat{\tau} + Z_\pi\int_0^\tau \frac{1-e^{-\kappa_\pi\hat{\tau}}}{\kappa_\pi}e^{-\hat{\kappa}_\beta(\tau-\hat{\tau})}d\hat{\tau}$$

which is positive as Z_r and Z_π are both positive. The proof for $A_\beta^R(\tau)$ is analogous.
□

Proof of Proposition 2. *The effect of the supply factor on duration risk is given by:*

$$\frac{\partial \lambda_{r,t}}{\partial \beta_t} = a\sigma_r^2 \int_0^T \theta^N(\tau) A_r^N(\tau) d\tau$$

which is positive as $\int_0^T \theta^N(\tau) A_r^N(\tau) d\tau > 0$, as shown in Proof of Proposition 1. The effect of the supply factor on interest rate risk is analogous.

□

Proof of Proposition 3. *This can be seen immediately from* (B11), (B12), (B18a) *and* (B18b) *as Z_r and Z_π both increase in a.*

□

Appendix C The ATSM of Abrahams et al. (2016)

The $K \times 1$ vector of pricing factors follows an autoregression under the physical measure ($\mathbb{P}$):

$$X_{t+1} - \mu_X = \Phi(X_t - \mu_X) + \nu_{t+1}, \qquad \nu_{t+1} \sim i.i.d.N(0_{K\times 1}, \Sigma) \tag{C24}$$

The stochastic discount factor is written as:

$$M_{t+1} = exp\Big(- r_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-1/2}\nu_{t+1}\Big)$$

where r_t is the nominal short rate and λ_t is $K \times 1$ the vector of risk prices. These are related to the pricing factors as: $\lambda_t = \Sigma^{-1/2}(\lambda_0 + \lambda_1 X_t)$. The short rate follows $r_t = \delta_0 + \delta_1 X_t$. Abrahams et al. (2016) define the parameters of the pricing factor dynamics under the risk neutral measure ($\mathbb{Q}$) as:

$$\tilde{\mu} = (I_K - \Phi)\mu_X - \lambda_0$$
$$\tilde{\Phi} = \Phi - \lambda_1$$

The model assumes that bond yields are affine functions of the state variables, which are assumed to be observable. Therefore, under the pricing measure, the log price, $P_t^{(\tau)}$, of a nominal zero-coupon risk-free bond with remaining time to maturity τ follows $\log P_t^{(\tau)} = A_\tau + B_\tau' X_t$. The log price of an inflation-linked bond follows similarly $\log P_{t,R}^{(\tau)} = A_{\tau,R} + B_{\tau,R}' X_t$. The price of such a bond also satisfies:

$$\begin{aligned} \log P_{t,R}^{(\tau)} &= E_t^{\mathbb{Q}}\Big[\exp(-r_t - \cdots - r_{t+\tau-1})\frac{Q_{t+\tau}}{Q_t}\Big] \\ &= E_t^{\mathbb{Q}}\Big[\exp(-r_t - \cdots - r_{t+\tau-1} + \pi_{t+1} + \cdots + \pi_{t+\tau})\Big] \end{aligned} \tag{C25}$$

where $E^{\mathbb{Q}}$ is the expectation operator under the risk neutral measure. Q_t is the price index at time t, and $\pi_t = \ln(Q_t/Q_{t-1})$ is the one period log inflation, related to the pricing factors as $\pi_t = \pi_0 + \pi_1' X_t$.

The system of recursive linear restrictions of the bond pricing parameters can be obtained once no-arbitrage is imposed (see Ang and Piazzesi (2003)):

$$\begin{aligned} A_\tau &= A_{\tau-1} + B_{\tau-1}'\tilde{\mu} + \frac{1}{2}B_{\tau-1}'\Sigma B_{\tau-1} - \delta_0 \\ B_\tau' &= B_{\tau-1}'\tilde{\Phi} - \delta_1' \\ A_0 &= 0, \quad B_0 = 0_{K\times 0} \end{aligned}$$

Risk neutral counterparts are obtained by setting the price of risk parameters λ_0 and λ_1 to zero. Then, the pricing recursion modifies to:

$$\begin{aligned} \tilde{A}_\tau &= \tilde{A}_{\tau-1} + \tilde{B}_{\tau-1}'(I_K - \Phi)\mu - \delta_0 \\ \tilde{B}_\tau' &= \tilde{B}_{\tau-1}'\Phi - \delta_1' \\ \tilde{A}_0 &= 0, \quad \tilde{B}_0 = 0_{K\times 0} \end{aligned}$$

The inflation-linked bond recursion can be obtained by writing Equation C25 in terms of an inflation-linked bond purchased one period ahead. Taking logs, calculating the expectation, and matching coefficients in the expression for the log bond price yields the recursion:[18]

$$\begin{aligned} A_{\tau,R} &= A_{\tau-1,R} + {B_{\tau-1,R}^{\pi}}'\tilde{\mu} + \frac{1}{2}{B_{\tau-1,R}^{\pi}}'\Sigma B_{\tau-1,R}^{\pi} - \delta_{0,R} \\ {B_{\tau,R}^{\pi}}' &= {B_{\tau-1,R}^{\pi}}'\tilde{\Phi} - \delta_1' \\ A_{0,R} &= 0, \quad B_{0,R} = 0_{K\times 0} \end{aligned}$$

where $\delta_{0,R} = \delta_0 - \pi_0$ and $B_{\tau,R}^{\pi} = B_{\tau,R} + \pi_1$. Similar to nominal bonds, the risk neutral counterparts are given by:

$$\begin{aligned} \tilde{A}_{\tau,R} &= \tilde{A}_{\tau-1,R} + {\tilde{B}_{\tau-1,R}^{\pi}}'(I_K - \Phi)\mu - \delta_{0,R} \\ {\tilde{B}_{\tau,R}^{\pi}}' &= {\tilde{B}_{\tau-1,R}^{\pi}}'\Phi - \delta_1' \\ \tilde{A}_{0,R} &= 0, \quad \tilde{B}_{0,R} = 0_{K\times 0} \end{aligned}$$

where $\tilde{B}_{\tau,R}^{\pi} = \tilde{B}_{\tau,R} + \pi_1$.

The elements of yields can be obtained as the following. We use the risk adjusted counterparts of the pricing recursion coefficients $\tilde{A}.$ and $\tilde{B}.$ to calculate the risk adjusted fitted yields. These yields are interpreted as the time t expectation of average future short rates over the next τ periods:

$$\frac{1}{\tau}\sum_{i=0}^{\tau} E_t r_{t+i} = -\frac{1}{\tau}\Big[\tilde{A}_\tau + \tilde{B}_\tau' X_t\Big], \qquad \frac{1}{\tau}\sum_{i=0}^{\tau} E_t r_{t+i,R} = -\frac{1}{\tau}\Big[\tilde{A}_{\tau,R} + \tilde{B}_{\tau,R}' X_t\Big]$$

The difference between the nominal and the real expected short rates is the average ex-

[18] For details see Abrahams et al. (2016)

pected future inflation over the next τ periods:

$$\frac{1}{\tau}\sum_{i=0}^{\tau} E_t(\pi_{t+i}) = -\frac{1}{\tau}\Big[\tilde{A}_\tau + \tilde{B}'_\tau X_t\Big] - \bigg(-\frac{1}{\tau}\Big[\tilde{A}_{\tau,R} + \tilde{B}'_{\tau,R} X_t\Big]\bigg)$$

Term premiums can be obtained by subtracting the expectation component from the fitted yields:

$$TP_t^{(\tau)} = -\frac{1}{\tau}\Big[A_\tau + B'_\tau X_t\Big] - \bigg(-\frac{1}{\tau}\Big[\tilde{A}_\tau + \tilde{B}'_\tau X_t\Big]\bigg)$$

for nominal term premium, and similarly for the real term premium $TP_{t,R}^{(\tau)}$. The inflation risk premium is obtained as the difference between the fitted breakeven inflation and the inflation expectation:

$$IRP_t^{(\tau)} = -\frac{1}{\tau}\Big[A_\tau + B'_\tau X_t\Big] - \bigg(-\frac{1}{\tau}\Big[A_{\tau,R} + B'_{\tau,R} X_t\Big]\bigg) - \frac{1}{\tau}\sum_{i=0}^{\tau} E_t(\pi_{t+i})$$

The model parameters are estimated following Adrian et al. (2013) and Abrahams et al. (2016). First, we estimate the risk neutral dynamics of the pricing factors by an autoregression. Then, we estimate the sensitivities of bond excess returns to current and past values of the pricing factors. Lastly, we do cross-sectional regressions of excess return sensitivities to lagged pricing factors onto excess return sensitivities to current pricing factors.

State variables are extracted principal components from yields. Following Abrahams et al. (2016), we extract three principal components from month-end zero coupon nominal yields and two principal components from orthogonalized real yields.[19] The factors are shown in Figure A14 in Appendix A. The short rate is the Bank of England's official bank rate, inflation is calculated with the monthly RPI series from the Office of National Statistics. We calculate excess returns on eleven nominal maturities of $\tau = 6, 12, 24, \ldots, 120$ months and eight real maturities of $\tau = 60, 66, 72, \ldots, 120$ months. For maturities that the Bank of England does not publish data, we interpolate it with cubic spline method. We do the decomposition up to 10 years, as the fit of the model deteriorates at higher maturities. In the baseline model, we do not account for the relative liquidity of inflation-linked bonds due to the lack of good liquidity proxies. Nevertheless, in the paper we also present a robustness exercise where we try also to take this into account.

The ATSM model parameters are estimated on monthly data from 03.1997 to 12.2019. Our goal is to decompose yields at the daily frequency, so we follow Adrian et al. (2013) and use the monthly model parameters on factors extracted from daily yield curve data from 31.03.1997 to 31.12.2019 to obtain the yield decomposition at the daily frequency. Model fit diagnostics are summarized in Tables A6 and A7 in the Appendix. The fit of the model

[19] Orthogonalized yields are obtained by purging inflation-linked yields from the nominal principal components, to reduce collinearity among the pricing factors.

at 10-years maturity is displayed in Figure A15 at the monthly frequency, and in Figure A16 at the daily frequency in Appendix A. The mean pricing errors are somewhat larger than in Abrahams et al. (2016), while the standard deviations are similar. Consistent with the relationship between yield and return pricing errors, we find a strong serial correlation in yield pricing errors but not in return pricing errors (see Adrian et al. (2013) for more details). The decomposed 10-year expected nominal short rate and nominal term premium are displayed in Figure A18. The decomposition of the 10-year breakeven inflation rate into inflation expectations and inflation risk premium are displayed in Figure A19. The trends in the estimated 10-year nominal term premium are in line with the estimates of Bianchi et al. (2009), Malik and Meldrum (2016), Abrahams et al. (2016) and Kaminska et al. (2018). The series fluctuates close to 1% at the beginning of the sample and rises after the Global Financial Crisis. It moves into negative territory towards the end of the sample, during the asset purchase programs of the Bank of England. Expected inflation and inflation risk premium are close to the estimates of Abrahams et al. (2016), Kaminska et al. (2018) and Bekaert and Ermolov (2023). Expected 10-year average inflation is rather stable, fluctuating close to 3%. The inflation risk premium shows more variation. It stays mostly within the 0-1% range but drops into negative territory in the early 2000s, the Global Financial Crisis, and the European debt crisis.

Chapter 2

Demand Shocks for Public Debt in the Eurozone*

2.1 Introduction

Government bonds of developed countries are usually considered the safest and most liquid assets. They have a key role in savers' portfolio decisions, investors' risk-management activities and banks' repo operations. Consequently, private sector demand for the largest European economies sovereign bonds was initially stable following the introduction of the single currency. Volatility was low, with yields exhibiting strong co-movements, up until the Global Financial Crisis.[1] The crisis broke this pattern and credit risk became an increasing problem in many euro area countries. Countries on the periphery of the euro area have seen private sector demand for their bonds drying up. The European Central Bank (ECB) took up an active role and entered the buy side of the market, in order to first facilitate liquidity with the Securities Markets Programme (SMP), and later to influence prices with the Public Sector Purchase Programme (PSPP). The aim of this latter program was to decrease long-term interest rates, by purchasing large quantities of long-term euro area public debt securities.

These shifts in the demand for public debt provide the motivation for our study. Our goal, in a broader sense, is to identify unexpected demand shocks for German and Italian government bonds and analyze how financial markets react to these shocks. The choice of these two countries stems from the fact that a shift in the demand for German and Italian debt is associated with markedly different sentiments in financial markets. German

* This chapter is based on Lengyel and Giuliodori (2022).

[1] See Figure 2.1.

government bonds are some of the safest traded securities, experiencing large inflows during times of high financial stress. Consequently, investors attach large “safety" and “liquidity" premiums to them (Ejsing and Sihvonen (2009) and Klingler and Sundaresan (2023)). Investors’ attitude towards Italian bonds, on the other hand, is substantially different. Italy has one of the highest public debt-to-GDP ratios in the euro area. Its sustainability is a topic of ongoing debate and investors require a substantial risk premium for holding Italian bonds.

The first contribution of our paper is to identify unexpected shifts in the demand for public debt in these two countries, by exploiting the institutional setup of government bond auctions. The identification strategy follows Gorodnichenko and Ray (2017), who proposed it in the context of US auctions. These auctions are important events where institutional investors can accommodate part of their security needs. Therefore, auction results can give hints about changes in market demand for these securities.[2] Furthermore, the prevailing demand at an auction shows investors’ perception of the health of a country’s economy and creditworthiness of its government.[3] Importantly, the timing and the setup of these events are such that they allow to capture price movements that can mainly be attributed to demand-side factors. Debt management agencies disclose information about the supplied securities and its quantity well in advance of an auction. On the auction day, therefore, investors are already well informed of the supply side of the market. The demand side is, however, unknown up until the agency releases the results of the auction. Hence, when these results are published, investors receive new information solely on the demand conditions. By looking at high-frequency price movements of government bond futures contracts around the first releases of the results, we can isolate price variations that are mainly attributed to the strength of the demand side.

The second contribution of the paper is on the estimation of the effects of the identified Treasury demand shocks on domestic bonds and the yield curve. Our estimates show that a one-standard-deviation demand shock in Germany decreases home yields by around 1.6 basis points. In Italy a similar shock has an effect of 3.3 basis points. These effects are found to last up to 30 trading days. Using our estimates we provide back-of-the-envelope calculations of the effects of the PSPP on German and Italian yields which are in line with the findings of the existing literature (Altavilla et al. (2015) and De Santis (2020)). We

[2] News agency Bloomberg wrote on the 27th December 2018: *“The Treasury in Rome plans to auction as much as 5 billion euros of debt Friday, including benchmark five-year and 10-year securities. (. . .) The results will provide an indication of the underlying demand for Italian bonds next year."*

[3] As news agency Reuters wrote on the 30th July 2018: *“Italy’s scheduled bond auctions, which included the sale of a new 10-year benchmark, was seen as a test of investor risk appetite amid political tensions in the euro zone’s third largest economy."*

also investigate the effects of a location-specific demand shock on the Treasury yield curve, distinguishing between short-term and long-term Treasury demand shocks. We show that in Germany the shocks have local effects, in the sense that nearby (i.e. similar maturity) yields respond stronger. Our findings for Germany are in line with Gorodnichenko and Ray (2017) who find local effects of demand shocks for the US Treasury yield curve. Our results are also consistent with studies bringing evidence of bond market segmentation, such as Boermans and Vermeulen (2018). In Italy, on the other hand, a positive demand shock always decreases short-term interest rates more, regardless of the shock location. We find that German and Italian demand shocks have spillover effects on the Treasury bond yields of other euro area countries. While the German shocks have more sizable effects on France and the Netherlands, the responses to the Italian demand shocks are mostly centered on the Spanish Treasury yields. We also show that Treasury demand shocks lead to reactions of the corporate bond markets. More specifically, euro area corporate bond yields drop in response to German demand shocks, whereas they are rather unaffected by Italian demand shocks.

Of particular interest are the results of the spillover effects of demand shocks on equity markets and Treasury CDS spreads. The seemingly similar information (increased demand for government debt) has vastly different effects between the two countries. More specifically the main euro area stock indices drop following German demand shocks, while Treasury CDS spreads increase. On the other hand, Italian demand shocks lead to positive responses of equity prices and decreases in CDS spreads. We reconcile these results as follows. The information about increased demand for bonds alters investors' risk preferences and expectations in two different ways. A sudden demand increase for German bonds is associated with financial markets turning to a "flight-to-safety" mode. Willingness to take risk decreases and investors re-balance their portfolios from equities to bonds. On the other hand, the market movements associated with a positive Italian demand shock are the result of a higher risk appetite and a better outlook for the Italian economy leading to positive effects on the Italian stock market and decreases in Treasury CDS spreads.

Finally, our last contribution is to test for the presence of state dependence. We show that during times of high financial stress the estimated responses documented above tend to be larger relative to periods of low financial stress. We also study if positive and negative demand shocks for Treasury bonds have asymmetric effects. We find that German responses are rather symmetric, while in the case of Italy the baseline results on Treasury markets seem to be mainly driven by positive demand shocks. This is particularly the case for the Treasury spillover effects to other euro area countries. Similarly, we find that both the equity and corporate bond indices react significantly only to positive Italian demand shocks.

The remainder of the paper is organized as follows. Section 2 summarizes the literature related to our study. Section 3 provides more details on the institutional setup of Treasury auctions in Germany and Italy. Section 4 describes the data used in the analysis. The high-frequency identification and the construction of the demand shocks are explained in Section 5. The main empirical results are shown in Section 6, which also contains an extensive robustness analysis. Section 7 examines the presence of state dependency and asymmetries in the effects. Finally, Section 8 concludes.

2.2 Literature review

This paper is related to several lines of existing research. On the one hand, it is connected to the literature that analyses Treasury market behavior in response to news. Macroeconomic announcements are typically found to cause large intraday movements in prices, traded volumes and bid-ask spreads in Treasury and foreign exchange futures markets (Ederington and Lee (1993) and Fleming and Remolona (1999)). Balduzzi et al. (2001) show that some releases do not affect the Treasury yield curve uniformly but have stronger effects on specific maturity segments. Fleming and Remolona (1997) find that not only macroeconomic announcements, but also monetary policy announcements and Treasury auction results are important drivers of bond prices. In line with these results, Huang et al. (2002) also find that following the release of auction results, trading activity increases in US government bond markets. This leads to the second line of research our paper relates to, that is, the literature examining market behavior around Treasury auctions.

Although Treasury auctions convey a substantial amount of information to financial markets, it is still an under-researched area. There is a long-standing literature documenting systematic price differential between the primary and the secondary Treasury markets. It suggests that the secondary market learns from the outcome of the auctions (see Bikhchandani and Huang (1989) and Cammack (1991), Goldreich (2007)). Others document predictable price and liquidity patterns in the secondary market around auctions (see Lou et al. (2013) and Beetsma et al. (2016), Fleming and Liu (2016)). Other papers show that auction effects have spillover effects internationally, e.g. Beetsma et al. (2018) and Eisl et al. (2019). Some recent papers look at auction results to identify changes in the demand for Treasuries. Klingler and Sundaresan (2023) analyze the liquidity premium and "safe-haven" status of Treasuries, Dobrev (2019) investigates how demand for US Treasuries has changed in recent years and Fuhrer and Giese (2021) study how demand shifts transmit across the yield curve in the UK. Gorodnichenko and Ray (2017) identify demand shocks by looking at high-frequency price changes around US auctions and analyze the effects on the yield curve and transmission into other markets.

Our paper also relates to the literature studying Treasury market segmentation in the maturity space, particularly in relation to large scale asset purchases by central banks. This is closely connected to the preferred-habitat view of interest rates, proposed by Culbertson (1957) and Modigliani and Sutch (1966) and recently picked up by Vayanos and Vila (2021). Empirical support for the theory has been provided by Krishnamurthy and Vissing-Jorgensen (2011) and Li and Wei (2013) in the US and Boermans and Vermeulen (2018) in Europe among others. Our study complements these findings by bringing further evidence of market segmentation in eurozone Treasury markets with a methodology that has not been used for this purpose.

To summarize, although public debt auctions convey a substantial amount of information to financial markets, it is still an under-researched area, especially within the euro area. The aim of this paper is to bridge this gap by employing the methodology of Gorodnichenko and Ray (2017) to identify demand shocks for German and Italian public debt and tracing their effects on financial markets. There is an important motive behind our country selection. Gorodnichenko and Ray (2017) focuses on the US Treasury market, which is highly safe and liquid. Italian and German bonds are also considerably liquid, but they are traded with substantial differences in terms of risk level. As a result, this study contributes to our understanding of the role of demand shocks in Treasury markets, in particular in relation to issuers with different risk characteristics. The paper also documents the transmission of the demand shock to other assets, namely corporate bonds and equities. Additionally, we look at spillovers of German and Italian Treasury demand shocks to other major euro area countries. The paper also contributes to the literature by studying the role of state-dependence (e.g. high versus low financial stress) and sign-dependence (positive versus negative demand shocks).

2.3 Data

The dataset we use in this study is collected from various sources. The sample starts on the 1st of January 1999 and ends on the 29th of December 2017. In the case of Italy a shorter sample was used, starting in September 2009. We use front contract intraday government bond futures prices data compiled by Refinitiv. There are four types of contracts for Germany, connected to four maturity segments. A short position arising from one of these contracts must be fulfilled by a German government bond with remaining maturity of 1.75 to 2.25, 4.5 to 5.5, 8.5 to 10.5 and 24 to 35 years, respectively. The longest contract was introduced in 2005, while the other three in 1999. The first day with sufficient data to construct the surprise measure is 17th March 1999 and 26th October 2005 for the longest contract. There are three types of Italian contracts, settled by Italian government bonds with remaining maturities of 2.0 to 3.25, 4.5 to 6.0 and 8.5 to 11.0 years. These were introduced to markets in 2010, 2011 and 2009 respectively. As the middle contract exhibits almost no trading activity, we omitted it from our analysis. The first date market depth allowed us to construct the high-frequency surprise is 26th October 2010 for the 2.0 to 3.25 maturity and 14th September 2009 for the 8.5 to 11.0 year maturity.

Besides intraday Treasury futures prices, our dataset consists of daily secondary market government bond yields, primary market auction data, daily stock market indices, daily corporate bond indices, daily credit default swap (CDS) premiums on Treasuries and a monthly country level financial stress index. These are displayed in Figures 2.1-2.4. Demand shocks are identified only for Germany and Italy, but the other largest euro area countries (namely France, Spain and the Netherlands) are considered in the spillover analysis.

Information on Treasury auction results is taken from the national debt management agency websites.[4] Our analysis covers all 2, 5, 10 and 30-year bond auctions available in the sample period. For Germany, there are 536 auctions, while for Italy the number is 247. All daily financial markets data was taken from Datastream with the exception of the corporate bond indices, which are sourced from FactSet. Government bond yields include $2, 3, 5, 7, 10, 20$ and 30-year maturities for both Germany and Italy, with the exception of 15 years instead of 20 for Italy. The country-level corporate bond indices are the corporate sub-indices of the Bloomberg Barclays Euro Aggregate Index. We use the five largest euro area countries' stock market indices: the German DAX, the French CAC40, the Italian FTSEMIB, the Spanish IBEX35 and the Dutch AEX. Finally, the Country Level Index of Financial Stress (CLIFS) was obtained from the ECB and it is based on Duprey et al. (2017). The index is constructed to identify regimes with financial stress that is associated with substantial negative impact on the real economy. The dummy variable we created from this index takes the value one when the index is above its historical 70th percentile,

[4] See https://www.deutsche-finanzagentur.de for Germany and http://www.dt.tesoro.it for Italy.

a threshold the authors mention for systemic financial stress events. The German and Italian CLIFS series within our sample period are displayed in Figure 2.4. The German index peaks in the 2009 market turmoil, with relatively high values also after the burst of the dot-com bubble in the early 2000s. The Italian index is relatively low prior to 2008, but fairly volatile thereafter, peaking during the euro area sovereign debt crisis.

2.4 Identification

In this section we briefly describe the institutional setup of German and Italian Treasury auctions. We then explain how we exploit it to capture unexpected demand shocks for public debt. A more detailed description of the auction procedures is available in the Appendix.

The Debt Management Office (DMO) of both countries publish a yearly issuance calendar at the end of every year to inform investors about the auction dates in the upcoming year. Then, at the end of every quarter, they publish an issuance schedule with information on the types of bonds and the volume to be issued at each auction day. A few days (e.g. 6 working days in Germany and 3 working days in Italy) prior to the auction, the agencies post the exact maturity and volume of the bonds, specify the coupon rate and provide additional details.[5]

On the day of German (Italian) auctions, bidding starts at 8:00 am Central European Time (CET). Primary dealers can place their bids until 11:30 a.m. (11.00 a.m.). At 11:30 a.m. (11.00 a.m.) the DMO collates the bids and decides on the allotment. The decision is made within roughly 5 (15) minutes, after which bidders are notified and the results are published. The published document contains information on the amount of bids received, the amount allotted, the resulting bid-to-cover ratio, the winning price(s) the agency chooses and much more.[6]

The information in auction results can be utilized in multiple ways to quantify demand shocks. The bid-to-cover ratio is the total amount of bids submitted by participants divided by the allotted volume. This measure has been used by Klingler and Sundaresan (2023), Dobrev (2019) and Fuhrer and Giese (2021). There are, however, some issues that make the headline bid-to-cover ratio problematic to identify unexpected changes in the demand for Treasuries. First of all, there is no easy and uncontroversial way to isolate its unexpected component, which is the main focus of this paper. Second, differently from the

[5] Figures 2.5 and 2.6 display two auction announcement documents for Germany and Italy.

[6] As an example Figures 2.7 and 2.8 display two auction result documents published by the German and Italian agencies.

US, German and Italian debt management agencies can adjust the final allotment volume during the auction, based on the prevailing demand conditions.[7] When the agencies judge that the demand conditions are weak, they do not allot all the bonds on offer.[8] Hence, the same headline bid-to-cover ratio could be the outcome of different demand conditions. For example, the German Treasury typically withholds around 14-23% of the issuance, to set aside for secondary market operations in the days following the auction.[9] The Italian agency also have this option, however, it rarely resorts to this measure. In general, agencies base their decision of retention and allotted volume on the prevailing demand conditions, and the resulting retained volume, which is communicated in the same document as the auction results, provides investors with additional information about the market demand.

To assess demand in an auction, market participants follow closely pricing data (see ITC Markets (2017)). An important price statistic in auctions is the average (or accepted) yield. Investors compare these with secondary market yields of comparable securities. When the demand is strong, auction participants tend to bid up prices and offer lower yields. The difference between the secondary and the primary market yields, which we call "yield gap" is another indicator of the strength of the demand for public debt.[10] However, similarly to the bid-to-cover ratio, interpreting the size and isolating the predictable and unpredictable components of the yield gap is again a difficult task.

To overcome these potential shortcomings, we employ a high-frequency identification. Nevertheless, as later described, we also utilize the above measures in an instrumental variable framework. We use the identification proposed by Gorodnichenko and Ray (2017). This relies on the idea that on the day of an auction, the debt management agencies have already disclosed virtually all relevant information about the supply side of the market, such as the issuance volume and security characteristics. Therefore, the press release with the auction results contains new information almost exclusively about the demand side of

[7] Also the UK debt management office reserves the right to withhold part of the gilts on offer, although this option is only considered in "exceptional circumstances."

[8] To bring empirical support to this, for Germany we regressed the retained amount (over the target volume) on the total volume of bids (over the target volume) submitted at each auctioned maturity. Results displayed in Table 2.1, which show that the amount of bids is significantly and negatively associated with the volume withheld by the agency. This shows that the higher the amount of bids submitted by the primary dealers, the smaller the amount retained by the German agency.

[9] It is important to note that the total volume of securities initially announced by the German Treasury is issued within days after the auction. In fact, the withheld securities enter the secondary market gradually in the days following the auction.

[10] More specifically, market participants focus on the "price gap", also called "concession", which is the difference between the secondary market price and the primary market accepted price. But given the relationship between prices and yields of coupon bonds, the yield gap offers the same amount of information. See ITC Markets (2017) for more details on how market participants assess results of Treasury auctions.

the market. Within a short event window around the release of the results, nearly all price movements can be attributed to unexpected changes in market beliefs about the demand for Treasury bonds. As such, this shock does not rely on a specific headline measure of the auction results, nor does it require specific assumptions to identify the unexpected variation of the demand conditions.

As it is standard in the high-frequency identification literature, we follow futures price movements in the event window. Futures contracts have many advantages compared to spot or when-issued prices, such as substantially higher liquidity and better data availability. Furthermore, futures markets tend to lead price discovery ahead of the spot (see Garbade and Silber (1979), Garbade and Silber (1983) and Upper and Werner (2007)).

Investors' reaction to the new information could potentially involve purchasing and selling securities throughout the entire maturity spectrum. Ideally, we would like to have time series that track these shifts in demand at every maturity point. However, using futures data, we can only proxy the shifts by price movements at the points where futures contracts are available.

The demand shock $D_t^{(m)}$ at an auction on day t for maturity m is measured as the difference between the (log) price after and before the release of the auction results. More explicitly:

$$D_t^{(m)} = \left(ln(P_{t,post}^{(m)}) - ln(P_{t,pre}^{(m)}) \right) \times 100 \qquad\qquad m \in \{2, 5, 10, 30\} \qquad (2.1)$$

where $P_{t,pre}^{(m)}$ and $P_{t,post}^{(m)}$ are the prices observed before and after the close of the auction. $D_t^{(m)}$ is calculated for all $m \in \{2, 5, 10, 30\}$ each day when an auction takes place, regardless of the maturity being auctioned. Therefore, on each auction day we record four data points in case of Germany and two in case of Italy. For example, the series $D_t^{(2Y)}$ is equal to the price difference from Equation 2.1 on auction days, and zero on non-auction days.

The frequency of the futures prices is at one minute, i.e. it displays the last traded price within a given minute. For some less liquid contracts there might not be a trade in every minute, therefore we use the 5-minute moving average of the observed traded prices. In case of the Italian contracts, there are periods of very infrequent trading. For those minutes when trading did not take place, we proxy the price with the average of the highest ask and the lowest bid price within the minute.

For Germany $P_{t,pre}$ is chosen to be 30-minutes before the closing of the tender and $P_{t,post}$ to be 30 minutes after, as the German Finance Agency releases the auction results within 5 minutes of the closing. The Italian agency indicated that their process might take up to 15 minutes, therefore we consider a window of 20 minutes before and 40 minutes after the closing of the tender. We experiment with alternative windows in both countries. The correlation coefficient among the resulting shock series is usually very high and the results of the analysis are qualitatively and quantitatively robust to alternative window sizes.[11]

As an illustration, Figure 2.9 shows the price movements of the four German futures contracts within the one-hour event window on the 23rd of November 2011. In this auction the German government was targeting to sell 6 billion euros of bonds with a maturity of 10 years, but primary dealers submitted bids for the total amount of around 3.9 billion. Facing this low demand, the Finanzagentur cut back on the supply and only sold 3.6 billion, resulting in the official bid-to-cover ratio of 1.1. Futures price movements during this specific auction show a large drop at the time the auction result was published (i.e. when investors learned how low market demand was). News agency Reuters commented on this auction as *"A "disastrous" sale of German benchmark bonds"*. As an illustration of a successful auction, we consider the 10-year bund tender on the 17th of April 2013. The 4 billion intended issuance volume met 5.2 billion of bids, resulting in a bid-to-cover ratio of 1.3. Figure 2.10 display sharp futures price increases following the release of the auction results. The Financial Times reported the following reaction by Rabobank analysts: *"Given the backdrop of a* [Euro Area] *peripheral rally and the very low yield available, this is a solid auction result."*

The time series of the identified high-frequency demand shocks are displayed in Figures 2.11 and 2.12. Due to limited availability of futures prices, for Italy the sample is constrained to the post-October-2010 period. The summary statistics of the shocks are displayed in Table 2.2. The means are close to zero (albeit all slightly negative) and the distribution is relatively symmetric, with standard deviations increasing with the maturity. The correlation coefficients among the shocks are generally very high (0.5-0.9) and even higher for shocks with a closer maturity.

There are two issues to consider regarding our identification. Firstly, unexpected changes in the demand for public debt are unobservable by nature. $D_t^{(m)}$ captures the equilibrium price change arising from the shift in the demand. Hence, it is reassuring to verify if $D_t^{(m)}$ is linked to observable movements in the demand. Tables 2.3-2.4 show that our demand shock is closely associated with the bid-to-cover ratio and the yield gap (and their unexpected component). In the robustness section below, we also show that our results are robust when these two indicators are used as instruments for our demand

[11] These results are available upon request.

shocks.

Second, our high-frequency surprises might be contaminated if other relevant events happen within our event window or earlier events have not yet been fully incorporated into asset prices at the start of the event window. Nonetheless, narrowing our event window or excluding days with relevant Eurostat data releases or ECB announcements yields qualitatively and quantitatively similar results.

2.5 Empirical analysis

We now use a regression analysis to study how financial markets react to the demand shocks identified above. The dependent variables in the regressions are daily price changes of different financial assets. By moving from intra-day to daily frequency, we intend to capture responses that might take longer to materialize then the one-hour length of the event window.

As shown in Table 2.2 and Figures 2.9 and 2.10, the demand shocks $D_t^{(m)}$ at different maturities are highly correlated. Following Gorodnichenko and Ray (2017), we compress these series into a single variable by taking the first principal component, denoted with D_t. As futures contracts were introduced at different points in time, we use probabilistic principal component analysis to deal with missing observations. D_t explains over 94% of the variation in the four shock series for Germany. In case of Italy the principal component is constructed from $D_t^{(2Y)}$ and $D_t^{(10Y)}$, due to the limited availability of $D_t^{(5Y)}$. The resulting series explains 98% of their variation. The interpretation of the shock D_t is an unexpected and non-maturity-specific shift in the demand for public debt.

Furthermore, we construct two additional series of shocks to give the analysis more granularity. A long-term shock series $D_t^{(long)}$ is constructed by taking the values of the 10- and 30-year surprise series on the days when 10- or 30-year maturity bonds were auctioned, and extracting the first principal component. It is meant to capture shocks that increase demand for longer maturity government debt. A short-term shock $D_t^{(short)}$ is constructed in a similar way, but with 2- and 5-year auctions. This series is meant to capture shocks that increase demand only for short maturities. The availability of the German shocks are from March 17th, 1999 onward. The futures contract on 30-year bonds does not exist for Italy and the 5-year contract is not liquid enough to be used in the analysis. Therefore short-term Italian demand shocks are proxied by the 2-year series $D_t^{(2Y)}$, while long-term demand shocks are based on the 10-year series $D_t^{(10Y)}$. The availability of these are from October 26th, 2010 and September 14th, 2009 onward respectively. All shock measures were then normalized to have zero mean and unit variance. Table 2.2 displays the summary statistics of these series. $D_t^{(short)}$ and $D_t^{(long)}$ are more correlated with shocks in their own maturity segment, while D_t shows a strong correlation with all maturities. D_t has means very close to zero in both countries, with similar standard deviation (0.26 in Germany and 0.22 in Italy).

2.5.1 Effects on the secondary Treasury market

How does a demand shock for public debt affect domestic Treasury yields? In order to answer this question, we estimate local projection (Jordà (2005)) specifications of following form:

$$\Delta^h R_{t+h}^{(m)} = \alpha_h + \beta_h D_t + \varepsilon_{t+h}. \qquad h = 0, 1, 2 \ldots 30 \tag{2.2}$$

Here $\Delta^h R_{t+h}^{(m)} = R_{t+h}^{(m)} - R_{t-1}^{(m)}$ is the difference between the yield of a bond with maturity $m \in \{2, 5, 10, 30\}$, h days after the auction relative to the day before the auction. D_t is the non-maturity specific demand shock. β_h are the coefficients of interest, while α_h can pick up any pattern in yields independent of the shock around the auction, e.g. documented by Lou et al. (2013). The responses of Treasury yields to the demand shock D_t are displayed in Figure 2.13, with Newey-West standard errors in parenthesis.[12] Not surprisingly, the shock (defined as a change in the futures price) and the bond yields move in opposite direction. What is more interesting is that a price movement in a very narrow intraday window has a large effect that even persists in the following days. The effects are strongest on impact, with bond yields decreasing by 1-2 basis points in Germany and gradually turning insignificant after around 15 days. In Italy the magnitude is larger, and yields drop by 3-4 basis points. The effect turns insignificant quicker, in about 7 days. The figures show that German long-term yields are more responsive to the shock, while in Italy 2- and 5-year bond yields decrease more.

A back-of-the-envelope application of these estimates is to assess the effects of the ECB government bond purchases, the Public Sector Purchase Programme (PSPP), on sovereign yields. On 22 January 2015 the ECB Governing Council announced the launch of its asset purchase program, which entailed the monthly purchases of €60 billion. Starting in March 2015 and carried out until (at least) September 2016, the announcement summed up to €1140 billion. This amount was to be allocated on the basis of the ECB's capital key,

[12] To address the serial correlation induced by the overlapping structure of the data, we set a lag truncation parameter of $2H$ throughout the paper. $H = 30$ is the max length of the local projection exercise. Our results, however, are robust to a wide spectrum of parameter values.

resulting in a share of €205.2 and €142.04 billion respectively. The average volume of submitted bids and the target in German (Italian) auctions has been €7.9 billion and €5.2 billion (€9 and €5.7). An unexpected increase of €205.2 and €142.04 billion in the submitted bids at a single auction would correspond to a bid-to-cover ratio of 35.8 in Germany and 50.7 in Italy. To translate this into futures price shocks, we scale these numbers by using the estimated coefficient of the regression of the surprise bid-to-cover ratios on the high-frequency futures price shocks (Tables 2.3 and 2.4). According to this, an unexpected increase in the demand for the Treasuries would be a 19 (Germany) and a 42.7 (Italy) standard-deviation event in terms of futures price shocks. Based on our estimated impact effects (Figure 2.13), this would decrease 10-year bond yields by 33.4 basis points in Germany and 141.2 in Italy.

There are two major caveats with the above back-of-the-envelope calculations. Firstly, the size of the shock makes it an enormous out-of-sample exercise. Secondly, the source of the demand shocks we identify is the change in *private* investors demand (as opposed to the *public* demand increase due to the PSPP program). Nevertheless, the effects we find are comparable to the findings of Altavilla et al. (2015) and De Santis (2020). According to the event study of Altavilla et al. (2015), the asset purchase announcements decreased 10-year government bond yields by 17 basis points in Germany and 75 basis points in Italy on the day of the announcement. De Santis (2020) calculated the effect of all news related to the ECB's asset purchases from September 2014 to February 2015, a period that includes the program announcement itself, as well as speculations about the possibility of asset purchases prior to the official announcement. De Santis (2020) estimates that these amounted to a decrease of 43 basis points in Germany and 80 basis points in Italy in the 10-year yields. The actual decrease in the secondary market yield on the day of the announcement was 56 basis points in Germany and 108 in Italy.

2.5.2 Impact responses of the yield curve

Up to now we looked at the effects of the demand shock over time and found that the strongest responses are mostly on impact. We now restrict our attention to the contemporaneous effects on the entire yield curve. In particular, we study whether demand shocks at a specific location move different parts of the yield curve. The standard no-arbitrage term structure asset pricing theory predicts that a demand shift unrelated to economic fundamentals, would not affect yields at all. On the other hand, when bond markets are perfectly segmented, interest rates are disconnected at different maturities and affected by local demand (and supply) conditions.

To assess these predictions, we regress $D_t^{(short)}$ and $D_t^{(long)}$ on elements of the Treasury yield curve:

$$\Delta R_t^{(m)} = \alpha^{(m)} + \beta^{(m)} D_t^{(m')} + \varepsilon_t^{(m)} \tag{2.3}$$

where $R_t^{(m)}$ is the yield of a bond with maturity $m \in \{2, 3, 5, 7, 10, 20, 30\}$ and $D_t^{(m')}$ is the demand shock for $m' \in \{short, long\}$.[13] Figure 2.14 displays the estimated $\beta^{(m)}$ coefficients from Equation 2.3. Panel A shows that a one-standard-deviation increase in the demand for short-term German bonds decreases 2-year yields by 1.3 basis points, while 30-year yields drop by 1.6. The demand shock for long-term maturities shows more local effects. Specifically, a positive one-standard-deviation long-term demand shock decreases 30-year bond yields by 2.2 basis points, whereas at the same time the 2-year yield only drops by 0.7 basis points. In the case of Italy, the responses show more disparity (see Panel B of Figure 2.14). A positive one-standard-deviation short-term demand shock decreases 2-year yields by 4.1 basis points and 30-year yields by around 1.4 basis points. Interestingly, the long-term demand shock has larger effects on the short end of the yield curve, then on the long end. While 3-year bond yields drop by 3.9 basis points, interest rates on 30-year bonds drop by only 2.2 basis points after a positive demand shock.

Our findings for Germany are in line with the results of Gorodnichenko and Ray (2017) and Fuhrer and Giese (2021) who find local effects of demand shocks in the US and the UK. It is also consistent with studies bringing evidence of bond market segmentation in the euro area, such as Boermans and Vermeulen (2018).

The strong reaction of short-maturity bonds to the long-term demand shock seems puzzling at first. One possible explanation is through a model where the government faces debt rollover risk and multiplicity of equilibria is present, such as Cole and Kehoe (2000). Here, one equilibrium is characterized by high interest rates and creditors not willing to roll over the debt inducing a default. The other equilibrium is with low interest rates and no default. The positive demand shock we identify might be taken by investors as a signal that Italy is heading towards the low-rate equilibrium. Debt rollover risks will not materialize in the near future, which drives short term interest rates down.

2.5.3 Spillover effects into other financial markets

In this section we study the spillover effects on Treasury bond, corporate bond and equity markets of Germany, Italy, France, Spain and the Netherlands. We intend to

[13] Due to data availability in the case of Italy we use the 15-year instead of the 20-year maturity.

uncover how different assets react to a general increase in the demand for public debt. For the analysis we use the demand shock D_t, that captures shifts in the demand for government bonds at various maturities. We regress this series on daily asset price changes:

$$\Delta Y_t = \mu + \delta D_t + \varsigma_t \tag{2.4}$$

where $\Delta Y_t = Y_t - Y_{t-1}$ is the change in the price of a financial asset on auction day t. Table 2.5 displays the reaction of Treasury yields and CDS spreads. Panel A (Panel B) shows the reactions to the German (Italian) demand shock.

The results indicate that regardless of the origin of the shock, bond yields tend to decrease in all the other countries and maturities following a positive demand shock. Nevertheless, some interesting differences are worth noticing. Increased demand for German debt is followed by an increase in the Italian-German sovereign spread. Increased demand for Italian debt, on the other hand, decreases this spread. Furthermore, a change in the demand for German debt has stronger effect on French and Dutch yields (especially at longer maturities) whereas the Italian demand shock affects mainly Spanish bonds (particularly at short maturities). These results can be explained by the fact that bond price movements are more correlated for securities with similar risk characteristics. Strong comovements between Italian and Spanish bonds after 2010 have been documented by many studies (e.g. Ehrmann and Fratzscher (2017)).[14]

Table 2.5 also shows the responses of CDS spreads. CDS contracts transfer the default risks of the bond from the buyer to the seller of the contract. According to a no-arbitrage argument, the CDS spread should match the yield spread of the corresponding bond with the risk-free rate (Duffie, Darrell (1999)). The CDS spread can be interpreted as an insurance premium payed to insure against the default of the bond issuing entity. The premium is widely used as a proxy for the credit risk of this entity. Results show interesting differences in the reaction of CDS spreads to demand shocks in Germany and Italy. As investors demand more Italian bonds, the credit risk priced in CDS spreads decrease in both countries. On the other hand, a positive German demand shock increases CDS spreads in both countries. To see rising credit risks priced by investors after an increase in the demand for bonds seems counter-intuitive at first. But it is consistent with the fact that a sudden increase in the (private) demand for German bonds is a reflection of a higher demand for “safe-haven" assets. This corresponds to a decrease in investors’ risk

[14] As an additional result, we find that German demand shocks have spillover effects to the US secondary Treasury yields, while the Italian demand shocks do not have any significant effect (see Table 2.6). This seems to be consistent with the fact that German and US yields are close substitutes and perceived as safe assets.

appetite and, in turn, higher insurance premiums, i.e. CDS spreads.

Table 2.7 shows the responses of equity and corporate bond markets to the German and Italian demand shocks. We find that corporate bond yields decrease, although these reactions are only statistically significant when German demand shocks occur. The reaction of German corporate bond yields to a positive German Treasury demand shock is comparable in size to its effect on the sovereign bond yields. These domestic responses of corporate bond markets are consistent with the findings of Gorodnichenko and Ray (2017) who focus on the US over the period 1995-2015. Interestingly, we find that the corporate bond yields of the other euro countries also drop between 1.1 and 1.4 basis points, with French corporate bonds being the most responsive. The effect of the Italian shock is also negative although largely insignificant. Altavilla et al. (2015) also find that the ECB's purchases had large spillover effects to corporate bonds, with movements comparable in size to the reaction of French sovereign yields.

Turning to equity markets, we document large and significant responses of the major euro area stock indices. This is in contrast with Gorodnichenko and Ray (2017), who found no response of the US equity market to Treasury demand shocks. The German shock decreases equity prices in all the major euro area economies. A one-standard-deviation increase in demand for German Treasuries is associated with a 0.26-0.30% drop of the stock prices, whereas a higher demand for Italian bonds leads to positive responses of the stock indices. These latter reactions are statistically significant only for the Italian and Spanish equity indices (0.25 and 0.13% respectively). These stock price movements are consistent with the responses of the CDS spreads. A sudden increase in the demand for German bonds is a signal of investors' lower risk appetite and lower willingness to hold risky assets in their portfolios, leading to a re-balancing from risky equities towards risk-free German public bonds. On the other hand, higher demand for Italian debt is a sign of investors' trust in the Italian economy and fiscal position, which leads to a higher willingness to hold riskier assets.

To bring some narrative evidence in support of this mechanism we refer to the 10-year Italian auction on the 30th of May 2018. During this auction the Italian Treasury allotted 1.8 billion euros, for which it received bids for over 2.7 billion euros, resulting in a bid-to-cover ratio of 1.48. This was considered an improvement relative to the previous auctions, where this measure ranged between 1.25 and 1.38. A global investment strategist at Principal Global Investors commented the outcome of the auction saying that this *"clearly indicates that investors still have faith in the Italian economy, if not the government (...) putting aside the political turmoil, Italy is enjoying a much improved economic and fiscal* *position"*. Commenting on the same auction, Investment Week summarized consequent market movements as: *"Following the auction, Italy's FTSE MIB, which slumped by 2.65% on Tuesday, is up 1.97% as at 12.30pm, and yields on the two-year government bond had fallen to 1.648% from Tuesday's high of 2.805%. Similarly, the five-year bond's yield has fallen to 2.246% from Tuesday's high of 3.074%"*. A German example is for the 10-year auction on the 4th of June 2019, when Deutsche Welle wrote: *"German 10-year government bond yields have fallen to an all-time low as investors scrambled to buy the safe haven asset amid worsening global economic outlook"*.

Summarizing, we find that German and Italian demand shocks for Treasury bonds have spillover effects on the Treasury bond yields of other euro area countries. While the German shocks have more sizable effects on France and the Netherlands, the responses to the Italian demand shocks are mostly centered on the Spanish Treasury yields. We also find that Treasury demand shocks lead to reactions of the corporate bond and equity markets. Namely, our estimates show that euro area corporate bond yields drop in response to German Treasury demand shocks, whereas they are rather unaffected by Italian demand shocks. Interestingly, we find that the main euro area stock indices drop following German demand shocks, whereas Italian Treasury demand shocks lead to positive responses of equity prices.

2.5.4 Robustness analysis

Before we extend our analysis, we briefly discuss a number of robustness checks. First, we test if the asymmetries we find between Germany and Italy depend on the different samples we use for the two countries. More specifically, the demand shock D_t is available from March 1999 onward for Germany, while for Italy only from September 2009. Tables 2.8-2.9 show our baseline results for Germany based on the full sample in comparison with the results obtained when imposing the same restricted sample available for Italy. We find that restricting the estimation sample of Germany to match the Italian one does not have major effects on the results.

In the second robustness exercise, we instrument the demand shocks $D_t^{(m)}$ with observable measures related to the strength of demand in the auctions. As discussed in Section 5, two such measures are the bid-to-cover ratio and the yield gap. Figure 2.15 and Tables 2.10-2.11 show the results when the high-frequency demand shocks are instrumented with the expected and the surprise components of the bid-to-cover ratios and the yield gaps.[15]

[15] Notice that the results of the first stage regressions were reported in the last column of Tables 2.3 and 2.4.

In case of Germany, the estimates are both qualitatively and quantitatively close to the baseline. For Italy the main results are qualitatively robust, but the estimates tend to be less statistically significant, which can be attributed to the weakness of the instruments (see the last column of Table 2.4).

In the third robustness exercise, we follow Gorodnichenko and Ray (2017) and rotate the German $D^{(short)}$ to be uncorrelated with the 30-year series $D_t^{(30Y)}$.[16] This exercise allows to better separate the shocks in the maturity space. Figure 2.16 shows that are main results are not affected.

Finally, we test if our results are robust to the inclusion of control variables. More specifically, we control for the lagged dependent variable, lagged changes in the domestic and the euro stock indices, lagged changes of the domestic 10-year government bond yield and the euro area government bond index and the lagged change of the domestic corporate bond index. The coefficients associated to these controls are found to be statistically insignificant in most cases and, not surprisingly, our baseline results are hardly affected (see Tables 2.12-2.13).

2.6 State dependence and asymmetries

During the sample period under examination, euro area countries went through times of high and low financial stress that may have affected the risk appetite of investors.[17] The theoretical model of Vayanos and Vila (2021) predicts that if investors risk aversion is high, the demand shock has more localized effects. On the other hand, if risk aversion is low, the shock will rather shift the entire yield curve. In order to proxy the risk appetite in markets, we use the monthly CLIFS indicator by Duprey et al. (2017) (see Section 2.3) and construct a dummy variable C_t, taking the value of one when the CLIFS index is over its 70th percentile and zero otherwise.[18] We then estimate the following extension of Equation 2.3:

$$\Delta R_t^{(m)} = C_t\Big(\alpha^{(m,H)} + \beta^{(m,H)} D_t^{(m')}\Big) + (1 - C_t)\Big(\alpha^{(m,L)} + \beta^{(m,L)} D_t^{(m')}\Big) + \eta_t^{(m)} \tag{2.5}$$

[16] More precisely, it was projected onto the space that is orthogonal to the space spanned by the 30-year shock: $D^{(short)} = \big[I - D^{30}\big((D^{30})'D^{30}\big)^{-1}(D^{30})'\big]D^{2,5}$, where $D^{2,5}$ is the first principal component of $D^{(2Y)}$ and $D^{(5Y)}$ on auction days, i.e. the non-rotated short-term shock.

[17] He and Krishnamurthy (2013) show that risk aversion and risk premium rise during a financial crisis.

[18] Slight modifications of this cutoff value do not affect our results.

for $m' \in \{short, long\}$ and for $m \in \{2, 3, 5, 7, 10, 20, 30\}$. The coefficients $\beta^{(m,H)}$ capture the impact of the demand shock $D^{(m')}$ at the maturity segment m' on Treasury yields at maturity m, during periods of high financial stress. Similarly, $\beta^{(m,L)}$ captures the effects of demand shocks during times of low financial stress. Figure 2.17 shows the main results. The contemporaneous response of the yield curve is strong and statistically significant in both countries. In the case of Germany (Panel A) the results for long-term shocks seem to be supportive of the prediction of the theoretical model and consistent with the findings of Gorodnichenko and Ray (2017) and Fuhrer and Giese (2021). Under the regime of no stress (or lower risk aversion), the responses to long-term shocks are flatter than under the regime of high stress (or higher risk aversion). When markets are under stress, domestic yields have stronger reactions, and the location of the demand shock bears importance. Figure 2.17 shows that under such market conditions, a one-standard-deviation demand increase for long-term German debt decreases 30-year bond yields by 3 basis points, while 2-year bonds are unaffected.

Panel B of Figure 2.17 shows the state-dependent effects of Italian yields to Treasury demand shocks. The reaction at the short end is very similar in both regimes. The results for the long-term shock, however, present some interesting features. At maturities longer then 10 years, the effects are quite similar. On the other hand, shorter maturity Treasuries react very strongly in the high stress state. A one-standard-deviation demand shock decreases 2-year bond yields by 5.3 basis points during times of high financial stress, which is more than three times as large as the reaction under the low stress regime.[19]

In Table 2.14 we show estimates of the state-dependent variant of Equation 2.4, that is:

$$\Delta Y_t = C_t\Big(\mu^{(H)} + \delta^{(H)} D_t\Big) + (1 - C_t)\Big(\mu^{(L)} + \delta^{(L)} D_t\Big) + \xi_t. \tag{2.6}$$

The coefficient $\delta^{(H)}$ captures the impact of the demand shock D_t, during periods of high financial stress, while $\delta^{(L)}$ captures its effect when financial stress is low. In general, we find stronger responses during times of higher financial stress. However, in most cases this difference is not statistically significant. While the German results seem fairly similar in the two states, the Italian results show larger differences. Increased demand for Italian debt is associated with large reductions in the credit risk priced in CDS spreads in both

[19] It is important to notice that the Italian sample runs from 2010 to 2017, a period generally classified as turbulent. The financial stress indicator is on a monthly frequency and can therefore identify stressful periods with some granularity (see Figure 2.4) and there is a sufficient number of months that fall below our threshold. Nevertheless, the limited number of available observations under each regime reduced the precision of the estimated coefficients.

Germany and Italy. At the same time, if the sudden shift in the demand occurs during calm periods, CDS spreads remain unaffected. The Italian and the Spanish equity indices display a similar behavior: unaffected during low stress periods, but strong and positive responses during market stress episodes (see Table 2.14).

So far, we have assumed that positive and negative demand shocks have symmetric effects. However, the behavioral finance literature and anecdotal evidence also suggest that markets might respond differently to positive and negative news.[20] In order to test the asymmetry of our results, we estimate variants of Equations 2.5 and 2.6 where we replace C_t with S_t, a dummy variable taking the value one when the identified demand shock is negative and a value zero when the demand shock is positive. The resulting estimated equations are:

$$\Delta R_t^{(m)} = S_t\Big(\alpha^{(m,N)} + \beta^{(m,N)} D_t^{(m')}\Big) + (1 - S_t)\Big(\alpha^{(m,P)} + \beta^{(m,P)} D_t^{(m')}\Big) + \nu_t^{(m)} \tag{2.7}$$

and

$$\Delta Y_t = S_t\Big(\mu^{(N)} + \delta^{(N)} D_t\Big) + (1 - S_t)\Big(\mu^{(P)} + \delta^{(P)} D_t\Big) + \zeta_t. \tag{2.8}$$

Here, $\beta^{(m,P)}$ and $\delta^{(P)}$ capture the effects of positive demand shocks, while $\beta^{(m,N)}$ and $\delta^{(N)}$ estimate the effects of negative demand shocks. The responses of the yield curve to short and long-term demand shocks are displayed in Figure 2.18. When German long-term demand shocks are negative, the localized effect on interest rates are stronger relative to the effects of positive demand shocks. The Italian yield curve, on the other hand, responds in a rather symmetric way. The results of the sign-dependent Equation 2.8 are shown in Table 2.15. Here, the German responses seem to be rather symmetric, while in the case of Italy the baseline results on Treasury markets appear to be mainly driven by positive demand shocks. This is particularly the case for the Treasury spillover effects to other euro area countries. Similarly, we find that that both the equity and corporate bond indices react significantly only to positive Italian demand shocks.

An interesting question is whether these sign-dependent effects are more or less pronounced during periods of high and low financial stress. In order to address this, we classified our shocks by their sign (positive versus negative) and by the level of financial stress when they occur. This exercise has the shortcoming that the number of observations for each of the four scenarios is limited.[21] Tables 2.16 and 2.17 show some interesting

[20] See Veronesi (1999) for an early reference.

[21] The number of negative shocks in high stress, positive shocks in high stress, negative shocks in low

findings, in particular with respect to the response of equity indices to demand shocks. We find that during high stress periods, German positive demand shocks lead to larger reduction of stock prices. Whereas in Italy the stock markets' reaction seem to be driven by positive demand shocks for Italian bonds during financial turmoil periods. When markets are under stress, investors react more positively to stronger demand conditions in the Treasury market. Seeing increased demand for Italian bonds assures markets about the soundness of public finances and the economic prospects, leading to higher stock prices. This is also apparent in CDS spreads. During high financial stress, a positive Italian demand shock decreases CDS spreads in both countries. The same shock in normal times, however, does not have significant effect.

2.7 Conclusions

In this paper we use high-frequency government bond futures price changes around German and Italian Treasury auctions to identify unexpected shifts in the demand for public debt. We first study their effects on secondary market yields of Treasury bonds. Our findings show that positive demand shocks for public debt lead to large negative movements in Treasury yields that can last up to 30 trading days. We test whether a location-specific demand shock moves interest rates at closer maturities. Our results show that shocks at a specific point of the German yield curve have stronger effects on nearby maturities. In Italy, a positive demand shock always decreases short-term interest rates more, regardless of the shock's location. We also document spillover effects into other euro area Treasury bond, corporate debt and equity markets. We find that German demand shocks have larger spillover effects on public debt yields in France and the Netherlands, whereas the Italian spillovers are mostly on Spain.

The most interesting differences we found are on the responses of equity markets and CDS spreads. An increase in the demand for German bonds is associated with drops in the stock prices and an increase in the credit risk priced in CDS spreads. This is in contrast with Italy, where a sudden increase of demand for its bonds is followed by stock price increases and decreases in CDS spreads. We believe that the divergent responses to the two countries demand shock is related to the difference in how investors perceive the seemingly similar information. Higher demand for German Treasuries is associated with a "flight-to-safety" behaviour with investors re-balancing from riskier equities to safer

stress and positive shocks in low stress is 59, 77, 187, 229 respectively in Germany, and 37, 45, 63 and 102 in Italy.

bonds and increases in Treasury CDS spreads. Italy, on the other hand, with its "high-debt" status, has been facing substantial credit risk, above all since the start of the euro area crisis. Higher demand for Italian Treasuries is perceived as a positive signal about its economy, eliminating fears of debt rollover issues. Increases of the stock market and decreases of Treasury CDS spreads indicate that investors are reassured and willing to take more risk. Most of these effects seem to be amplified when markets experience high financial stress. Furthermore, we document that for both countries, stock prices are more responsive to a sudden increase in the demand for Treasuries compared to a decrease, especially during market stress.

2.8 Figures and tables

Figure 2.1: Government bond yields of the five largest euro area member country

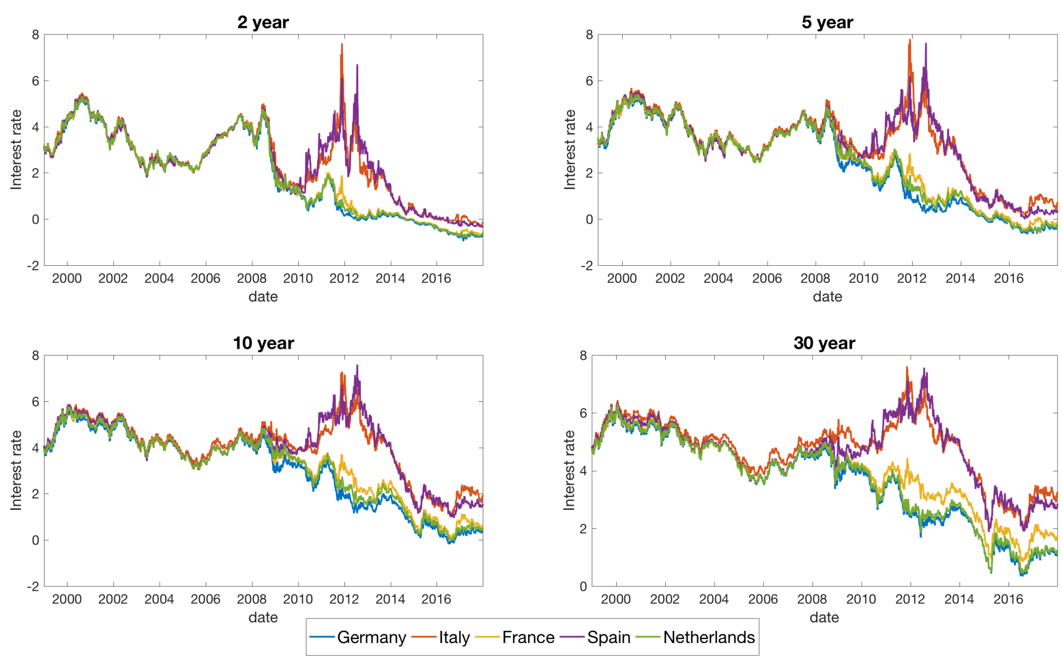
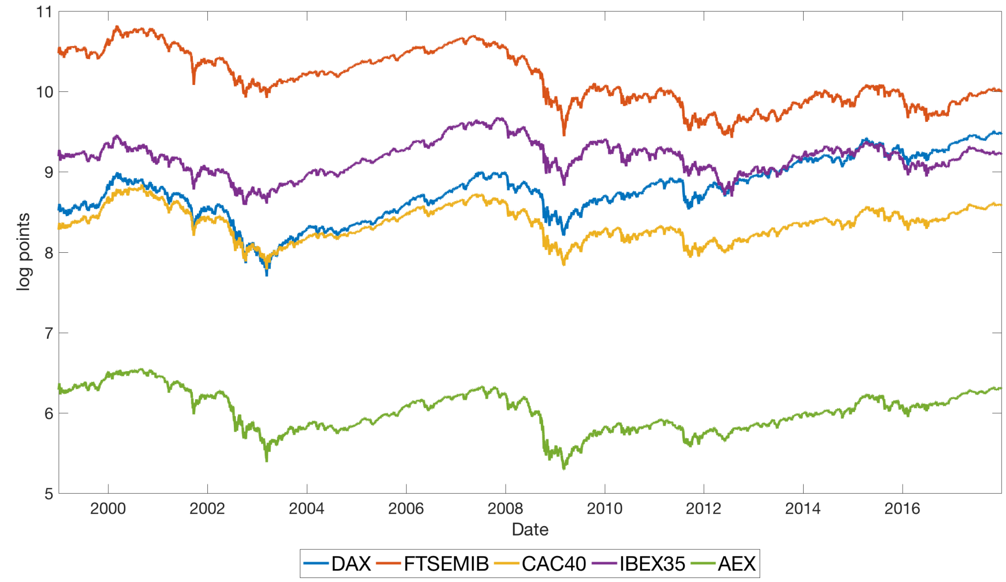


Figure 2.2: Stock indices of the five largest euro area member country



Note: Germany: DAX, Italy: FTSEMIB, France: CAC40. Spain: IBEX35, Netherlands: AEX. Values in logarithm

Figure 2.3: CDS spread on German and Italian 10-year government bond

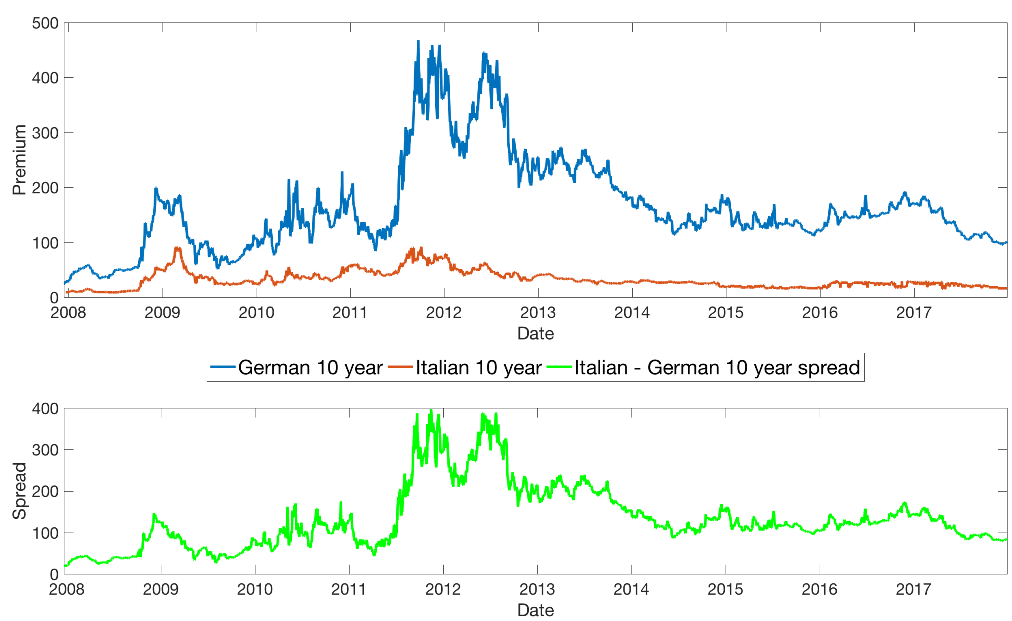


Figure 2.4: Country Level Index of Financial Stress (CLIFS) and its historical 70th percentile in Germany and Italy

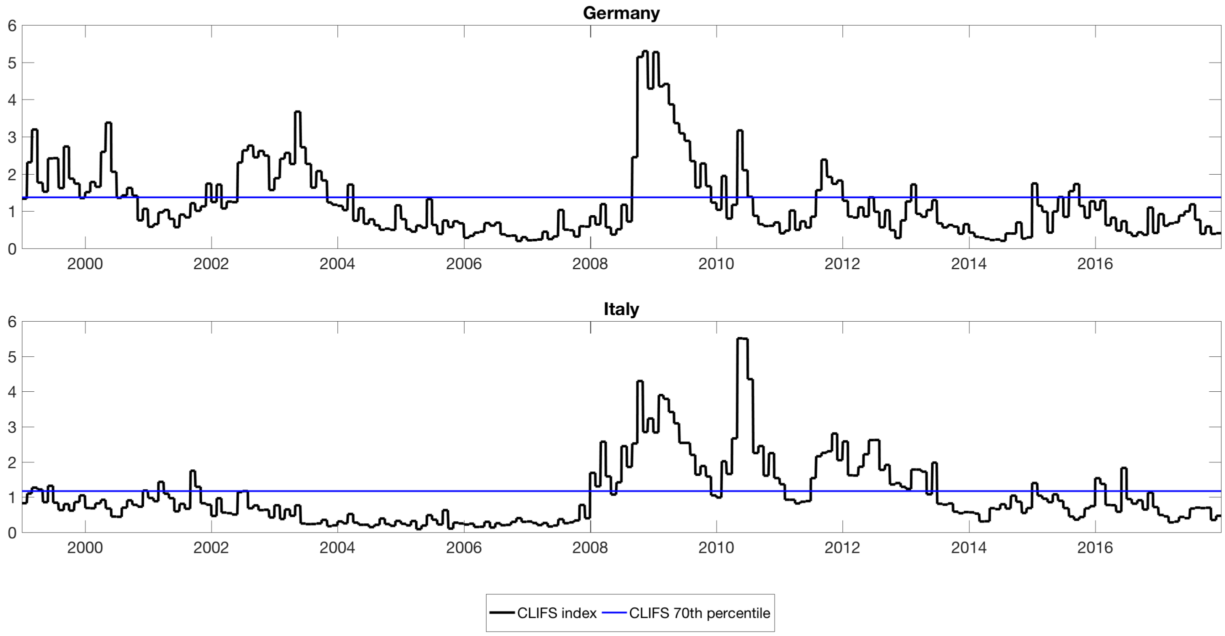


Figure 2.5: Announcement of the 11 January 2017 10-year German bond auction

DEUTSCHE BUNDESBANK EUROSYSTEM

Federal Republic of Germany
Finance Agency

Press release

Frankfurt am Main
3 January 2017
Page 1 of 1

Announcement of auction
New 10-year Federal bond issue

As already announced in the issuance calendar for the first quarter of 2017, the Federal Government will launch a new bond issue (maturity: 15 February 2027) by auction on 11 January 2017. An issue volume (auction allotment and amount set aside for secondary market operations) of € 5 billion is envisaged. Members of the Bund Issues Auction Group are entitled to bid.

Time schedule of the auction procedure:

Date of invitation to bid:	Tuesday, 10 January 2017
Bidding period:	Wednesday, 11 January 2017 from 8.00 a.m. until 11.30 a.m. Frankfurt time
Stock exchange listing:	Wednesday, 11 January 2017
Value date:	Friday, 13 January 2017

Characteristics of the Federal bond:

Maturity:	15 February 2027
Interest payment:	annually on 15 February, interest begins to accrue as of 13 January 2017
First interest payment:	15 February 2018 for 398 days
ISIN	DE0001102416

The nominal interest rate of the Federal bond will be published on the date of invitation to bid. In case of a nominal interest rate higher than zero the separate trading of registered interest and principal („stripping“) will be possible.

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Figure 2.6: Announcement of the 30 January 2017 10-year Italian bond auction

Ministero dell'Economia e delle Finanze

PRESS RELEASE

Medium-Long Term Issuances

The Ministry of Economy and Finance announces the following bonds' issuance and the relative subscription calendar:

Underwriting Deadline date for the Public	Deadline date for Presentation of bids in auction strictly prior to 11.00 am	Submission of bids for the supplementary auction no later than 3.30 pm on	Settlement date
January 27, 2017	**January 30, 2017**	**January 31, 2017**	**February 1, 2017**

Type	**CCTeu**	**BTP 5 year**	**BTP 10 year**
Year residual maturity	On the Run	On the Run	On the Run
ISIN Code	IT0005218968	IT0005216491	to be assigned
Tranche	7th	9th	1st
Issue date	Aug 15, 2016	Oct 03, 2016	Feb 01, 2017
Maturity date	Feb 15, 2024	Nov 01, 2021	Jun 01, 2027
Coupon	-	0.35%	2.20%
Nominal yield	0.561%	-	-
Spread	0.75%	-	-
Current coupon	0.287%	-	-
Coupon payment date	Feb 15, 2017	May 01, 2017	Jun 01, 2017(*)
Amount Min offered (mln. €)	1,750	2,250	3,500
Amount Max offered (mln. €)	2,250	2,750	4,000
Accrued coupon days	170	92	0
Placement fee	0.25%	0.25%	0.35%
% add. amount specialists	15%	15%	30%
Amount offered to Specialists (mln. €)	337.500	412.500	1,200.000

(*) First short coupon: 0,725275% with accrual period: 1 Feb 2017 - 1 Jun 2017 (120 days in a 182 semester).
- After the first, the ordinary cycle will be: 1 Dec - 1 Jun.

The placement mechanism for the above mentioned bonds will be that of a uniform-price (marginal) auction with discretionary determination of allotment price and issued amount within the indicated issuance range.

The issued amount will be set excluding all the bids submitted at prices deemed not to be convenient given market conditions.

The following subjects are allowed to participate in the auction: Italian, EU and non-EU banks, financial brokers and EU and non-EU investment companies as indicated in each issuance decree. They submit bids for their own property or on their clients behalf.

Any bid submitted must contain the reference price. Every dealer can submit a maximum of five bids, which can differ from each other. The minimum bid is 500,000 euro. Any bid inferior to the minimum amount won't be considered. Any bid more than the whole amount offered will be allowed only up to that amount. Bid prices can vary by at least one cent of euro and different changes will be rounded up. Medium and long-term bonds can be subscribed for a minimum amount of 1,000 euro.

They are offered through a uniform-price (marginal) auction referred to the price, without any initial price reference. Dealers' bids have to be transmitted to Bank of Italy within the deadline, described in the "subscription calendar", using the National Interbanking Network with the technical modalities indicated by Bank of Italy itself and well-known by the dealers.

1

Figure 2.7: Press release of the 11 January 2017 10-year German bond auction results

DEUTSCHE BUNDESBANK
EUROSYSTEM

Federal Republic of Germany
Finance Agency

Press release

Frankfurt am Main
11 January 2017
Page 1 of 1

Federal bond issue - Auction result -

The result of the auction of 11 January 2017 for the

0.25 % **bond of the Federal Republic of Germany of 2017 (2027)**
due on 15 February 2027
annual coupon date 15 February
interest begins to accrue on 13 January 2017
first interest payment on 15 February 2018 for 398 days
ISIN DE0001102416

was as follows:

Bids		**€ 7,134.00 mn**
Competitive bids	€ 1,570.00 mn	
Non-competitive bids	€ 5,564.00 mn	
Allotment		**€ 4,017.70 mn**
- Lowest accepted price	98.90 %	
- Weighted average price	98.91 %	
- Average yield	0.36 %	
- Allotment		
- for bids at the lowest accepted price	65 %	
- for non-competitive bids	55 %	
Cover ratio	1.8	
Amount set aside for secondary market operations (Own account of the Federal Government) 1)		**€ 982.30 mn**
Issue volume		**€ 5,000.00 mn**

1) Placing by the German Finance Agency in the secondary market

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Figure 2.8: Press release of the 30 January 2017 10-year Italian bond auction results

Ministero dell'Economia e delle Finanze

Auction Results: 10 YEAR BTP

Date: January 30, 2017 - January 31, 2017

ISIN Code	IT0005240830
Tranche	1st - 2nd
Coupon(*)	2.20%
Issue Date	February 01, 2017
Maturity Date	June 01, 2027
Auction Date	January 30, 2017
Settlement Date	February 01, 2017
Amount Max Offered	4,000.000
Amount Min Offered	3,500.000
Amount Bid	5,147.463
Amount Allotted	4,000.000
Allotment Price	98.58
Bid To Cover Ratio	1.29
Gross Yield	2.37%
Accrued Coupon Days	-
Placement Fee	0.35%
Price for Individual Investors	98.580000
Price for fiscal purpose	98.580
Amount Offered to Specialists	1,200.000
Amount Bid to Specialists	2,982.040
Amount Allotted to Specialists	1,200.000

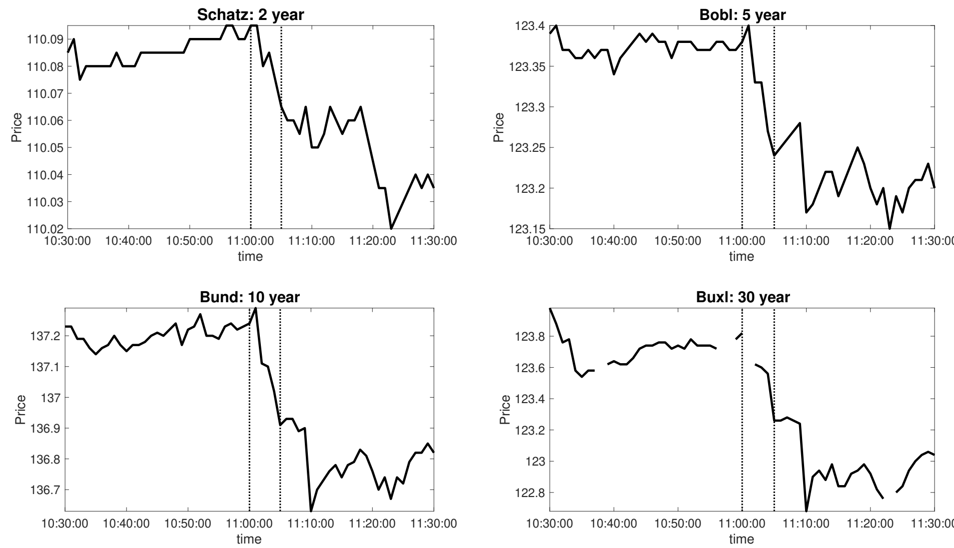
Issue Volume

Outstanding	5,200.000
Amounts allotted to Specialists in supplementary placements and/or syndacated	1,200.000

(*) First short coupon: 0.725275% with accrual period: 1 Feb 2017 - 1 Jun 2017 (120 days) - After the first, the ordinary cycle will be: 1 Dec - 1 Jun

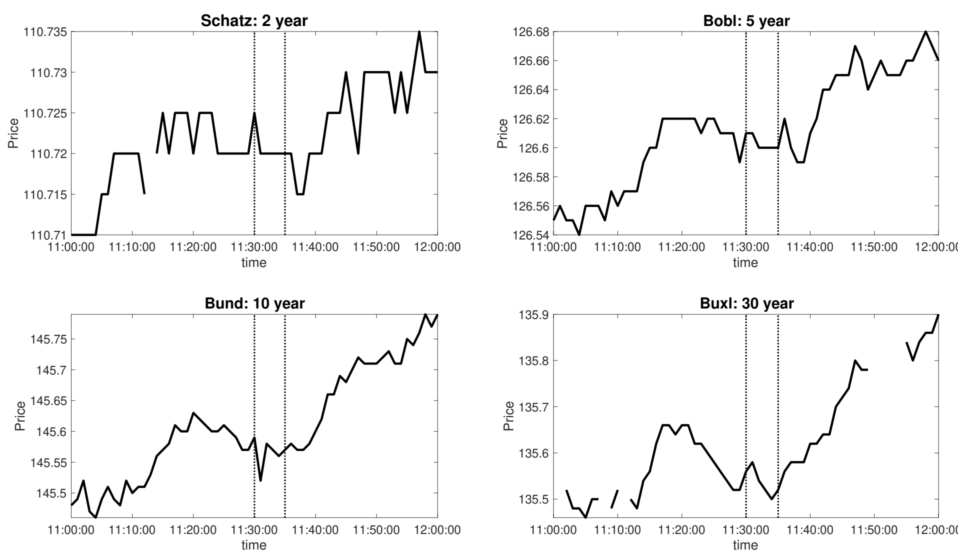
Nominal Amounts are expressed in millions of Euros.
Gross Yields are calculated on 365 days basis.

Figure 2.9: German futures price movements in the event window on the 23rd of November 2011



Note: Auction results are published between 11:00 and 11:05, indicated by the dotted lines.

Figure 2.10: German futures price movements in the event window on the 17th of April 2013



Note: Auction results are published between 11:30 and 11:35, indicated by the dotted lines.

Figure 2.11: Time series of the German demand shock

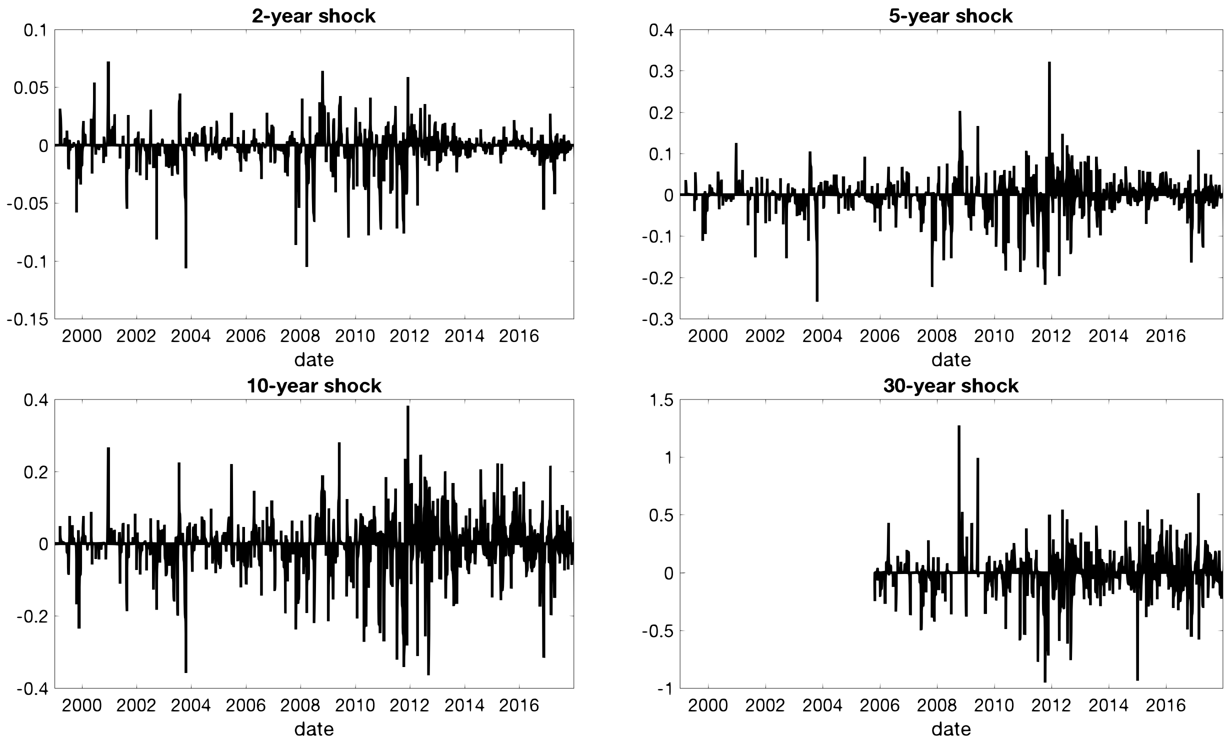


Figure 2.12: Time series of the Italian demand shock

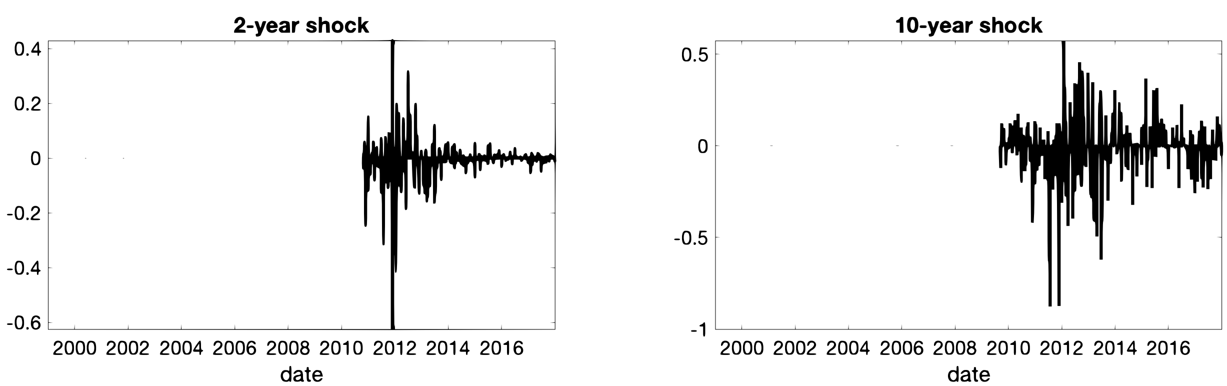
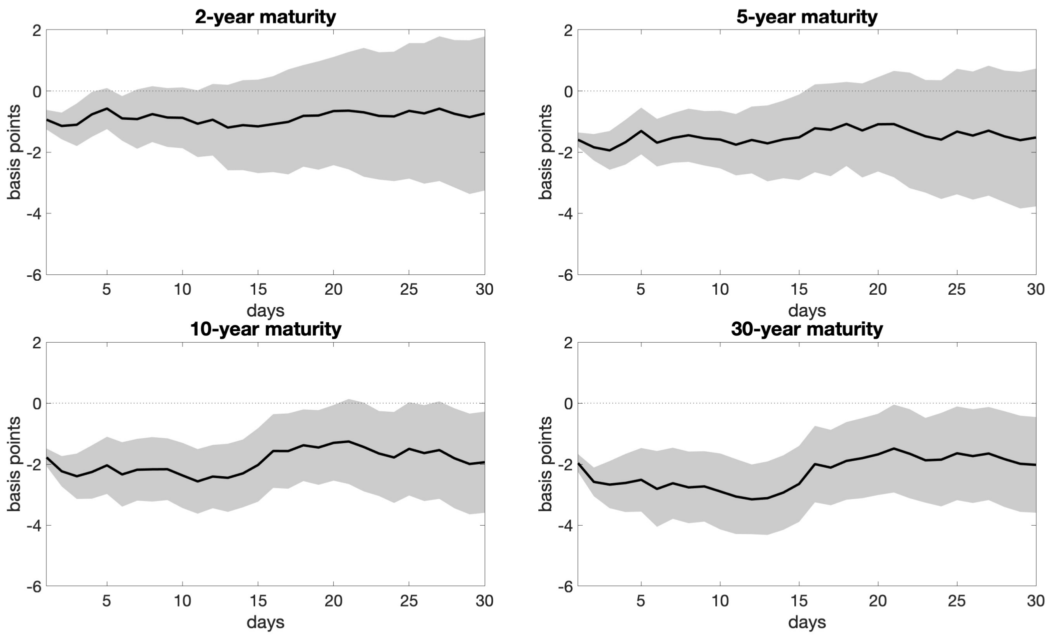
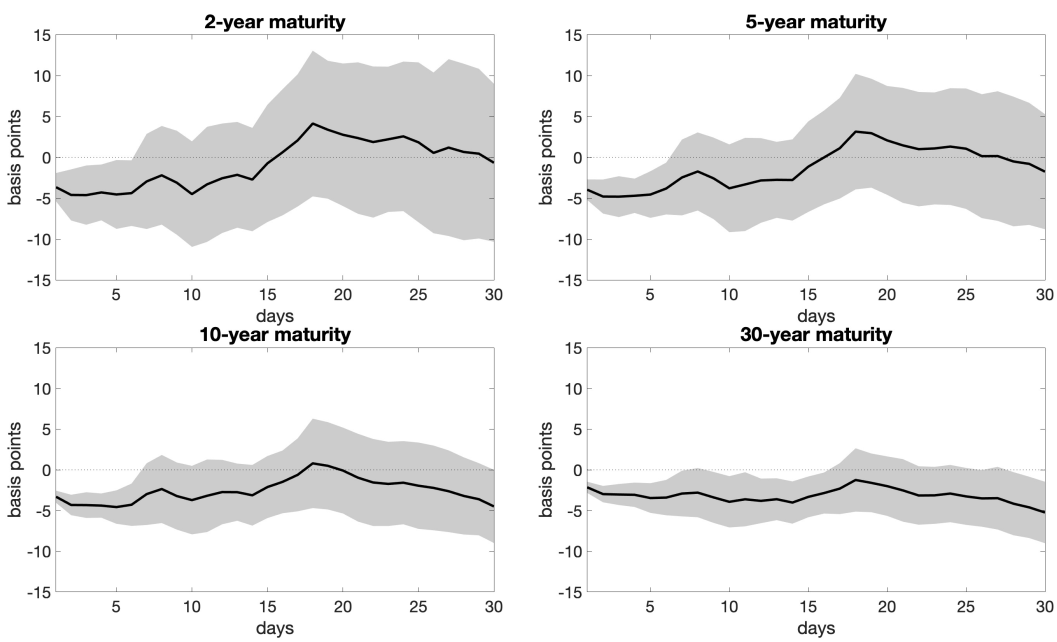


Figure 2.13: Impulse responses of secondary market yields to the Treasury demand shock

(a) Panel A: Germany



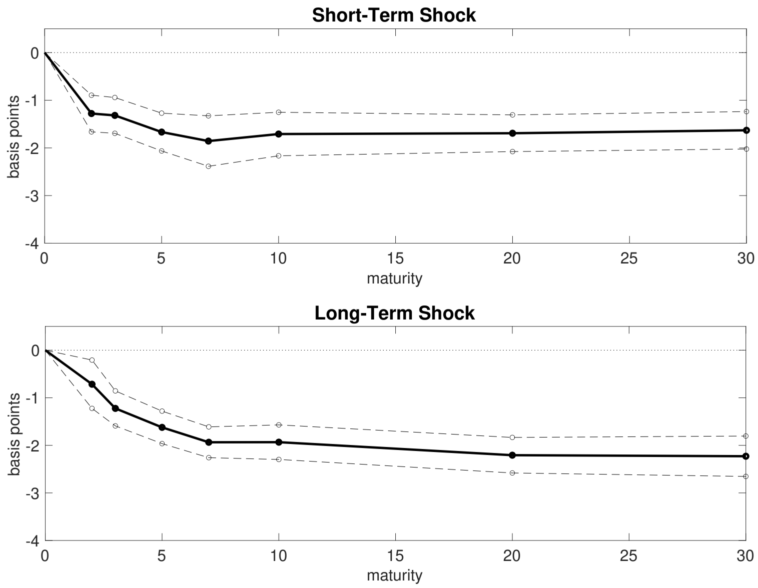
(b) Panel B: Italy



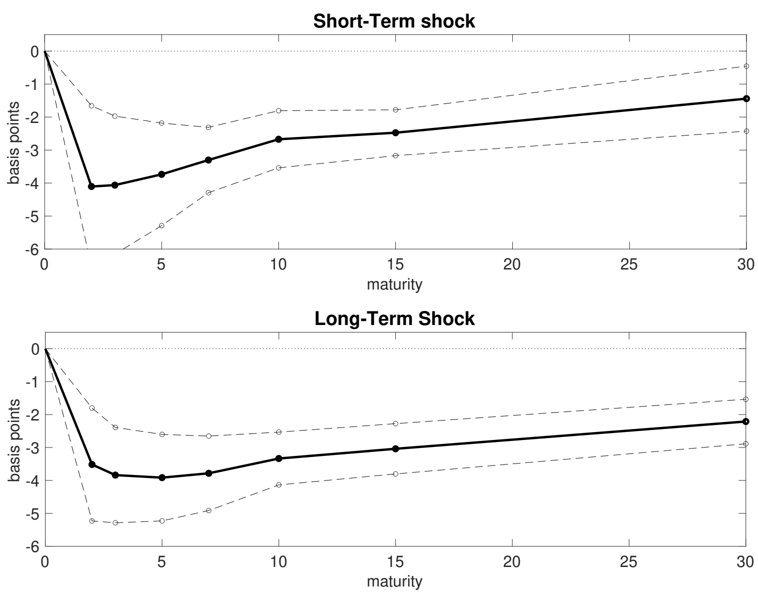
Note: Estimated β_h coefficients from $\Delta^h R_{t+h}^{(m)} = \alpha_h + \beta_h D_t + resid$. Panel (A) shows the impulse responses of 2, 5, 10 and 30-year benchmark German government bonds to the non-maturity specific German Treasury demand shock. Panel (B) shows the impulse responses of 2, 5, 10 and 30-year benchmark Italian government bonds to the non-maturity specific Italian Treasury demand shock. Shaded areas are 90% Newey-West confidence intervals.

Figure 2.14: On impact response of the Treasury yield curve

(a) Panel A: Germany



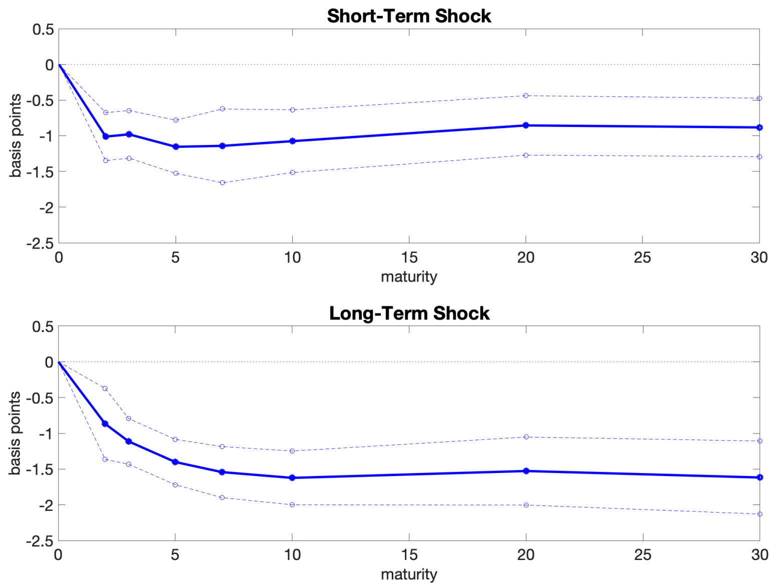
(b) Panel B: Italy



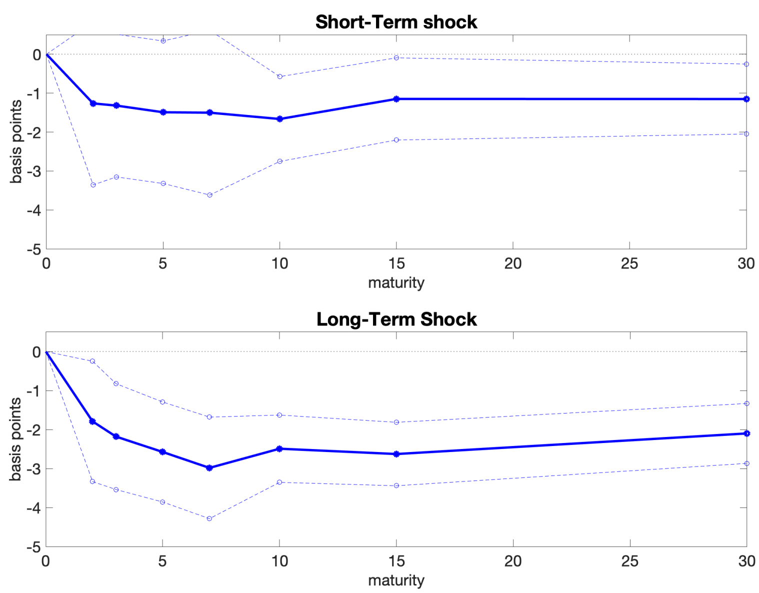
Note: Nodes are the estimated β coefficients from the equation $\Delta R_t = \alpha + \beta D_t^{(m)} + \epsilon_t^{(m)}$, for $m \in \{short, long\}$. $D_t^{(long)}$ is the first principal component of the 10 and 30 year shock, $D_t^{(short)}$ is the first principal component of the 2 and 5 year shock. In case of Italy $D_t^{(short)} = D_t^{(2Y)}$ and $D^{(long)} = D^{(10Y)}$ due to data limitations. Dashed lines are 90% Newey-West confidence intervals.

Figure 2.15: On impact response of the Treasury yield curve (Bid-to-cover IV regression)

(a) Panel A: Germany

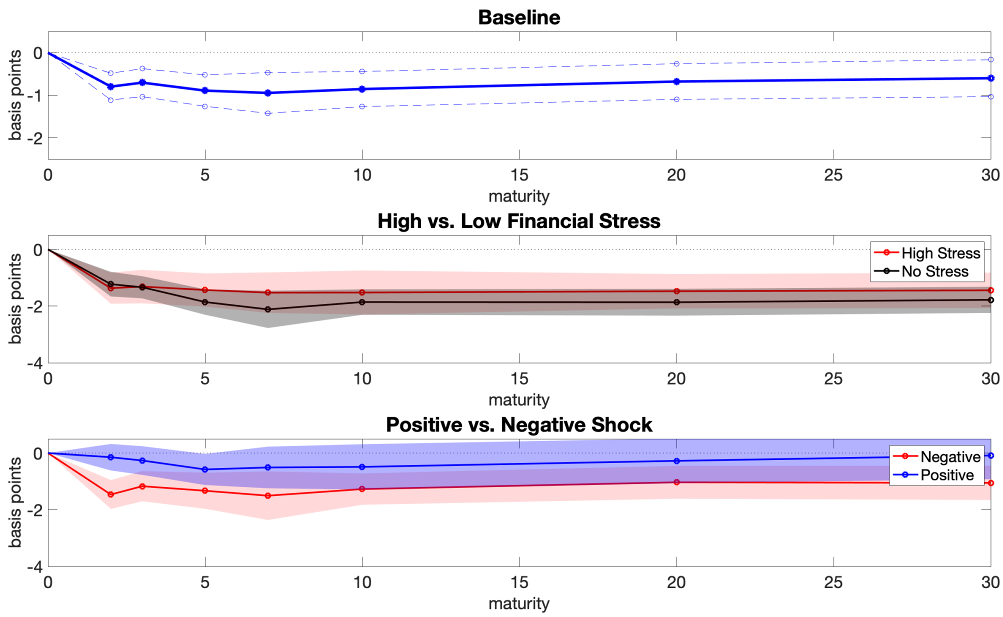


(b) Panel B: Italy



Note: Nodes are the estimated β coefficients from the equation $\Delta R_t = \alpha + \beta \hat{D}_t^{(m)} + \epsilon_t$, for $m \in \{short, long\}$. $\hat{D}_t^{(long)}$ is $D_t^{(long)}$ instrumented by 10 and 30-year auction bid-to-cover and yield gap expected and surprise values, controlling for two lags of the total bid-to-cover ratio. $\hat{D}_t^{(short)}$ is instrumented similarly, using 2 and 5-year auctions. The surprise components obtained as the residuals of univariate AR(4) models. For Italy $D_t^{(short)} = D_t^{(2Y)}$ and $D_t^{(long)} = D_t^{(10Y)}$. $\hat{D}_t^{(m)}$ is then normalized to zero mean, unit variance. Dashed lines are 90% Newey-West confidence intervals.

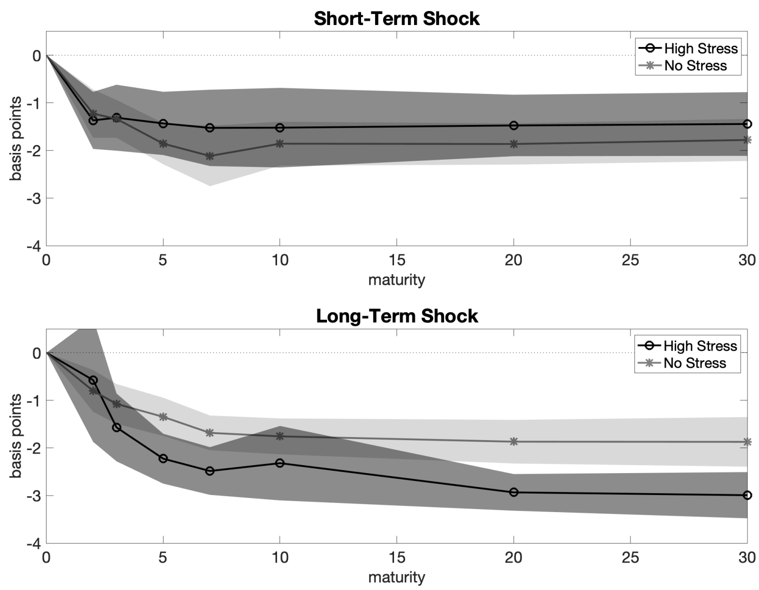
Figure 2.16: On impact response of the German Treasury yield curve to the rotated short-term shock



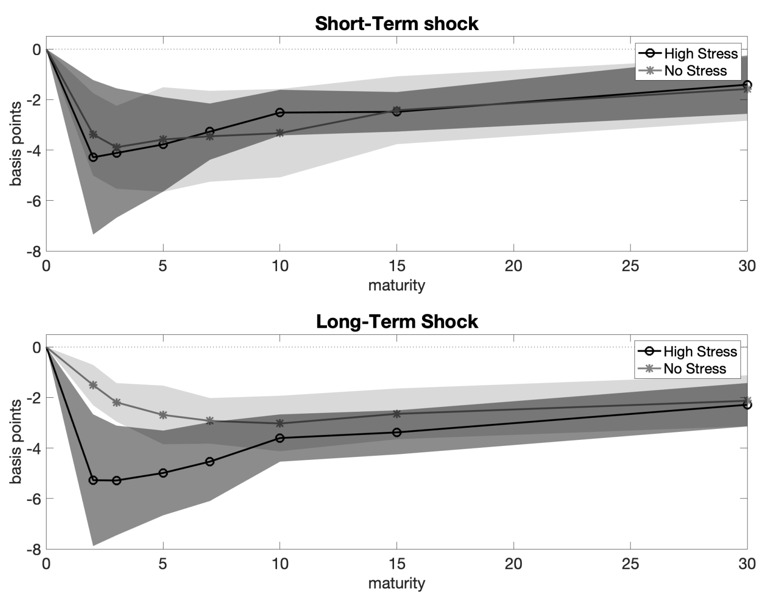
Note: See Figure 2.14, 2.17 and 2.18. $D_t^{(short)}$ is rotated to be uncorrelated with $D_t^{(30Y)}$ and normalized to have zero mean and unit variance.

Figure 2.17: On impact response of the Treasury yield curve in periods of high and low financial stress

(a) Panel A: Germany



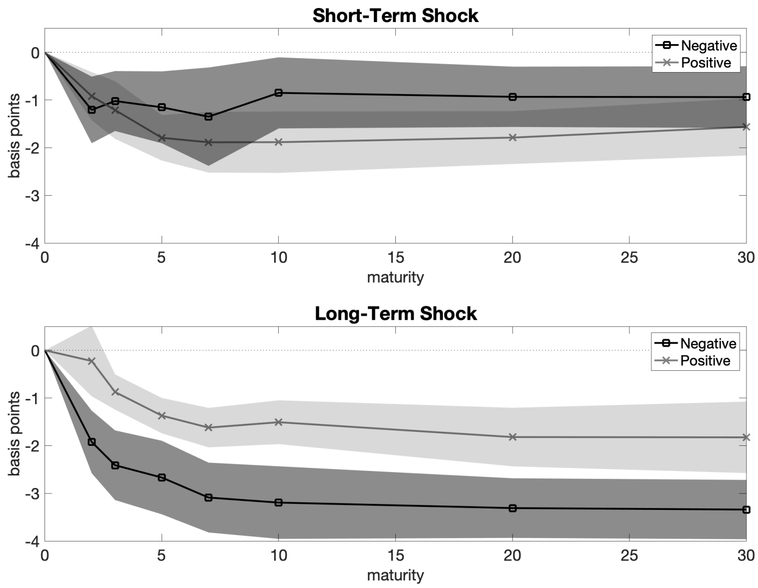
(b) Panel B: Italy



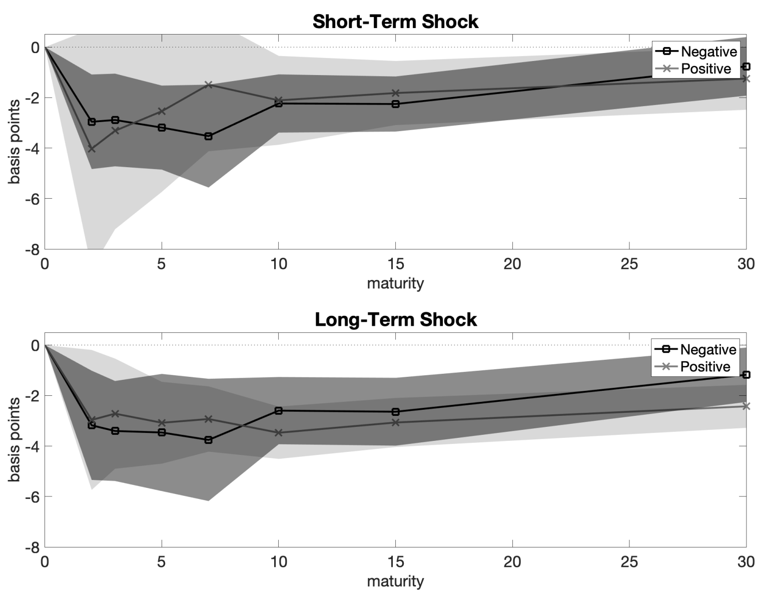
Note: Nodes are the estimated $\beta^{(m,L)}$ and $\beta^{(m,H)}$ coefficients from the equation $\Delta R_t^{(m)} = C_t\big(\alpha^{(m,H)} + \beta^{(m,H)} D_t^{(m')}\big) + (1-C_t)\big(\alpha^{(m,L)} + \beta^{(m,L)} D_t^{(m')}\big) + \eta_t^{(m)}$, for $m' \in \{short, long\}$. $D_t^{(long)}$ is the first principal component of the 10 and 30 year shock, $D_t^{(short)}$ is the first principal component of the 2 and 5 year shock. In case of Italy $D^{(short)} = D_t^{(2Y)}$ and $D^{(long)} = D_t^{(10Y)}$ due to data limitations. C_t is a high financial stress dummy, indicating a CLIFS index higher than its historical 70th percentile value. Shaded areas are 90% Newey-West confidence intervals.

Figure 2.18: On impact response of the Treasury yield curve to negative vs. positive demand shock

(a) Panel A: Germany



(b) Panel B: Italy



Note: Nodes are the estimated $\beta^{(m,N)}$ and $\beta^{(m,P)}$ coefficients from the equation $\Delta R_t^{(m)} = S_t\big(\alpha^{(m,N)} + \beta^{(m,N)} D_t^{(m')}\big) + (1-S_t)\big(\alpha^{(m,P)} + \beta^{(m,P)} D_t^{(m')}\big) + \nu_t^{(m)}$, for $m' \in \{short, long\}$. $D_t^{(long)}$ is the first principal component of the 10 and 30 year shock. $D_t^{(short)}$ is the first principal component of the 2 and 5 year shock. S_t is a dummy variable taking one when $D_t^{(m')} < 0$. In case of Italy $D_t^{(short)} = D_t^{(2Y)}$ and $D_t^{(long)} = D_t^{(10Y)}$ due to data limitations. Shaded areas are 90% Newey-West confidence intervals.

Table 2.1: Retained volume in German auctions explained by demand

	(1) **2 year**	(2) **5 year**	(3) **10 year**	(4) **30 year**	(5) **All**
Intercept	0.214***	0.246***	0.247***	0.338***	0.243***
	(0.009)	(0.014)	(0.012)	(0.032)	(0.006)
Bid-to-cover	-0.029***	-0.042***	-0.046***	-0.115***	-0.043***
	(0.005)	(0.009)	(0.008)	(0.024)	(0.004)
Observations	172	142	167	59	540
R^2	0.162	0.135	0.155	0.286	0.172

Note: Regressions of total amount of bids on the retained amount, both scaled by the targeted volume. Column (1) restricts the sample to include only auctions of 2-year bonds, Column (2) restricts the sample to include only auctions of 5-year bonds, (3) restricts the sample to include only auctions of 10-year bonds, (4) restricts the sample to include only auctions of 30-year bonds. Column (5) includes auction with all maturities. Standard errors in parentheses, (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively.

Table 2.2: Summary statistics of the high-frequency demand shock

Germany	Sample	N	Mean	Med.	Std. dev.	Correlations						
						$D_t^{(2Y)}$	$D_t^{(5Y)}$	$D_t^{(10Y)}$	$D_t^{(30Y)}$	D_t	$D_t^{(short)}$	$D_t^{(long)}$
$D_t^{(2Y)}$	03.1999-12.2017	536	-0.003	0.000	0.020	1.000						
$D_t^{(5Y)}$	03.1999-12.2017	536	-0.004	0.002	0.057	0.837	1.000					
$D_t^{(10Y)}$	03.1999-12.2017	536	-0.005	0.001	0.099	0.674	0.915	1.000				
$D_t^{(30Y)}$	10.2005-12.2017	414	-0.008	0.000	0.238	0.512	0.742	0.858	1.000			
D_t	03.1999-12.2017	536	0.000	0.016	0.251	0.564	0.797	0.903	0.995	1.000		
$D_t^{(short)}$	10.2005-12.2017	536	0.000	0.000	0.046	0.703	0.772	0.627	0.450	0.543	1.000	
$D_t^{(long)}$	03.1999-12.2017	536	0.000	0.000	0.185	0.386	0.550	0.678	0.726	0.731	0.000	1.000
Italy	**Sample**	**N**	**Mean**	**Med.**	**Std. dev.**	**Correlations**						
						$D_t^{(2Y)}$	$D_t^{(10Y)}$	D_t				
$D_t^{(2Y)}$	10.2010-12.2017	208	-0.011	-0.002	0.092	1.000						
$D_t^{(10Y)}$	09.2009-12.2017	247	-0.022	-0.003	0.193	0.760	1.000					
D_t	10.2010-12.2017	247	0.000	0.020	0.203	0.832	0.993	1.000				

Note: Shocks are the recorded high-frequency futures price movements in the event window on auction days. $D_t^{(short)}$ is the first principal component of the 2 and 5 year shock, recorded on days of auctions of 2 and 5-year bonds. $D_t^{(long)}$ is the first principal component of the 10 and 30 year shock recorded on days of auctions of 10 and 30-year bonds. In case of Italy $D_t^{(short)} = D_t^{(2Y)}$ and $D_t^{(long)} = D_t^{(10Y)}$ due to data limitations. D_t is the first principal component of the surprise series at all maturities.

Table 2.3: Auction results and high-frequency surprises: Germany

Panel (A): Total bid-to-cover ratio and yield gap					
	(1)	(2)	(3)	(4)	(5)
	$D_t^{(2Y)}$	$D_t^{(5Y)}$	$D_t^{(10Y)}$	$D_t^{(30Y)}$	D_t
Bid-to-Cover	0.007**	0.045***	0.106***	0.141	0.059***
	(0.003)	(0.013)	(0.020)	(0.124)	(0.021)
Yield gap	-0.063**	-0.192*	-0.608***	-2.946**	-0.033***
	(0.031)	(0.099)	(0.215)	(1.376)	(0.005)
Observations	170	140	163	48	536
R^2	0.099	0.146	0.252	0.165	0.085
Panel (B): Expected and surprise components					
	(1)	(2)	(3)	(4)	(5)
	$D_t^{(2Y)}$	$D_t^{(5Y)}$	$D_t^{(10Y)}$	$D_t^{(30Y)}$	D_t
Bid-to-cover (exp.)	-0.003	0.019	0.123**	0.007	0.045
	(0.010)	(0.033)	(0.056)	(0.282)	(0.029)
Bid-to-cover (surp.)	0.010**	0.057***	0.106***	0.185	0.113***
	(0.004)	(0.015)	(0.024)	(0.141)	(0.031)
Yield gap (exp.)	0.171	0.261**	-1.410	29.381	-0.341
	(0.107)	(0.457)	(2.320)	(20.819)	(0.600)
Yield gap (surp.)	-0.096***	-0.224	-0.605***	-3.346**	-1.387***
	(0.033)	(0.104)	(0.226)	(1.370)	(0.248)
Observations	168	138	163	48	524
R^2	0.127	0.176	0.252	0.230	0.092

Note: Panel (A) shows the estimated coefficient of the regression of the bid-to-cover ratio (total bids over targeted volume) and the yield gap (yield at the action minus the secondary market yield the previous day) series on the high frequency demand shocks. Panel (B) shows the estimated coefficient when the expected and the surprise component of the bid-to-cover and the yield gap series enter separately. The expected component is defined as the fitted values, while the surprise is the residual series from an AR(4) model. All regressions include two lagged values of the total ratio, omitted from the tables. Column (1) restricts the sample to include only auctions of 2-year bonds, Column (2) restricts the sample to include only auctions of 5-year bonds, Column (3) restricts the sample to include only auctions of 10-year bonds, Column (4) restricts the sample to include only auctions of 30-year bonds. Column (5) pools auction with all maturities. Standard errors in parentheses. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively.

Table 2.4: Auction results and high-frequency surprises: Italy

Panel (A): Total bid-to-cover ratio and yield gap				
	(1)	(2)	(3)	(4)
	$D_t^{(2Y)}$	$D_t^{(5Y)}$	$D_t^{(10Y)}$	D_t
Bid-to-cover	0.023	0.093	0.184*	0.022
	(0.024)	(0.126)	(0.093)	(0.036)
Yield gap	0.085*	-0.390*	-0.079	-0.018*
	(0.045)	(0.199)	(0.065)	(0.010)
Observations	80	20	102	247
R^2	0.148	0.309	0.104	0.025

Panel (B): Expected and surprise components				
	(1)	(2)	(3)	(4)
	$D_t^{(2Y)}$	$D_t^{(5Y)}$	$D_t^{(10Y)}$	D_t
Bid-to-cover (exp.)	-0.013	0.046	0.243	0.011
	(0.063)	(0.321)	(0.228)	(0.050)
Bid-to-cover (surp.)	0.026	0.155	0.160	0.034
	(0.026)	(0.151)	(0.106)	(0.043)
Yield gap (exp.)	-0.238	-1.344	-1.453	-0.106**
	(0.428)	(0.933)	(1.085)	(0.043)
Yield gap (surp.)	0.088*	-0.379*	-0.073	-0.046
	(0.046)	(0.200)	(0.066)	(0.053)
Observations	80	20	102	247
R^2	0.155	0.405	0.121	0.042

Note: See Table 2.3.

Table 2.5: Reaction of Treasuries yields and CDS spreads

	DE	IT	FR	ES	NL
Panel (A): German demand shock					
	DE	**IT**	**FR**	**ES**	**NL**
2 year	-0.941***	-0.549*	-0.921***	-0.289	-0.964***
	(0.194)	(0.382)	(0.190)	(0.449)	(0.155)
5 year	-1.596***	-0.616*	-1.388***	-0.354	-1.445***
	(0.145)	(0.385)	(0.167)	(0.457)	(0.146)
10 year	-1.780***	-0.596**	-1.508***	-0.392	-1.594***
	(0.177)	(0.322)	(0.176)	(0.374)	(0.163)
30 year	-1.968***	-0.569**	-1.655***	-0.541*	-1.877***
	(0.179)	(0.273)	(0.184)	(0.337)	(0.167)
2-year CDS	0.211*	1.313***	0.353***	1.933***	0.211**
	(0.129)	(0.497)	(0.135)	(0.487)	(0.096)
10-year CDS	0.207*	1.175***	0.386**	1.798***	0.202**
	(0.238)	(0.463)	(0.202)	(0.468)	(0.114)
Panel (B): Italian demand shock					
	DE	**IT**	**FR**	**ES**	**NL**
2 year	-0.004	-3.650***	-0.409**	-3.297***	-0.174
	(0.159)	(1.056)	(0.179)	(0.832)	(0.171)
5 year	-0.256	-3.962***	-0.519**	-3.209***	-0.559***
	(0.236)	(0.768)	(0.265)	(0.608)	(0.239)
10 year	-0.373*	-3.304***	-0.734***	-2.491***	-0.682***
	(0.252)	(0.454)	(0.264)	(0.519)	(0.246)
30 year	-0.276	-2.151***	-0.651***	-1.811***	-0.307
	(0.255)	(0.418)	(0.253)	(0.422)	(0.241)
2-year CDS	-0.295***	-1.565***	-0.119	-1.109**	-0.406***
	(0.116)	(0.651)	(0.186)	(0.519)	(0.155)
10-year CDS	-0.505***	-1.377**	-0.384**	-0.980***	-0.434***
	(0.176)	(0.650)	(0.170)	(0.411)	(0.155)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta D_t + \varsigma_t$. D_t is the first principal component of the shock measures, normalized to zero mean and unit variance. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively. Panel (A) displays the estimates of the German demand shock, Panel (B) displays the estimates of the Italian shock. The columns correspond to German (DE), Italian (IT), French (FR), Spanish (ES) and Dutch (NL) assets. The rows correspond to 2-year, 5-year, 10-year and 30-year Treasuries and credit defaults swaps (CDS) written on 2 and 10-year Treasuries.

Table 2.6: Spillover effects into the US Treasury market

	German shock	**Italian shock**
2 year	-0.507***	0.052
	(0.153)	(0.085)
5 year	-1.117***	-0.011
	(0.472)	(0.329)
10 year	-1.118***	-0.128
	(0.461)	(0.332)
30 year	-0.984***	-0.125
	(0.326)	(0.300)

Note: See Table 2.5.

Table 2.7: Reaction of equity indices and corporate bond indices

	Germany		**Italy**	
	Equities	**Corp. bonds**	**Equities**	**Corp. bonds**
Germany	-0.260***	-1.129***	0.039	-0.206
	(0.090)	(0.185)	(0.066)	(0.205)
Italy	-0.300***	-1.177***	0.245***	-0.342
	(0.082)	(0.187)	(0.093)	(0.343)
France	-0.246***	-1.355***	0.074	-0.149
	(0.091)	(0.285)	(0.083)	(0.347)
Spain	-0.291***	-1.091***	0.128*	-0.226
	(0.092)	(0.320)	(0.087)	(0.490)
Netherlands	-0.257***	-1.106***	0.037	-0.335**
	(0.104)	(0.285)	(0.055)	(0.176)
Euro area	-0.266***	-1.254***	0.076	-0.221
	(0.093)	(0.192)	(0.076)	(0.269)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta D_t + \varsigma_t$. The left two columns display the estimates of the German demand shock, the right two columns display the estimates of the Italian shock. D_t is the first principal component of the shock measures, normalized to zero mean and unit variance. Equity indices are the DAX, FTSEMIB, CAC40, IBEX35, AEX, EUROSTOXX in logarithm. Corporate bond indices are the corporate sub-index of the country level Bloomberg Barclays Euro Aggregate Index. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively.

Table 2.8: Reaction of Treasury yields and CDS spreads - German sub-sample

			Panel (A): Full sample		
	DE	**IT**	**FR**	**ES**	**NL**
2 year	-0.941***	-0.549*	-0.921***	-0.289	-0.964***
	(0.194)	(0.382)	(0.190)	(0.449)	(0.155)
5 year	-1.596***	-0.616*	-1.388***	-0.354	-1.445***
	(0.145)	(0.385)	(0.167)	(0.457)	(0.146)
10 year	-1.780***	-0.596**	-1.508***	-0.392	-1.594***
	(0.177)	(0.322)	(0.176)	(0.374)	(0.163)
30 year	-1.968***	-0.569**	-1.655***	-0.541*	-1.877***
	(0.179)	(0.273)	(0.184)	(0.337)	(0.167)
2-year CDS	0.211*	1.313***	0.353***	1.933***	0.211**
	(0.129)	(0.497)	(0.135)	(0.487)	(0.096)
10-year CDS	0.207*	1.175***	0.386**	1.798***	0.202**
	(0.238)	(0.463)	(0.202)	(0.468)	(0.114)
			Panel (B): Sub-sample		
	DE	**IT**	**FR**	**ES**	**NL**
2 year	-0.787***	-0.188	-0.648***	0.300	-0.782***
	(0.144)	(0.556)	(0.190)	(0.614)	(0.149)
5 year	-1.583***	-0.213	-1.274***	0.282	-1.380***
	(0.166)	(0.578)	(0.210)	(0.644)	(0.160)
10 year	-1.994***	-0.239	-1.598***	0.111	-1.667***
	(0.175)	(0.466)	(0.203)	(0.518)	(0.168)
30 year	-2.144***	-0.099	-1.723***	0.064	-2.008***
	(0.185)	(0.371)	(0.216)	(0.438)	(0.177)
2-year CDS	0.085*	1.426***	0.348**	1.946***	0.132*
	(0.062)	(0.570)	(0.150)	(0.559)	(0.082)
10-year CDS	0.095	1.266***	0.394**	1.794***	1.267
	(0.092)	(0.534)	(0.232)	(0.538)	(0.109)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta D_t + \varsigma_t$. D_t is the first principal component of the shock measures, normalized to zero mean and unit variance. Panel (A) shows the estimates for the full sample (1999-2017), Panel (B) displays the estimates after restricting the sample to match the Italian sample (2009-2017). Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively. The columns correspond to German (DE), Italian (IT), French (FR), Spanish (ES) and Dutch (NL) assets. The rows correspond to 2-year, 5-year, 10-year and 30-year Treasuries and credit defaults swaps (CDS) written on 2 and 10-year Treasuries.

Table 2.9: Reaction of equity indices and corporate bond indices - German sub-sample

	Full sample		Sub-sample	
	Equities	**Corp. bonds**	**Equities**	**Corp. bonds**
Germany	-0.260***	-1.129***	-0.131**	-1.188***
	(0.090)	(0.185)	(0.061)	(0.206)
Italy	-0.300***	-1.177***	-0.267***	-1.176***
	(0.082)	(0.187)	(0.081)	(0.416)
France	-0.246***	-1.355***	-0.138***	-1.387***
	(0.091)	(0.285)	(0.058)	(0.230)
Spain	-0.291***	-1.091***	-0.198***	-1.195***
	(0.092)	(0.320)	(0.082)	(0.442)
Netherlands	-0.257***	-1.106***	-0.124**	-1.407***
	(0.104)	(0.285)	(0.052)	(0.176)
Euro area	-0.266***	-1.254***	-0.152***	-1.325***
	(0.093)	(0.192)	(0.061)	(0.221)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta D_t + \varsigma_t$. The left two columns display the estimates for the full sample (1999-2017), the right two columns display the estimates after restricting the sample to match the Italian sample (2009-2017). D_t is the first principal component of the shock measures, normalized to zero mean and unit variance. Equity indices are the DAX, FTSEMIB, CAC40, IBEX35, AEX, EUROSTOXX in logarithm. Corporate bond indices are the corporate sub-index of the country level Bloomberg Barclays Euro Aggregate Index. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively.

Table 2.10: Reaction of Treasury yields and CDS spreads - IV regression

		Panel (A): Germany			
	DE	IT	FR	ES	NL
2 year	-1.039***	-0.236	-0.816***	-0.261	-0.970***
	(0.202)	(0.406)	(0.194)	(0.470)	(0.156)
5 year	-1.409***	-0.355	-1.159***	-0.291	-1.217***
	(0.152)	(0.398)	(0.174)	(0.474)	(0.153)
10 year	-1.412***	-0.357	-1.115***	-0.281	-1.224***
	(0.196)	(0.331)	(0.192)	(0.381)	(0.173)
30 year	-1.277***	-0.204	-1.031***	-0.205	-1.172***
	(0.235)	(0.281)	(0.227)	(0.344)	(0.221)
2-year CDS	0.174*	1.315***	0.146	1.646***	0.169**
	(0.130)	(0.497)	(0.144)	(0.499)	(0.096)
10-year CDS	0.177	1.309***	0.261	1.481***	0.133
	(0.138)	(0.462)	(0.207)	(0.482)	(0.113)
		Panel (B): Italy			
	DE	IT	FR	ES	NL
2 year	-0.182	-1.231	-0.194	-0.842	-0.242*
	(0.183)	(1.221)	(0.174)	(1.067)	(0.177)
5 year	-0.287	-1.357*	-0.238	-0.808	-0.361*
	(0.239)	(0.972)	(0.264)	(0.834)	(0.226)
10 year	-0.361*	-1.158**	-0.355*	-0.775*	-0.344*
	(0.252)	(0.660)	(0.259)	(0.621)	(0.236)
30 year	-0.364*	-0.744**	-0.363*	-0.673*	-0.336*
	(0.266)	(0.452)	(0.249)	(0.434)	(0.245)
2-year CDS	-0.071	-0.070	-0.020	-0.511	-0.019
	(0.145)	(0.750)	(0.190)	(0.551)	(0.199)
10-year CDS	0.011	0.055	-0.040	-0.382	-0.025
	(0.249)	(0.725)	(0.202)	(0.454)	(0.200)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta \hat{D}_t + \varsigma_t$. $\hat{D}_t$ is the first principal component of the high-frequency shock measures, instrumented by the surprise component of the bid-to-cover ratio and yield-gap expected and surprise components pooled together at all maturities and normalized to zero mean and unit variance. The surprise component is obtained as the residuals of a univariate AR(4) models. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively. Panel (A) displays the estimates of the German demand shock, Panel (B) displays the estimates of the Italian shock. The columns correspond to German (DE), Italian (IT), French (FR), Spanish (ES) and Dutch (NL) assets. The rows correspond to 2-year, 5-year, 10-year and 30-year Treasuries and credit defaults swaps (CDS) written on 2 and 10-year Treasuries.

Table 2.11: Reaction of equity indices and corporate bond indices - IV regression

	Germany		**Italy**	
	Equities	**Corp. bonds**	**Equities**	**Corp. bonds**
Germany	-0.259***	-0.914***	-0.101*	-0.067
	(0.093)	(0.188)	(0.074)	(0.193)
Italy	-0.262***	-0.854***	-0.059	-0.399
	(0.086)	(0.298)	(0.113)	(0.347)
France	-0.231***	-1.038***	-0.082	-0.251
	(0.094)	(0.198)	(0.090)	(0.357)
Spain	-0.303***	-0.801***	-0.033	-0.362
	(0.095)	(0.329)	(0.092)	(0.503)
Netherlands	-0.255***	-0.931***	-0.053	-0.164
	(0.108)	(0.294)	(0.057)	(0.167)
Euro area	-0.252***	-0.903***	-0.080	-0.198
	(0.096)	(0.202)	(0.081)	(0.267)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta \hat{D}_t + \varsigma_t$. $\hat{D}_t$ is the first principal component of the high-frequency shock measures, instrumented by the surprise component of the bid-to-cover ratio and yield-gap expected and surprise components pooled together at all maturities and normalized to zero mean and unit variance. The surprise component is obtained as the residuals of a univariate AR(4) models. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively. Equity indices are the DAX, FTSEMIB, CAC40, IBEX35, AEX, EUROSTOXX in logarithm. Corporate bond indices are the corporate sub-index of the country level Bloomberg Barclays Euro Aggregate Index.

Table 2.12: Reaction of Treasury yields and CDS spreads - control variables

			Panel (A): Full sample		
	DE	IT	FR	ES	NL
2 year	-0.930***	-0.373	-0.873***	-0.116	-0.930***
	(0.188)	(0.360)	(0.176)	(0.397)	(0.146)
5 year	-1.573***	-0.475*	-1.334***	-0.219	-1.405***
	(0.141)	(0.329)	(0.173)	(0.398)	(0.398)
10 year	-1.759***	-0.546**	-1.476***	-0.323	-1.569***
	(0.188)	(0.299)	(0.193)	(0.333)	(0.177)
30 year	-1.921***	-0.518**	-1.617***	-0.491*	-1.844***
	(0.196)	(0.264)	(0.185)	(0.308)	(0.181)
2-year CDS	0.210*	1.308***	0.370**	2.102***	0.231**
	(0.130)	(0.457)	(0.172)	(0.487)	(0.102)
10-year CDS	0.179	1.175***	0.406**	1.886***	0.220**
	(0.151)	(0.433)	(0.202)	(0.462)	(0.103)
			Panel (B): Sub-sample		
	DE	IT	FR	ES	NL
2 year	0.001	-3.343***	-0.354**	-2.954***	-0.166
	(0.143)	(0.925)	(0.211)	(0.734)	(0.166)
5 year	-0.245	-3.807***	-0.470**	-3.143***	-0.516***
	(0.226)	(0.808)	(0.238)	(0.706)	(0.706)
10 year	-0.368*	-3.289***	-0.685***	-2.516***	-0.655***
	(0.262)	(0.571)	(0.244)	(0.647)	(0.233)
30 year	-0.271	-2.106***	-0.628***	-1.809***	-0.278
	(0.249)	(0.542)	(0.236)	(0.623)	(0.237)
2-year CDS	-0.282***	-1.603**	-0.081	-1.004*	-0.368***
	(0.113)	(0.757)	(0.217)	(0.662)	(0.148)
10-year CDS	-0.508***	-1.398**	-0.377**	-0.902**	-0.413***
	(0.168)	(0.703)	(0.180)	(0.538)	(0.145)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta D_t + controls + \varsigma_t$. D_t is the first principal component of the shock measures, normalized to zero mean and unit variance. Control variables include lagged dependent variable, lagged change in the domestic stock index, lagged change in the euro are stock index, lagged change in the domestic 10-year government bond yield, lagged change in the euro area government bond index, lagged change in the domestic corporate bond index. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively. Panel (A) displays the estimates of the German demand shock, Panel (B) displays the estimates of the Italian shock. The columns correspond to German (DE), Italian (IT), French (FR), Spanish (ES) and Dutch (NL) assets. The rows correspond to 2-year, 5-year, 10-year and 30-year Treasuries and credit defaults swaps (CDS) written on 2 and 10-year Treasuries.

Table 2.13: Reaction of equity indices and corporate bond indices - control variables

	Germany		**Italy**	
	Equities	**Corp. bonds**	**Equities**	**Corp. bonds**
Germany	-0.258***	-1.087***	0.039	-0.178
	(0.077)	(0.228)	(0.075)	(0.224)
Italy	-0.291***	-1.041***	0.245***	-0.404
	(0.074)	(0.282)	(0.088)	(0.337)
France	-0.235***	-1.295***	0.071	-0.171
	(0.074)	(0.198)	(0.082)	(0.343)
Spain	-0.286***	-0.947***	0.116*	-0.275
	(0.073)	(0.303)	(0.085)	(0.464)
Netherlands	-0.247***	-1.079***	0.033	-0.339**
	(0.083)	(0.257)	(0.059)	(0.180)
Euro area	-0.256***	-1.164***	0.075	-0.226
	(0.077)	(0.204)	(0.079)	(0.253)

Note: Estimated δ coefficients from $\Delta Y_t = \mu + \delta D_t + controls + \varsigma_t$. D_t is the first principal component of the shock measures, normalized to zero mean and unit variance. Control variables include lagged dependent variable, lagged change in the domestic stock index, lagged change in the euro are stock index, lagged change in the domestic 10-year government bond yield, lagged change in the euro area government bond index, lagged change in the domestic corporate bond index. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at 10, 5 and 1% respectively. The left two columns display the estimates for the full sample (1999-2017), the right two columns display the estimates after restricting the sample to match the Italian sample (2009-2017).

Table 2.14: Asset price reaction in low and high financial stress

	Germany			Italy		
	Low Stress	**High Stress**	**Test**	**Low Stress**	**High Stress**	**Test**
10-year Treasury yields						
German	-1.870***	-1.640***		-0.618	-0.201	
	(0.161)	(0.390)		(0.516)	(0.257)	
Italian	-0.270	-1.148***		-3.097***	-3.461***	
	(0.425)	(0.365)		(0.720)	(0.486)	
French	-1.517***	-1.499***		-0.644*	-0.811***	
	(0.188)	(0.357)		(0.481)	(0.236)	
Spanish	-0.134	-0.834**		-1.886***	-2.942***	
	(0.491)	(0.504)		(0.630)	(0.627)	
Dutch	-1.636***	-1.538***		-0.805**	-0.600**	
	(0.172)	(0.330)		(0.470)	(0.264)	
CDS on 10-year Treasuries						
German	0.145*	0.330		-0.084	-0.823***	† † †
	(0.104)	(0.347)		(0.096)	(0.239)	
Italian	1.673***	0.202	†	-0.206	-2.246***	† †
	(0.532)	(0.631)		(0.486)	(0.876)	
Equity indices						
German	-0.163***	-0.430**		0.033	0.042	
	(0.069)	(0.205)		(0.070)	(0.107)	
Italian	-0.266***	-0.361**		0.113	0.339***	
	(0.082)	(0.169)		(0.092)	(0.142)	
French	-0.143**	-0.425**		0.056	0.085	
	(0.068)	(0.204)		(0.088)	(0.139)	
Spanish	-0.242***	-0.379**		0.053	0.181	
	(0.085)	(0.199)		(0.083)	(0.142)	
Dutch	-0.115**	-0.505**		0.041	0.033	
	(0.059)	(0.243)		(0.064)	(0.089)	
Corporate bond indices						
German	-1.258***	-0.918**		-0.490**	-0.003	
	(0.165)	(0.403)		(0.288)	(0.270)	
Italian	-1.164***	-1.186**		-0.836***	0.033	
	(0.235)	(0.619)		(0.218)	(0.530)	
French	-1.344***	-1.380***		-0.645***	0.217	
	(0.146)	(0.440)		(0.243)	(0.526)	
Spanish	-1.154***	-0.995**		-0.403*	-0.083	
	(0.373)	(0.592)		(0.314)	(0.825)	
Dutch	-1.387***	-0.634		-0.660***	-0.098	
	(0.137)	(0.702)		(0.266)	(0.216)	

Note: Estimated $\delta^{(L)}$ and $\delta^{(H)}$ coefficients from $\Delta Y_t = C_t(\mu^{(H)} + \delta^{(H)} D_t) + (1 - C_t)(\mu^{(L)} + \delta^{(L)} D_t) + \xi_t$. D_t is the first principal component of the shock measures normalized to zero mean, unit variance. C_t is a financial stress dummy, indicating a CLIFS index value higher than its historical 70th percentile value. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at the 10, 5 and 1% level. (†), (††) and († † †) in the test columns indicate statistically different estimates in the two regimes at the 10, 5 and 1% level.

Table 2.15: Asset price reaction to positive and negative demand shock

	Germany			Italy		
	Negative	**Positive**	**Test**	**Negative**	**Positive**	**Test**
10-year Treasury yields						
German	-1.997***	-1.498***		0.060	-1.295***	† † †
	(0.369)	(0.245)		(0.391)	(0.322)	
Italian	-0.486	-0.634**		-2.611***	-3.445***	
	(0.794)	(0.365)		(0.755)	(0.645)	
French	-1.627***	-1.357***		-0.220	-1.365***	††
	(0.398)	(0.223)		(0.406)	(0.407)	
Spanish	-0.220	-0.178		-1.866***	-2.773***	
	(0.727)	(0.457)		(0.765)	(0.758)	
Dutch	-1.702***	-1.519***		-0.344	-1.629***	† † †
	(0.354)	(0.221)		(0.383)	(0.303)	
CDS on 10-year Treasuries						
German	0.244*	0.197		-0.264	-0.566**	
	(0.165)	(0.243)		(0.252)	(0.313)	
Italian	1.125	0.835*		-1.383**	-1.070*	
	(1.076)	(0.625)		(0.801)	(0.716)	
Equity indices						
German	-0.215**	-0.393***		-0.019	0.172**	
	(0.108)	(0.162)		(0.118)	(0.077)	
Italian	-0.269***	-0.365***		0.144*	0.418***	††
	(0.105)	(0.143)		(0.102)	(0.116)	
French	-0.152**	-0.408***		0.007	0.254***	
	(0.092)	(0.165)		(0.142)	(0.091)	
Spanish	-0.203**	-0.443***		0.056	0.294**	
	(0.100)	(0.162)		(0.102)	(0.129)	
Dutch	-0.138*	-0.462***	†	0.004	0.135**	
	(0.099)	(0.189)		(0.097)	(0.062)	
Corporate bond indices						
German	-1.165***	-0.929***		0.222	-0.999***	† † †
	(0.355)	(0.249)		(0.257)	(0.262)	
Italian	-1.303**	-1.067***		0.006	-1.033**	
	(0.583)	(0.310)		(0.689)	(0.480)	
French	-1.393***	-1.215***		0.584	-1.180***	††
	(0.348)	(0.258)		(0.623)	(0.380)	
Spanish	-1.285**	-0.816***		0.056	-0.712	
	(0.631)	(0.284)		(1.071)	(0.625)	
Dutch	-1.442***	-0.646		-0.024	-0.939***	† † †
	(0.320)	(0.514)		(0.244)	(0.236)	

Note: Estimated $\delta^{(N)}$ and $\delta^{(P)}$ coefficients from $\Delta Y_t = S_t(\mu^{(N)} + \delta^{(N)} D_t) + (1 - S_t)(\mu^{(P)} + \delta^{(P)} D_t) + \zeta_t$. D_t is the first principal component of the shock measures normalized to zero mean, unit variance. S_t is a dummy variable, taking the value 1 when $D_t < 0$. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance at the 10, 5 and 1% level. (†), (††) and († † †) in the test columns indicate statistically different estimates in the two regimes at the 10, 5 and 1% level.

Table 2.16: Positive and negative shock estimates in high and low financial stress - Germany

	Low stress			High stress				
	Negative	**Positive**	**Test(1)**	**Negative**	**Positive**	**Test(1)**	**Test(2)**	**Test(3)**
10-year Treasury yields								
German	-2.232***	-1.543***		-1.765***	-1.423***			
	(0.338)	(0.260)		(0.702)	(0.503)			
Italian	0.949	-0.811**	†	-2.369***	-0.284	†	- - -	
	(0.903)	(0.384)		(0.827)	(0.690)			
French	-1.642***	-1.454***		-1.534**	-1.175**			
	(0.450)	(0.202)		(0.733)	(0.511)			
Spanish	0.952	-0.322		-1.759***	0.101		- -	
	(0.989)	(0.377)		(0.663)	(1.148)			
Dutch	-1.772***	-1.552***		-1.695***	-1.469***			
	(0.402)	(0.193)		(0.663)	(0.527)			
CDS on 10-year Treasuries								
German	0.426**	-0.006	†	-0.123	0.564		-	
	(0.199)	(0.114)		(0.246)	(0.538)			
Italian	3.242***	0.687*	††	-2.729***	1.104	††	- - -	
	(0.968)	(0.494)		(0.952)	(1.390)			
Equity indices								
German	-0.303***	-0.099		-0.143	-0.964***	††		+++
	(0.114)	(0.106)		(0.193)	(0.261)			
Italian	-0.437***	-0.143	†	-0.106	-0.797***	††		+++
	(0.114)	(0.119)		(0.193)	(0.215)			
French	-0.223***	-0.111		-0.094	-0.987***	† † †		+++
	(0.095)	(0.108)		(0.181)	(0.258)			
Spanish	-0.391***	-0.194*		0.015	-0.929***	† † †	- -	++
	(0.118)	(0.127)		(0.130)	(0.260)			
Dutch	-0.157**	-0.122*		-0.144	-1.124***	††		+++
	(0.095)	(0.087)		(0.215)	(0.327)			
Corporate bond indices								
German	-1.380***	-1.083***		-0.900*	-0.640			
	(0.435)	(0.197)		(0.580)	(0.577)			
Italian	-1.293**	-1.091***		-1.230	-1.009**			
	(0.654)	(0.329)		(1.054)	(0.567)			
French	-1.579***	-1.157***		-1.067 *	-1.336**			
	(0.323)	(0.215)		(0.678)	(0.645)			
Spanish	-1.215*	-1.045***		-1.380*	-0.385			
	(0.871)	(0.229)		(0.941)	(0.633)			
Dutch	-1.596***	-1.182***		-1.272**	0.396			
	(0.275)	(0.219)		(0.623)	(1.189)			

Note: Estimated $\delta^{\cdot}$ coefficients from $\Delta Y_t = C_t S_t(\mu^{(H,N)} + \delta^{(H,N)} D_t) + (1 - C_t) S_t(\mu^{(L,N)} + \delta^{(L,N)} D_t) + C_t(1-S_t)(\mu^{(H,P)} + \delta^{(H,P)} D_t) + (1-C_t)(1-S_t)(\mu^{(L,P)} + \delta^{(L,P)} D_t) + \zeta_t$. D_t is the first principal component of the shock measures normalized to zero mean, unit variance. C_t is a high financial stress indicator, S_t is a negative shock indicator. Newey-West standard errors in parenthesis. (*), (**) and (***) denote statistical significance, (†), (††) and († † †) indicate statistically different estimates within high and low stress states, (-), (–) and (—) indicates statistically different estimates of negative shocks between high and low stress regimes, (+), (++) and (+++) indicate statistically different estimates of positive shocks between high and low stress regimes, at the 10, 5 and 1% level.

Table 2.17: Positive and negative shock estimates in high and low financial stress - Italy

	Low stress			High stress				
	Negative	**Positive**	**Test(1)**	**Negative**	**Positive**	**Test(1)**	**Test(2)**	**Test(3)**
10-year Treasury yields								
German	0.218	-1.675***	† † †	-0.424	-0.911**			
	(0.539)	(0.495)		(0.533)	(0.392)			
Italian	-1.684**	-4.383***	††	-2.737***	-2.636***			
	(0.739)	(0.776)		(0.781)	(0.980)			
French	0.237	-1.667***	† † †	-0.495	-1.056**			
	(0.407)	(0.595)		(0.566)	(0.547)			
Spanish	-1.044*	-2.850***		-1.707*	-2.725***			
	(0.679)	(0.868)		(1.059)	(1.136)			
Dutch	-1.167	-1.717***	††	-0.859**	-1.503***			
	(0.535)	(0.538)		(0.509)	(0.324)			
CDS on 10-year Treasuries								
German	0.156	0.009		-0.480*	-1.042***			++
	(0.165)	(0.223)		(0.368)	(0.430)			
Italian	-0.416	0.647		-1.786	-2.594***			+++
	(0.563)	(0.832)		(1.465)	(0.850)			
Equity indices								
German	0.107	0.070		-0.161	0.270**	††		
	(0.111)	(0.119)		(0.151)	(0.127)			
Italian	0.063	0.191*		0.144	0.644***	† † †		+
	(0.106)	(0.135)		(0.196)	(0.174)			
French	0.178*	0.073		-0.173	0.427***	† † †	-	+
	(0.120)	(0.140)		(0.174)	(0.130)			
Spanish	0.163**	0.062		-0.107	0.517***	† † †		+
	(0.091)	(0.164)		(0.167)	(0.180)			
Dutch	0.100	0.042		-0.079	0.229***	††		
	(0.081)	(0.109)		(0.152)	(0.084)			
Corporate bond indices								
German	-0.210	-0.802***		0.135	-1.114***	††		
	(0.404)	(0.325)		(0.403)	(0.411)			
Italian	-1.114***	-0.695*		0.528	-1.358**		-	
	(0.318)	(0.438)		(0.876)	(0.783)			
French	-0.486*	-0.914***		1.080*	-1.378**	††	-	
	(0.320)	(0.344)		(0.768)	(0.621)			
Spanish	-0.600	-0.234		0.440	-1.193			
	(0.484)	(0.662)		(1.645)	(0.989)			
Dutch	-0.450	-0.805***		-0.145	-1.020***			
	(0.397)	(0.298)		(0.412)	(0.366)			

Note: See Table 2.16.

Appendix: Institutional Setup of Government Bond Auctions

In this section we briefly discuss the institutional setup of German and Italian government bond auctions, also known as the primary markets. For more details, we refer to AFME (2020). The German Finance Agency is responsible for all public debt management functions in Germany. In Italy the government bond auctions are carried out by the Bank of Italy in collaboration with the Treasury of the Ministry of Economy and Finance. Only a specific group of investors is allowed to participate in the auction process. In Germany this group is the “Bund Issues Auction Group" and in Italy they are the “Authorized Dealers." These so-called primary dealers have to meet specific requirements such as a minimum amount of successfully submitted bids within a year.[22] The German Treasury uses a multi-price auction process where the winning bids are allotted at the price specified in the bid. For the maturities we consider, the Italian Treasury employs a uniform price auction where the Treasury discretionarily sets the clearing price of the auction and the quantity issued within a range previously announced.

Both governments publish a yearly issuance calendar at the end of every year to inform investors about the auction dates in the upcoming year. Then, at the end of every quarter they publish an issuance schedule with information on the types of bonds and the volume to be issued. A few days (e.g. 6 working days in Germany and 3 working days in Italy) prior to the auction, the agencies post the exact maturity and volume of the bonds, specify the coupon rate and provide additional details. Figures 2.5 and 2.6 show a press release of an auction announcement of a 10-year German bond and a 10-year Italian bond. Both documents clearly communicate the issuance volumes, coupon payments, the time frame when bids are accepted and the settlement date among other pieces of information. Finally, the German agency posts an invitation to bid the weekday prior to the tender, to inform about the timing of the auction.

On the day of German auctions, bidding starts at 8:00 am Central European Time (CET). Primary dealers can place their bids until 11:30 a.m., but before 2012 this was 11:00 a.m. Multiple bids at different prices can be placed, but bids must be of par value of no less than one million EUR. Bidders may also choose to issue non-competitive bids with no upper limit on the demanded amount. The price that these bidders must pay is the weighted average of the winning competitive bids. Whether these bids will be

[22] In 2017 the number of primary dealers was 36 in Germany and 18 in Italy. For the complete list of members and the membership requirements see AFME (2020).

filled completely or not will be decided by the agency. Non-competitive bids accounts for around 30% of all bids. At 11:30 a.m. CET the agency collates the bids and decides on the allotment. The decision is made within roughly 5 minutes, after which bidders are notified and the results are published. This document contains information on the amount of competitive and non-competitive bids received, the allotted volume, the resulting bid-to-cover ratio and the lowest and the average price of the allotted bonds (see Figure 2.7). In each auction, the German Treasury sets aside a part of the initially announced issuance volume for future secondary market operations. This amount is communicated in the same document.

The Italian auctions are organized in a very similar fashion.[23] On the day of the auction, all authorized dealers are allowed to submit their bids through the Italian electronic interbank network. These bids can be continuously adjusted until the closing time, which is 11.00 a.m. After that, the bids collected and sent out to the Treasury Officer who sets the results, publishes the outcome and communicates it to the Bank of Italy. According to the Treasury, this process can take up to 15 minutes. A press statement is then uploaded to the agency's website. An example of these press releases is shown in Figure 2.8. The document contains the amount of bids received, the amount allotted, the resulting bid-to-cover ratio, the market clearing price the agency chooses and much more. The settlement of the securities is two working days following the auction. After posting the press release with the issuance volume, the Italian Treasury did not have any discretion regarding the allotment volume prior to 2008. It issued 100% of the quantity announced. After 2008 the Treasury allowed itself some flexibility, and now only announces a minimum and maximum amount to be allotted at the auctions. Nonetheless it rarely exercises this option, and most cases the maximum amount is allotted.

[23] For details see Dipartimento Del Tresoro (2017).

Chapter 3

Changing Patterns of Risk-Sharing Channels in the United States and the Euro Area*

3.1 Introduction

The debate on how to improve the absorption of macroeconomic shocks has gained attention in the aftermath of the Great Financial Crisis (GFC) and, more recently, in the context of the Covid-19 pandemic and the economic crisis which followed Russia's invasion of Ukraine. Both in the United States (US) and in the Euro Area (EA), the discussion has often developed around the concept of "risk-sharing", which refers to the idea that states within a federation, or countries in a monetary union, share "risks" to insure their future consumption or income streams against negative shocks to local output, i.e., "idiosyncratic shocks" (see, e.g., Canova and Ravn, 1996). The literature has generally identified two main categories of risk-sharing channels, broadly defined as "private" and "public" channels. The former category includes the capital channel, which mainly operates via portfolio diversification in the international financial markets, and the credit channel, through which private and public agents borrow from international banks and which also includes smoothing via domestic savings.[1] The second category mainly comprises the fiscal channel, which operates through transfers from the federal government (in the US), from a common budget (in the EA) or via international transfers.

* This chapter is based on joint work with Jacopo Cimadomo, Massimo Giuliodori, and Haroon Mumtaz.

[1] Loans from international institutions such as the International Monetary Fund or the European Stability Mechanism, or bilateral cross-country loans, fall also typically under the credit channel (see, e.g., Cimadomo et al., 2020).

So far, the literature has not provided clear answers on whether the share of smoothed relative to unsmoothed output shocks has varied over time in these two regions, and the role of the three risk-sharing channels in explaining time variation. Indeed, there are historical and economic reasons that would explain why these channels operate in different ways over time. First, risk-sharing via the capital channel may have changed due to varying cross-ownership of productive assets and different degrees in the synchronization of financial cycles, and among other factors. Second, looser financial regulation may have facilitated cross-state bank lending, which in turn may have reinforced inter-state risk-sharing via the credit channel. At the same time, cross-border lending is often pro-cyclical, especially in recessions, therefore financial de-regulation may have amplified shocks in bad times (see, e.g., Albertazzi and Bottero, 2014). Third, the role of federal fiscal policy as a shock absorber may have also changed, as also reflected in a different policy stance and fiscal activism by subsequent (Republican and Democratic) US governments, or new common initiatives at the European Union (EU) level, such as the "Next Generation EU" (i.e., NGEU).[2]

In addition, it is unclear whether an increase in the effectiveness of one channel could reinforce other channels (complementarity hypothesis), or vice versa would weaken them (substitutability hypothesis). Some authors make the case of complementarity between risk-sharing channels. Notably, Farhi and Werning (2017) strongly argue in favour of fiscal insurance as a necessary complement to private risk-sharing in a currency union. They argue that even when markets are complete, international fiscal transfers are necessary to achieve the desired allocation, because agents do not fully internalise the macroeconomic stabilisation effects of private insurance. On the other hand, other authors support the idea of substitutability. For instance, Belke and Gros (2015) claim that a fully-fledged banking union would reinforce the credit channel and may *de facto* operate as a substitute for a fiscal union, therefore weakening the role of the fiscal channel. In addition, the effects of a stronger credit channel on the effectiveness of the capital channel are unclear. On the one hand, easier access to credit from foreign banks may reduce the scope for portfolio diversification on international financial markets, thus suggesting substitutability between these two channels. On the other hand, the presence of foreign banks and financial intermediaries in a given region may facilitate investment opportunities in foreign financial assets. In this case, we would observe complementarity between the two channels.

[2] The NGEU is a fiscal stimulus programme approved in the EU on 19 February 2021 in the context of the Covid-19 pandemic. It was launched to finance reforms and investments in EU Member States and is set to run until 31 December 2026. It has made €723.8 billion available to EU countries in total, of which €385.8 billion in loans and €338 billion in grants.

Our paper offers three main contributions. First, we extend Asdrubali and Kim (2004) and propose a time-varying parameter panel VAR model, with stochastic volatility, which allows us to estimate how the share of smoothed vs. unsmoothed idiosyncratic output shocks have changed over time in the US and the EA. In this context, we provide evidence on whether the contribution of the different risk-sharing channels have varied over time, or have remained stable. Second, we develop a new test of the complementarity hypothesis which is based on the correlation between the time-varying impulse responses of the various channels to the output shock. Third, for the US, we test if some key macroeconomic and financial variables (e.g., output gap, financial development, interest rates and the public debt-to-GDP ratio) may have contributed to explain the time-varying dynamics of risk-sharing channels estimated in the first stage.[3]

We find substantial time variation in the risk-sharing mechanism. The overall level of shock absorption has increased since the 1970s in the US. This improvement has been mainly driven by private risk-sharing channels. Notably, the capital channel has improved substantially in the 1970s and since then it smooths on average around 40% of an output shock on impact. We also find that credit markets attenuates around 30% of an output shock, and its contribution has tended to increase over the sample. Finally, we show that the federal tax/transfer system absorbs around 10% of the shock. However, its smoothing effect has decreased to around 5% in recent years.

We document that the overall level of risk-sharing in the EA is much lower. Only around 30% of an output shock is smoothed on impact, with a slight improvement over the sample. The credit channel has a predominant and increasing role, while the capital channel turns out to be less important, on average over the sample. Our time-varying analysis allows us to uncover that the capital channel had a sizeable smoothing effect in the leading up to the GFC of 2008-2009 and weakened afterward. At the same time, the importance of credit-market smoothing has gone up after the GFC. Not surprisingly, we find that the role of the fiscal channel is negligible, although it shows a slight improvement after the European Sovereign Debt Crisis (ESDC) of 2010-2012, and in particular during Covid-19 pandemic crisis.

When looking at the interactions between the channels, we find clear evidence of substitution between the capital and the credit channels. This result is evident in both regions, and it has been relatively stable over time. This suggests crowding out between these two channels. This finding can also be interpreted as confirming the the "spare-tire" hypoth-

[3] This part of the analysis is not carried out for the EA given the shorter size of the impulse responses estimated in the first stage, i.e., only 25 years.

esis, postulating that stock markets can mitigate the effects of a banking crisis (Levine et al., 2016). Interestingly, we also find that, in the US, the fiscal and private risk-sharing channels have reinforced each other. This result underpins the theoretical findings of Farhi and Werning (2017), who suggested potential complementarity between public and private risk-sharing channels. In the EA, we find similar qualitative results, although the they are estimated less precisely, also due to a smaller sample size.

Finally, for the US, our results suggest that the federal tax/transfer system and the capital channel provide more smoothing during weak economic conditions. In addition, we find that stronger financial integration lead to better functioning fiscal and credit channels. We also show that a higher government debt-to-GDP ratio is associated with a weaker fiscal and credit channel.

The rest of the paper is structured as follows. Section 3.2 reviews the related literature on risk-sharing; Section 3.3 presents the empirical methodology and describes the dataset used in the empirical analysis; Section 3.4 reports and discusses our empirical results. Finally, Section 3.5 concludes.[4]

3.2 Related Literature

The literature on consumption and income risk-sharing has soared since the 1990s, focusing on both the US and on European countries.[5] Analyzing US state-level data between 1963 and the early 1990s, the seminar paper by Asdrubali et al. (1996) find that 75% of a local shock to per capita gross product of individual states has been smoothed. Cross-state asset ownership contributed to 39% of smoothing, the federal tax/transfer system smoothed 13%, while 22% was accounted by cross-state lending and borrowing (see also Del Negro, 1998, Asdrubali and Kim, 2004 and Mélitz and Zumer, 1999 for early studies on the US). More recent empirical investigations highlighted an increase in the overall level of risk-sharing in advanced economies (see, e.g., Nikolov, 2016). Some authors estimate risk-sharing over rolling windows of data. The general finding is that total shock absorption varies between 70-85% and has gone up since the 1960s. Some papers showed that the role of the fiscal channel has been relatively stable, while the capital and the credit channels have increased in importance (Alcidi et al., 2017, Stempel

[4] Supplementary material containing estimation details, a Monte-Carlo experiment on the estimation of the time-varying panel VAR model, robustness analysis, and additional empirical results are included in the Appendix.

[5] See Cimadomo et al. (2022) for a review of the literature and a comparison of the effectiveness of risk-sharing channels across countries and regions.

et al., 2021).

Results for the EA indicate that a much larger share of shocks is not smoothed, i.e., around 70%. As opposed to the US, the role of capital markets is limited, while the majority of the absorption is via credit markets. Smoothing via the fiscal channel is negligible, which is attributable to the small size of the EU common budget (Asdrubali and Kim, 2004, Afonso and Furceri, 2008 and Furceri and Zdzienicka, 2015). Some studies document a significant drop in risk-sharing during the GFC and the ESDC. During these periods the capital market channel collapsed, and even acted as a shock amplifier (Kalemli-Ozcan et al., 2014 and Alcidi et al., 2017). More recent studies find some improvement after 2012, with the help of official assistance programs of the ESM and the IMF (Milano and Reichlin, 2017 and Cimadomo et al., 2020). The latest evidence from the Covid-19 pandemic shows some additional improvement through the credit channel, most likely due to the NGEU and its Recovery and Resilience Facility (RRF) (Cimadomo et al., 2022 and ECB, 2022).

Overall, these studies suggest that risk-sharing estimates may be heterogeneous across countries and regions and have varied over the years. Nevertheless, to the best of our knowledge, there are no papers analysing how the importance of the three risk-sharing channels evolved over time. Two papers related to ours are Asdrubali et al. (2023) and Foresti and Napolitano (2022). The former paper proposes a heterogeneous panel VAR approch for 21 OECD countries where estimates of risk-sharing channels are allowed to change across countries, but not over time. Foresti and Napolitano (2022) add time variation to a framework which also allows for country heterogeneity. They analyse how the total amount of risk-sharing has changed over time, without disentangling the effects of the single channels. Moreover, they employ a static panel framework, i.e., they can only analyse the contemporaneous response to shocks. Our paper, on the other hand, proposes a fully-fledged dynamic panel VAR model with stochastic volatility. Our approach allows us to estimate shock absorption through the different channels in a fully-fledged time-varying framework, and to trace out how these channels react to the idiosyncratic output shock contemporaneously and up to four years after the shock.

The empirical literature on the interaction between risk-sharing channels is scarce. This literature documents either no interaction or substitution effects between the channels. Evidence from the US from Asdrubali and Kim (2004) shows crowding-out effects between credit-market and capital-market channels. For EU and OECD countries, some papers also highlight substitution effects between capital markets and fiscal risk-sharing (Asdrubali and Kim, 2004, Alcidi et al., 2017 and Asdrubali et al., 2023). The possibility of complementarity between the credit and the capital channels is analysed in the context of the literature on the banking union and the capital market union for the EA (see, e.g., Hoffmann et al., 2018). In this paper, we contribute to the debate by bringing new empirical evidence on complementarity between the fiscal channel and the credit channel, based on a new test.

As a third contribution of the paper, we provide a narrative behind the observed time variation in risk-sharing in the US, based on some macroeconomic and financial determinants motivated by earlier literature. Previous studies identified several factors that could influence the risk-sharing mechanism. In particular, papers have analysed whether risk-sharing has been counter-cyclical, thus providing stronger absorption of local shocks when it is more needed (i.e., during economic downturns), or if it has been instead pro-cyclical, thus amplifying the effects of shocks. Focusing on the period 1963-2005, Hoffmann and Shcherbakova-Stewen (2011) find that inter-state risk-sharing in the US has been pro-cyclical, i.e., increasing in booms and decreasing during downturns. They show that income smoothing through capital income flows tends to be counter-cyclical, whereas the credit-saving channel is strongly pro-cyclical, and this latter effect turns out to dominate. This is confirmed by results from the EA, where a significant drop in risk-sharing has been documented around the GFC, due to the collapse of the capital market channel (Kalemli-Ozcan et al., 2014 and Alcidi et al., 2017).

Financial deregulation and integration are also found to influence risk-sharing (see, e.g., Athanasoulis and Wincoop, 2001 and Demyanyk et al., 2007). Furthermore, several empirical studies document that greater financial globalization tends to increase risk-sharing, at least among industrial countries. The underlying intuition is that more internationally diversified investment portfolios generate income flows that are unrelated to fluctuations in domestic income, therefore better isolating agents from idiosyncratic shocks that hit locally their economies (see Kose et al., 2009, Demyanyk et al., 2008, Pierucci and Ventura, 2010 and Rangvid et al., 2016). Nevertheless, differences in regulation and accounting standards across countries may generate home bias, resulting in sub-optimal shares of foreign assets in domestic portfolios and lower-than-optimal international risk-sharing. Indeed, Sørensen et al. (2007) show that international home bias in debt and equity holdings declined during the period 1993-2003, and this decline was accompanied by an increase in international risk-sharing.

The above findings generally refer to periods of a financial upturn, while the effects of more financial market integration may be reversed during financial market downturns. In addition, if globalization leads to stronger co-movements between international stock markets, the benefits of cross-border holdings of financial assets might be limited (see, e.g., Beine et al., 2010). This is sometimes referred to as the "knife-edge" property of the financial markets: financial interconnections work as a shock absorber (i.e., leading to more risk-sharing) in certain states of the world. In others, interconnections tend to generate shock amplification, i.e., risk-spreading (see Tasca and Battiston, 2014; Balli et al., 2013).

The state of public finances may also have an effect on risk-sharing. A more prudent government and with a larger fiscal space has more capacity to finance regional counter-cyclical policies, therefore reinforcing shock absorption. Stempel et al. (2021) tests this hypothesis at the intra-state US level. He clusters US states based on their risk-sharing profiles. He finds that states differ both in the overall level of risk-sharing and also in the relative importance of each channel. State with stronger fiscal rules are likely to have a better government financial position (Grembi et al., 2016), and, therefore, more capacity for fiscal smoothing. This is likely to be valid also at the inter-state, i.e., federal level.

Another key determinant of risk-sharing could be monetary policy. A change in the interest rate environment directly affects credit costs and risk premia (Gertler and Karadi, 2015). This influences the marginal cost of smoothing via credit markets and fiscal policy. If monetary policy affects risk pricing, it could also have an effect on capital market smoothing.[6] Monetary policy is also likely to have strong effects during periods of financial market stress, in particular, if it backstops a financial meltdown, and a freeze of the capital and credit markets. The last section of this paper is devoted to investigate if these variables, in particular the business cycle, indices of financial development, the debt-ratio as measure of the state of public finances, and interest rates as measures of the monetary policy stance, had a role in influencing the effectiveness of the different risk-sharing channels.

3.3 Methodology

Our starting point is the framework proposed by Asdrubali et al. (1996), who quantify risk-sharing based on the cross-sectional variance decomposition of shocks to output. Empirically, this amounts to running regressions of each risk-sharing channel onto changes in output, using National Accounts data (for a brief derivation of this decomposition, see Appendix A). However, this static model does not capture the dynamic behaviour of con-

[6] In a recent paper, Hauptmeier et al. (2022) analyse the other direction of causality, i.e., they show that risk-sharing can influence the regional transmission of monetary policy.

sumption smoothing or feedback effects among the three channels. Asdrubali and Kim (2004) overcome this by generalizing the framework and estimating risk-sharing in a panel VAR model. They combine output and the smoothing variables in a single system of endogenous variables and apply a recursive identification scheme, motivated by the nature of the National Accounts variables. We follow the same approach in constructing our vector of endogenous variables Y_{it}:

$$Y_{it} = \begin{pmatrix} \Delta\log GDP_{it} \\ \Delta\log GDP_{it} - \Delta\log GNP_{it} \\ \Delta\log GNP_{it} - \Delta\log GDI_{it} \\ \Delta\log GDI_{it} - \Delta\log C^{it} \end{pmatrix} \tag{3.1}$$

where, in the case of the EA, GDP_t^i is the real per capita gross domestic product of country i in year t, GNP_t^i is the real per capita gross national product, GDI_t^i is real per capita gross disposable income, and C_t^i is rela per capita total consumption (both private and public). For the US, we use the state-level equivalents corresponding to the gross state product, state income, disposable state income, and state consumption. As standard in the literature, we interpret the changes in $\Delta\log GDP_t^i - \Delta\log GNP_t^i$, $\Delta\log GNP_t^i - \Delta\log GDI_t^i$ and $\Delta\log GDI_t^i - \Delta\log C_t^i$ in response to orthogonalised shocks to $\Delta\log GDP_t^i$, as measures of risk-sharing achieved by capital markets (capital channel), international transfers (fiscal channel), and credit markets (credit channel), respectively. The response to GDP which is not absorbed by these three channels is labeled as "unsmoothed". In order to identify the effects of idiosyncratic shocks, all variables are expressed in log-deviations from their respective weighted average values, where the weight of each country (state) in each year is based on the size of the real GDP (GSP in the case of the US) in the previous year.[7]

We extend the framework of Asdrubali and Kim (2004) by adding time-varying parameters and stochastic volatility to the baseline fixed-coefficient panel VAR model. In particular, Y_{it} denotes the matrix of N endogenous variables for country i and X_{it} collects

[7] We demean by the weighted averages to account for the different size of the countries/states in the respective region. However, below we show that demeaning by the simple averages (which effectively amounts to introducing time-fixed effects) produces very similar results.

all the right-hand-side variables $X_{it} = [Y_{it-1}, .., Y_{it-p}, 1]$.[8] The panel VAR is given by:

$$\begin{aligned} Y_{it} &= X_{it}B_{it} + A_{it}^{-1}H_t^{1/2}e_{it} \\ e_{it} &\sim N(0, \sigma_i^2) \end{aligned} \tag{3.2}$$

where the intercept is the last column of X_{it}. The entries of the lower triangular contemporaneous impact matrix A_{it} have two components. The first component a_i is idiosyncratic, but fixed over time. The second component α_t captures time-variation that is common across countries. The coefficient matrix B_{it} has the same structure. Its first component b_i is constant, but unit specific, while the second component β_t varies over time, but is common across countries. We focus on this component, as the interest of our analysis is on the variation of risk-sharing channels over time.[9]

$$\begin{aligned} B_{it} &= b_i + \beta_t & \qquad A_{it} &= a_i + a_t \\ \beta_t &= \beta_{t-1} + \eta_t & a_t &= a_{t-1} + v_t \\ \eta_t &\sim N(0, Q_B) & v_t &\sim N(0, Q_A) \end{aligned}$$

We assume that the time-varying components β_t and a_t evolve as random walks, with variances Q_B and Q_A. The lower triangular form of A_{it} imposes contemporaneous feedback restrictions, as in Asdrubali and Kim (2004). Variables in the system are contemporaneously exogenous with respect to the variables ordered below it: $\Delta\log GDP_t^i$ is the most exogenous, as GDP relates to the value added in production, before tax payments, and financial income or credit flows. And it is important to note that the responses of the three variables (each identifying one of the three risk-sharing channels) to an output shock (ordered first) are invariant to the ordering of these variables (see Christiano et al., 1999).

The residuals of the model are heteroscedastic with the stochastic volatility given by H_t, which is a diagonal matrix. The diagonal elements are assumed to evolve as geometric

[8] Due to presence of large outliers in the stochastic volatilities of the CRE series for the US, we also include five time dummies in X_t, namely for the years 1977, 1979, 1984, 1997 and 1999. The inclusion of these dummies does not materially affect the main results, but allows to obtain smoother impulse response functions.

[9] The unit-specific component of the coefficients, a_i and b_i, allows countries (or states in the US) to have unique risk-sharing profiles. However, as our interest is risk-sharing within a region, when showing our results, we produce them using the weighted average of b_i and a_i. Details of the estimation are shown in Appendix B.

random walks. The volatility can also differ across units.

$$H_t = diag\,(h_t)$$
$$\ln h_t = \ln h_{t-1} + \varepsilon_t$$
$$\varepsilon_t \sim N(0, g)$$

The model is closely related to the panel VAR proposed in Canova and Ciccarelli (2004). However, the structure of the model is more parsimonious in our setting and the estimation is simpler as a result.[10] Our model generalises the threshold panel VAR used in Mumtaz and Sunder-Plassmann (2021) by allowing for the possibility of changes in coefficients and variances at each point in time. Finally, compared to time-varying VARs proposed in Cogley and Sargent (2005), the cross-sectional dimension of the model can help to obtain more precise estimates of the parameters of the state-transition equation. The model is estimated with a Gibbs sampler, based on the sampler for panel VARs described in Canova and Ciccarelli (2004) and Jarocinski (2010) and the sampler for TVP-VARs with stochastic volatility described in Cogley and Sargent (2005) and Primiceri (2005). The estimation details are described in Appendix B, together with a Monte-Carlo experiment to test the estimation algorithm. The model is estimated with one lag, but below we show that our results are robust to the inclusion of two lags.

The dataset for the US covers the period from 1963 to 2020, all fifty states and it is sourced from the Bureau of Economic Analysis (BEA) and the US Census Bureau. We use the first ten years to train our priors. State income is constructed as state personal income plus federal nonpersonal taxes, contributions, and state and local nonpersonal taxes minus direct transfers to individuals. Disposable state income is state income plus federal grants to state governments, federal transfers to individuals minus federal nonpersonal taxes and contributions, and federal personal taxes. State consumption is PCE-deflated retail sales plus state and local government consumption.[11] The cross-sectional mean and standard deviation of the four series used in the panel VAR are displayed in Figure C1 of the Appendix.

The dataset for the EA is from the AMECO database (2022 Spring vintage) and covers the sample 1990-2023. We have decided to include in the analysis also 2022 and 2023,

[10] Canova and Ciccarelli (2004) allow for the possibility of dynamic interdependencies amongst countries. This feature is not needed in our application and a simpler specification can be employed. In contrast to Canova and Ciccarelli (2004), we model time-variation directly in each coefficient and incorporate a stochastic volatility specification for the disturbances.

[11] The variables definition follows Asdrubali et al. (1996). We, therefore, refer to their paper for a detailed description of how each variable is constructed.

although these years were still preliminary at the time of the analysis, because this allows us to derive some insight into how risk-sharing has evolved during the Covid-19 pandemic and the subsequent economic crisis.[12] The sample includes the EA-19 countries (excluding Ireland[13]): Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Greece, Italy, Latvia, Lithuania, Luxembourg, Malta, the Netherlands, Portugal, Slovakia, Slovenia, and Spain. We use the first seven years to train our priors. Gross national product is calculated as GDP plus net factor income from abroad. Gross disposable income is gross national product plus net transfers from abroad. Total consumption is gross disposable income minus gross national savings. The cross-sectional mean and standard deviation of the four series of the panel VAR for the EA are displayed in Figure C2 of the Appendix.

3.4 Results

In this section, we first estimate the evolution of the overall degree of risk-sharing and how the relative importance of the different risk-sharing channels has evolved over time in the US and the EA. We focus on both the risk-sharing achieved on impact, as well as four years after the shock. Then, we move on to explore whether the risk-sharing channels have acted as complements or substitutes over time. In order to do this, we develop a new test of the complementarity hypothesis based on the correlation between the time-varying impulse responses to the output shock. Finally, exploiting the length of the US sample, we explore potential determinants of each of the three risk-sharing channels.

3.4.1 Time variation in risk-sharing channels

We first analyse how the absorption of an output shock via the three channels has evolved over time in the US and the EA. This is represented by the impulse responses to the first shock in model (3.2). The reaction of the other endogenous variables tells us how a shock to the per capita output of an individual country or state is absorbed by the three risk-sharing channels. The total impact of the shock is normalised at every horizon, such that the reactions of the three channels plus the un-smoothed part add up to one hundred. Figure 3.1 displays the median impulse responses on impact (in blue), and the cumulative effect four years after the shock (in red) in the US. Shaded areas are the 16^{th}

[12] The results for the pre-2022 sample do not change significantly when we drop 2022 and 2023 from the analysis. Note that these years were not available in the US dataset at the time of the analysis.

[13] Ireland is excluded from the analysis owing to unusually large revisions of some of the country's main macroeconomic statistics for 2015 that were undertaken in July 2016. These revisions affected real GDP, some of its components and balance of payments figures.

and 84^{th} percentiles of the posterior distribution.[14]

The bottom right panel of the figure shows that the total amount of shock absorption has increased since the 1970s, which is reflected in a declining share of unsmoothed shocks over the sample. At the start of the sample, a drop in a state's output translated into a 35% drop in consumption in the year of the shock. In the mid-2010s, this decreased to around 20%, before picking up again somewhat at the end of the sample. We find less smoothing in the long run. Only 40% of an output shock was smoothed four years after the shock in the mid-1970s, which has increased to around 60% towards the end of our sample. This increase was mostly due to private risk-sharing channels, and, in particular, to the capital channel, which operates through cross-state factor income. On impact, the capital channel shows a sharp increase until the mid-1980s, and large fluctuations around 40% afterward. The last decade of the sample highlights a increase in shock absorption via this channel.

Figure 3.1 (blue line) also shows that the role of the credit channel has decreased from about 30% to 10% in the mid-1980s, and starts increasing again afterward, although with significant fluctuations. Interestingly, this positive trend, which is even more evident in the long run (red line), seems to accelerate with the Riegle-Neal Interstate Banking and Branching Efficiency Act, signed in 1994 and implemented in 1997, allowing the opening of bank branches across state lines. One explicit goal of the de-regulation was to allow banks to diversify geographic risk. The Act improved bank efficiency and increased state-level per capita growth of personal income and GDP (Jayaratne and Strahan, 1997 and Aguirregabiria et al., 2010). The effectiveness of the credit channel then decreases in the last part of the sample, coinciding with the Fed's tapering phase, which is generally considered to start in 2014.

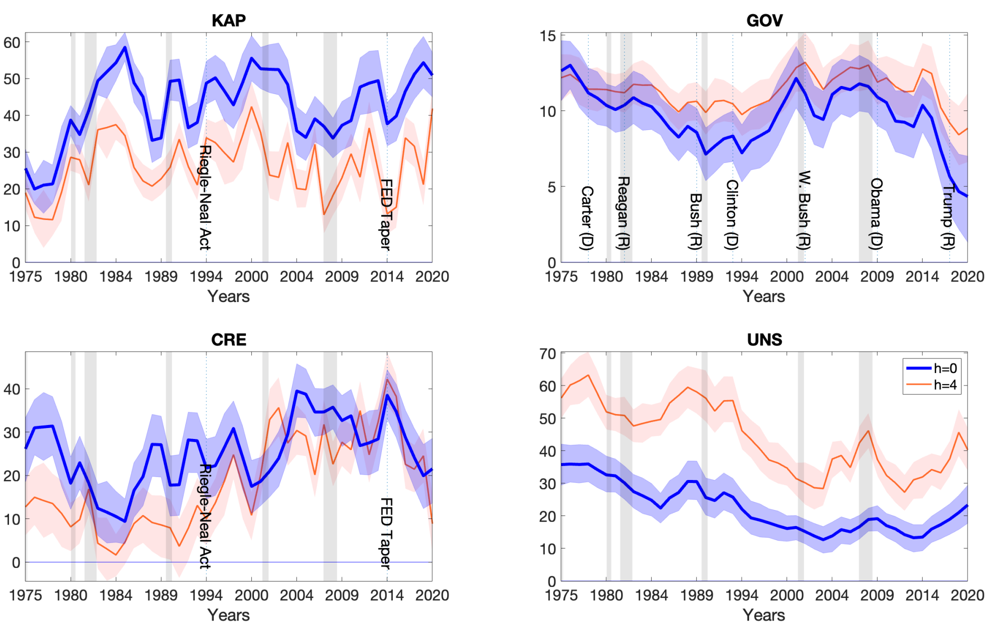
The absorption capacity of the fiscal channel is above 10% in the beginning of the sample and then decreases to around 7% by the mid-1990s. The fiscal channel improves again until the GCF in 2008-2009 and then declines, in particular as of 2015, possibly due to a deterioration of the fiscal space in the US.[15] Finally, it is worth noticing that the overall effectiveness of the three risk-sharing channels is higher on impact relative to the long run, but there are interesting differences. For instance, the capital channel, and,

[14] Figure C3 of the Appendix displays the surface plots of the median cumulative impulse responses over time. In this section, we report the normalized responses to facilitate the interpretation of the results, and directly quantify the relative importance of each risk-sharing channel. Figure C5 shows the estimated stochastic volatility series for the US.

[15] In Section 3.4.3, we explore potential determinants of the risk-sharing channels, and find a negative association between the fiscal channel and the federal debt-to-GDP ratio. This could be related to the fact that a higher debt implies less fiscal space and capacity to finance counter-cyclical transfer policies.

although to a lesser extent, the credit channel, have a larger smoothing role on impact. On the other hand, the fiscal channel tend to have a larger role in the long run, possibly due to lagged and persistent effects of the federal fiscal transfers.

Figure 3.1: Impulse responses to a state-specific output shock in the US

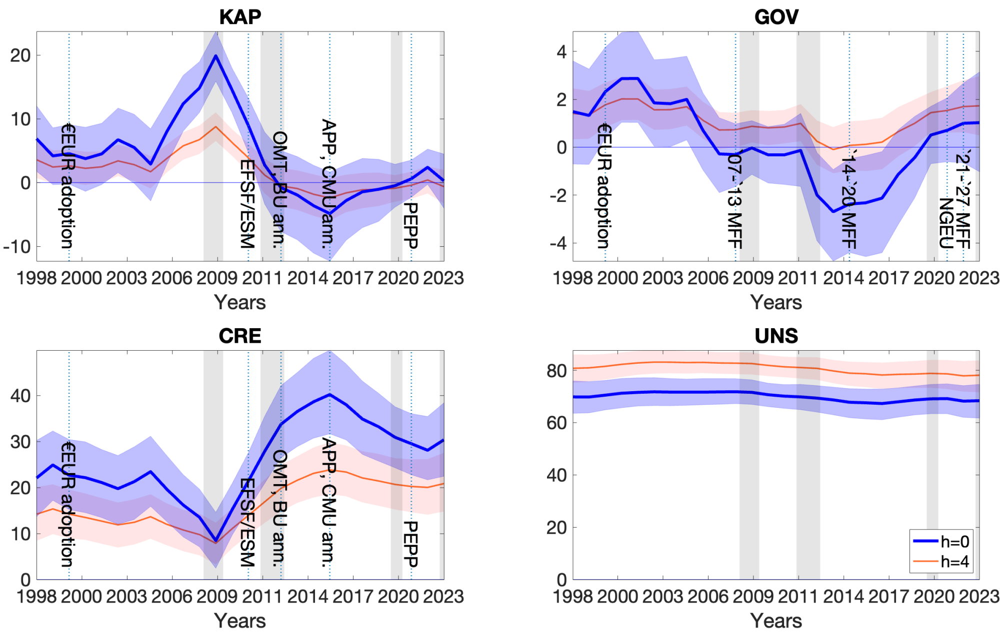


Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. Blue line: impact ($h = 0$) effect. Red line: cumulative effect four years ($h = 4$) after the shock. The effects of the shock are normalized to one hundred in each horizon. Shaded areas are the 16^{th} and 84^{th} percentile posterior bands around the median. Vertical shaded (grey) areas indicate NBER recessions, vertical dotted lines indicate the dates of the presidential elections, the start of the FED tapering, and the Riegle–Neal Interstate Banking and Branching Efficiency Act.

The normalized impulse responses on impact and at the four years horizon for the EA are displayed in Figure 3.2.[16] First, we observe less risk-sharing among EA countries than among US states (see also ECB, 2022). Around 70% of a country's output contraction translated into a drop in consumption in the year of the shock, and around 80% four years after the shock. Second, there is evidence of less time variation in the EA than in the US. Some improvement is noticeable, however, since the GFC of 2008-2009, as reflected in a slightly declining share of unsmoothed shocks.

[16] The surface plots of the cumulative impulse responses are provided in Figure C4 of the Appendix. Figure C6 shows the stochastic volatility series.

Figure 3.2: Impulse responses to a country-specific output shock in the EA



Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. Blue line: impact ($h = 0$) effect. Red line: cumulative effect four years ($h = 4$) after the shock. The effects of the shock are normalized to one hundred in each horizon. Shaded areas are the 16^{th} and 84^{th} percentile posterior bands around the median. Vertical shaded (grey) areas indicate recessions. Dotted lines indicate the dates of the activation of the EFSF/ESM support programs, the APP and PSPP of the ECB, the NGEU program and the years of the EU multi-annual financial framework (MFF) budget.

The most striking result is in the development of the capital and the credit channels, which evolve in opposite directions. On impact, the capital channel smooths around 5% at the beginning of the sample, and then increases remarkably until the peak of the 2008-2009 GFC, which probably reflects the stronger capital integration among countries following the introduction of the monetary union. This also possibly shows the ex-ante nature of the capital channel, as portfolio allocations are made prior to the realization of a (negative) output shock, and may revert after such shock. The capital channel then collapses during the GFC and thereafter declines sharply to even negative (although statistically insignificant) values in mid-2010s. The steady decline after 2009 can be due to flight-to-safety capital flows. Indeed, countries that were hit harder by the crisis (such as EA periphery countries) also experienced capital outflow towards countries that were less affected (such as EA core countries).

The credit channel resembles the mirror image of the capital channel. It accounts for the largest share of smoothing on impact (around 20%) at the beginning of the sample, and, after a drop ending in 2009, it shows a steep increase reaching 40% in 2015 and stabilizing at around 30% towards the end of our sample. The improvement after 2009 could be at least in part attributed to official assistance programs of the International Monetary Fund (IMF) and the ESM (see, also, Cimadomo et al., 2020). There is some evidence of a strengtening of the credit channel during Covid-19, which could be due to the NGEU funds and its Recovery and Resilience Facility (RRF) programs. The effect of these official programs is reflected in both the credit and the fiscal channel, as they have a loan and as well as a grant component (Giovannini et al. (2022)). The ECB's pandemic emergency purchase programme (PEPP) may have also prevented a freeze in the interbank credit market in this period. The role of the fiscal channel is much smaller. Interestingly, it has dropped to sligthly negative values between 2012 and 2017, which is evidence of dis-smoothing. Some improvement is visible from the mid-2010s, which could be attributed to a more fiscally ambitious EU multi-annual financial framework (MFF) for the period 2014-2020, entailing larger fiscal transfers between countries, and to fiscal grants under the EU's RRF. We also find that, similarly to the results for the US, whereas the private risk-sharing channels show the strongest smoothing role on impact, the effects of the fiscal channel are somewhat larger in the long run.

All in all, we uncover significant time variation in the risk-sharing channels and some improvement in smoothing in both regions.[17] These effects are mainly driven by private risk-sharing channels both in the US and the EA. However, while in the US the capital channel is the most important, we find that the credit channel tends to dominate in the EA, especially in the last part of the sample. We also find that the fiscal channel plays a more important role in the US relatively to the EA, but it has lost power in the last half of the 2010s, possibly due to a reduced fiscal space in this country.

3.4.2 Complementarity vs. substitutability

This section is devoted to study the interaction between the three risk-sharing channels. Uncovering these relationships is important for two main reasons. First, our results show that capital markets can freeze during a financial crisis, which leads to fragmentation and flight to safety, seriously impeding risk-sharing. Therefore, it is useful to understand how the other channels respond, when one channel is less operative. Second, the interactions

[17] These results are robust to estimating the model with endogenous variables demeaned by their respective simple averages (see Figures C7 and C8 of the Appendix), and using two lags instead of one (Figures C9 and C10 of the Appendix).

between the risk-sharing channels are specifically important in the context of a monetary union, where the single states or countries lose their monetary independence and exchange rate flexibility to counteract asymmetric shocks. In this context, it is crucial to understand how institutional reinforcing one channel (e.g., introduction of a Fiscal Union in the EA, or expansion of the federal budget in the US) affect the functioning of the other channels.

The hypothesis of complementary vs. substitutability between risk-sharing channels has not been formally tested in the literature, except for Asdrubali and Kim (2004). Using a panel of the US and other OECD countries for the period 1960-1990, they test complementarity by estimating the effect of a shock to one risk-sharing channel on the other channels. Here, we propose an alternative approach. In particular, our approach is based on the correlations between the impulse responses of the three risk-sharing channels to the idiosyncratic output shock. The interpretation is also different from the test proposed by Asdrubali and Kim (2004): in our case, we test how the various channels have co-moved, or diverged, *conditional* on an output shock, instead of looking how a channel responds to a shock to another channel, whose intuition is less clear in our view.

More specifically, we take each unscaled impulse response draw, and calculate the correlations among the three risk-sharing channels.[18] A negative correlation indicates that when one channel improves (deteriorates), the other channel tends to lose (gain) importance. This can be interpreted as evidence of substitution between the two channels. On the other hand, a positive correlation between the two channels is evidence of complementarity: one channel tends to improve when the effectiveness of the other channel(s) also increases. Such a test based on the impulse response correlations is a novelty which is allowed by our time-varying parameter framework.

Table 3.1 shows the correlations among the risk-sharing channels in the US. The top panel refers to the full sample, while below we present the results for three sub-samples of 14 years each. The left side of the table reports the correlations of the impulse responses in the year of the shock ($h = 0$), while the right side refers to the cumulative impulse responses four years after the shock ($h = 4$). The correlation between the capital and the credit channels is negative in the full sample both in impact and in the long run. This suggests substitution between the two channels and potential crowding-out effects. An alternative explanation for a negative correlation is the spare-tire hypothesis suggested by Levine et al. (2016), who postulates that stock markets can mitigate the effects of a banking crisis through equity financing. Interestingly, the substitution effect between

[18] The Gibbs sampler allows us to derive a empirical distribution of the impulse responses, therefore also of the correlations among risk sharing channels.

these two channels seem to be attenuated over time, as it was very strong in the 1976-1990 window, and becomes progressively less strong over the subsequent two windows, ending in 2020.

Table 3.1: Correlations among the risk-sharing channels in the United States

	On impact ($h = 0$) KAP	GOV	CRE		Long run ($h = 4$) KAP	GOV	CRE
Full sample							
KAP	1			KAP	1		
GOV	0.23* (0.07, 0.38)	1		GOV	0.34* (0.15, 0.51)	1	
CRE	-0.48* (-0.63, -0.32)	0.40* (0.20, 0.58)	1	CRE	-0.60* (-0.72, -0.44)	0.06 (-0.18, 0.31)	1
1975 - 1990							
KAP	1			KAP	1		
GOV	0.10 (-0.17, 0.38)	1		GOV	0.31* (0.00, 0.56)	1	
CRE	-0.72* (-0.84, -0.54)	0.30 (-0.03, 0.58)	1	CRE	-0.58* (-0.78, -0.27)	0.16 (-0.20, 0.49)	1
1991 - 2005							
KAP	1			KAP	1		
GOV	0.32* (0.01, 0.59)	1		GOV	0.19 (-0.20, 0.51)	1	
CRE	-0.54* (-0.74, -0.25)	0.14 (-0.21, 0.45)	1	CRE	-0.61* (-0.78, -0.36)	0.09 (-0.32, 0.50)	1
2006 - 2020							
KAP	1			KAP	1		
GOV	0.07 (-0.27, 0.40)	1		GOV	0.05 (-0.27, 0.36)	1	
CRE	-0.03 (-0.38, 0.32)	0.87* (0.76, 0.93)	1	CRE	-0.56* (-0.73, -0.31)	0.68* (0.45, 0.83)	1

Notes: The left (right) panel shows the correlations of the unscaled impact (four-year cumulative) impulse responses to a state-specific output shock. For each draw of coefficients, we calculate the time series of the impulse response of each channel to the state-specific output shock, and then calculate the correlation between the three time series. The values reported are the medians, and the 16^{th} and 84^{th} percentiles (in parenthesis) of these correlations. Star denotes if zero is outside of the 16^{th}-84^{th} credible interval.

The correlation between the capital and the fiscal channels is found to be positive and statistically significant both on impact and in the long run in the full sample. Interestingly, for the full sample, we also find a positive correlation between the credit and the fiscal channel on impact. This relationship seems to be dominated by the last sub-sample (2006-2020) which shows a statistically significant correlation both on impact and in the long run. To the best of our knowledge, this is the first empirical evidence supporting the complementarity hypothesis between private and public risk-sharing channels outlined by Farhi and Werning (2017), Buti and Carnot (2018) and Giovannini et al. (2022).

The correlations between the risk-sharing channels in the EA are displayed in Table 3.2. These findings are somewhat less clear-cut than for the US, also due to a smaller sample size, but some observations are still worth noticing. First, we still find a substitution between the capital and the credit channels, although this relationship is statistically weak. We also document a positive correlation between the credit and the fiscal channels in the long run, although this is statistically significant only in the post-2010 sub-sample. This was the period in which the EFSF/ESM support programs where lunched for some EA countries under financial stress, and where the 2014-2020 multi-annual financial framework (MFF) budget was approved in the EU. Both programmes had a loan component (which is reflected in the credit channel) and a grant component (which affects the fiscal channel), therefore the two channels appeared to reinforce each other.

Overall, in this section we provide two main findings. Firstly, in the US and, although to a lesser extent, in the EA, the capital and credit risk-sharing channels act as substitutes. Secondly, for the US we also document some complementarity between the private and public risk-sharing channels, a finding which is consistent with Farhi and Werning (2017), Buti and Carnot (2018) and Giovannini et al. (2022).

Table 3.2: Correlations among the risk-sharing channels in the euro area

Full sample							
	On impact ($h = 0$)				Long run ($h = 4$)		
	KAP	GOV	CRE		KAP	GOV	CRE
KAP	1			KAP	1		
GOV	0.01 (-0.27, 0.29)	1		GOV	0.21 (-0.12, 0.51)	1	
CRE	-0.32 (-0.57, 0.00)	-0.23 (-0.48, 0.06)	1	CRE	-0.15 (-0.47, 0.21)	0.12 (-0.22, 0.45)	1
1998 - 2010							
	On impact ($h = 0$)				Long run ($h = 4$)		
	KAP	GOV	CRE		KAP	GOV	CRE
KAP	1			KAP	1		
GOV	-0.42 (-0.72, 0.00)	1		GOV	0.00 (-0.49, 0.53)	1	
CRE	-0.30 (-0.70, 0.33)	-0.04 (-0.48, 0.41)	1	CRE	0.17 (-0.46, 0.66)	0.10 (-0.36, 0.54)	1
2010 - 2023							
	On impact ($h = 0$)				Long run ($h = 4$)		
	KAP	GOV	CRE		KAP	GOV	CRE
KAP	1			KAP	1		
GOV	0.09 (-0.32, 0.46)	1		GOV	0.26 (-0.16, 0.59)	1	
CRE	-0.11 (-0.43, 0.26)	-0.01 (-0.35, 0.33)	1	CRE	0.03 (-0.33, 0.40)	0.45* (0.08, 0.70)	1

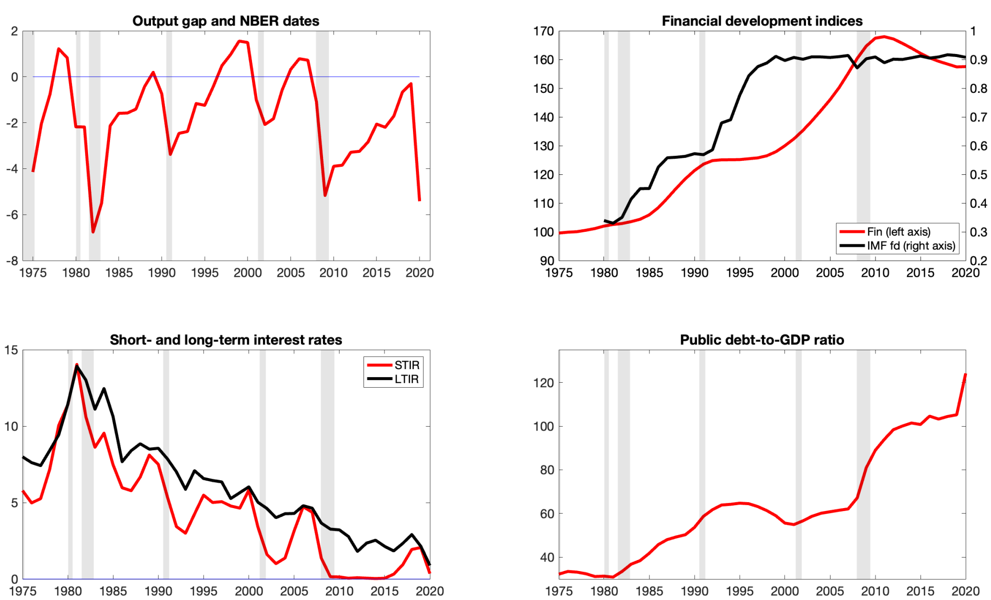
Notes: The left (right) panel shows the correlations of the unscaled impact (four-year cumulative) impulse responses to a country-specific output shock. For each draw of coefficients, we calculate the time series of the impulse response of each channel to the country-specific output shock, and then calculate the correlation between the three time series. The values reported are the medians, and the 16^{th} and 84^{th} percentiles (in parenthesis) of these correlations. Star denotes if zero is outside of the 16^{th}-84^{th} credible interval.

3.4.3 Determinants of risk-sharing channels

In this final section, we attempt to explain the time variation in the responses of the risk-sharing channels to output shocks with some macroeconomic and financial determinants, motivated by the existing literature. The goal of this exercise is not to provide causal evidence of the drivers of risk-sharing, but rather to help us in building a narrative and to complement, within our time-varying framework, the findings of the existing literature. We focus on the US as this allows us to exploit 45 years of observations, whereas, for the EA, we have only two decades, which limit the feasibility of such an analysis. We investigate the role of some key macroeconomic and financial determinants, described at length in Section 3.2, in affecting the effectiveness of our three risk-sharing channels based a system of seemingly unrelated regressions (SUR), estimated with Bayesian methods. Our dependent variables are the unscaled median impulse responses of the three risk-sharing channels to the idiosyncratic output shock. Estimating these relationships in a system allows us to account for the correlation among the risk-sharing channels documented in the previous section, and improve the efficiency of the estimates. In order to account for the estimation uncertainty of the dependent variables in the system, we follow Mumtaz and Sunder-Plassmann (2021), and weight the observations with the inverse of the posterior variance of the risk-sharing channels. Intuitively, this amounts to a weighted least squares, that down-weights observations where the estimation uncertainty is large.

We focus on four main (sets of) explanatory variables, which are displayed in Figure 3.3: business cycle and financial development indices, short- and long-term interest rates (as proxies for the stance of monetary policy), and the public debt-to-GDP ratio (as a synthetic indicator for the state of public finances). The rationale of using this set of variables has been discussed in Section 3.2. First, as previous studies find that risk-sharing tends to vary with economic activity (Hoffmann and Shcherbakova-Stewen, 2011 and Furceri and Zdzienicka, 2015), we take the output gap as a measure of the stance of business cycle. If risk-sharing through a specific channel increases in weak (strong) economic conditions, the coefficient on the output gap variable is expected to be negative (positive). The existing literature also argues that financial integration, deregulation, and innovation are expected to influence the overall level of risk-sharing (Athanasoulis and Wincoop, 2001, Kose et al., 2009, Demyanyk et al., 2007 and Demyanyk et al., 2008). Therefore, following Schularick and Taylor (2012) who identify financial progress as the main determinant of the long-run trending behaviour of credit in advanced economies, we add the long-run trend of the credit-to-GDP series from the BIS as a proxy of financial development. In Figure 3.3 we observe that this series is increasing during most of the sample, albeit with a plateau in the early 1990s and a decrease after 2010.[19] In order to capture the stance of monetary policy and to proxy the overall interest-rate environment in the US, we use the the 3-month Treasury yield (STIR), and, as an alternative measure, the 10-year Treasury yield (LTIR). A higher interest-rate environment may raise the marginal cost of smoothing a negative shock to a state's domestic product, as households and firms will need to borrow at higher rates. Similarly, it may also raise the cost of debt-financed fiscal stabilization. Lastly, as a measure of fiscal space, we include the US federal nominal debt-to-GDP ratio (Debt). Fiscal space may affect the capacity of the federal government to finance counter-cyclical policies (Stempel et al., 2021). As a result, we expect that with a lower (higher) debt-to-GDP ratio, the federal government will have more (less) fiscal space to absorb shocks. Figure 3.3 shows that, with the only exception of the period from 1995 to 2001, the US federal debt-to-GDP ratio steadily increased over the sample, from just above 30% in the mid-1970s to around 120% in 2020.

Figure 3.3: Time series of the determinants



Note: OG: output gap; NBER: annualised NBER recession indicator; Fin: Long-run credit-to-GDP trend obtained from the BIS; IMF fd: IMF financial development index; STIR: 3-month Treasury yield; LTIR: 10-year Treasury yield; Debt-to-GDP ratio: US federal nominal debt-to-GDP ratio. Vertical shaded areas indicate NBER recessions.

[19] We also experimented with an alternative measure for financial innovation, that is the Financial Markets Development composite index of the IMF, which, as Figure 3.3 shows, is only available as of 1980. These two financial innovation measures are highly positively correlated with each other (0.83).

Our baseline regression results are displayed in Table 3.3. To address the potential simultaneity bias between the dependent and the independent variables, we use the lagged values of the predictors, which in practice can be interpreted as the initial conditions before a state-specific output shock hits. The first three columns show the regression results where the dependant variable is the impact ($h = 0$) response of the risk-sharing channels to the output shock, whereas the last three columns display the estimates for the long-run (cumulated) responses four years ($h = 4$) after the shock. We find that the output gap is negatively associated with both the capital and fiscal risk-sharing channels both on impact and after four years. This result is consistent with the capital markets and the federal government providing additional risk-sharing during downturns.[20] We also find that financial integration/innovation seems to have a positive relationship with the the fiscal and the credit channels. Our interpretation is that financial innovation (e.g., securitization of mortgage loans) improves agents' access to credit, which in turn strengthens risk-sharing through credit markets. The stance of the monetary policy does not seem to play a large role, with the only exception of the credit channel in the long run. More specifically, a higher interest-rate environment, which is generally associated with tighter credit conditions, leads to a lower role of the credit channel in the long run. Finally, we find that the debt-to-GDP ratio is negatively associated with the effectiveness of the government and the credit channels, both on impact and in the long run. The negative coefficient of the fiscal channel implies that a higher initial debt level (or less fiscal space) reduces risk-sharing via the federal tax and transfer system. This is consistent with the idea that more fiscal space leads to more capacity to implement counter-cyclical stabilization. A higher debt level also drives up interest rates in the economy, which makes smoothing through credit markets more expensive and results in less risk-sharing via that channel.

Our results are robust to the use of alternative measures of the stance of the business cycle, the interest-rate environment, and the financial development. More specifically, in Table C1 of Appendix C we replace the output gap with a dummy based on NBER recession dates. The dummy tracks the number of quarters within the year that are labeled as a recessions, e.g., it takes a value of 1 is the economy was in recession of the full year, 0.75 if the economy was in recession in three out of four quarters in a given year, etc. In Table C2, we use long-term Treasury yield instead of the short-term interest rate, and in Table C3 we test for the Financial Markets Development composite index of the IMF instead of the BIS credit-to-GDP trend. Results are qualitatively similar to what we find

[20] This finding is also consistent with the empirical literature documenting larger fiscal multipliers during recessions (see e.g. Auerbach and Gorodnichenko (2012b) among others).

in Table 3.3.

Table 3.3: Determinants of the risk-sharing channels

	Impact ($h = 0$)			Long-run ($h = 4$)		
	KAP	GOV	CRE	KAP	GOV	CRE
Gap	-0.15*	-0.06*	0.07	-0.28*	-0.11*	-0.09
	(-0.26, -0.05)	(-0.08, -0.04)	(-0.01, 0.15)	(-0.41, -0.16)	(-0.14, -0.07)	(-0.17, 0.01)
Fin	0.01	0.01*	0.03*	0.01	0.01	0.08*
	(-0.00, 0.03)	(0.01, 0.02)	(0.01, 0.05)	(-0.01, 0.03)	(-0.00, 0.01)	(0.07, 0.10)
STIR	0.02	0.02	0.07	0.08	0.02	-0.19*
	(-0.08, 0.13)	(-0.00, 0.04)	(-0,.00 0.16)	(-0.05, 0.22)	(-0.02, 0.06)	(-0.29, -0.10)
Debt	-0.01	-0.02*	-0.04*	-0.01	-0.02*	-0.09*
	(-0.03, -0.00)	(-0.02, -0.01)	(-0.05, -0.02)	(-0.02, 0.01)	(-0.02, -0.01)	(-0.11, -0.08)

Notes: Bayesian weighted SUR regression with flat priors, and weights given by the inverse of the posterior variance of the risk-sharing channels. Dependent variables are the unscaled median impulse responses of the risk-sharing channels to the idiosyncratic output shocks, at horizon $h = 0$ (left panel) or cumulated at horizon $h = 4$ (right panel). Explanatory variables are the following. Gap: the output gap; Fin: long-run credit-to-GDP trend; STIR: 3-months Treasury yield; Debt: debt-to-GDP ratio. they are lagged at time $t - 1$. Constant omitted from the table. Stars denote if posterior median is outside of the 16^{th}-84^{th} percentile credible interval.

3.5 Conclusions

Focusing on the US and the EA, this paper offers new insights on how idiosyncratic output shocks, i.e., shocks which are not common to all states (US) or countries (EA), are absorbed by private and public risk-sharing channels. In particular, we look at the capital channel (income from cross-ownership of productive assets), the credit channel (cross-state or cross-border lending and borrowing), and the fiscal channel (federal or cross-country transfers). We assess how these three risk-sharing channels have operated over time within each region, using a novel time-varying parameter panel VAR model with stochastic volatility, which generalises Mumtaz and Sunder-Plassmann (2021).

Our analysis allows us to uncover that, over the last decades, the overall level of risk-sharing has clearly increased in the US and also, though to a smaller extent, in the EA. This improvement is mostly due to private risk-sharing channels. However, while in the US the role of the capital channel is the most sizable, we find that in the EA the credit channel dominates. Due to the presence of a larger federal system, we also find that the fiscal channel plays a more important role in the US relatively to the EA. Interestingly, we document that smoothing role of the private risk-sharing channels is larger on impact than in the long run, wheares the opposite is found for the fiscal channel.

In the second part of the paper, we study the degree of substitution and complementarity between the three risk-sharing channels. We find strong substitution effects between capital market smoothing and credit market smoothing in both regions. This can imply either crowding out between these two channels, or support for the spare-tire hypothesis, postulating that stock markets can mitigate the effects of a banking crisis. In the case of the US, we also find evidence for complementarity between the private (e.g. credit and capital) and the public (e.g. fiscal) channels, which supports the argument of Farhi and Werning (2017).

Finally, for the US, we attempt to explain the time variation in the responses of the risk-sharing channels to output shocks with some macroeconomic and financial determinants. We show that the effectiveness of both the capital and the fiscal risk-sharing channels improves during weak economic conditions. We also find that the fiscal and credit channels work better under stronger financial integration, and when a country is characterised by more fiscal space. At the same time, monetary policy does not seem to be very powerful in influencing the functioning of risk-sharing channels.

Our results have important implications for the improvement of the European Monetary Union's governance framework. We show that there is substantial room to strengthen the shock absorption capacity of the EMU. The Capital Market Union and the Credit Market Union could reinforce risk-sharing and improve the resilience of the EA to macroeconomic shocks by contributing to smooth output shocks through the capital channel and the credit channel. The US experience also points to positive externalities between the fiscal channel and the private risk-sharing channels. Looking at this finding from the perspective of the EA, this suggests that progress on the Fiscal Union could also have a beneficial effect on private risk-sharing mechanisms in the EMU.

Appendix A Deriving the risk-sharing coefficients

Our starting point is the methodology of Asdrubali et al. (1996), who decompose the cross-sectional variance of shocks to output. Consider the identity for the EA, that holds for each period t, suppressing the time index:

$$GDP_i = \frac{GDP_i}{GNP_i}\frac{GNP_i}{GDI_i}\frac{GDI_i}{C_i}C_i$$

where all the variables are in per-capita deviations from their respective regional aggregate values. Taking logs, differencing both sides and taking expectations leads to the cross-sectional decomposition of the variance in GDP:

$$\begin{aligned} var\{\Delta logGDP_i\} &= cov\{\Delta logGDP_i - \Delta logGNP_i, \Delta logGDP_i\} \\ &+ cov\{\Delta logGNP_i - \Delta logGDI_i, \Delta logGDP_i\} \\ &+ cov\{\Delta logGDI_i - \Delta logC_i, \Delta logGDP_i\} \\ &+ cov\{\Delta logC_i, \Delta logGDP_i\} \end{aligned}$$

Then, dividing by $var\{\Delta logGDP_i\}$ yields:

$$1 = \beta_{KAP} + \beta_{GOV} + \beta_{CRE} + \beta_{UNS}$$

where for example

$$\beta_{KAP} = \frac{cov\{\Delta logGDP_i - \Delta logGNP_i, \Delta logGDP_i\}}{var\{\Delta logGDP_i\}}$$

is the cross-sectional OLS slope coefficient in the regression of $\Delta logGDP_i - \Delta logGNP_i$ on $\Delta logGDP_i$. It is interpreted as the percentage of smoothing of an output shock, achieved through international factor income.

Appendix B Estimation

B.1 Priors and starting values

The following steps describe the setting of the priors and starting values.

1. $p\left(b_i|\bar{b}, \lambda\right)$. We assume a hierarchical prior for b_i centered on the weighted cross-

section average $\bar{b}$

$$p\left(b_i|\bar{b},\lambda\right) \sim N\left(\bar{b},\lambda\Lambda_i\right)$$

where Λ_i is set according to the Minnesota procedure. The parameter λ controls the degree of pooling in the model. As $\lambda \to 0$ the heterogeneity across states declines. In order to set the variances Λ_i, we use dummy observations as in Banbura et al. (2010), setting the overall prior tightness parameter to 1.

2. $p(\lambda)$. This prior is inverse Gamma: $p(\lambda) \sim IG(S_0, V_0)$ where $S_0 = 0$ and $V_0 = -1$. As discussed in Gelman (2006), this prior corresponds to a uniform prior on the standard deviation.

3. $p(Q_B)$. This prior is inverse Wishart $IW(Q_0, T_0)$ As common in the literature on time-varying VARs, the scale matrix is set based on a pre-sample of T_0 observations. Let Q_{ols} denote the average of the OLS estimate of coefficient covariance across countries using the pre-sample. We set $Q_0 = Q_{ols} \times T_0 \times \kappa$ where $\kappa = 0.003$. Let β_{ols} denote the average OLS estimate of the coefficients using the pre-sample. Then the initial state is given as: $\beta_{0\backslash 0} \sim N(\beta_{ols}, Q_{ols})$

4. $p(a_i|\bar{a},\delta)$. We set a hierarchical prior:

$$p(a_i|\bar{a},\delta) \sim N(\bar{a},\delta\Xi_i)$$

where $\bar{a}$ are weighted cross-sectional averages and Ξ_i equals a matrix that reflects the relative scale of the residuals of the model. The degree of pooling is controlled by δ.

5. $p(\delta)$. As in step 3 the prior is inverse Gamma $p(\delta) \sim IG(s_0, v_0)$ where $s_0 = 0$ and $v_0 = -1$.

6. $p(Q_A)$. This prior is inverse Wishart $IW(Q_{A,0}, T_{A,0})$. $Q_{A,0}$ is set as a block diagonal matrix. Each block corresponds the time-varying elements in each row of A_{it}. The diagonal elements for each block are set equal to 0.001. The degrees of freedom $T_{A,0}$ are set equal to the rows of the block plus 1 indicating a non-informative prior. The initial values $a_{0\backslash 0}$ are set using OLS estimates as in step 3 above.

7. Starting values for H_t are obtained as $\left(e_t^{ols}\right)^2 + 0.001$ where e_t^{ols} denote the OLS estimates of the residuals of the panel VAR assuming fixed coefficients. A small scaling factor is added to the squared residuals to remove zeros. The initial conditions

are $\ln h_{0\backslash 0} \sim N(\bar{\mu}_i, s)$ where $\bar{\mu}_i$ are the diagonal elements of the VAR error covariance estimated via OLS on the pre-sample explained in step 3.

8. $p(g_i)$. The prior is inverse Gamma: $IG(g_0, d_0)$ where $g_0 = 0.01$ and $d_0 = 1$.

9. $p(\sigma_i^2)$. The prior is inverse Gamma: $IG(\sigma_0, D_0)$ where $\sigma_0 = 0.01$ and $D_0 = 1$.

B.2 Gibbs sampling algorithm

The Gibbs sampling algorithm involves sampling from the conditional posterior distributions described below. The Gibbs sampler is based on the sampler for panel VARs described in Canova and Ciccarelli (2004) and Jarocinski (2010). In addition, it features elements of the sampler for TVP-VARs with stochastic volatility described in Cogley and Sargent (2005) and Primiceri (2005). Note that Ψ denotes all remaining parameters.

1. $H(a_t|\Psi)$. Given the coefficients B_{it} and the stochastic volatilities H_{it}, the model can be written as

$$A_{it}u_{it} = H_{it}^{1/2}\tilde{e}_{it}$$
$$\tilde{e}_{it} \sim N(0,1)$$

where $u_{it} = Y_{it} - X_{it}B_{it}$ and $u_{it} = [u_{1,it}, .., u_{N,it}], \tilde{e}_{it} = [\tilde{e}_{1,it}, .., \tilde{e}_{N,it}], H_{it} = diag(h_{it})$ where $h_{it} = [h_{1,it}, .., h_{N,it}] = [\sigma_{1,i}^2 h_{1,t}, .., \sigma_{N,i}^2 h_{N,t}]$. For each cross-sectional unit, this represents a system of equations:

$$\begin{aligned}
u_{1,it} &= (h_{1,it})^{1/2}\,\tilde{e}_{1,it} \\
u_{2,it} &= -a_{(2,1),it}u_{1,it} + (h_{2,it})^{1/2}\,\tilde{e}_{2,it} \\
u_{3,it} &= -a_{(3,1),it}u_{1,it} - a_{(3,2),it}u_{2,it} + (h_{3,it})^{1/2}\,\tilde{e}_{3,it} \\
&. \\
&. \\
u_{N,it} &= \sum_{j=1}^{N-1} -a_{(N,j),it}u_{j,it} + (h_{4,it})^{1/2}\,\tilde{e}_{N,it}
\end{aligned}$$

The first equation is an identity and can be ignored. Note that as we condition on the unit specific coefficients a_i, this system can be written as:

$$u_{1,it} = \left(h_{1,it}\right)^{1/2} \tilde{e}_{1,it} \tag{B1}$$

$$u_{2,it} - \left(-a_{(2,1),i} u_{1,it}\right) = -a_{(2,1),t} u_{1,it} + \left(h_{2,it}\right)^{1/2} \tilde{e}_{2,it} \tag{B2}$$

$$u_{3,it} - \left(-a_{(3,1),i} u_{1,it} - a_{(3,2),i} u_{2,it}\right) = -a_{(3,1),t} u_{1,it} - a_{(3,2),t} u_{2,it} + \left(h_{3,it}\right)^{1/2} \tilde{e}_{3,it} \tag{B3}$$

$$.$$

$$.$$

$$u_{N,it} - \left(\sum_{j=1}^{N-1} -a_{(N,j),i} u_{j,it}\right) = \sum_{j=1}^{N-1} -a_{(N,j),t} u_{j,it} + \left(h_{4,it}\right)^{1/2} \tilde{e}_{N,it} \tag{B4}$$

That is, unit specific effects can be subtracted out leaving the common time-varying component. Given the assumption that Q_A is block diagonal, each equation can then be stacked across the cross-sectional units (as the time-varying coefficients are common across units) and has the following state space representation:

$$\tilde{U}_{m,t} = \sum_{j=1}^{m-1} -a_{(m,j),t} U_{j,t} + V_{m,t} \tag{B5}$$

$$a_{m,t} = a_{m,t-1} + v_{m,t}$$

where m denotes the mth equation and : $\tilde{U}_{m,t} = \begin{pmatrix} u_{m,1t} - \left(\sum_{j=1}^{m-1} -a_{(m,j),1} u_{j,1t}\right) \\ . \\ . \\ u_{m,Mt} - \left(\sum_{j=1}^{m-1} -a_{(m,j),M} u_{j,Mt}\right) \end{pmatrix}$, $U_{j,t} = \begin{pmatrix} u_{j,1t} \\ . \\ . \\ u_{j,Mt} \end{pmatrix}$, $V_{m,t} = \begin{pmatrix} \left(h_{m,1t}\right)^{1/2} \tilde{e}_{m,1t} \\ . \\ . \\ \left(h_{m,Mt}\right)^{1/2} \tilde{e}_{m,Mt} \end{pmatrix}$ and $a_{m,t} = [a_{(1,j),t}, .., a_{(m-1,j),t}]$ is the vector of coefficients. The variance of $v_{m,t}$ is the mth diagonal block of Q_A. Given the conditionally linear and Gaussian state-space model in equation B5, the Carter and Kohn (1994) algorithm can be used to draw from the conditional posterior of $a_{m,t}$. This is repeated for each equation in the system.

2. $H\left(Q_A|\Psi\right)$. Note that vector $a_{j,t}$ denotes the time-varying coefficients of the jth equation. For example $a_{3,t} = [a_{(3,1),t}, a_{(3,2),t}]$ in equation B3. Given a draw for $a_{j,t}$the

jth diagonal block of Q_A has an inverse Wishart conditional posterior:

$$IW((a_{j,t} - a_{j,t-1})' (a_{j,t} - a_{j,t-1}) + Q_{j,A,0}, T + T_{A,0})$$

where $Q_{j,A,0}$ denotes the jth diagonal block of Q_A.

3. $H(\beta_t|\Psi)$. Given a draw for the elements of A,the stochastic volatility H_{it} and the unit specific coefficients b_i, the panel VAR can be written as:

$$y_{it} = x_{it}\beta_t + \bar{e}_{it}$$

where $y_{it} = vec(Y_{it} - X_{it}b_i)$, x_{it} denotes the RHS variables stacked: $\begin{pmatrix} X_{1,t} \\ . \\ . \\ X_{M,t} \end{pmatrix}$ and $var(\bar{e}_{it}) = blkdiag\left([A_{1t}^{-1}H_{1t}^{1/2}A_{1t}^{-1\prime}, .., A_{Mt}^{-1}H_{Mt}^{1/2}A_{Mt}^{-1\prime}\right)$ where $H_{it} = diag(h_t\sigma_i^2)$. Given the transition equation $\beta_t = \beta_{t-1} + \eta_t, var(\eta_t)$, this represents a conditionally linear Gaussian state-space system and the Carter and Kohn (1994) algorithm can be used to draw β_t.

4. $H(Q_B|\Psi)$. Given a draw for β_t, the conditional posterior for Q_B is inverse Wishart:

$$IW((\beta_t - \beta_{t-1})' (\beta_t - \beta_{t-1}) + Q_0, T + T_0)$$

5. $H(b_i|\Psi)$ for $i = 1, 2, .., M$. Given a draw for the time-varying VAR coefficients β_t and the matrix A_{it} and the volatility H_{it}, the model for ith unit can be written as

$$\bar{Y}_{it} = X_{it}b_{it} + \bar{e}_{it}$$

where $\bar{Y}_{it} = Y_{it} - X_{it}\beta_t$ and $var(\bar{e}_{it}) = \Sigma_{it} = A_{it}^{-1}H_{it}A_{it}^{-1\prime}$. This is a VAR with heteroscedastic disturbances. After a GLS transformation, the conditional posterior is normal: $N(M, V)$

$$M = V\left(vec\left(\sum_{t=1}^{T}\left(X_{it}\bar{Y}_{it}'\Sigma_{it}^{-1}\right)\right) + (\lambda\Lambda_i)^{-1}\bar{b}\right)$$
$$V = \left(\sum_{t=1}^{T}\left(\Sigma_{it}^{-1} \otimes X_{it}X_{it}'\right) + (\lambda\Lambda_i)^{-1}\right)^{-1}$$

6. $H\left(a_i|\Psi\right)$ for $i=1,2,..,M$. For each unit, the model can be written in terms of the residuals:

$$A_{it}u_{it} = H_{it}^{1/2}\tilde{e}_{it}$$
$$\tilde{e}_{it} \sim N(0,1)$$

where $u_{it} = Y_{it} - X_{it}B_{it}$ and $u_{it} = [u_{1,it},..,u_{N,it}], \tilde{e}_{it} = [\tilde{e}_{1,it},..,\tilde{e}_{N,it}], H_{it} = diag\left(h_{it}\right)$ where $h_{it} = [h_{1,it},..,h_{N,it}] = [\sigma^2_{1,i}h_{1,t},..,\sigma^2_{N,i}h_{N,t}]$. As noted above, this is a system of equations:

$$\begin{aligned}
u_{1,it} &= h_{1,it}^{1/2}\tilde{e}_{1,it} \\
u_{2,it} &= -a_{(2,1),it}u_{1,it} + h_{2,it}^{1/2}\tilde{e}_{2,it} \\
u_{3,it} &= -a_{(3,1),it}u_{1,it} - a_{(3,2),it}u_{2,it} + h_{3,it}^{1/2}\tilde{e}_{3,it} \\
&. \\
&. \\
u_{N,it} &= \sum_{j=1}^{N-1} -a_{(N,j),it}u_{j,it} + h_{N,it}^{1/2}\tilde{e}_{N,it}
\end{aligned}$$

We can subtract out the impact of the common time-varying coefficients:

$$u_{1,it} = h_{1,it}^{1/2}\tilde{e}_{1,it} \tag{B6}$$
$$u_{2,it} - \left(-a_{(2,1),t}u_{1,it}\right) = -a_{(2,1),i}u_{1,it} + h_{2,it}^{1/2}\tilde{e}_{2,it} \tag{B7}$$
$$u_{3,it} - \left(-a_{(3,1),t}u_{1,it} - a_{(3,2),t}u_{2,it}\right) = -a_{(3,1),i}u_{1,it} - a_{(3,2),i}u_{2,it} + h_{3,it}^{1/2}\tilde{e}_{3,it} \tag{B8}$$
$$.$$
$$.$$
$$u_{N,it} - \left(\sum_{j=1}^{N-1} -a_{(N,j),t}u_{j,it}\right) = \sum_{j=1}^{N-1} -a_{(N,j),i}u_{j,it} + h_{N,it}^{1/2}\tilde{e}_{N,it} \tag{B9}$$

The first equation is redundant, while the remaining equations represent linear regressions with heteroscedasticity. For the mth equation let the dependent variable be denoted by $\breve{y}_m$ and the independent variables by $\breve{x}_m$, then the regression can be written as

$$\breve{y}_m = \breve{x}_m\breve{b}_m + h_{m,it}^{1/2}\tilde{e}_{m,it}$$

Let $\breve{y}_m^* = \frac{\breve{y}_m}{h_{m,it}^{1/2}}, \breve{x}_m^* = \frac{\breve{x}_m}{h_{m,it}^{1/2}}$ and the model can be written in terms of a unit variance error term. The conditional posterior for this linear regression is normal: $N(m, v)$

$$m = \left((\delta\Xi_i)^{-1} + \breve{x}_m^{*\prime}\breve{x}_m^{*}\right)^{-1}\left((\delta\Xi_i)^{-1}\bar{a} + \breve{x}_m^{*\prime}\breve{y}_m^{*}\right)$$
$$v = \left((\delta\Xi_i)^{-1} + \breve{x}_m^{*\prime}\breve{x}_m^{*}\right)^{-1}$$

7. $H\left(\sigma_i^2|\Psi\right)$ for $i = 1, 2, .., M$. For each unit, the model can be written in terms of the orthogonal residuals

$$A_{it}u_{it} = \breve{e}_{it}$$

where $\breve{e}_{it} = [\breve{e}_{1,it}, .., \breve{e}_{N,it}]$. Note that $var\left(\breve{e}_{it}\right) = diag\left(h_t\sigma_i^2\right)$ with $h_t = [h_{1,t}, .., h_{N,t}]$ and $\sigma_i^2 = [\sigma_{1,i}^2, .., \sigma_{N,i}^2$. A GLS transformation can remove the influence of h_t

$$\breve{e}_{it}^{*} = \frac{\breve{e}_{it}}{h_t^{1/2}}$$

and $\sigma_{m,i}^2$ can be drawn from the inverse Gamma distribution as :

$$IG\left(\breve{e}_{m,it}^{*\prime}\breve{e}_{m,it}^{*} + \sigma_0, T + D_0\right)$$

for $m = 1, .., N$.

8. $H\left(H_t|\Psi\right)$. The model for the ith unit can be written as

$$A_{it}u_{it} = H_t^{1/2}e_{it}$$

where $e_{it} = [e_{1,it}, .., e_{Nit}]$, $H_t = diag\left(h_t\right)$ with $h_t = [h_{1,t}, .., h_{N,t}]$ and $e_{it} \sim N(0, \sigma_i^2)$ with $\sigma_i^2 = [\sigma_{1,i}^2, .., \sigma_{N,i}^2]$. Stacking the mth orthogonal residual across units $m = 1, .., N$, we get the non-linear state space system:

$$\begin{pmatrix} h_{m,t}^{1/2}e_{m,1t} \\ . \\ . \\ h_{m,t}^{1/2}e_{m,Mt} \end{pmatrix}, var\begin{pmatrix} e_{m,1t} \\ . \\ . \\ e_{m,Mt} \end{pmatrix} = diag\left(\sigma_i^2\right)$$

$$\ln h_{m,t} = \ln h_{m,t-1} + \varepsilon_{m,t}, var\left(\varepsilon_{m,t}\right) = g_m$$

Thus each orthogonalised residual features common stochastic volatility. As in Cogley and Sargent (2005), we use the independence Metropolis algorithm of Jacquier et al. (2004) to sample each column of $h_{m,t}$.

9. $H(g|\Psi)$. Given a draw for $\ln h_t$, the variances $g = [g_1.., g_N]$ can be drawn one by one from the inverse Gamma distribution:

$$IG\left((\ln h_{m,t} - \ln h_{m,t-1})' (\ln h_{m,t} - \ln h_{m,t-1}) + g_0, T + d_0\right)$$

10. $H(\lambda|\Psi)$. The form of the conditional posterior is inverse Gamma with scale parameter $\sum_{i=1}^{M}\left(b_i - \bar{b}\right)\Lambda_i^{-1}\left(b_i - \bar{b}\right)' + S_0$ and degrees of freedom $\left(M \times \bar{K}\right) + V_0$ where $\bar{K} = N(Np+1)$

11. $H(\delta|\Psi)$. The form of the conditional posterior is inverse Gamma with scale parameter $\sum_{i=1}^{M}(a_i - \bar{a})\,\Xi\,(a_i - \bar{a})' + s_0$ and degrees of freedom $\left(M \times \tilde{K}\right) + v_0$ where $\tilde{K} = \frac{N(N-1)}{2}$

12. $H\left(\bar{b}|\Psi\right)$. By the Bayes Theorem, $H\left(\bar{b}|b_i, \lambda\right) \propto p\left(b_i|\bar{b}, \lambda\right) p\left(\bar{b}\right)$. This density is normal as $p\left(b_i|\bar{b}, \lambda\right)$ is normal and product of the normal priors for each i. With a flat prior for $\bar{b}$ this density is given by $N\left(\bar{M}, \bar{V}\right)$:

$$\bar{V} = \left(\frac{1}{\lambda}\sum_{i=1}^{M}\Lambda_i^{-1}\right)^{-1}$$
$$\bar{M} = \bar{V}\left(\frac{1}{\lambda}\sum_{i=1}^{M}\Lambda_i^{-1}b_i\right)$$

13. $H(\bar{a}|\Psi)$. As in step 11, this conditional posterior is normal $N(\bar{m}, \bar{v})$

$$\bar{v} = \left(\frac{1}{\delta}\sum_{i=1}^{M}\Xi_i^{-1}\right)^{-1}$$
$$\bar{m} = \bar{v}\left(\frac{1}{\delta}\sum_{i=1}^{M}\Xi_i^{-1}a_i\right)$$

B.3 Monte-Carlo experiment

To test the algorithm and c de, we run a small Monte-Carlo experiment. We generate data from the following panel VAR model:

$$Y_{it} = X_{it}B_{it} + A_{it}^{-1}H_t^{1/2}e_{it}$$

where Y_{it} contains $N = 3$ endogenous variables for $i = 1, 2, .., M = 40$ units. The time series is assumed to be $t = 1, 2, .., 160$. The first 100 observations are discarded to remove the influence of initial conditions leaving a time-series of $T = 60$.

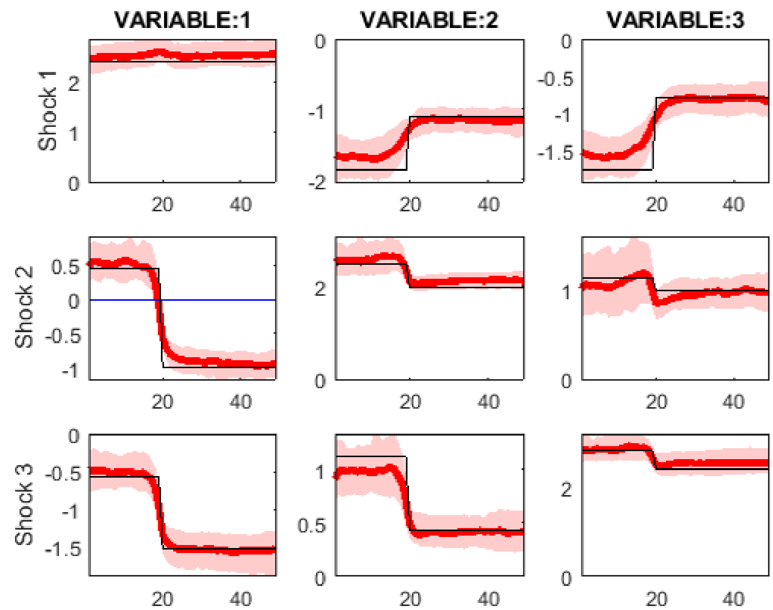
The components for the VAR coefficients $B_{it} = b_i + \beta_t$ are generated as follows. b_i is assumed to be $N\left(vec\left(\bar{b}\right), \bar{v}\right)$ where $\bar{b} = \begin{pmatrix} 0.7 & -0.1 & -0.1 \\ 0.1 & 0.7 & 0.1 \\ -0.1 & 0.1 & 0.7 \\ 0 & 0 & 0 \end{pmatrix}$ and $\bar{v} = 0.01$. β_t is assumed to follow a simple process. For the first 30 observations $\beta_t = vec\begin{pmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \end{pmatrix}$ and for the remainder $\beta_t = vec\begin{pmatrix} -0.1 & -0.1 & -0.1 \\ -0.1 & -0.1 & -0.1 \\ -0.1 & -0.1 & -0.1 \\ -0.1 & -0.1 & -0.1 \end{pmatrix}$. This formulation, thus induces a one-time change in the coefficients.

A similar process is assumed for the elements of $A_{it} : a_{it} = a_i + a_t$. a_i is generated from $N\left(\bar{a}, \bar{w}\right)$ where $\bar{a} = \begin{pmatrix} 0.2 \\ 0.1 \\ -0.2 \end{pmatrix}$, $\bar{w} = 0.01$. $a_t = \begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}$ for the first 30 observations and $a_t = \begin{pmatrix} -0.3 \\ -0.3 \\ -0.3 \end{pmatrix}$ for the remainder.

$\ln h_t$ is assumed to be zero for the first 30 observations and then assumed to increase to one. σ_i^2 is drawn from $U(0, 1)$.

Given the artificial data, the Gibbs sampling algorithm described above is used to approximate the posterior distribution. We use 10,000 iterations discarding the first 5,000 as burn-in. The saved draws are used to compute the time-varying impulse response to three shocks identified via a recursive scheme. The experiment is repeated 100 times.

Figure B1: Estimated and true cumulative impulse responses

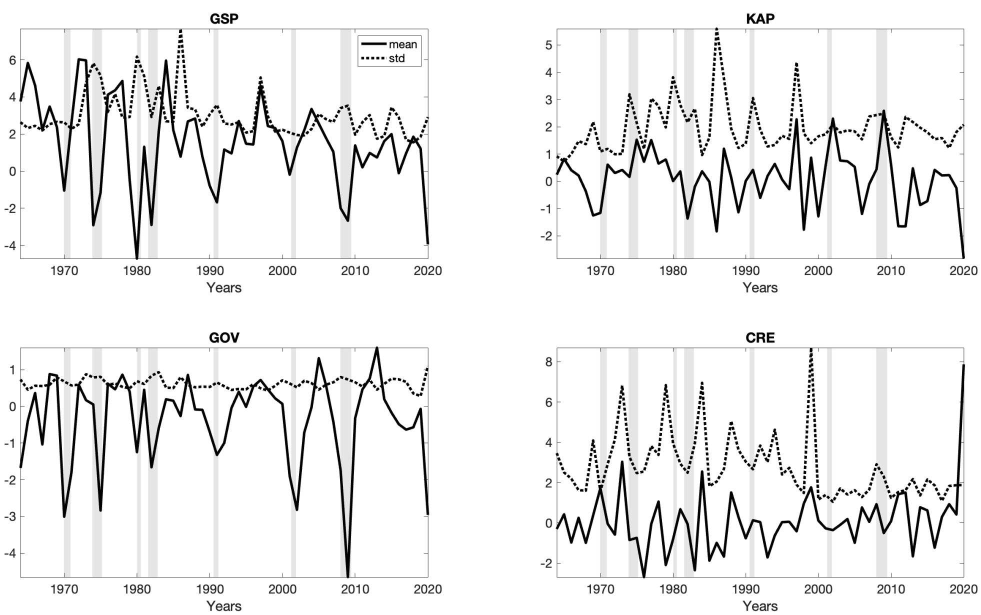


Note: The red line and the shaded area show the median estimated response, the 16^{th} and the 84^{th} bands. The black line shows the true response. The responses are cumulated and the plotted for horizon 40.

Figure B1 shows the estimated cumulated responses at horizon 40 to the three shocks for the average unit (median and 1-standard-error bands across the replications). The black line displays the true responses. The figures show that the estimated responses track the structural change reasonably well.

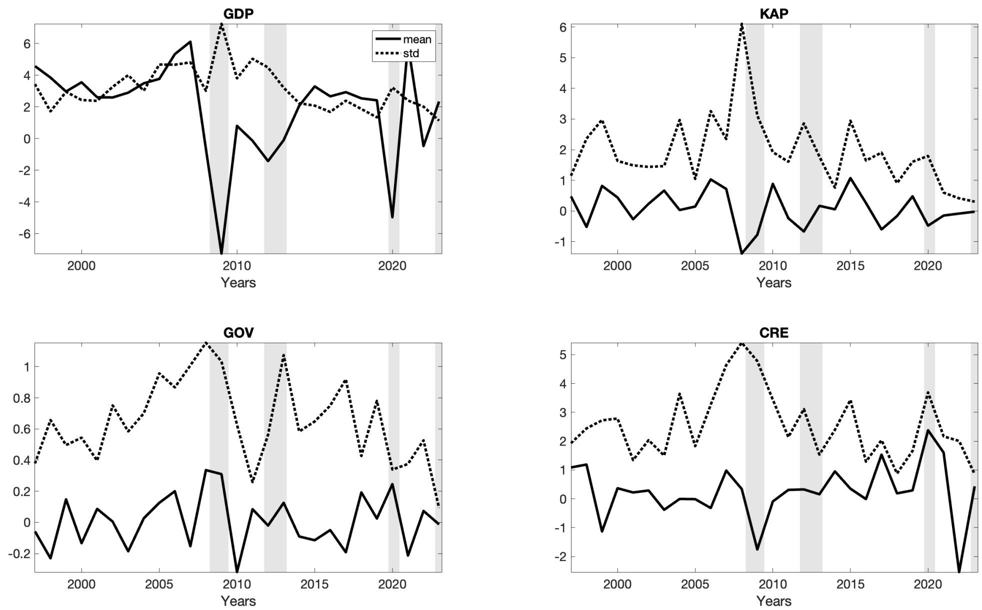
Appendix C Figures and Tables

Figure C1: Cross-sectional mean and std. deviation of the series in the US



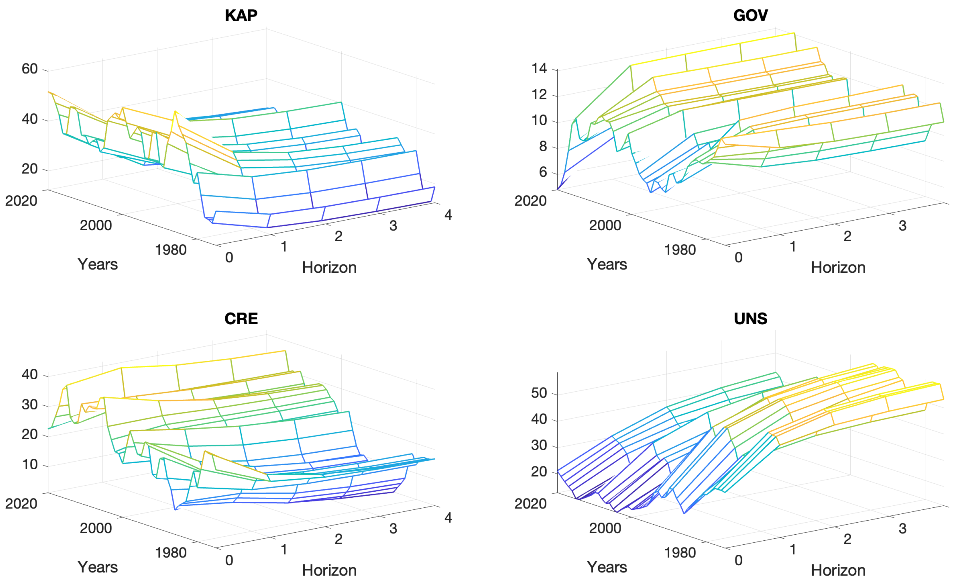
Notes: $GSP = \Delta\log GSP_t$, $KAP = \Delta\log GSP_t - \Delta\log SI_t$, $GOV = \Delta\log SI_t - \Delta\log DSI_t$, $CRE = \Delta\log DSI_t - \Delta\log C_t$. Vertical shaded areas indicate NBER recessions.

Figure C2: Cross-sectional mean and std. deviation of the series in the EA



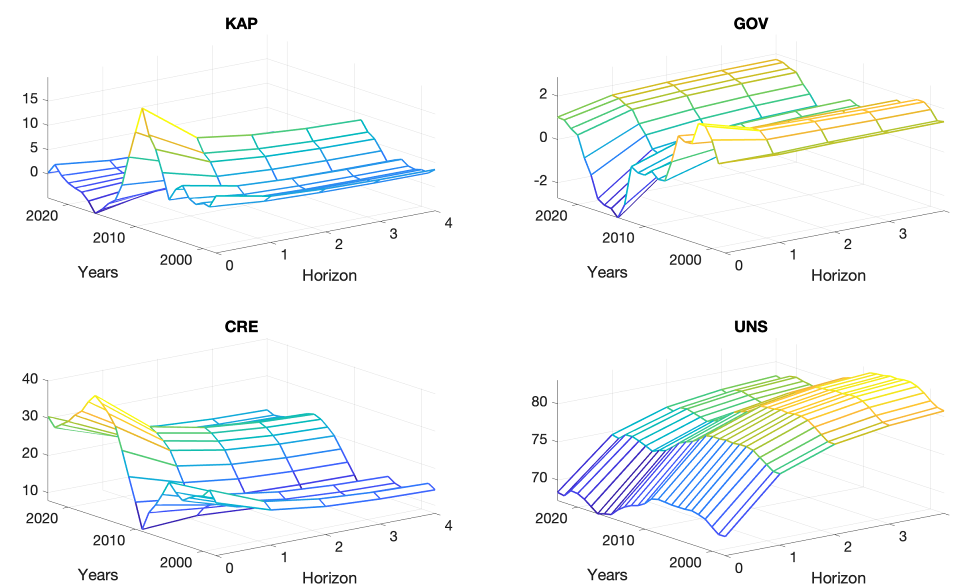
Notes: $GDP = \Delta\log\,GDP_t$, $KAP = \Delta\log\,GDP_t - \Delta\log\,GNP_t$, $GOV = \Delta\log\,GNP_t - \Delta\log\,GDI_t$, $CRE = \Delta\log\,GDI_t - \Delta\log\,C_t$. Vertical shaded areas indicate recessions.

Figure C3: Median cumulative impulse responses to a state-specific output shock in the US



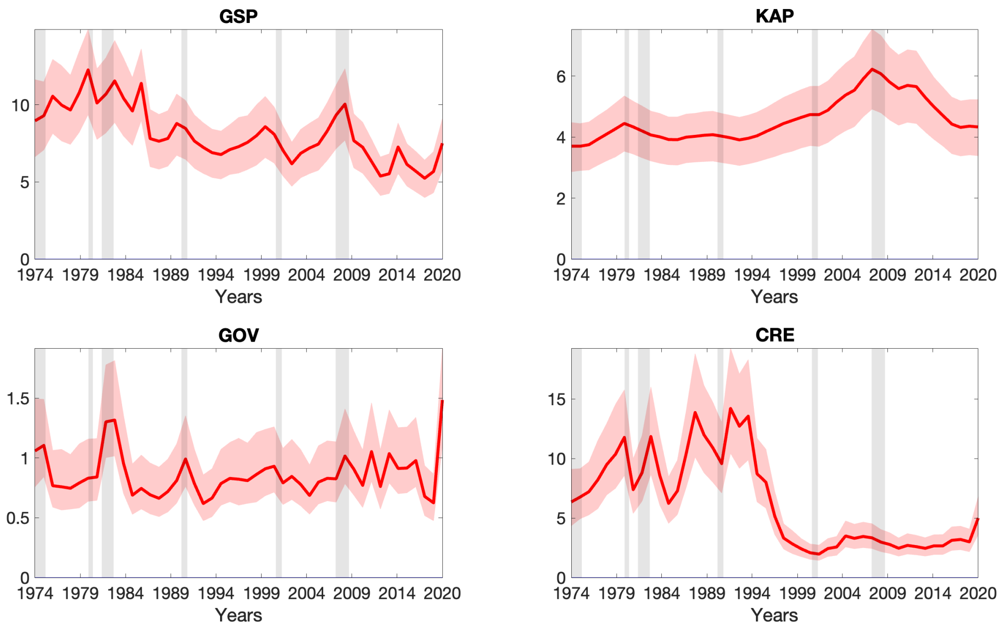
Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. The effects of the shock are normalized to one hundred in each horizon.

Figure C4: Median cumulative impulse responses to a country-specific output shock in the EA



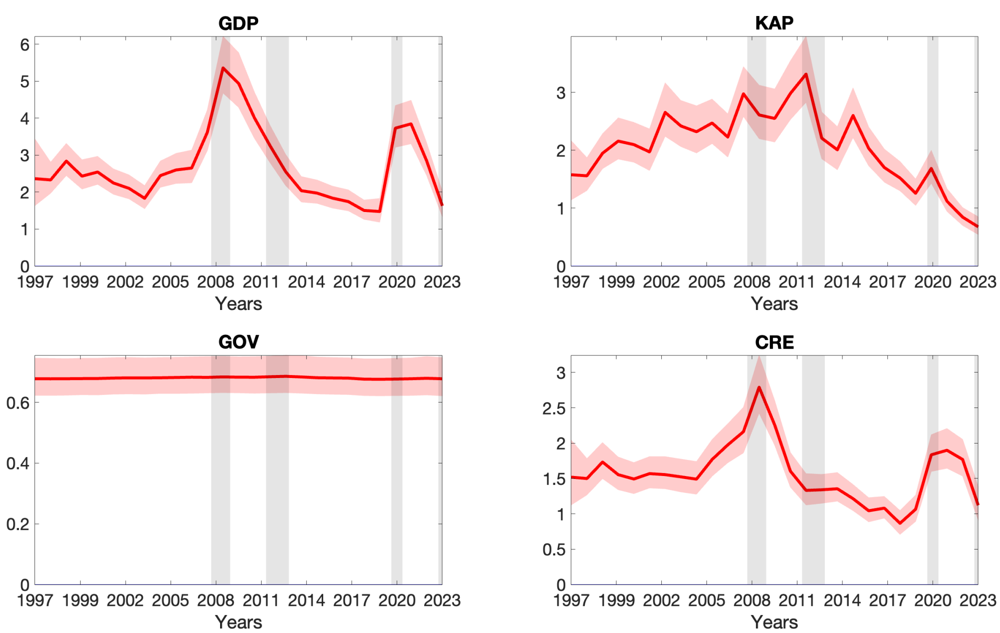
Notes: See Figure C3.

Figure C5: Time series of stochastic volatilities in the US



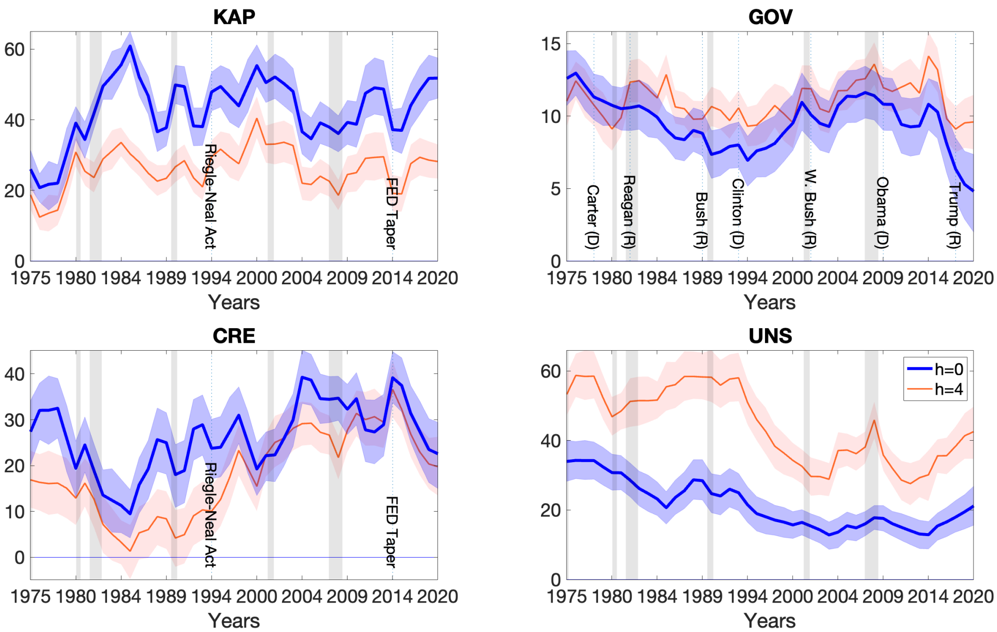
Notes: Vertical shaded areas indicate NBER recessions.

Figure C6: Time series of stochastic volatilities in the EA



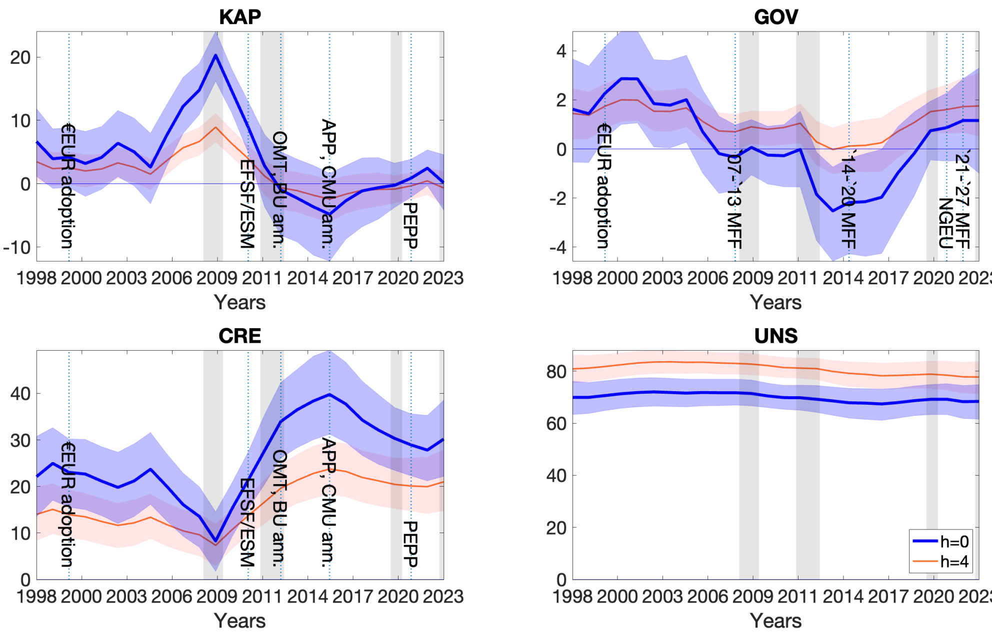
Notes: Vertical shaded areas indicate recessions.

Figure C7: Impulse responses to a state-specific output shock in the US - simple demeaning



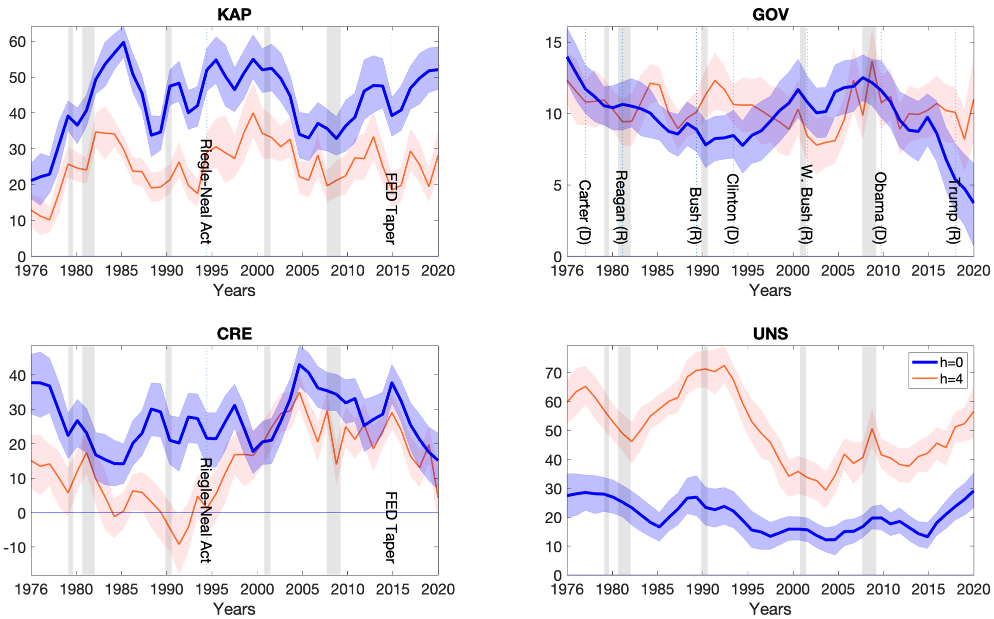
Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. Blue line: impact ($h = 0$) effect. Red line: cumulative effect four years ($h = 4$) after the shock. The effects of the shock are normalized to one hundred in each horizon. Shaded areas are the 16^{th} and 84^{th} percentile posterior bands. Vertical shaded (grey) areas indicate NBER recessions, lines indicate the dates of the presidential elections (black), and the Riegle–Neal Interstate Banking and Branching Efficiency Act. Series are demeaned by the simple averages, instead of weighted averages.

Figure C8: Impulse responses to a state-specific output shock in the EA - simple demeaning



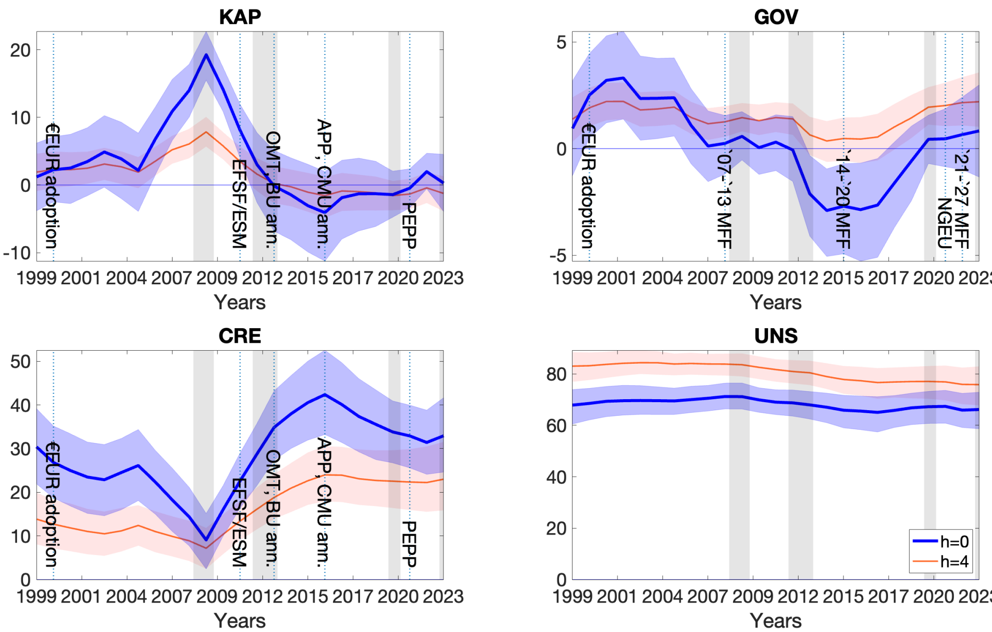
Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. Blue line: impact ($h = 0$) effect. Red line: cumulative effect four years ($h = 4$) after the shock. The effects of the shock are normalized to one hundred in each horizon. Shaded areas are the 16^{th} and 84^{th} percentile posterior bands. Series are demeaned by the simple averages, instead of weighted averages. Vertical shaded (grey) areas indicate recessions. Dotted lines indicate the dates of the activation of the EFSF/ESM support programs, the APP and PSPP of the ECB, the NGEU program and the years of the EU multi-annual financial framework (MFF) budget.

Figure C9: Impulse responses to a state-specific output shock in the US - VAR(2)



Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. Blue line: impact ($h = 0$) effect. Red line: cumulative effect four years ($h = 4$) after the shock. The effects of the shock are normalized to one hundred in each horizon. Shaded areas are the 16^{th} and 84^{th} percentile posterior bands. Vertical shaded (grey) areas indicate NBER recessions, lines indicate the dates of the presidential elections (black), and the Riegle–Neal Interstate Banking and Branching Efficiency Act.

Figure C10: Impulse responses to a state-specific output shock in the EA - VAR(2)



Notes: KAP: capital channel, GOV: fiscal channel, CRE: credit channel, and UNS: unsmoothed part. Blue line: impact ($h = 0$) effect. Red line: cumulative effect four years ($h = 4$) after the shock. The effects of the shock are normalized to one hundred in each horizon. Shaded areas are the 16^{th} and 84^{th} percentile posterior bands. Vertical shaded (grey) areas indicate recessions. Dotted lines indicate the dates of the activation of the EFSF/ESM support programs, the APP and PSPP of the ECB, the NGEU program and the years of the EU multi-annual financial framework (MFF) budget.

Table C1: Determinants of the risk-sharing channels - NBER dates

	Impact ($h = 0$)			Long-run ($h = 4$)		
	KAP	GOV	CRE	KAP	GOV	CRE
NBER	1.01*	0.39*	-0.59	0.32	0.61*	0.36
	(0.06, 2.00)	(0.18, 0.61)	(-1.29, 0.17)	(-0.61, 1.41)	(0.22, 1.04)	(-0.52, 1.31)
Fin	-0.01	0.01	0.03*	-0.01	0.00	0.06*
	(-0.03, 0.01)	(-0.00, 0.01)	(0.01, 0.05)	(-0.03, 0.01)	(-0.01, 0.01)	(0.04, 0.08)
STIR	-0.01	-0.02	0.19*	-0.08	-0.03	-0.17*
	(-0.11, 0.09)	(-0.04, 0.01)	(0.10, 0.29)	(-0.21, 0.06)	(-0.08, 0.02)	(-0.27, -0.07)
Debt	0.01	-0.02*	-0.03*	0.00	-0.02*	-0.07*
	(-0.01, 0.03)	(-0.02, -0.01)	(-0.04, -0.01)	(-0.01, 0.02)	(-0.02, -0.01)	(-0.09, -0.06)

Notes: Bayesian weighted SUR regression with flat priors, and weights given by the inverse of the posterior variance of the risk-sharing channels. Dependent variables are the median risk-sharing channels. Explanatory variables are lagged. NBER: NBER recession dates; Fin: long-run credit-to-GDP trend; STIR: 3-months Treasury yield; Debt: debt-to-GDP ratio. Left panel: on impact ($h = 0$). Right panel: four years after the shock ($h = 4$). Constant omitted from the table. Star denotes if posterior median is outside of the 16^{th}-84^{th} percentile credible interval.

Table C2: Determinants of the risk-sharing channels - Long-term interest rates

	Impact ($h = 0$)			Long-run ($h = 4$)		
	KAP	GOV	CRE	KAP	GOV	CRE
Gap	-0.17*	-0.05*	0.06	-0.24*	-0.11*	-0.16*
	(-0.27, -0.06)	(-0.07, -0.03)	(-0.02, 0.15)	(-0.35, -0.11)	(-0.14, -0.07)	(-0.26, -0.06)
Fin	0.02	0.01*	0.02	0.02	0.00	0.09*
	(-0.00, 0.04)	(0.01, 0.02)	(-0.00, 0.05)	(-0.00, 0.04)	(-0.01, 0.01)	(0.07, 0.12)
LTIR	0.06	0.01	0.05	0.18	-0.02	-0.15
	(-0.08, 0.23)	(-0.02, 0.04)	(-0.07, 0.18)	(-0.00, 0.38)	(-0.08, 0.04)	(-0.29, -0.00)
Debt	-0.02	-0.02*	-0.03*	-0.01	-0.02*	-0.1*
	(-0.03, -0.00)	(-0.02, -0.01)	(-0.04, -0.02)	(-0.02, 0.01)	(-0.02, -0.01)	(-0.11, -0.08)

Notes: Bayesian weighted SUR regression with flat priors, and weights given by the inverse of the posterior variance of the risk-sharing channels. Dependent variables are the median risk-sharing channels. Explanatory variables are lagged. Gap: the output gap; Fin: long-run credit-to-GDP trend; LTIR: 10-year Treasury yield; Debt: debt-to-GDP ratio. Left panel: on impact ($h = 0$). Right panel: four years after the shock ($h = 4$). Constant omitted from the table. Star denotes if posterior median is outside of the 16^{th}-84^{th} percentile credible interval.

Table C3: Determinants of the risk-sharing channels - IMF financial development index

	Impact ($h = 0$)			Long-run ($h = 4$)		
	KAP	GOV	CRE	KAP	GOV	CRE
Gap	-0.46*	-0.08*	0.00	-0.39*	-0.07*	-0.31*
	(-0.78, -0.13)	(-0.11, -0.05)	(-0.16, 0.17)	(-0.58, -0.19)	(-0.13, -0.01)	(-0.46, -0.15)
FinIMF	10.08*	0.45*	2.28	0.60	-0.74	5.21*
	(5.24, 15.63)	(0.04, 0.94)	(-0.29, 4.83)	(-2.22, 3.98)	(-1.62, 0.19)	(3.20, 7.42)
STIR	0.28	0.00	0.09	0.00	-0.04	-0.23*
	(-0.01, 0.62)	(-0.02, 0.03)	(-0.07, 0.26)	(-0.18, 0.20)	(-0.09, 0.02)	(-0.36, -0.10)
Debt	-0.05*	-0.01*	-0.02*	-0.02*	-0.01*	-0.06*
	(-0.07, -0.02)	(-0.01, -0.01)	(-0.03, -0.01)	(-0.04, -0.01)	(-0.02, -0.01)	(-0.07, -0.05)

Notes: Bayesian weighted SUR regression with flat priors, and weights given by the inverse of the posterior variance of the risk-sharing channels. Dependent variables are the median risk-sharing channels. Explanatory variables are lagged. Gap: the output gap; FinIMF: IMF financial development index; STIR: 3-month Treasury yield; Debt: debt-to-GDP ratio. Left panel: on impact ($h = 0$). Right panel: four years after the shock ($h = 4$). Constant omitted from the table. Star denotes if posterior median is outside of the 16^{th}-84^{th} percentile credible interval. Estimation sample starts in 1980.

Chapter 4

Fiscal Multipliers in Pandemics*

4.1 Introduction

In early 2020, the world faced an exponential growth in cases and deaths resulting from a novel respiratory illness: the Coronavirus Disease 2019 (COVID-19). To slow down the spread of the disease, policymakers around the world closed borders, schools, and workplaces, and recommended social distancing. Episodes of health crises have often hampered economic activity (Ma et al. (2023)) and the recession resulting from the COVID-19 pandemic was not different. The pandemic led to an unprecedented loss in output, greater than during the global financial crisis.

Policymakers used a range of policy tools to lessen the adverse economic impact of the COVID-19 pandemic, among which fiscal policy was key and center. Many countries responded to the health crisis through substantial countercyclical fiscal policy through support to the health care sector and assistance to businesses and households, in particular those impacted by the pandemic. With many parts of the world gradually moving from pandemic to endemic phase, a pertinent question is how impactful these expansionary fiscal policies have been and what could be their impact over time?

There is a large literature on state-dependent fiscal multipliers. In their seminal work, Auerbach and Gorodnichenko (2012b) relied on a nonlinear empirical model for the US to show that fiscal multipliers are larger during recession or periods of economic slack. Numerous subsequent studies (Auerbach and Gorodnichenko (2012a); Fazzari et al. (2015); Caggiano et al. (2015); Cohen-Setton et al. (2019)) corroborated these findings.[1] The

* This chapter is based on Kinda et al. (2022).

[1] Several studies found that the results could be sensitive to changes in specifications or estimation methods. Model-based studies find different results, including Canzoneri et al. (2016) that show higher but relatively short-lived multipliers during recessions; and Sims and Wolff (2018) that highlight that

literature also highlighted that fiscal multipliers tend to be larger when interest rates are near the zero lower bound or when monetary policy accommodates government spending (Farhi and Werning (2016); Christiano et al. (2011); Coenen et al. (2012)).[2] Ilzetzki et al. (2013) show that fiscal multipliers are greater in industrial countries, in presence of fixed exchange rates and in closed economies.

A recent literature has emerged on macroeconomic policies during health crises, with a focus on fiscal policy. Eichenbaum et al. (2021) and Glover et al. (2023) embedded epidemiology models in real business cycles models to study optimal health policy responses and find that severe recessions generated by agents' optimal decision to cut back on consumption and hours worked help reduce the severity of epidemics. Elenev et al. (2022) found that fiscal support to distressed firms during the pandemic were effective in preventing corporate bankruptcies, although the authors did not compute specific multipliers. Focusing on fiscal policy, Auerbach et al. (2022) estimated larger employment multipliers during the 2020 lockdowns in US states, particularly in states with less stringent stay-at-home orders. Bayer et al. (2023) show that fiscal multipliers in the US were larger (around 1.5) for targeted transfers such as unemployment insurance compared to untargeted lump-sum transfers (around 0.25) during the COVID-19 pandemic. Drawing on similarities with the Great Recession, Wilson (2020) suggested that fiscal multipliers were around 1.5 during the health crisis. Using TANK model calibrated with the US data during the COVID-19 pandemic, Faria-e Castro (2021) shows that lump-sum transfers have a multiplier of 0.65, while government consumption has a multiplier of about 1.25. Opposite to the main findings of most studies, Guerrieri et al. (2022) present a theory that illustrates lower fiscal multipliers (below one) in presence of COVID-19 type shocks that lead to a shutdown of specific sectors in the economy, compared to supply shocks that affect all segments of the economy and are associated with larger fiscal multipliers (above one).

This paper investigates fiscal multipliers during episodes of health crises, such as the COVID-19 pandemic. It contributes to the literature by (i) providing additional evidence on the state-dependent effects fiscal multipliers, and (ii) advancing the recent and growing literature on macroeconomic policies during health crises by providing cross-country estimates of fiscal multipliers and assessing potential transmission channels at play.

The paper shows that fiscal multipliers are larger in the near-term after a pandemic.

multiplier can be mildly procyclical.

[2] For instance, recent empirical findings highlighted evidence of higher multipliers, ranging from 1.5 to 2.5 at the zero lower bound for Japan (Miyamoto et al. (2018)) and around 1.5 for in the United States (Ramey and Zubairy (2018)).

Based on local projections (Jordà (2005)), the results highlight that cumulative fiscal multipliers one year after a health crisis are about twice larger than during normal times, particularly in advanced economies. During health crises, multipliers associated with public investments are significantly larger than those associated with public consumption. While higher debt levels tend to decrease the effectiveness of fiscal policy in normal times, public debt seems to be less of a concern for an effective fiscal response during a health crisis. The presence of a fiscal rule can further enhance the output effect of a fiscal expansion at the onset of a pandemic, most likely due to the credibility channel. Potential factors or transmission channels underpinning the differences between pandemic and non-pandemic fiscal multipliers can be categorized into three groups: uncertainty, suppressed demand, and supply bottlenecks. Controlling for these channels, the results illustrate that fiscal multipliers in the pandemic and the non-pandemic regimes are no longer statistically different.

The rest of the paper is organized as follows. Section II discusses potential channels through which health crises could impact output and how fiscal policy could play a role in cushioning these effects. Section III presents the data, Section IV discusses the empirical strategy and Section V presents the baseline results. Section VI investigates the potential role of transmission channels through which the pandemic could impact output. Section VII presents some robustness checks and Section V concludes.

4.2 Health Crises and Economic Output: Potential Transmission Channels

The literature has highlighted various factors or transmission channels that can interfere with the effectiveness of fiscal policy during episodes of health crises. The section below focuses on three categories that came to the force during the COVID-19 pandemic: heightened uncertainty, supply bottlenecks, and suppressed demand.

4.2.1 Heightened Uncertainty

A rise in uncertainty could dampen the stimulative effect of fiscal policy. Economic agents tend to take a more cautious approach in the presence of increased uncertainty, including by postponing hiring and investment decisions, which in turn reduces the effect of a policy stimulus. While heightened uncertainty tends to dampen the impact of policies, some theoretical work suggest that policy stimulus becomes more effective once the uncertainty subsides (Bloom (2014), Bloom et al. (2018)). Empirical studies bring evidence both for and against these theoretical findings. Alloza (2022) finds that the response of output to fiscal stimulus is low or insignificant during periods of high uncertainty, defined as unusually high implied stock market volatility. Berg (2019) on the other hand derives a business uncertainty measure using firm-level data and finds that the response of output to fiscal stimulus increases with the level of uncertainty. The author argues that as uncertain times coincide with tight financial conditions, fiscal policy helps ease the latter.

4.2.2 Supply Bottlenecks

The need to reduce in-person interactions and enhance social distancing at the beginning of the pandemic led to disruptions in the workplace. In addition, the pandemic also led to a large drop in employment and weaker investment, worsening supply capacities. While labor markets in most countries have started to recover after the initial shock, the pandemic caused longer lasting scarring and disruptions in the international supply chains, both in production and shipping. This had led to reduced availability of goods and services (Bonadio et al. (2021); Zhang (2021); Mahajan and Tomar (2021); Lafrogne-Joussier et al. (2023)). Recent theoretical work on state-dependent fiscal policy shows that demand stimulating measures tend to be less effective when inadequate supply is the source of the prevailing economic downturn (Ghassibe and Zanetti (2022); Jo and Zubairy (2022)). Auerbach et al. (2021) modelled the impact of pandemic containment measures, which constrained supply and the set of goods available for trade and found that fiscal stimulus has a reduced effect in presence of restrictions, reflecting a muted response of consumption. However, the re-opening of the economy is associated with a surge in consumption, output, and inflation due to pent-up demand. Recent empirical studies also show that fiscal stimulus has been less effective in presence of lockdown policies (Coibion et al. (2020); Auerbach et al. (2022); Brunet and Hlatshwayo (2021)).

4.2.3 Suppressed Demand

The COVID-19 pandemic and associated spike in unemployment led to a sharp decline in affected households' income, an increase of income at risk and higher uncertainty faced by households. Consumer confidence and household consumption expenditures also dropped rapidly at the onset of the health crisis. While fiscal measures such as cash transfers can be effective at stimulating demand and spurring economic activity (Eggertsson and Krugman (2012); Kaplan and Violante (2014); Ghassibe and Zanetti (2022)), their impact can be reduced in the presence of a health crisis. Guerrieri et al. (2022) highlight that fiscal transfers could be less effective in sectors that are impacted by health crises, in particular when these sectors are closed or significantly constrained by containment measures. Recent experience in the U.S. illustrates that fiscal expenditures did not significantly stimulate consumption or employment in areas with strict lockdown measures as cash transfers received in 2020 in form of stimulus checks were mostly saved (Coibion et al. (2020)). Regions with less stringent lockdowns however experienced significant increases in employment in response to higher fiscal spending (Auerbach et al. (2022)).

4.3 Data

The empirical analyses rely on a sample of 91 countries with annual data spanning from 1980 to 2020. The sample excludes small countries with a population of less than 1 million. Output and fiscal related variables (total public expenditure, public consumption, and public investment) are from the IMF's World Economic Outlook (WEO) database and CEIC's Global Economic Monitor dataset.

Disease outbreak data are from the World Health Organization and encompass the following health crises: Ebola Virus Disease (EVD), Severe Acute Respiratory Syndrome (SARS), Influenza A (H1N1), Middle East Respiratory Syndrome Coronavirus (MERS-CoV), and Coronavirus Disease 2019 (COVID-19). The data of case numbers per thousand population were aggregated at annual frequencies for each of the health crises to create an epidemic/pandemic dummy variable, taking the value 1 when the case number per thousand population is greater than the median value, and the value of 0 otherwise. Therefore, we have health crisis specific case number thresholds.

The analyses on transmission channels uses the World Pandemic Uncertainty Index from Ahir et al. (2022), gross private savings from the WEO, the number of air passengers from the International Civil Aviation Organization (ICAO) and container traffic in ports from the United Nations Conference on Trade and Development (UNCTAD) to construct measures of uncertainty, supply bottlenecks, and suppressed demand. A fiscal rule dummy captures the presence or not of the rule (Schaechter et al. (2012)) to gauge the role of fiscal credibility. The paper also uses WEO forecast errors of public expenditures to identify unexpected expenditure shocks and provide an alternative method to estimate fiscal multipliers.[3]

[3] For certain years in the past the WEO was released in September instead of October which were included.

4.4 Empirical strategy

The paper captures the effect of fiscal policy on output through impulse responses produced by local projections (LP) as introduced by Jordà (2005). This method allows the estimation of impulse responses without specifying an approximation of the underlying multivariate dynamic system and is a solid alternative to the traditional method of vector autoregression (VAR). The LP method has recently emerged as one of the foremost ways of studying the transmission of structural shocks in macroeconomics (Miranda-Agrippino and Ricco (2021)). In this method, the weaker assumptions on data dynamics allow for more adaptable and robust impulse response estimations compared to those obtained from VARs (Ramey (2016)). In addition, the LP method: (1) is not constrained by the curse of dimensionality, which is an intrinsic feature of VARs (Ramey (2016)); (2) better captures nonlinearities (Auerbach and Gorodnichenko (2012b)); (3) avoids misspecification errors and biases that can be compounded at each horizon when using VAR (Jordà (2005)); and (4) facilitates the accommodation of state dependency (Auerbach and Gorodnichenko (2012a)).[4]

The paper estimates linear and state-dependent panel local projections in the form of:

$$Z_{i,t+h} = FE_i^h + FE_t^h + \beta^h Shock_{i,t} + \gamma^h(L)\mathbf{X_{i,t}} + \varepsilon_{i,t+h} \tag{4.1}$$

where for $Z_{i,t+h}$ we use real GDP (Y) or real government expenditure (G); $FE_i^h +$ and FE_t^h are country and time fixed-effects; $Shock_{i,t}$ is the fiscal shock identified via Blanchard and Perotti (2002) in the baseline case and with an alternative identification later on. $\gamma^h(L)$ is a polynomial in the lag operator and $\mathbf{X_{i,t}}$ is a vector of control variables. Control variables include one lag of real GDP growth, one lag of government expenditure growth and the lagged output gap.[5] Driscoll and Kraay (1998) standard errors, which correct for potential heteroskedasticity, autocorrelation, or correlated errors across countries are reported.[6]

The state-dependent equivalent of equation (4.1) is:

$$\begin{aligned} Z_{i,t+h} = FE_i^h + FE_t^h + I_{i,t}\Big[\beta^{P,h} Shock_{i,t} + \gamma^{P,h}(L)\mathbf{X_{i,t}}\Big] \\ + (1 - I_{i,t})\Big[\beta^{NP,h} Shock_{i,t} + \gamma^{NP,h}(L)\mathbf{X_{i,t}}\Big] + \varepsilon_{i,t+h} \end{aligned} \tag{4.2}$$

[4] See Auerbach and Gorodnichenko (2012a) and Ramey and Zubairy (2018) for a discussion of LP's use with state dependence and how it compares with smooth transition VARs.

[5] Obtained as deviation from the HP-filter extracted trend.

[6] Including a lagged depended variable and unit fixed effects can lead to biased estimates. However, our sample size (T=40) implies that this is quantitatively small.

where $I_{i,t}$ is the pandemic indicator function. We use the variable definition of Hall (2009) and Owyang et al. (2013) and scale our dependent variable by lagged real GDP, to convert both our dependent variables to the same units. Our two dependent variables $Z_{i,t}$ are $(Y_{i,t+h} - Y_{i,t-1})/Y_{i,t-1}$ and $(G_{i,t+h} - G_{i,t-1})/Y_{i,t-1}$.

Our baseline identification assumes no contemporaneous response of government expenditures to macroeconomic aggregates, following Blanchard and Perotti (2002). This identification scheme was proposed in the context of quarterly time-series data, as the assumption is more plausible at this frequency. Nevertheless, several authors applied it at an annual or semi-annual frequency (see Beetsma et al. (2006), Bénétrix and Lane (2013), Huidrom et al. (2020) and others). The advantages of using annual data are that they allow for a wider country coverage (as non-interpolated fiscal data is not available historically for most countries) and identified shocks do not reflect changes in the timing of spending (see Auerbach and Gorodnichenko (2016)). Furthermore, as Born and Müller (2012) shows, impulse responses obtained on annual data and on (annualized) quarterly data show high similarities. Nevertheless, we also show the robustness of our results when using forecasting errors to identify unpredictable innovations to government spending.

We rely on the one-step IV methodology proposed by Ramey and Zubairy (2018) to calculate cumulative fiscal multipliers and the corresponding confidence bands. Therefore, instead of estimating Equation (4.1) separately for GDP and government expenditure, we estimate the following equation:

$$\sum_{h=0}^{H} \frac{Y_{i,t+h} - Y_{i,t-1}}{Y_{i,t-1}} = FE_i^h + FE_t^h + m^h \sum_{h=0}^{H} \frac{G_{i,t+h} - G_{i,t-1}}{Y_{i,t-1}} + \delta^h(L)\mathbf{X_{i,t}} + \varepsilon_{i,t+h} \tag{4.3}$$

where we instrument $\frac{G_{i,t+h} - G_{i,t-1}}{Y_{i,t-1}}$ with current period government spending $G_{i,t}$. This isolates the variation in future government spending that is due to the fiscal shock in the current period. The estimated coefficient m^h is the cumulative multiplier at horizon h, with the corresponding estimated standard error. The one-step IV version of Equation (4.2) is:

$$\begin{aligned}\sum_{h=0}^{H} \frac{Y_{i,t+h} - Y_{i,t-1}}{Y_{i,t-1}} = FE_i^h + FE_t^h + I_{i,t}\Big[m^{P,h} \sum_{h=0}^{H} \frac{G_{i,t+h} - G_{i,t-1}}{Y_{i,t-1}} + \delta^{P,h}(L)\mathbf{X_{i,t}}\Big] \\ + (1 - I_{i,t})\Big[m^{NP,h} \sum_{h=0}^{H} \frac{G_{i,t+h} - G_{i,t-1}}{Y_{i,t-1}} + \delta^{NP,h}(L)\mathbf{X_{i,t}}\Big] + \varepsilon_{i,t+h}\end{aligned} \tag{4.4}$$

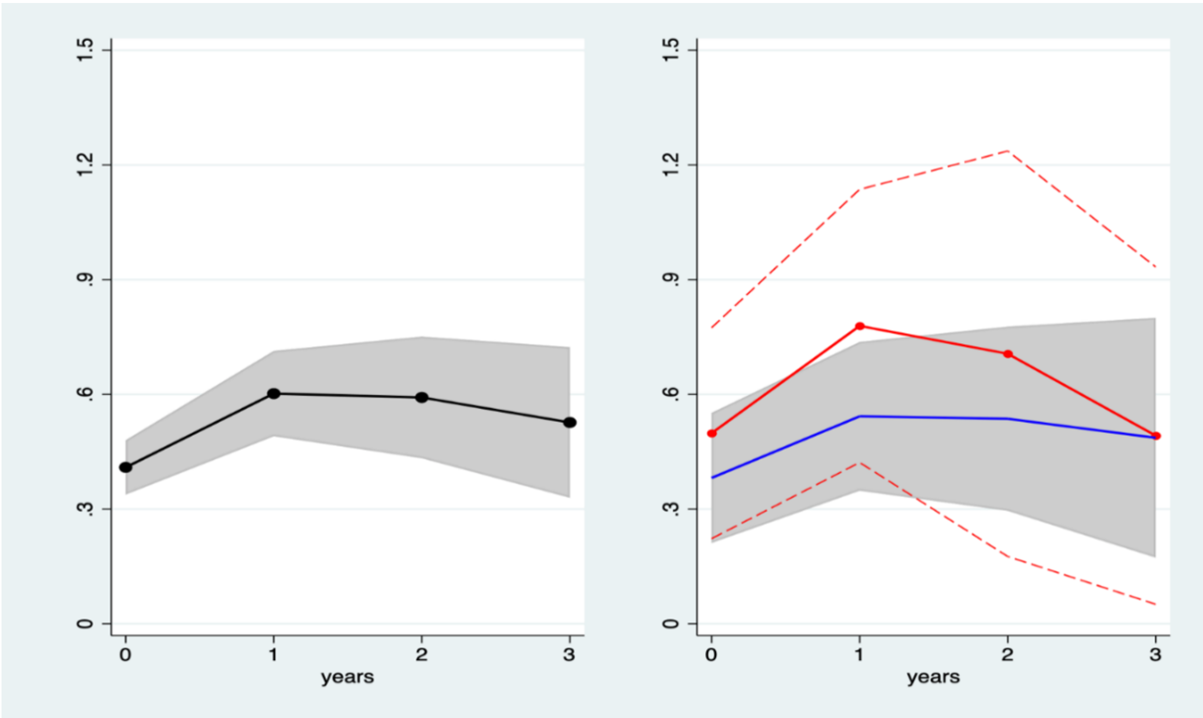
Where we instrument the cumulative government spending growths in the pandemic and the non-pandemic regime by $I_{i,t} \times G_{i,t}$ and $(1 - I_{i,t}) \times G_{i,t}$ respectively. The estimated coefficient $m^{P,h}$ is the cumulative multiplier at horizon h in the pandemic regime, while $m^{NP,h}$ is the estimated cumulative multiplier at horizon h in the non-pandemic regime.

4.5 Main Results

This section first presents the main results, focusing first on total public expenditure, before breaking it down to public consumption and investment. Second, it analyzes advanced and developing countries separately. Last, it investigates the role of the three potential transmission channels discussed above: uncertainty, supply bottlenecks and suppressed demand. Our baseline sample excludes 2020, the year of COVID-19. However, as discussed below, the results are qualitatively and quantitatively similar when we include 2020 in the sample.

Figure 4.1 summarizes the baseline results using total public expenditure and the Blanchard and Perotti (2002) identification. The left panel shows the estimated m^h coefficients from Equation (4.3). In line with the literature, expansionary fiscal policy has a positive and significant effect on output. The estimated cumulative multiplier is 0.4 in the year of the shock, and 0.5-0.6 in the medium term. The right panel of Figure 4.1 shows the estimated $m^{P,h}$ in red and $m^{NP,h}$ in blue from Equation (4.4). The multipliers in the pandemic states are above the multipliers in the non-pandemic states up until the third year after the shock. In the pandemic state, the contemporaneous multiplier is 0.5, and cumulates to 0.7 in the first two years after the shock. In non-pandemic states, the contemporaneous multiplier is lower at 0.38, and cumulates to 0.55 in the first two years following the shock. Table 4.1 shows the p-value of the Chi-squared test for the difference between the estimated multipliers in the two states. The largest difference between the estimates is in $t = 1$, but the p-value of 0.12 suggests that the estimated fiscal multipliers between pandemic and non-pandemic states are not statistically different.

Figure 4.1: Cumulative fiscal multipliers – total expenditures, all countries



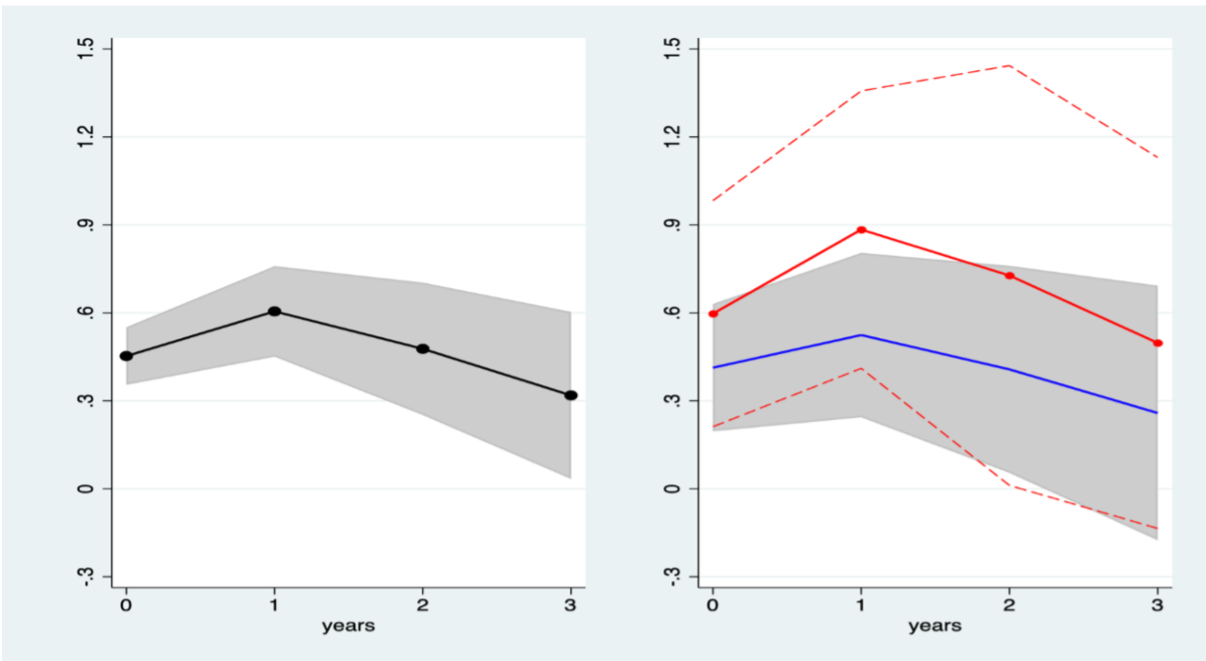
Notes: Cumulative fiscal multipliers from Equation (4.3) (Left panel) and Equation (4.4) (Right panel). The black line represents the average (non-state dependent) fiscal multiplier. The red and blue lines represent multipliers in pandemic and non-pandemic states respectively. The shaded areas and dashed lines are corresponding 95% Driscoll and Kraay (1998) confidence intervals.

Table 4.1: P-values of the Chi-squared test of the difference between cumulative fiscal multipliers in pandemic and non-pandemic states

		T=0			T=1			T=2			T=3	
	Total	Cons.	Inv.	Total	Cons.	Inv.	Total	Cons.	Inv.	Total	Cons.	Inv.
All	0.37	0.28	0.10	0.13	0.09	0.07	0.36	0.30	0.24	0.68	0.40	0.40
Advanced	0.08	0.12	0.01	0.02	0.08	0.00	0.11	0.23	0.03	0.34	0.52	0.01
Non-Adv.	0.84	0.88	0.73	0.59	0.32	0.57	0.99	0.76	0.96	0.96	0.71	0.92

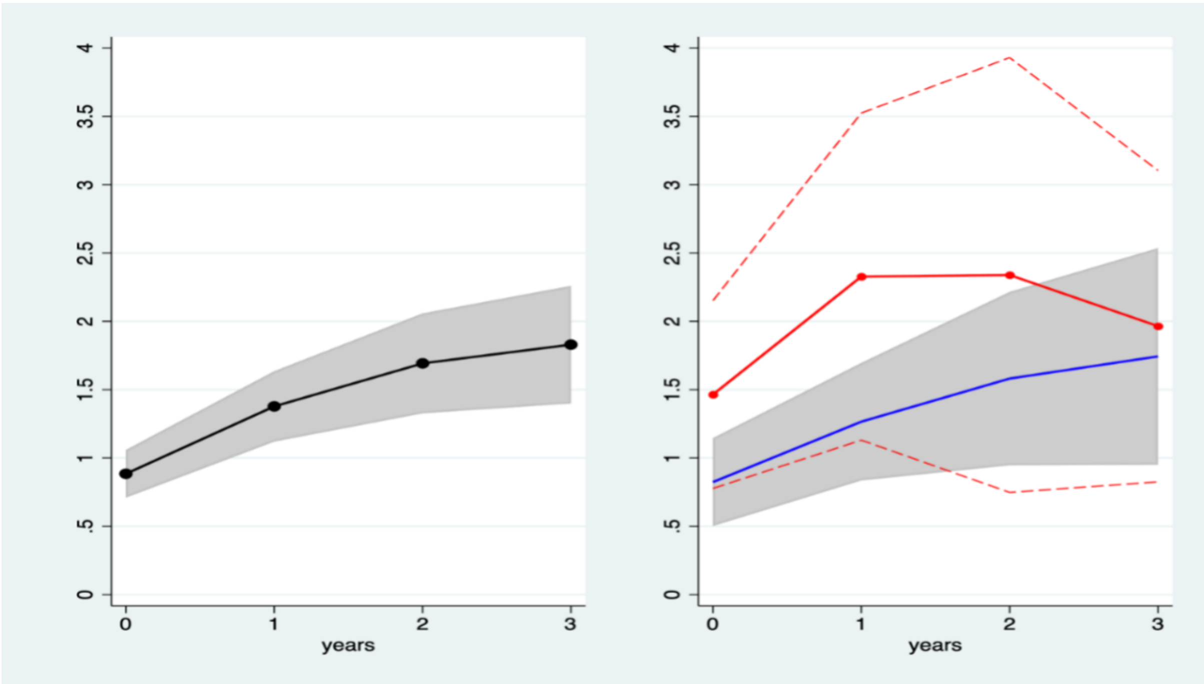
Figure 4.2 illustrates our baseline results when focusing only on public consumption, instead of total public expenditure. In this case, the fiscal multiplier is lower over the medium term and stands at 0.32 in the third year after the shock. Notably, the fiscal multipliers in the pandemic and non-pandemic states are statistically different, with the multiplier in the pandemic state reaching 0.9 in the first year after the shock. Focusing on investment highlights significantly larger multipliers in both pandemic and non-pandemic states, with the multipliers during pandemics exceeding 2 one year after shock, statistically higher than multipliers in non-pandemic states (Figure 4.3 and Table 4.1).

Figure 4.2: Cumulative fiscal multipliers – public consumption, all countries



Notes: See Figure 4.1.

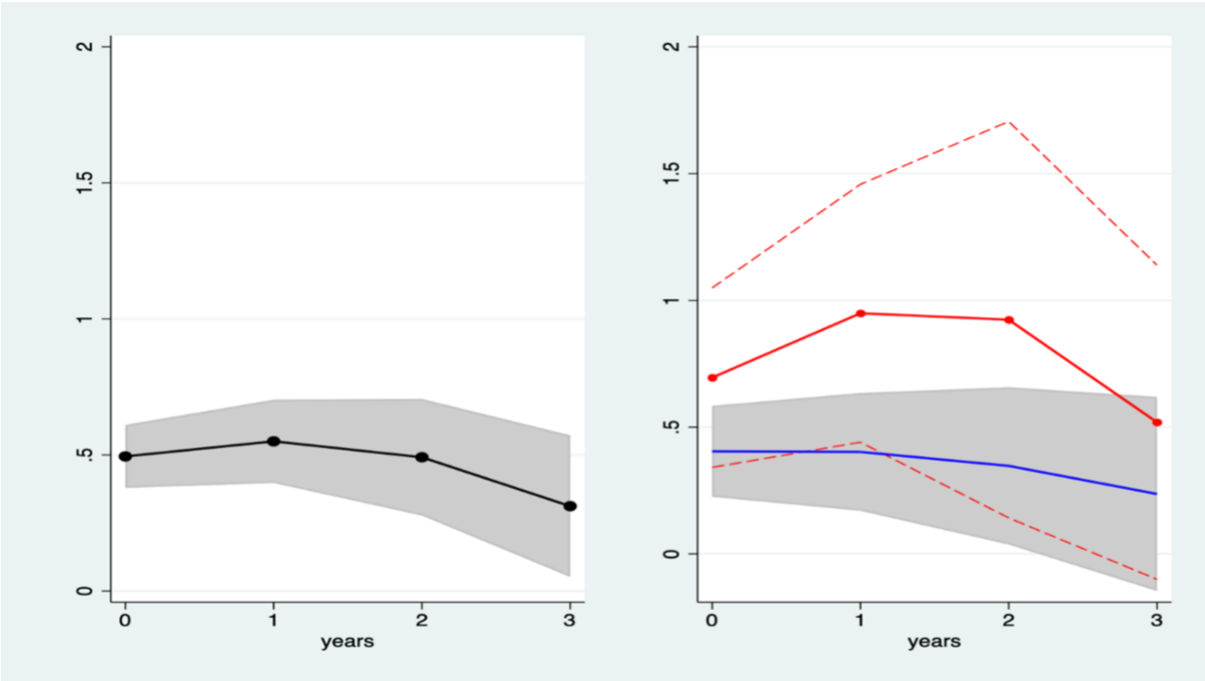
Figure 4.3: Cumulative fiscal multipliers – fixed capital formation, all countries



Notes: See Figure 4.1.

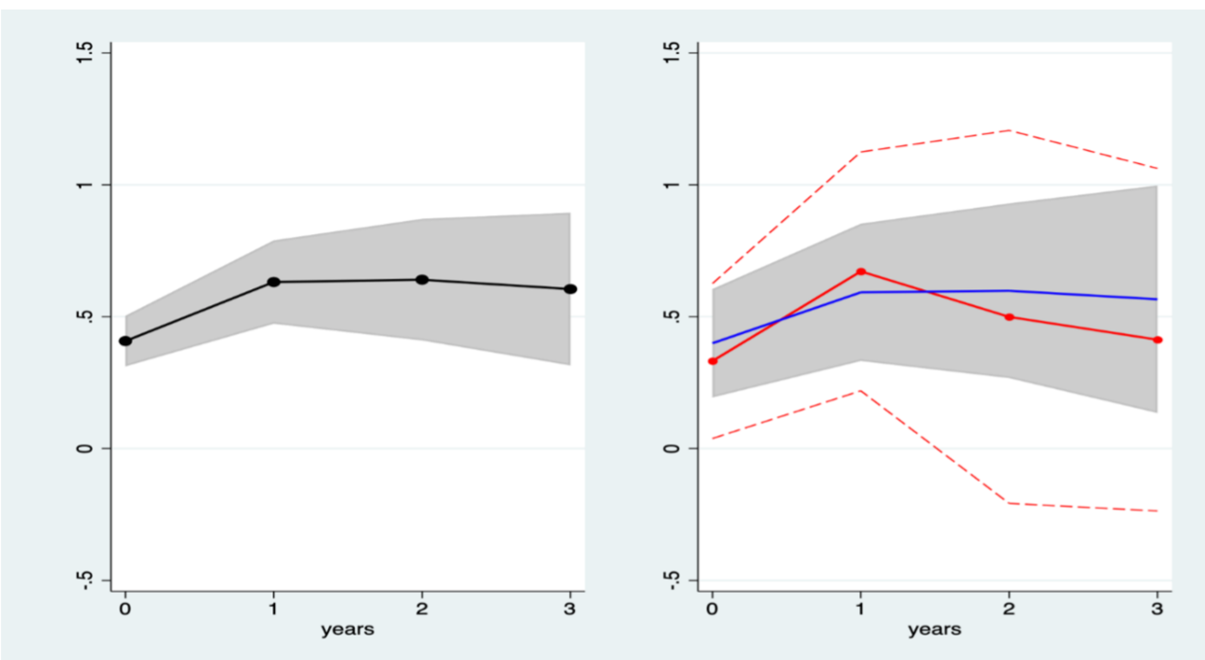
Focusing on advanced economies, illustrates that fiscal multipliers are significantly larger during pandemics, particular one year after the shock when the multiplier is close to 1, about double the multiplier in the non-pandemic state (Figure 4.4). In developing economies, fiscal multipliers reach about 0.6 one year after the shock and are not statistically different between the pandemic and non-pandemic states, possibly reflecting that spending inefficiencies could worsen during a pandemic and constraint the effectiveness of fiscal policy (Figure 4.5).

Figure 4.4: Cumulative fiscal multipliers – total expenditures, advanced economies



Notes: See Figure 4.1. intervals.

Figure 4.5: Cumulative fiscal multipliers – total expenditures, developing economies



Notes: See Figure 4.1.

4.6 Transmission Channels

This section investigates the role of potential transmission channels discussed in Section 4.2: heightened uncertainty, supply chain disruptions, and suppressed demand. These channels could be amplified during pandemics, for instance social distancing can hamper the effectiveness of fiscal policy, and impact economic activity.

The paper relies on four proxies to capture these channels: an uncertainty index related to pandemic events, excess private savings, the number of air passengers and container traffic in ports. Deviations from their filtered trends are used for the last three variables. Excess private savings captures forgone consumption by households due to the pandemic and is a proxy for suppressed demand. The (de-trended) number of air passengers, which proxy mobility trends, and below trend container traffic in ports can indicate impediments in transportation and production.

While these proxies can broadly capture the potential transmission channels, they do not necessarily match these perfectly. For instance, consumer may decide to postpone planned purchases of goods and resort to excess saving if goods and services they intended to purchase are not available because of production or shipping impediments (supply bottleneck channel). Lockdowns can constraint mobility and lead to lower consumption of goods and services (suppressed demand channel), and also result in higher excess savings. Similarly, container traffic in ports could be below trend due to subdued demand of goods or delays in production. Considering the intertwined relationships between the various proxies and transmission channels, the paper jointly includes all proxies together to control for likely correlations. In this regard, we extend the specification of Equation (4.4) by including the proxies and their interactions with the shock as illustrated in the following equations:

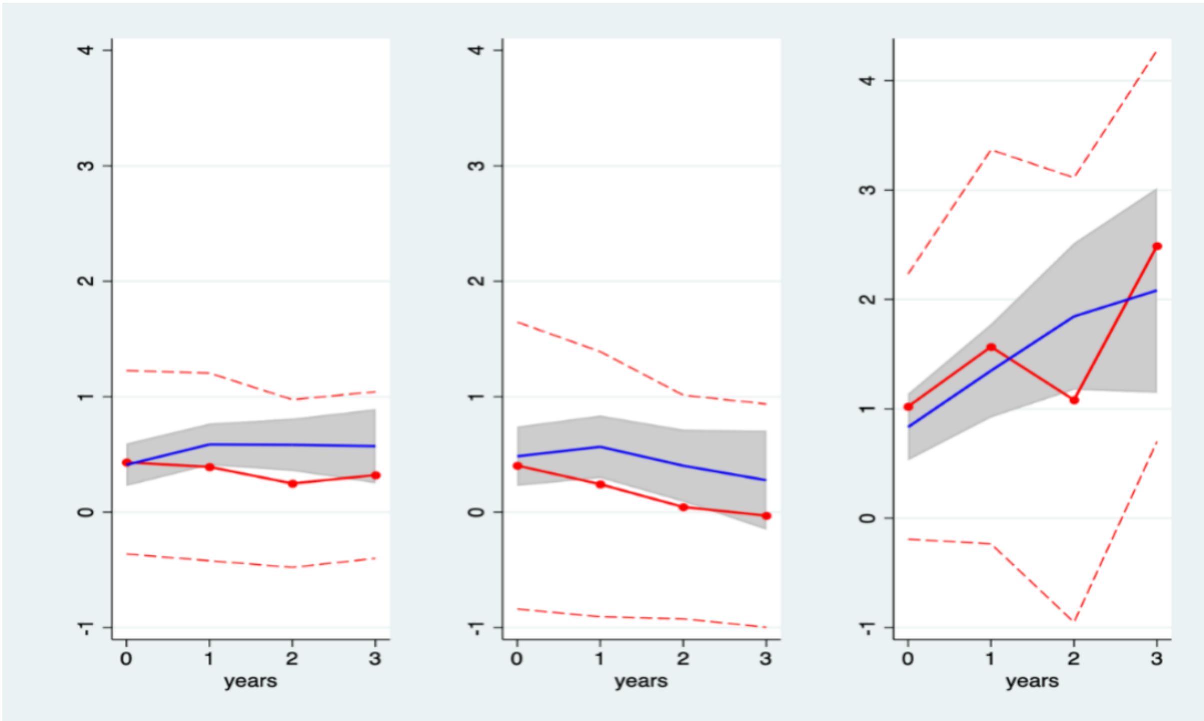
$$\begin{aligned}
\sum_{h=0}^{H}\frac{Y_{i,t+h}-Y_{i,t-1}}{Y_{i,t-1}} &= FE_i^h + FE_t^h \\
&+ I_{i,t}\Big[m^{P,h}\sum_{h=0}^{H}\frac{G_{i,t+h}-G_{i,t-1}}{Y_{i,t-1}} + \delta^{P,h}(L)\mathbf{X_{i,t}} + \theta^{P,h}\mathbf{W_{i,t}} + \phi^{P,h}\mathbf{W_{i,t}}Shock_{i,t}\Big] \\
&+ (1-I_{i,t})\Big[m^{NP,h}\sum_{h=0}^{H}\frac{G_{i,t+h}-G_{i,t-1}}{Y_{i,t-1}} + \delta^{NP,h}(L)\mathbf{X_{i,t}} + \theta^{NP,h}\mathbf{W_{i,t}} + \phi^{NP,h}\mathbf{W_{i,t}}Shock_{i,t}\Big] \\
&+ \varepsilon_{i,t+h} \qquad (4.5)
\end{aligned}$$

Where $\mathbf{W_{i,t}}$ is a vector containing the proxies. We focus mostly on the p-values of the test of difference between the estimated multipliers in the two states (pandemic and non-pandemic) after including the proxies in the specification. If the differences in fiscal multipliers between pandemic and non-pandemic states are no longer present after the introduction of the proxies, we can attest that one or more of these proxies act as transmission channels. In addition, the statistical significance of the $\phi^{P,h}$ coefficients provide a sense of the relative importance of each channel in the pandemic state.

The results are summarized in Figure and Table 4.2. Figure 4.6 plots the unconditional multiplier for total expenditures, i.e., the multiplier when the uncertainty index is at zero, while private savings, air passengers and container traffic are at their trend value. The multipliers in the two states have visibly shifted closer to each other for all three expenditure statistics. This is also confirmed when looking at the statistical test of difference in fiscal multipliers in the two states for total public expenditures, consumption, and investment (Table 4.2). The p-values associated with the test of differences increase markedly. At $h = 0$ and $h = 1$, and the tests very confidently reject a statistical difference between the multipliers in the two states, including in cases where a statistical difference was previously established. This suggests that the tested transmission channels could underpin the larger fiscal multipliers during health crises.[7] An additional observation in Figure 4.6 is that the multipliers in the non-pandemic regime show virtually no differences compared to the normal regime multipliers estimated without the proxies (Figures 4.1-4.3).

[7] It is worth noting that the wide confidence bands for the pandemic regime multipliers also influence the test statistics.

Figure 4.6: Cumulative multipliers after adding proxies: total expenditures, public consumption and public investment, all countries



Notes: Cumulative fiscal multipliers from Equation 4.5. The red and blue lines represent multipliers in pandemic and non-pandemic states respectively. The shaded areas and dashed lines are corresponding 95% Driscoll and Kraay (1998) confidence intervals. $\mathbf{W_{i,t}}$ includes the deviation of real private savings over GDP from its trend, the deviation of the number of air passengers from its trend and the deviation of the container traffic in ports over its trend. Trends obtained via the HP filter.

Table 4.2: P-Value of the Chi-Squared Test of the Difference between the Cumulative Multipliers in the Two States, after Adding Proxies for Transmission Channels

	Total expenditures		**Gov. consumption**		**Gov. investment**	
	Baseline	Proxies incl.	Baseline	Proxies incl.	Baseline	Proxies incl.
h=0	0.37	0.93	0.29	0.98	0.10	0.91
h=1	0.13	0.71	0.09	0.79	0.08	0.99
h=2	0.35	0.26	0.30	0.54	0.24	0.23
h=3	0.68	0.39	0.41	0.61	0.39	0.59

Notes: Left columns are the baseline specifications, the right columns include the proxies (Equation (4.5)).

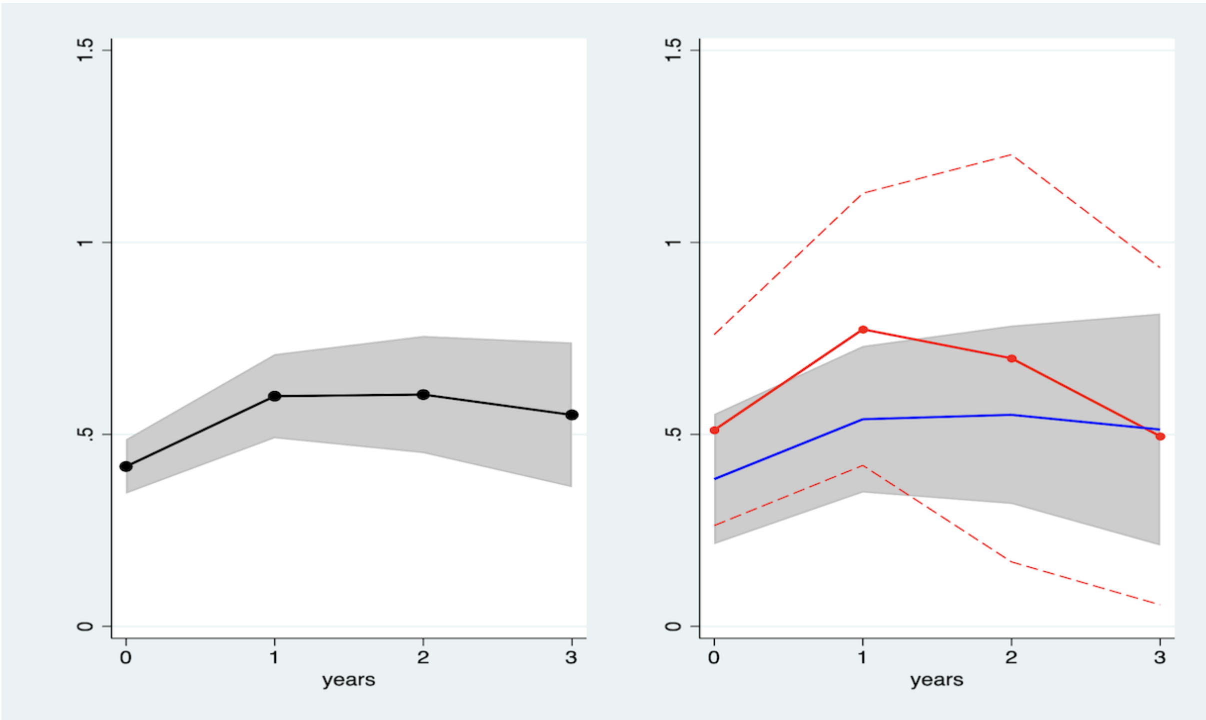
Table A3 in the Appendix illustrates the results of Equation (4.5), estimated via the one-step IV methodology, for $h = 1, 2, 3, 4$. The table only displays the estimated coefficients of the proxies. It shows that among the three transmission channels, uncertainty has the largest impact on the effectiveness of fiscal policy immediately at the onset of a health crisis. The positive coefficient implies that a high level of uncertainty is associated with a higher fiscal multiplier, in line with the findings of Berg (2019). This result suggests for instance that fiscal policy can be more effective when private activity (e.g., investment and consumption) is dampened by heightened uncertainty. The positive estimate of the excess savings' triple interaction suggests that (a given level of) fiscal policy is more effective when private consumption is impaired due to a health crisis. Above or below trend container traffic in ports also appears to have an important effect, with long lasting implications for the effectiveness of fiscal policy. The estimated positive coefficients on the triple interaction terms indicate that during and after pandemics, lower container traffic increases the multipliers. The result suggests that this proxy may be more prone at capturing deficiencies in demand, rather than supply bottlenecks.

4.7 Robustness and Additional Exercises

Several countries have implemented strict lockdowns and social distancing policies during the COVID-19 pandemic. In addition, the crisis has brought about long-lasting disruptions in the production of many goods and in international shipping. The analysis above suggests that fiscal multipliers would be relatively higher in this environment. To assess whether the results could apply to the current COVID-19 crisis, we extended the estimation period to include the year 2020. The estimated multipliers are both qualitatively and quantitatively similar to our baseline results (Figure 4.7). The main results are also robust to adding several control variables, such as additional lags of the existing control variables and further control variables (domestic credit growth and the current account balance).[8]

[8] These results are available upon request.

Figure 4.7: Cumulative fiscal multipliers – including 2020



Notes: See Figure 4.1.

Fiscal multipliers tend to be low in countries with a weak fiscal position (Sutherland (1997); Perotti (1999); Nickel and Tudyka (2014); Huidrom et al. (2020)). One reason behind this evidence is the interest rate channel: fiscal expenditure at an already high-level of debt or low fiscal credibility can further increase the credit risk and borrowing costs across the economy. This also limits a country's potential to fight a pandemic. To assess this, we first add the debt-to-GDP ratio to our baseline specification, in the form of Equation (4.5), estimated via the one-step IV method. Panel A of Table A.4 shows the estimated θ and ϕ coefficients. In line with the literature, a high level of debt tends to decrease the effectiveness of fiscal policy. Interestingly, we do not find this to be the case in the pandemic regime, supporting the priority given to health responses and immediate economic support by most countries at the onset of the COVID-19 pandemic. In Panel B, we introduce a dummy variable that takes the value one if the country had a fiscal rule in place at the given year (Schaechter et al. (2012); Davoodi et al. (2022)) and zero otherwise. The results highlight that fiscal spending is more effective at boosting aggregate activity in countries with at least one fiscal rule at the onset of the pandemic, suggesting that the rule could boost fiscal credibility and help anchor medium-term expectations. We do note, however, that many countries have relaxed their fiscal rules or activated escape clauses due to the pandemic.

Our baseline identification, the traditional Blanchard and Perotti (2002) identification has been criticized by the literature on two main grounds. First, fiscal policies are usually anticipated by the public long before their actual implementation, which can lead to inconsistent estimation of the effect of the shock (see Leeper et al. (2013)). Second, as the Blanchard and Perotti (2002) shock can be forecasted by professional forecasters, it may include endogenous response of the government to economic conditions. To overcome these issues, an alternative identification has been proposed, namely, to identify unexpected expenditure shocks through public expenditure forecasts errors (Auerbach and Gorodnichenko (2012a); Forni and Gambetti (2016); Abiad and Topalova (2016); among others). We use the public expenditure forecasts for the upcoming year from the October edition of the IMF World Economic Outlook to compute forecast errors of public spending. To control for any change in expenditure due to change in output, we also include real GDP forecasts errors in the estimation. The estimation sample for this analysis is 2002-2019 and 49 countries due to limited availability of forecast data. The top panel of Figure A5 confirms our main findings, that is short- and medium-term fiscal multipliers are higher in the pandemic regime and that the difference between the estimated multipliers in the two regimes are statistically significant. The bottom panel of Figure A5 displays the results using the baseline identification and matching estimation sample for comparison.

4.8 Conclusion

This paper assessed how fiscal multipliers vary during health crises, particularly how factors such as social distancing and uncertainty could lower contemporaneous multipliers and increase near-term multipliers.

It showed that fiscal multipliers are larger in the immediate years that follow the onset of a health crisis. The baseline results confirm that an expansionary fiscal policy has a positive and significant effect on output with an estimated multiplier of 0.4 in year T and a cumulative multiplier of 0.5-0.6 in the medium term. Comparing pandemic and non-pandemic states highlights significantly larger multipliers during pandemics, especially in advanced economies. The cumulative multiplier in the pandemic rises to almost 1 one year after the pandemic shock in advanced economies, compared to about 0.4 in the non-pandemic regime. The paper also showed that the differences in the pandemic and non-pandemic fiscal multipliers can be explained by three main factors: uncertainty, suppressed demand and supply bottlenecks. These results are supported by a variety of robustness checks such as (i) including 2020, the height of the COVID-19 pandemic, in the estimation period; (ii) controlling for additional variables such domestic credit growth, the current account balance, and the debt-to-GDP ratio; and (iii) testing for a fiscal credibility channel by controlling for the presence of fiscal rule in the country. In addition, the results are also confirmed with use of an alternative identification strategy that capture unexpected expenditure shocks through public expenditure forecasts errors.

The findings in the paper suggest that the growth impact of fiscal stimulus packages during health crises and pandemics could be larger and longer lasting than often assumed. As such, the large fiscal support deployed at the onset of the COVID-19 pandemic could have a larger impact than in non-pandemic times. Notably, the impact could be at least 50 percent larger one year after the large-scale fiscal stimulus, a noticeable difference when compared to expectations based on more traditional (non-pandemic) fiscal multipliers.

4.9 Appendix

Table A1: Summary statistics

	Mean	Std. Dev.	Minimum	Maximum	Observations
Dummy Variable for Deaths Above Median	0.01	0.11	0.00	1.00	3380
Gross Domestic Product at 2014p (mn. USD)	904760.80	2188649	3411.64	19700000	1947
Gross Domestic Product at current prices (mn. USD)	568781.20	1688126	568.21	21400000	2756
Log of GDP at current prices	4.55	0.20	1.84	6.18	2321
Log of GDP at 2014p	12.36	1.67	8.13	16.80	1947
Total Government Expenditure	142167.30	406480.50	136.84	4625097	2295
Gov. Gross Nominal Fixed Capital Formation (mn. USD)	32935.11	130854.70	-30393.35	2229804	2538
Gov. Consumption Expenditure, current price (mn. USD)	94461.13	266176.30	0.00	2973918	2789
Tax revenue (mn. USD)	125604.40	329061.70	15.86	3411509	1957
Consolidated Fiscal Balance (mn. USD)	-18773.35	90847.98	-1471297	254848	2050
Government Debt (% of GDP)	0.51	0.34	0.02	2.01	1943
Government Debt (mn. USD)	548445.70	1853147	293.59	23200000	1787
Private Consumption Expenditure, current price (mn. USD)	325973.80	1045287	0.00	14400000	2771
Private Nominal Saving and Investment Saving (mn. USD)	40800000	315000000	-1041097	6530000000	2845
Total Imports (mn. USD)	123085.10	260540.80	478.26	2537730	2484
Current Account Balance (mn. USD)	247.95	57578.74	-816646.00	420568.50	2769
Domestic Credit, Y-o-Y growth (%)	17.55	142.29	-4079.33	4093.33	1961
Unemployment Rate (%)	7.98	5.56	0.05	38.40	2389
Consumer Price Index, Period Avg., Y-o-Y growth (%)	38.03	359.28	-8.53	11749.63	2937
Population (mn. persons)	59.90	176.54	1.31	1410.08	3322
Tourist arrival (persons)	10100000	15600000	700.00	190000000	2220
GDP Deflator	96.68	19.08	6.29	482.49	2321
GDP Deflator, Y-o-Y growth (%)	11.36	52.55	-26.87	1489.50	2321

Table A2: Mean Values, Standard Deviations, and P-Value T-Tests of Difference in Mean of Transmission Channel Proxies in Pandemic and Non-pandemic States

	Pandemic		**Non-pandemic**		**P-value**
	Mean	Std.	Mean	Std.	
Uncertainty	5.247	10.210	0.182	2.110	0.000
Excess private saving (mill.)	-0.40	0.154	0.003	0.093	0.000
Air passengers (detrended)	-0.928	0.019	0.001	0.014	0.000
Container traffic (detrended)	-3.780	4.132	0.300	5.111	0.000

Notes: Uncertainty is the World Pandemic Uncertainty index. Excess private savings is the deviation of real private savings over GDP from its trend. Air passengers is the deviation of the number of air passengers over population from its trend. Container traffic is the deviation of the container traffic in ports in one hundred thousand over its trend. Trends obtained via the HP filter.

Table A3: Transmission Channels Regressions, Total Expenditure, All Countries

	(1) h=0	(2) h=1	(3) h=2	(4) h=3
Uncert. $\times I_t$	0.001*	0.002**	0.002**	0.003**
	(0.000)	(0.001)	(0.001)	(0.001)
Uncert. $\times(1-I_t)$	-0.000	0.000	0.001	0.001**
	(0.000)	(0.000)	(0.001)	(0.001)
Exc. Save $\times I_t$	0.051**	0.063**	0.064	0.067
	(0.026)	(0.030)	(0.040)	(0.056)
Exc. Save$\times(1-I_t)$	0.019	0.037**	0.040*	0.037
	(0.012)	(0.018)	(0.022)	(0.048)
Psgr. $\times I_t$	0.055	0.346	0.138	-0.081
	(0.181)	(0.290)	(0.368)	(0.459)
Psgr.$\times(1-I_t)$	-0.098	-0.025	0.037	-0.002
	(0.136)	(0.181)	(0.236)	(0.341)
Cont. traff.$\times I_t$	0.000	-0.003	-0.004	-0.003
	(0.002)	(0.003)	(0.004)	(0.005)
Cont. traff.$\times(1-I_t)$	0.000	-0.000	-0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.003)
Shock$\times$Uncert.$\times I_t$	0.113***	0.011	-0.054	-0.158
	(0.032)	(0.063)	(0.089)	(0.129)
Shock$\times$Uncert.$\times(1-I_t)$	0.243	0.298	0.234	0.205
	(0.263)	(0.315)	(0.326)	(0.347)
Shock$\times$Exc. Save$\times I_t$	0.854	0.877*	1.653**	0.812
	(0.776)	(0.496)	(0.690)	(0.802)
Shock$\times$Exc. Save$\times(1-I_t)$	-0.559	-0.341	-0.336	-0.739
	(0.639)	(0.837)	(1.037)	(1.246)
Shock$\times$Psgr.$\times I_t$	19.884	20.466	13.388	33.084**
	(15.317)	(24.344)	(17.177)	(16.073)
Shock$\times$Psgr$\times(1-I_t)$	-3.922	-8.219	-11.414	-12.275
	(3.986)	(5.794)	(7.177)	(8.868)
Shock$\times$Contain. traffic$\times I_t$	-0.199***	-0.284**	-0.274***	-0.348***
	(0.069)	(0.117)	(0.088)	(0.086)
Shock$\times$Contain. traffic$\times(1-I_t)$	0.001	0.015	0.022	0.041
	(0.018)	(0.025)	(0.028)	(0.038)
R-squared	0.707	0.783	0.817	0.842

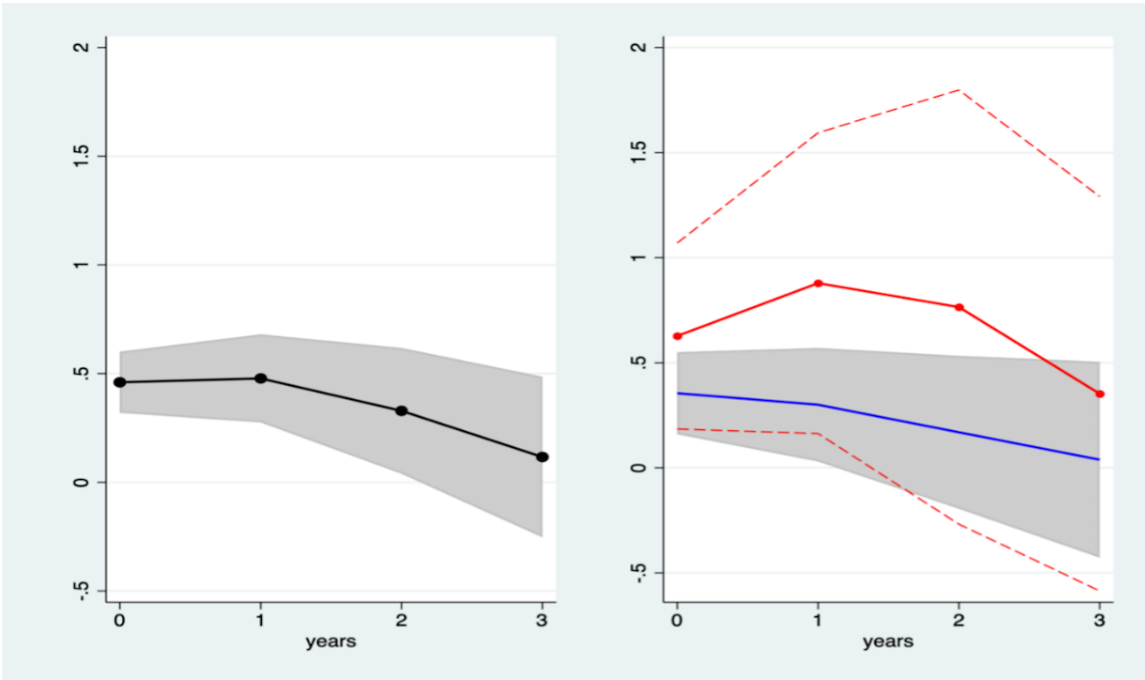
Notes: Output tables of Equation (4.5), estimated via the 1-step IV method. Fixed-effects and control variables omitted. I_t is the pandemic dummy, shock is the fiscal policy shock, Uncert. is the Pandemic Uncertainty Index, Exc. Save is de-trended private savings, Cont. Traff. is de-trended container traffic, Psgr is the de-trended number of air passengers. Robust standard errors in parenthesis, stars denote statistical significance at 1, 5 and 10%.

Table A4: The Role of Fiscal Position and Credibility

Panel (A): Debt-to-GDP ratio				
	T=0	T=1	T=2	T=3
Debt $\times I_t$	0.006	0.038***	0.056***	0.080***
	(-0.006)	(-0.011)	(-0.014)	(-0.017)
Debt $\times(1-I_t)$	0.009	0.039***	0.070***	0.096***
	(-0.006)	(-0.012)	(-0.016)	(-0.019)
Shock $\times$ Debt $\times I_t$	0.012	-0.122	-0.285	-0.11
	(-0.313)	(-0.433)	(-0.538)	(-0.485)
Shock $\times$ Debt $\times(1-I_t)$	-0.302***	-0.095	0.010	-0.087
	(-0.109)	(-0.152)	(-0.165)	(-0.213)
Panel (B): Fiscal Rule				
	T=0	T=1	T=2	T=3
FiscalRule $\times I_t$	-0.002	-0.003	-0.002	0.004
	(-0.007)	(-0.009)	(-0.012)	(-0.014)
FiscalRule $\times(1-I_t)$	-0.003	-0.004	-0.005	-0.002
	(-0.004)	(-0.006)	(-0.007)	(-0.008)
Shock $\times$ FiscalRule $\times I_t$	0.522**	0.119	-0.019	-0.2
	(-0.260)	(-0.421)	(-0.425)	(-0.375)
Shock $\times$ FiscalRule $\times(1-I_t)$	0.198	0.233	0.395	0.336
	(-0.179)	(-0.266)	(-0.292)	(-0.392)

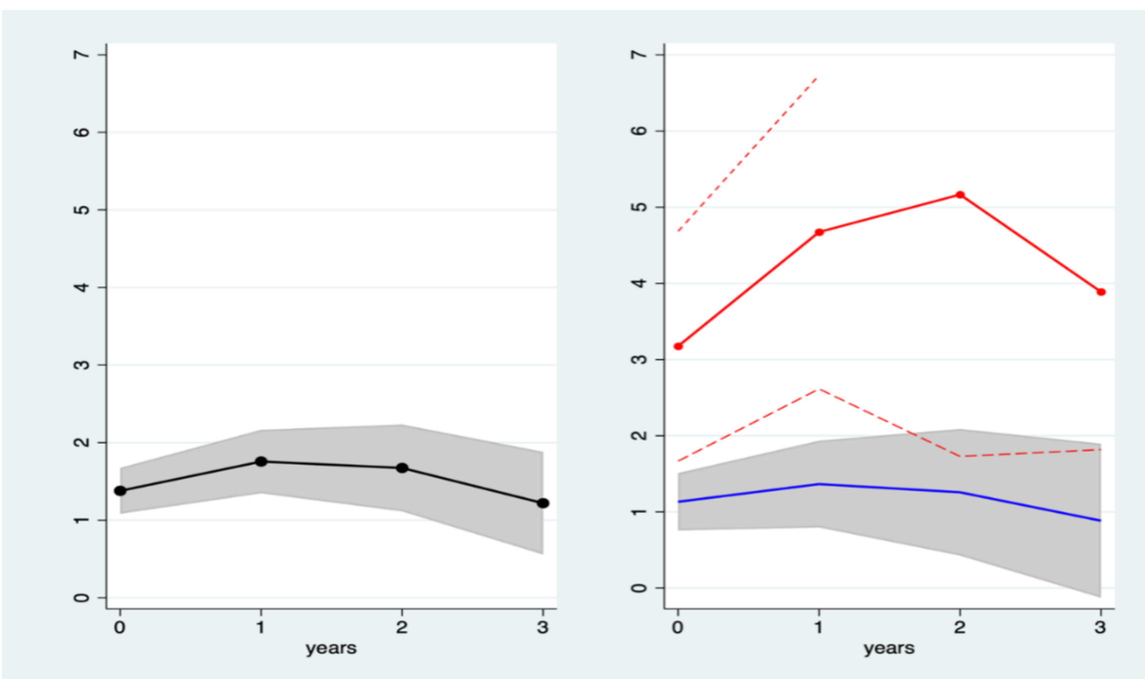
Notes: Output tables of Equation (4.5), estimated via the 1-step IV method. Fixed-effects and control variables omitted. Shock is the fiscal policy shock, Debt is the debt-to-GDP ratio, Fiscal Rule is a dummy variable equal to one if a country has a fiscal rule in place at a given year.

Figure A1: Cumulative Multipliers – Public Consumption, Advanced Countries



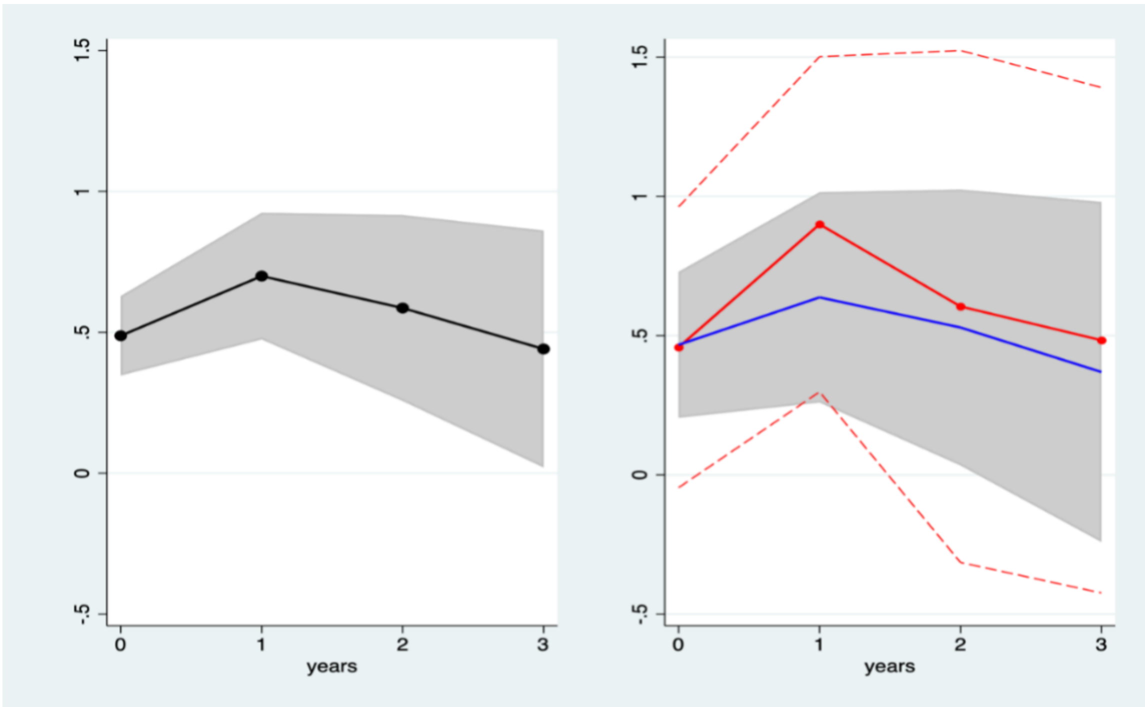
Notes: See Figure 4.1.

Figure A2: Cumulative Multipliers – Fixed Capital Formation, Advanced Countries



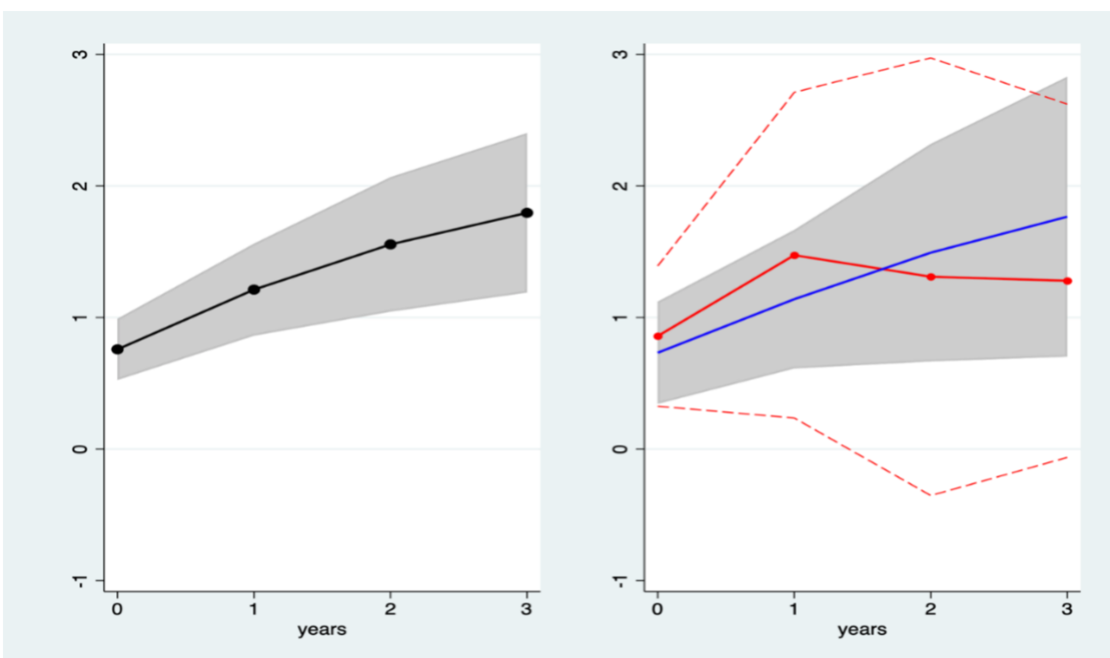
Notes: See Figure 4.1.

Figure A3: Cumulative Multipliers – Public Consumption, Non-advanced Countries



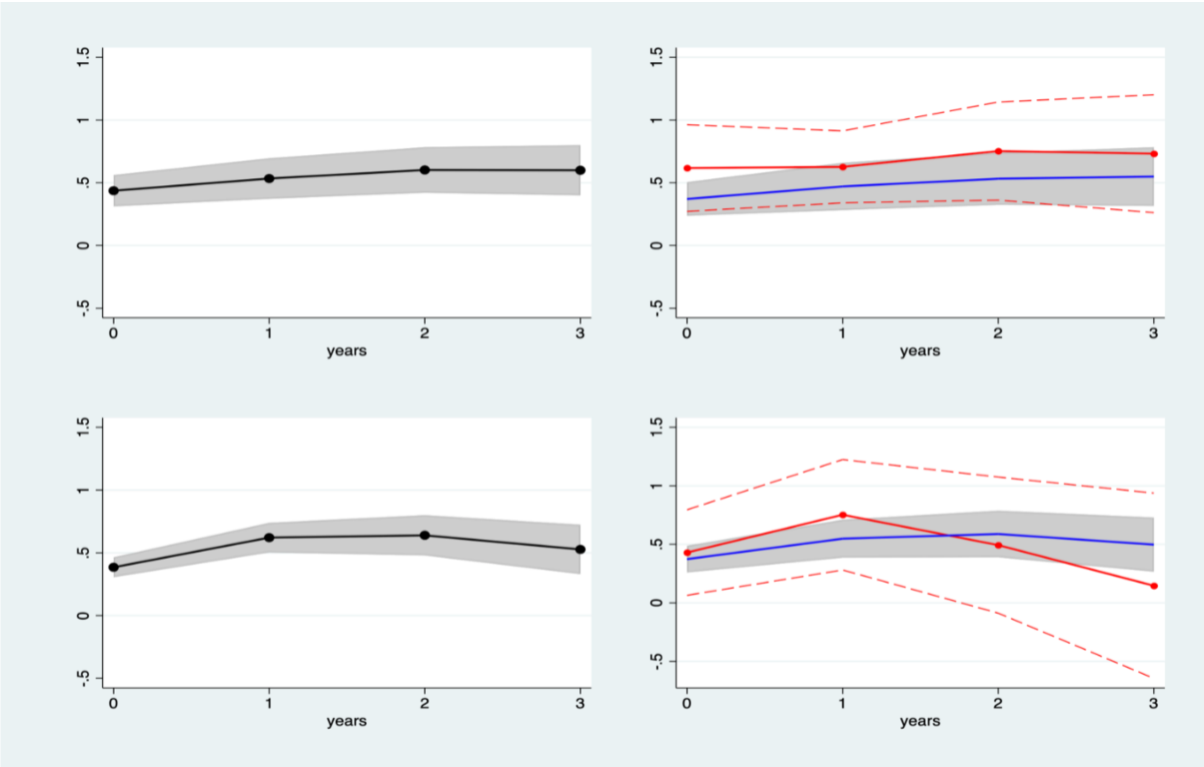
Notes: See Figure 4.1.

Figure A4: Cumulative Multipliers – Fixed Capital Formation, Non-advanced Countries



Notes: See Figure 4.1.

Figure A5: Multipliers with the Forecast Error Identification



Notes: Cumulative fiscal multipliers from Equation (4.3) (Left panel) and Equation (4.4) (Right panel). The black line represents the average (non-state dependent) fiscal multiplier. The red and blue lines represent multipliers in pandemic and non-pandemic states respectively. The shaded areas and dashed lines are corresponding 95% Driscoll and Kraay (1998) confidence intervals. The top panel shows the results of the forecast error identification, while the bottom panel shows the results of the baseline specification, with constraining the sample to match the forecast error identification sample.

Summary

This thesis aims to identify the effects of macroeconomic stabilization of the recent past, with a particular focus on the effect on government bond markets. The first two chapters study the reaction of Treasury markets to shifts in the demand or supply of government debt. The remaining two chapters explore macroeconomic stabilization in a broader context.

Chapter 1 identifies government bond supply shocks by following the price movements of government bonds around the announcements of auction volumes. We focus on the UK Debt Management Office's announcements and analyze how the issuance of additional debt impacts the term structure of interest rates. Our findings indicate that a standard deviation bond supply shock leads to a nominal yield increase of approximately 1-1.5 basis points. Additionally, real rates also rise by around 1-1.2 basis points, indicating a moderate reaction in inflation compensation. To uncover the transmission of this shock, we break down yields into expected short rates and risk premia. The results show that the supply shock predominantly affects the risk premia components, with minor effects on future expected average short-term rates, and no impact on expected inflation. Both the real term premium and the inflation risk premium respond positively to higher bond supply. To explain these observations, we propose an equilibrium term structure model, wherein risk-averse investors absorb shocks related to the supply of nominal bonds. Consequently, their equilibrium portfolio becomes more sensitive to duration and inflation risks, leading to increased prices for these risk factors. Consequently, risk premia and yields also rise. Additionally, since inflation-linked bonds remain unaffected by inflation risk, the breakeven inflation rate increases. The model further predicts that during periods of high risk aversion, the effects of the supply shock are more pronounced. Empirically, we confirm this by observing stronger yield reactions to the supply shock during times of financial market stress and when the effective lower bound is reached. In such situations, the increase in yields is primarily driven by higher risk premia, aligning with the predictions of the equilibrium model. Moreover, we find evidence suggesting that the impact of

the shock becomes more localized during market stress periods.

Chapter 2 uses high-frequency government bond futures price changes around German and Italian Treasury auctions to identify unexpected shifts in the demand for public debt. The study focuses on the effects of these demand shocks on secondary market yields of Treasury bonds. Positive demand shocks lead to significant and lasting negative movements in Treasury yields, up to 30 trading days. We examine location-specific demand shocks and find that in Germany, shocks at specific points on the yield curve have stronger effects on nearby maturities, while in Italy, positive demand shocks consistently decrease short-term interest rates, irrespective of the shock's location. Spillover effects into other euro area Treasury bond, corporate debt, and equity markets are also observed. German demand shocks have larger spillover effects on public debt yields in France and the Netherlands, whereas Italian spillovers are more prominent on Spain. Notably, the responses of equity markets and CDS spreads differ between the two countries. Increased demand for German bonds is associated with drops in stock prices and an increase in credit risk priced in CDS spreads. In contrast, Italy experiences stock price increases and decreases in CDS spreads following a sudden increase in demand for its bonds. These divergent responses are believed to stem from differences in how investors perceive the seemingly similar information. Higher demand for German Treasuries is linked to a "flight-to-safety" behavior, with investors rebalancing from riskier equities to safer bonds, resulting in increased Treasury and CDS spreads. For Italy, higher demand for its Treasuries signals a positive outlook for the economy, alleviating debt rollover concerns, leading to stock market increases, and decreasing Treasury and CDS spreads. These effects are amplified during times of high financial stress.

Chapter 3 provides fresh insights into the absorption of idiosyncratic output shocks in the US and the EA (Euro Area) through private and public risk-sharing channels. The study focuses on the capital channel (income from cross-ownership of assets), credit channel (cross-border lending and borrowing), and fiscal channel (federal or cross-country transfers) to understand how these risk-sharing mechanisms have evolved over time within each region. The analysis employs a novel time-varying parameter panel VAR model with stochastic volatility. The findings reveal that over the last few decades, risk-sharing has increased in both the US and the EA, primarily driven by private risk-sharing channels. In the US, the capital channel plays a dominant role, while in the EA, the credit channel is more significant. The presence of a larger federal system in the US leads to a more important role for the fiscal channel compared to the EA. Interestingly, the private risk-sharing channels have a more immediate impact in smoothing shocks, while the fiscal channel's effectiveness is more pronounced in the long run. The study also explores the degree of substitution and complementarity between the three risk-sharing channels. Strong substitution effects are observed between capital market smoothing and credit market smoothing in both regions, which could imply crowding out or support the spare-tire hypothesis. In the case of the US, there is evidence for complementarity between private (credit and capital) and public (fiscal) channels, supporting the argument from previous studies. Finally, for the US, the paper attempts to explain the time variation in the responses of risk-sharing channels to output shocks based on macroeconomic and financial determinants. The effectiveness of the capital and fiscal risk-sharing channels improves during weak economic conditions. Additionally, the fiscal and credit channels perform better under stronger financial integration and when a country has more fiscal space. However, monetary policy does not seem to have a significant influence on the functioning of risk-sharing channels.

Finally, Chapter 4 paper examines contemporaneous and near-term multipliers fiscal multipliers during health crises episodes. It reveals that fiscal expansions are effective at the onset of the pandemic and in the years following the health crisis. The cumulative multiplier is about fifty percent higher in the year after a pandemic, compared to non-pandemic times. The differences in the multipliers are attributed to three main factors: uncertainty, suppressed demand, and supply bottlenecks. Our findings are supported by robustness checks, including the inclusion of the height of the COVID-19 pandemic (2020) in the estimation period, controlling for additional variables, and testing for a fiscal credibility channel. Additionally, an alternative identification strategy using public expenditure forecast errors confirms the results. The paper suggests that fiscal stimulus packages during health crises and pandemics may have a larger and more prolonged growth impact than commonly assumed. The significant fiscal support deployed during the COVID-19 pandemic could have an even more substantial impact than in non-pandemic times.

Samenvatting

Dit proefschrift heeft tot doel de effecten van macro-economische stabilisatie in het recente verleden te identificeren, met speciale aandacht voor de impact op overheidsobligatiemarkten. De eerste twee hoofdstukken onderzoeken de reactie van de schatkistmarkten op verschuivingen in de vraag naar of het aanbod van overheidsschuld. De overige twee hoofdstukken verkennen macro-economische stabilisatie in een breder kader.

Hoofdstuk 1 identificeert schokken in het aanbod van overheidsobligaties door de prijsbewegingen van overheidsobligaties rondom aankondigingen van veilingvolumes te volgen. We concentreren ons op aankondigingen van het Britse Bureau voor Schuldbeheer en analyseren hoe de uitgifte van extra schuld de rentestructuur van de rente beïnvloedt. Onze bevindingen geven aan dat een standaardafwijking in het aanbod van obligaties leidt tot een stijging van de nominale rente van ongeveer 1-1,5 basispunten. Bovendien stijgen de reële rentes ook met ongeveer 1-1,2 basispunten, wat wijst op een gematigde reactie in de inflatiecompensatie. Om de overdracht van deze schok te ontrafelen, splitsen we de rentes op in verwachte korte rentes en risicopremies. De resultaten tonen aan dat de aanbodschok voornamelijk van invloed is op de risicopremiecomponenten, met geringe effecten op de verwachte gemiddelde korte rentes in de toekomst, en geen invloed op de verwachte inflatie. Zowel de reële termijnpremie als de inflatierisicopremie reageren positief op een groter aanbod van obligaties. Om deze observaties te verklaren, stellen we een evenwichtsmodel voor van de rentestructuur, waarin risicomijdende beleggers schokken met betrekking tot het aanbod van nominale obligaties absorberen. Hierdoor wordt hun evenwichtsportefeuille gevoeliger voor looptijdrisico en inflatierisico, wat leidt tot hogere prijzen voor deze risicofactoren. Dit resulteert op zijn beurt in hogere risicopremies en rentes. Bovendien, aangezien inflatie-gekoppelde obligaties niet worden beïnvloed door inflatierisico, neemt het inflatieverwachtingstarief toe. Het model voorspelt verder dat de effecten van de aanbodschok tijdens periodes van hoge risico-aversie sterker zijn. Empirisch bevestigen we dit door sterkere rentereacties op de aanbodschok te observeren tijdens perioden van financiële marktspanning en wanneer de effectieve ondergrens wordt bereikt. In dergelijke situaties wordt de stijging van de rentes voornamelijk veroorzaakt door hogere risicopremies, wat overeenkomt met de voorspellingen van het evenwichtsmodel. Bovendien vinden we bewijs dat de impact van de schok tijdens periodes van marktspanning meer gelokaliseerd is.

Hoofdstuk 2 maakt gebruik van prijsveranderingen van overheidsobligatiefutures rondom Duitse en Italiaanse schatkistveilingen om onverwachte verschuivingen in de vraag naar overheidsschuld te identificeren. De studie richt zich op de effecten van deze vraagschokken op de rendementen op de secundaire markt van schatkistobligaties. Positieve vraagschokken leiden tot aanzienlijke en langdurige negatieve bewegingen in de rendementen van schatkistobligaties, tot wel 30 handelsdagen. We onderzoeken vraagschokken op specifieke locaties en constateren dat in Duitsland schokken op specifieke punten op de yield curve sterkere effecten hebben op nabijgelegen looptijden, terwijl in Italië positieve vraagschokken consistent leiden tot een daling van de korte rentetarieven, ongeacht de locatie van de schok. Ook spillovereffecten naar andere eurogebied-schatkistobligaties, bedrijfsschulden en aandelenmarkten worden waargenomen. Duitse vraagschokken hebben grotere spillovereffecten op de rendementen van schatkistobligaties in Frankrijk en Nederland, terwijl de Italiaanse spillovers vooral van invloed zijn op Spanje. Opvallend is dat de reacties van aandelenmarkten en CDS-spreads verschillen tussen de twee landen. Een toename van de vraag naar Duitse obligaties gaat gepaard met een daling van de aandelenkoersen en een toename van het kredietrisico, zoals weerspiegeld in CDS-spreads. In tegenstelling hiermee leidt een plotselinge toename van de vraag naar Italiaanse obligaties tot stijgingen van de aandelenkoersen en een afname van CDS-spreads. Deze uiteenlopende reacties worden toegeschreven aan verschillen in hoe beleggers de ogenschijnlijk vergelijkbare informatie waarnemen. Een hogere vraag naar Duitse schatkistobligaties wordt gezien als een teken van "vlucht naar veiligheid", waarbij beleggers hun portefeuilles herbalanceerden van risicovollere aandelen naar veiliger obligaties, wat resulteert in hogere CDS-spreads. Voor Italië wordt een grotere vraag naar schatkistobligaties gezien als een positief signaal over de economie, waardoor zorgen over de rollover van schulden worden weggenomen. Dit leidt tot stijgingen van aandelenkoersen en dalingen van CDS-spreads. Deze effecten worden versterkt tijdens periodes van hoge financiële stress.

Hoofdstuk 3 biedt nieuwe inzichten in de absorptie van idiosyncratische outputschokken in de VS en het Eurogebied (EA) via private en publieke risicodelingskanalen. De studie richt zich op het kapitaalkanaal (inkomsten uit het wederzijdse eigendom van activa), het kredietkanaal (grensoverschrijdende leningen en leningen) en het begrotingskanaal (federale of grensoverschrijdende overdrachten) om te begrijpen hoe deze risicodelingsmechanismen in de loop van de tijd binnen elke regio zijn geëvolueerd. De analyse maakt gebruik van een nieuw model met parameters die in de tijd variëren en een stochastische volatiliteit. De bevindingen tonen aan dat in de afgelopen decennia de risicodeling is toegenomen in zowel de VS als het EA, voornamelijk gedreven door private risicodelingskanalen. In de VS speelt het kapitaalkanaal een dominante rol, terwijl in het EA het kredietkanaal belangrijker is. Het grotere federale systeem in de VS zorgt ervoor dat het begrotingskanaal een belangrijkere rol speelt in vergelijking met het EA. Opvallend is dat de private risicodelingskanalen een directer effect hebben bij het dempen van schokken, terwijl de effectiviteit van het begrotingskanaal op de lange termijn sterker is. Het onderzoek onderzoekt ook de mate van substitutie en complementariteit tussen de drie risicodelingskanalen. Sterke substitutie-effecten worden waargenomen tussen het gladstrijken van kapitaalmarkten en het gladstrijken van de kredietmarkt in beide regio's, wat kan wijzen op verdringing of het ondersteunen van de reserveband-hypothese. In het geval van de VS is er ook bewijs voor complementariteit tussen private (krediet en kapitaal) en publieke (begrotings) kanalen, wat het argument van eerdere studies ondersteunt. Tot slot probeert het paper voor de VS de tijdvariatie in de reacties van risicodelingskanalen op outputschokken te verklaren op basis van macro-economische en financiële determinanten. De effectiviteit van het kapitaal- en begrotingsrisicodelingskanaal neemt toe tijdens zwakke economische omstandigheden. Bovendien presteren het begrotings- en kredietkanaal beter bij sterkere financiële integratie en wanneer een land meer begrotingsruimte heeft. Daarentegen lijkt monetair beleid geen grote invloed te hebben op het functioneren van risicodelingskanalen.

Ten slotte onderzoekt hoofdstuk 4 de multipliers van de overheidsuitgaven tijdens episodes van gezondheidscrises, zowel contemporaine als op de korte termijn. Het toont aan dat expansief begrotingsbeleid effectief is bij het begin van de pandemie en in de jaren na de gezondheidscrisis. De cumulatieve multiplier is ongeveer vijftig procent hoger in het jaar na een pandemie, vergeleken met niet-pandemische tijden. De verschillen in de multipliers worden toegeschreven aan drie belangrijke factoren: onzekerheid, onderdrukte vraag en aanbodverstoringen. Onze bevindingen worden ondersteund door robuustheidstests, waaronder de opname van de hoogte van de COVID-19-pandemie (2020) in de schatting, controle voor aanvullende variabelen en testen voor een begrotingsgeloofwaardigheidskanaal. Bovendien bevestigt een alternatieve identificatiestrategie met behulp van fouten in publieke uitgavenprognoses de resultaten. Het paper suggereert dat stimuleringspakketten tijdens gezondheidscrises en pandemieën een grotere en langduriger groei-impact kunnen hebben dan vaak wordt aangenomen. De significante begrotingssteun die werd ingezet tijdens de COVID-19-pandemie kan een nog grotere impact hebben dan in niet-pandemische tijden.

List of Authors

- Chapter 1 is independent work and is available as MNB Working Papers 2022/6.

- Chapter 2 is joint work with Massimo Giuliodori and is published in the Journal of Money, Credit and Banking Volume 54, Issue 7, Pages 1997-2028, October 2022.

- Chapter 3 is joint work with Jacopo Cimadomo, Massimo Giuliodori and Haroon Mumtaz.

- Chapter 4 is joint work with Tidiane Kinda and Kaustubh Chahande and is available as IMF Working Paper No. 2022/149.

- During the preparation of parts of this work the author used ChatGPT in order to improve language and readability. After using this tool, the author reviewed and edited the content as needed and take full responsibility for the content of the publication.

Bibliography

Abiad, Abduland Furceri, D. and P. Topalova (2016): "The macroeconomic effects of public investment: evidence from advanced economies," *Journal of Macroeconomics*, 50, 224–240.

Abrahams, M., T. Adrian, R. K. Crump, E. Moench, and R. Yu (2016): "Decomposing real and nominal yield curves," *Journal of Monetary Economics*, 84, 182–200.

Adrian, T., R. K. Crump, and E. Moench (2013): "Pricing the term structure with linear regressions," *Journal of Financial Economics*, 110, 110–138.

AFME (2020): "European Primary Dealers Handbook 2019/2020," .

Afonso, A. and D. Furceri (2008): "EMU enlargement, stabilization costs and insurance mechanisms," *Journal of International Money and Finance*, 27, 169–187.

Aguirregabiria, V., R. Clark, and H. Wang (2010): "Bank expansion after the Riegle-Neal Act: the role of diversification of geographic risk," *University of Toronto Department of Economics Working Paper*.

Ahir, H., N. Bloom, and D. Furceri (2022): "The world uncertainty index," Tech. rep.

Albertazzi, U. and M. Bottero (2014): "Foreign bank lending: evidence from the Global Financial Crisis," *Journal of International Economics*, 92, 22–35.

Alcidi, C., P. D'Imperio, and G. Thirion (2017): "Risk-sharing and consumption-smoothing patterns in the US and the Euro Area: a comprehensive comparison," *CEPS Working Paper No. 2017/04*.

Alesina, A., M. De Broeck, A. Prati, and G. Tabellini (1992): "Default risk on government debt in OECD countries," *Economic policy*, 7, 427–463.

Alloza, M. (2022): "Is fiscal policy more effective during recessions?" *International Economic Review*, 63, 1271–1292.

Altavilla, C., G. Carboni, and R. Motto (2015): "Asset purchase programmes and financial markets: lessons from the euro area," *ECB Working Paper Series No. 1864*.

Ang, A., G. Bekaert, and M. Wei (2008): "The term structure of real rates and expected inflation," *The Journal of Finance*, 63, 797–849.

Ang, A. and M. Piazzesi (2003): "A no-arbitrage vector autoregression of term structure dynamics with macroeconomic and latent variables," *Journal of Monetary Economics*, 50, 745–787.

Arslanalp, S. and T. Tsuda (2014): "Tracking global demand for advanced economy sovereign debt," *IMF Economic Review*, 62, 430–464.

Asdrubali, P. and S. Kim (2004): "Dynamic risk-sharing in the United States and Europe," *Journal of Monetary Economics*, 51, 809–836.

Asdrubali, P., S. Kim, F. M. Pericoli, and P. Poncela (2023): "Risk sharing channels in OECD countries: a heterogeneous panel VAR approach," *Journal of International Money and Finance*, 131.

Asdrubali, P., B. E. Sørensen, and O. Yosha (1996): "Channels of inter-state risk-sharing: United States 1963–1990," *The Quarterly Journal of Economics*, 111, 1081–1110.

Athanasoulis, S. G. and E. v. Wincoop (2001): "Risk sharing within the United States: what do financial markets and fiscal federalism accomplish?" *Review of Economics and Statistics*, 83, 688–698.

Auerbach, A., Y. Gorodnichenko, P. B. McCrory, and D. Murphy (2022): "Fiscal multipliers in the COVID19 recession," *Journal of International Money and Finance*, 126, 102669.

Auerbach, A. J. and Y. Gorodnichenko (2012a): "Fiscal multipliers in recession and expansion," in *Fiscal Policy After the Financial Crisis*, University of Chicago Press, pages 63–98.

——— (2012b): "Measuring the output responses to fiscal policy," *American Economic Journal: economic Policy*, 4, 1–27.

——— (2016): "Effects of fiscal shocks in a globalized world," *IMF Economic Review*, 64, 177–215.

Auerbach, A. J., Y. Gorodnichenko, and D. Murphy (2021): "Inequality, fiscal policy and COVID19 restrictions in a demand-determined economy," *European Economic Review*, 137, 103810.

Balduzzi, P., E. J. Elton, and C. T. Green (2001): "Economic news and bond prices: evidence from the US Treasury market," *Journal of Financial and Quantitative Analysis*, 36, 523–543.

Balli, F., S. A. Basher, and H. O. Balli (2013): "International income risk-sharing and the Global Financial Crisis of 2008–2009," *Journal of Banking & Finance*, 37, 2303–2313.

Banbura, M., D. Giannone, and L. Reichlin (2010): "Large bayesian vector auto regressions," *Journal of Applied Econometrics*, 25, 71–92.

Bayer, C., B. Born, and R. Luetticke (2020): "The liquidity channel of fiscal policy," *CEPR Discussion Paper No. DP14883*.

Bayer, C., B. Born, R. Luetticke, and G. J. Müller (2023): "The coronavirus stimulus package: how large is the transfer multiplier," *The Economic Journal*, 133, 1318–1347.

Beetsma, R., M. Giuliodori, F. de Jong, and D. Widijanto (2016): "Price effects of sovereign debt auctions in the euro-zone: the role of the crisis," *Journal of Financial Intermediation*, 25, 30–53.

Beetsma, R., M. Giuliodori, J. Hanson, and F. de Jong (2018): "Bid-to-cover and yield changes around public debt auctions in the euro area," *Journal of Banking & Finance*, 87, 118–134.

Beetsma, R., M. Giuliodori, and F. Klaassen (2006): "Trade spill-overs of fiscal policy in the European Union: a panel analysis," *Economic Policy*, 21, 640–687.

Beine, M., A. Cosma, and R. Vermeulen (2010): "The dark side of global integration: Increasing tail dependence," *Journal of Banking & Finance*, 34, 184–192.

Bekaert, G. and A. Ermolov (2023): "International yield comovements," *Journal of Financial and Quantitative Analysis*, 58, 250–288.

Belke, A. and D. Gros (2015): "Banking Union as a shock absorber," *ROME Working Papers, No 2015/02.*

Bénétrix, A. S. and P. R. Lane (2013): "Fiscal shocks and the real exchange rate," *International Journal of Central Banking*, 9.

Berg, T. O. (2019): "Business uncertainty and the effectiveness of fiscal policy in Germany," *Macroeconomic Dynamics*, 23, 1442–1470.

Bianchi, F., H. Mumtaz, and P. Surico (2009): "The great moderation of the term structure of UK interest rates," *Journal of Monetary Economics*, 56, 856–871.

Bikhchandani, S. and C. Huang (1989): "Auctions with resale markets: an exploratory model of Treasury Bill markets," *The Review of Financial Studies*, 2, 311–339.

Blanchard, O. (2023): *Fiscal Policy Under Low Interest Rates*, MIT press.

Blanchard, O. and R. Perotti (2002): "An empirical characterization of the dynamic effects of changes in government spending and taxes on output," *The Quarterly Journal of Economics*, 117, 1329–1368.

Bloom, N. (2014): "Fluctuations in uncertainty," *Journal of Economic Perspectives*, 28, 153–176.

Bloom, N., M. Floetotto, N. Jaimovich, I. Saporta-Eksten, and S. J. Terry (2018): "Really uncertain business cycles," *Econometrica*, 86, 1031–1065.

Boermans, M. and R. Vermeulen (2018): "Quantitative easing and preferred habitat investors in the euro area bond market," *De Nederlandsche Bank Working Paper No. 586.*

Bonadio, B., Z. Huo, A. A. Levchenko, and N. Pandalai-Nayar (2021): "Global supply chains in the pandemic," *Journal of International Economics*, 133, 103534.

Born, B. and G. J. Müller (2012): "Government spending shocks in quarterly and annual time series," *Journal of Money, Credit and Banking*, 44, 507–517.

Bretscher, L., A. Hsu, and A. Tamoni (2020): "Fiscal policy driven bond risk premia," *Journal of Financial Economics*, 138, 53–73.

Brunet, G. and S. Hlatshwayo (2021): "War, pandemic, and household saving: the COVID-19 pandemic through the lens of WWII," *Manuscript, Wesleyan Universit.*

Buti, M. and N. Carnot (2018): “The case for a central fiscal capacity in EMU,” *VOX CEPR Policy Portal, 7th Dec, 2018*.

Caggiano, G., E. Castelnuovo, V. Colombo, and G. Nodari (2015): “Estimating fiscal multipliers: news from a non-linear world,” *The Economic Journal*, 125, 746–776.

Cammack, E. B. (1991): “Evidence on bidding strategies and the information in Treasury Bill auctions,” *Journal of Political Economy*, 99, 100–130.

Canova, F. and M. Ciccarelli (2004): “Forecasting and turning point predictions in a Bayesian panel VAR model,” *Journal of Econometrics*, 120, 327 – 359.

Canova, F. and M. O. Ravn (1996): “International consumption risk sharing,” *International Economic Review*, 37, 573–601.

Canzoneri, M., F. Collard, H. Dellas, and B. Diba (2016): “Fiscal multipliers in recessions,” *The Economic Journal*, 126, 75–108.

Carter, C. K. and R. Kohn (1994): “On Gibbs sampling for state space models,” *Biometrika*, 81, 541–553.

Chadha, J., P. Turner, and F. Zampolli (2013): “The interest rate effects of government debt maturity,” *BIS Working Paper No. 415*.

Christiano, L., M. Eichenbaum, and S. Rebelo (2011): “When is the government spending multiplier large?” *Journal of Political Economy*, 119, 78–121.

Christiano, L. J., M. Eichenbaum, and C. L. Evans (1999): “Monetary policy shocks: what have we learned and to what end?” *Handbook of Macroeconomics*, 1, 65–148.

Cimadomo, J., G. Ciminelli, O. Furtuna, and M. Giuliodori (2020): “Private and public risk sharing in the Euro Area,” *European Economic Review*, 121, 103347.

Cimadomo, J., E. Gordo Mora, and A. A. Palazzo (2022): “Enhancing private and public risk sharing,” *ECB Occasional Paper No. 2022/306*.

Coenen, G., C. J. Erceg, C. Freedman, D. Furceri, M. Kumhof, R. Lalonde, D. Laxton, J. Lindé, A. Mourougane, D. Muir, et al. (2012): “Effects of fiscal stimulus in structural models,” *American Economic Journal: Macroeconomics*, 4, 22–68.

Cogley, T. and T. J. Sargent (2005): “Drifts and Volatilities: Monetary Policies and Outcomes in the Post WWII US,” *Review of Economic Dynamics*, 8, 262–302.

Cohen-Setton, J., E. Gornostay, and C. Ladreit (2019): "Aggregate effects of budget stimulus: evidence from the large fiscal expansions database," *Peterson Institute for International Economics Working Paper*.

Coibion, O., Y. Gorodnichenko, and M. Weber (2020): "How did US consumers use their stimulus payments?" *NBER Working Paper 27693*.

Cole, H. L. and T. J. Kehoe (2000): "Self-fulfilling debt crises," *The Review of Economic Studies*, 67, 91–116.

Corhay, A., T. Kind, H. Kung, and G. Morales (2021): "Discount rates, debt maturity, and the fiscal theory," *SAFE Working Paper No. 323*.

Culbertson, J. M. (1957): "The term structure of interest rates," *The Quarterly Journal of Economics*, 71, 485–517.

Dai, Q. and T. Philippon (2005): "Fiscal policy and the term structure of interest rates," *NBER Working Papers 11574*.

D'Amico, S. and T. Seida (2020): "Unexpected supply effects of uantitative easing and tightening," *FRB Chicago Working Paper No. 2020-17*.

Davoodi, M. H. R., P. Elger, A. Fotiou, M. D. Garcia-Macia, X. Han, A. Lagerborg, W. R. Lam, and M. P. A. Medas (2022): *Fiscal rules and fiscal councils: recent trends and performance during the COVID-19 pandemic*, International Monetary Fund.

De Santis, R. A. (2020): "Impact of the asset purchase programme on euro area government bond yields using market news," *Economic Modelling*, 86, 192–209.

Del Negro, M. (1998): "Aggregate risk and risk sharing across US states and across European countries," *Manuscript, Yale University 1998*.

Demyanyk, Y., C. Ostergaard, and B. E. Sørensen (2007): "US banking deregulation, small businesses, and interstate insurance of personal income," *The Journal of Finance*, 62, 2763–2801.

——— (2008): "Risk sharing and portfolio allocation in EMU," *European Economy Discussion Papers No. 334, July 2008*.

Di Gangi, D., V. S. Lazarov, A. Mankodi, and L. Silvestri (2022): "Links between government bond and futures markets: dealer-client relationships and price discovery in the UK," *Bank of England Working Paper No. 991.*

Diez de los Rios, A. (2020): "A portfolio-balance model of inflation and yield curve determination," *Bank of Canada Working Paper 2020-6.*

Dipartimento Del Tresoro (2017): "Guidelines for Public Debt Management," .

DMO, United Kingdom Debt Management Office. (2021): "Official Operations in the Gilt Market: An Operational Notice," .

Dobrev, P. (2019): "Demand for US Treasuries after the 2008-09 financial crisis: empirical evidence and potential policy implications," *Manuscript, Drexel University.*

Driscoll, J. C. and A. C. Kraay (1998): "Consistent covariance matrix estimation with spatially dependent panel data," *Review of Economics and Statistics*, 80, 549–560.

Droste, M., Y. Gorodnichenko, and W. Ray (2021): "Unbundling quantitative easing: taking a cue from Treasury auctions," Manuscript, university of california, berkeley.

Duffie, Darrell (1999): "Credit swap valuation," *Financial Analysts Journal*, 55, 73–87.

Duprey, T., B. Klaus, and T. Peltonen (2017): "Dating systemic financial stress episodes in the EU countries," *Journal of Financial Stability*, 32, 30–56.

ECB, European Central Bank. (2018): "Review of recent developments in market-based measures of inflation expectations and their drivers," *Economic Bulletin, Issue 2018/6.*

——— (2022): "Financial Integration and Structure in the Euro Area, 2022 - April," Tech. rep.

Ederington, L. H. and J. H. Lee (1993): "How markets process information: news releases and volatility," *The Journal of Finance*, 48, 1161–1191.

Eggertsson, G. B. and P. Krugman (2012): "Debt, deleveraging, and the liquidity trap: a Fisher-Minsky-Koo approach," *The Quarterly Journal of Economics*, 127, 1469–1513.

Ehrmann, M. and M. Fratzscher (2017): "Euro area government bonds – fragmentation and contagion during the sovereign debt crisis," *Journal of International Money and Finance*, 70, 26–44.

Eichenbaum, M. S., S. Rebelo, and M. Trabandt (2021): "The macroeconomics of epidemics," *The Review of Financial Studies*, 34, 5149–5187.

Eisl, A., C. Ochs, N. Osadchiy, and M. G. Subrahmanyam (2019): "The linkage between primary and secondary markets for Eurozone sovereign debt: free flow or bottleneck?" *Manuscript, Paper presented at The 2nd International Conference on Sovereign Bond Markets*.

Ejsing, J. and J. Sihvonen (2009): "Liquidity premia in German government bonds," *ECB Working Paper No. 1081*.

Elenev, V., T. Landvoigt, and S. Van Nieuwerburgh (2022): "Can the COVID bailouts save the economy?" *Economic Policy*, 37, 277–330.

Engen, E. M. and R. G. Hubbard (2004): "Federal government debt and interest rates," *NBER macroeconomics annual*, 19, 83–138.

Evans, M. D. (1998): "Real rates, expected inflation, and inflation risk premia," *The Journal of Finance*, 53, 187–218.

Farhi, E. and I. Werning (2016): "Fiscal multipliers: liquidity traps and currency unions," in *Handbook of macroeconomics*, Elsevier, vol. 2, 2417–2492.

——— (2017): "Fiscal unions," *American Economic Review*, 107, 3788–3834.

Faria-e Castro, M. (2021): "Fiscal policy during a pandemic," *Journal of Economic Dynamics and Control*, 125, 104088.

Fazzari, S. M., J. Morley, and I. Panovska (2015): "State-dependent effects of fiscal policy," *Studies in Nonlinear Dynamics & Econometrics*, 19, 285–315.

Fleming, M. J. and W. Liu (2016): "Intraday pricing and liquidity effects of US Treasury auctions," *Manuscript, Federal Reserve Bank of New York*.

Fleming, M. J. and E. M. Remolona (1997): "What moves the bond market?" *Economic policy review*, 3.

——— (1999): "Price formation and liquidity in the US Treasury market: the response to public information," *The Journal of Finance*, 54, 1901–1915.

Foresti, P. and O. Napolitano (2022): "Risk sharing in the EMU: a time-varying perspective," *Journal of Common Market Studies*, 60, 319–336.

Forni, M. and L. Gambetti (2016): "Government spending shocks in open economy VARs," *Journal of International Economics*, 99, 68–84.

Fuhrer, L. M. and J. Giese (2021): "Gilt auctions and secondary market dynamics," *Finance Research Letters*, 38, 101400.

Furceri, D. and A. Zdzienicka (2015): "The Euro Area crisis: need for a supranational fiscal risk sharing mechanism?" *Open Economies Review*, 26, 683–710.

Gale, W. G. and P. R. Orszag (2003): "Economic effects of sustained budget deficits," *National Tax Journal*, 463–485.

Garbade, K. and W. L. Silber (1979): "Dominant and satellite markets: a study of dually-traded securities," *The Review of Economics and Statistics*, 61, 455–60.

Garbade, K. D. and W. L. Silber (1983): "Price movements and price discovery in futures and cash markets," *The Review of Economics and Statistics*, 65, 289–297.

Gelman, A. (2006): "Prior distributions for variance parameters in hierarchical odels," *Bayesian Analysis*, 1, 1–19.

Gertler, M. and P. Karadi (2015): "Monetary Policy Surprises, Credit Costs, and Economic Activity," *American Economic Journal: Macroeconomics*, 7, 44–76.

Ghassibe, M. and F. Zanetti (2022): "State dependence of fiscal multipliers: the source of fluctuations matters," *Journal of Monetary Economics*, 132, 1–23.

Giovannini, A., D. Ioannou, and L. Stracca (2022): "Public and private risk sharing: friends or foes? The interplay between different forms of risk sharing," *ECB Occasional Paper No. 2022-295.*

Glover, A., J. Heathcote, D. Krueger, and J.-V. Ríos-Rull (2023): "Health versus wealth: on the distributional effects of controlling a pandemic," *Journal of Monetary Economics.*

GOLDREICH, D. (2007): "Underpricing in discriminatory and uniform-price treasury auctions," *The Journal of Financial and Quantitative Analysis*, 42, 443–466.

GOMES, F., A. MICHAELIDES, AND V. POLKOVNICHENKO (2013): "Fiscal policy and asset prices with incomplete markets," *The Review of Financial Studies*, 26, 531–566.

GORODNICHENKO, Y. AND W. RAY (2017): "The effects of quantitative easing: taking a cue from treasury auctions," .

GREENWOOD, R. AND D. VAYANOS (2014): "Bond supply and excess bond returns," *Review of Financial Studies*, 27, 663–713.

GREMBI, V., T. NANNICINI, AND U. TROIANO (2016): "Do fiscal rules matter?" *American Economic Journal: applied Economics*, 8, 1–30.

GUERRIERI, V., G. LORENZONI, L. STRAUB, AND I. WERNING (2022): "Macroeconomic implications of COVID-19: Can negative supply shocks cause demand shortages?" *American Economic Review*, 112, 1437–1474.

GÜRKAYNAK, R. S., B. SACK, AND E. SWANSON (2005): "Do actions speak louder than words? The response of asset prices to monetary policy actions and statements," *International Journal of Central Banking*, 1, 55–93.

HALL, R. E. (2009): "By how much does GDP rise if the government buys more output?" *NBER Working Paper 15496.*

HAMILTON, J. D. AND J. C. WU (2012): "The effectiveness of alternative monetary policy tools in a zero lower bound environment," *Journal of Money, Credit and Banking*, 44, 3–46.

HANSON, S. G. AND J. C. STEIN (2015): "Monetary policy and long-term real rates," *Journal of Financial Economics*, 115, 429–448.

HAUPTMEIER, S., F. HOLM-HADULLA, AND T. RENAULT (2022): "Risk sharing and monetary policy transmission," *ECB Working Paper No. 2746.*

HE, Z. AND A. KRISHNAMURTHY (2013): "Intermediary asset pricing," *American Economic Review*, 103, 732–70.

HOFFMANN, M., E. MASLOV, I. STEWEN, AND B. E. SØRENSEN (2018): "Are banking and capital markets union complements? Evidence from channels of risk sharing in the Eurozone," *CEPR Discussion Paper No. DP13254.*

Hoffmann, M. and I. Shcherbakova-Stewen (2011): "Consumption risk sharing over the business cycle: the role of small firms' access to credit markets," *Review of Economics and Statistics*, 93, 1403–1416.

Hollo, D., M. Kremer, and M. Lo Duca (2012): "CISS - a composite indicator of systemic stress in the financial system," *ECB Working paper No. 1426*.

Horvath, R., L. Kaszab, and A. Marsal (2021): "Fiscal policy and the nominal term premium," *Journal of Money, Credit and Banking*.

Huang, R. D., J. Cai, and X. Wang (2002): "Information-based trading in the Treasury note interdealer broker market," *Journal of Financial Intermediation*, 11, 269–296.

Huidrom, R., M. A. Kose, J. J. Lim, and F. L. Ohnsorge (2020): "Why do fiscal multipliers depend on fiscal positions?" *Journal of Monetary Economics*, 114, 109–125.

Ilzetzki, E., E. G. Mendoza, and C. A. Végh (2013): "How big (small?) are fiscal multipliers?" *Journal of monetary economics*, 60, 239–254.

ITC Markets (2017): "Government Bond Auctions: How to Read Bond Auction Results," .

Jacquier, E., N. G. Polson, and P. E. Rossi (2004): "Bayesian analysis of stochastic volatility models with fat-tails and correlated errors," *Journal of Econometrics*, 122, 185–212.

Jarocinski, M. (2010): "Responses to monetary policy shocks in the east and the west of Europe: a comparison," *Journal of Applied Econometrics*, 25, 833–868.

Jayaratne, J. and P. E. Strahan (1997): "The benefits of branching deregulation," *Economic Policy Review*, 3.

Jo, Y. J. and S. Zubairy (2022): "State dependent government spending multipliers: downward nominal wage rigidity and sources of business cycle fluctuations," *NBER Working Paper No. 30025*.

Jordà, Ò. (2005): "Estimation and inference of impulse responses by local projections," *American Economic Review*, 95, 161–182.

Joyce, M. A., P. Lildholdt, and S. Sorensen (2010): “Extracting inflation expectations and inflation risk premia from the term structure: a joint model of the UK nominal and real yield curves,” *Journal of Banking & Finance*, 34, 281–294.

Kalemli-Ozcan, S., E. Luttini, and B. Sørensen (2014): “Debt crises and risk-sharing: the role of markets versus sovereigns,” *The Scandinavian Journal of Economics*, 116, 253–276.

Kaminska, I., Z. Liu, J. Relleen, and E. Vangelista (2018): “What do the prices of UK inflation-linked securities say on inflation expectations, risk premia and liquidity risks?” *Journal of Banking & Finance*, 88, 76–96.

Känzig, D. R. (2021): “The macroeconomic effects of oil supply news: evidence from OPEC announcements,” *American Economic Review*, 111, 1092–1125.

——— (2023): “The unequal economic consequences of carbon pricing,” *NBER Working Paper No. 31221*.

Kaplan, G. and G. L. Violante (2014): “A model of the consumption response to fiscal stimulus payments,” *Econometrica*, 82, 1199–1239.

Kerssenfischer, M. and M. Schmeling (2022): “What moves markets?” *Deutsche Bundesbank Discussion Paper No. 16/2022*.

Kinda, T., A. Lengyel, and K. Chahande (2022): “Fiscal multipliers during pandemics,” *IMF Working Paper No. 2022/149*.

Klingler, S. and S. Sundaresan (2023): “Diminishing treasury convenience premiums: effects of dealers’ excess demand and balance sheet constraints,” *Journal of Monetary Economics*, 135, 55–69.

Kose, M. A., E. S. Prasad, and M. E. Terrones (2009): “Does financial globalization promote risk sharing?” *Journal of Development Economics*, 89, 258–270.

Krishnamurthy, A. and A. Vissing-Jorgensen (2011): “The effects of quantitative easing on interest rates: channels and implications for policy,” *NBER Working Papers No. 17555*.

——— (2012): “The aggregate demand for treasury debt,” *Journal of Political Economy*, 120, 233–267.

Kučera, A., E. Kočenda, and A. Maršál (2022): "Yield curve dynamics and fiscal policy shocks," *IES Working Paper No. 4/2022*.

Kuttner, K. N. (2001): "Monetary policy surprises and interest rates: evidence from the Fed funds futures market," *Journal of monetary economics*, 47, 523–544.

Lafrogne-Joussier, R., J. Martin, and I. Mejean (2023): "Supply shocks in supply chains: evidence from the early lockdown in China," *IMF economic review*, 71, 170–215.

Laubach, T. (2011): "Fiscal policy and interest rates: the role of sovereign default risk," *NBER International Seminar on Macroeconomics*, 7, 7–30.

Leeper, E. M., T. B. Walker, and S.-C. S. Yang (2013): "Fiscal foresight and information flows," *Econometrica*, 81, 1115–1145.

Lengyel, A. (2022): "Treasury supply shocks and the term structure of interest rates in the UK," *MNB Working Papers 2022/6, Magyar Nemzeti Bank (Central Bank of Hungary).*

Lengyel, A. and M. Giuliodori (2022): "Demand shocks for public debt in the Eurozone," *Journal of Money, Credit and Banking*, 54, 1997–2028.

Levine, R., C. Lin, and W. Xie (2016): "Spare tire? Stock markets, banking crises, and economic recoveries," *Journal of Financial Economics*, 120, 81–101.

Li, C. and M. Wei (2013): "Term structure modeling with supply factors and the Federal Reserve's Large-Scale Asset Purchase Progarms," *International Journal of Central Banking*, 9, 3–39.

Liu, Y. (2023): "Government debt and risk premia," *Journal of Monetary Economics*, 136, 18–34.

Lou, D., H. Yan, and J. Zhang (2013): "Anticipated and repeated shocks in liquid markets," *The Review of Financial Studies*, 26, 1891–1912.

Lucas Jr, R. E. (2003): "Macroeconomic priorities," *American Economic Review*, 93, 1–14.

Ma, C., J. Rogers, and S. Zhou (2023): "Modern pandemics: recession and recovery," *Journal of the European Economic Association*, jvad019.

Mahajan, K. and S. Tomar (2021): "COVID-19 and supply chain disruption: evidence from food markets in India," *American Journal of Agricultural Economics*, 103, 35–52.

Malik, S. and A. Meldrum (2016): "Evaluating the robustness of UK term structure decompositions using linear regression methods," *Journal of Banking & Finance*, 67, 85–102.

McLaren, N., R. N. Banerjee, and D. Latto (2014): "Using changes in auction maturity sectors to help identify the impact of QE on Gilt yields," *The Economic Journal*, 124, 453–479.

Mélitz, J. and F. Zumer (1999): "Interregional and international risk-sharing and lessons for EMU," in *Carnegie-Rochester Conference Series on Public Policy*, Elsevier, vol. 51, 149–188.

Milano, V. and P. Reichlin (2017): "Risk sharing across the US and eurozone: the role of public institutions," *VoxEU.org, 23rd Jan. 2017*.

Miranda-Agrippino, S. and G. Ricco (2021): "The transmission of monetary policy shocks," *American Economic Journal: Macroeconomics*, 13, 74–107.

Miyamoto, W., T. L. Nguyen, and D. Sergeyev (2018): "Government spending multipliers under the zero lower bound: evidence from Japan," *American Economic Journal: Macroeconomics*, 10, 247–277.

Modigliani, F. and R. Sutch (1966): "Innovations in interest rate policy," *The American Economic Review*, 56, 178–197.

Mumtaz, H. and L. Sunder-Plassmann (2021): "Nonlinear effects of government spending shocks in the USA: evidence from state-level data," *Journal of Applied Econometrics*, 36, 86–97.

Nakamura, E. and J. Steinsson (2018): "High-frequency identification of monetary non-neutrality: the information effect," *The Quarterly Journal of Economics*, 133, 1283–1330.

Nguyen, T. T. (2018): "Public debt and the slope of the term structure," *Fisher College of Business Working Paper*, 23.

Nickel, C. and A. Tudyka (2014): "Fiscal stimulus in times of high debt: reconsidering multipliers and twin deficits," *Journal of Money, Credit and Banking*, 46, 1313–1344.

Nikolov, P. (2016): "Cross-border risk sharing after asymmetric shocks: evidence from the Euro Area and the United States," *Quarterly Report on the Euro Area (QREA)*, 15, 7–18.

Owyang, M. T., V. A. Ramey, and S. Zubairy (2013): "Are government spending multipliers greater during periods of slack? Evidence from twentieth-century historical data," *American Economic Review*, 103, 129–134.

Perotti, R. (1999): "Fiscal policy in good times and bad," *The Quarterly Journal of Economics*, 114, 1399–1436.

Pflueger, C. E. and L. M. Viceira (2016): "Return predictability in the Treasury market: real rates, inflation, and liquidity," *Handbook of Fixed-Income Securities*, 191–209.

Phillot, M. (2021): "US Treasury auctions: a high frequency identification of supply shocks," *Manuscript*.

Piazzesi, M. and E. Swanson (2008): "Futures prices as risk-adjusted forecasts of monetary policy," *Journal of Monetary Economics*, 55, 677–691.

Pierucci, E. and L. Ventura (2010): "Risk sharing: a long run issue?" *Open Economies Review*, 21, 705–730.

Primiceri, G. E. (2005): "Time varying structural vector autoregressions and monetary policy," *The Review of Economic Studies*, 72, 821–852.

Ramey, V. (2016): "Chapter 2. Macroeconomic shocks and their propagation," *Handbook of Macroeconomics*, 2.

Ramey, V. A. and S. Zubairy (2018): "Government spending multipliers in good times and in bad: evidence from US historical data," *Journal of Political Economy*, 126, 850–901.

Rangvid, J., P. Santa-Clara, and M. Schmeling (2016): "Capital market integration and consumption risk sharing over the long run," *Journal of International Economics*, 103, 27–43.

Risa, S. (2001): "Nominal and inflation indexed yields: separating expected inflation and inflation risk premia," *Manuscript, Columbia University.*

Romer, C. D. (2021): "The fiscal policy response to the pandemic," *Brookings Papers on Economic Activity*, 89–110.

Saúl, J. (2012): "The US breakeven inflation rates in a preferred-habitat model," *Manuscript, Universidad Carlos III de Madrid*.

Schaechter, M. A., M. T. Kinda, M. N. Budina, and A. Weber (2012): "Fiscal rules in response to the Crisis: toward the "next-generation" rules: a new dataset," *IMF Working Paper No. 2012/187*.

Schularick, M. and A. M. Taylor (2012): "Credit booms gone bust: monetary policy, leverage cycles, and financial crises, 1870-2008," *American Economic Review*, 102, 1029–61.

Simon, D. P. (1991): "Segmentation in the Treasury Bill market: evidence from cash management Bills," *The Journal of Financial and Quantitative Analysis*, 26, 97–108.

Sims, E. and J. Wolff (2018): "The output and welfare effects of government spending shocks over the business cycle," *International Economic Review*, 59, 1403–1435.

Sørensen, B. E., Y.-T. Wu, O. Yosha, and Y. Zhu (2007): "Home bias and international risk sharing: twin puzzles separated at birth," *Journal of International Money and Finance*, 26, 587–605.

Stempel, D. et al. (2021): "Risk sharing heterogeneity in the United States," *Economics Bulletin*, 41, 1223–1240.

Stock, J. H. and M. W. Watson (2002): "Has the business cycle changed and why?" *NBER Macroeconomics Annual*, 17, 159–218.

Summers, P. M. et al. (2005): "What caused the Great Moderation? Some cross-country evidence," *Economic Review-Federal Reserve Bank of Kansas City*, 90, 5.

Sutherland, A. (1997): "Fiscal crises and aggregate demand: can high public debt reverse the effects of fiscal policy?" *Journal of Public Economics*, 65, 147–162.

Tasca, P. and S. Battiston (2014): "Diversification and financial stability," *SRC Discussion Paper No. 10*.

Upper, C. and T. Werner (2007): "The tail wags the dog: time-varying information shares in the Bund market," *BIS Working Papers No. 224*.

VAYANOS, D. AND J.-L. VILA (2021): "A preferred-habitat model of the term structure of interest rates," *Econometrica*, 89, 77–112.

VERONESI, P. (1999): "Stock market overreaction to bad news in good times: a rational expectations equilibrium model," *Review of Financial Studies*, 12, 975–1007.

WILSON, D. J. (2020): "The covid-19 fiscal multiplier: lessons from the great recession," *FRBSF Economic Letter*, 13, 1–5.

ZHANG, H. (2021): "The impact of COVID-19 on global production networks: evidence from Japanese multinational firms," *ERIA Discussion Paper Series No. 364*.

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